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Sowing Seeds of Mobility: The Gendered Impact of Land Reform^{*}

Abstract

We study how land market frictions affect men and women differently during structural transformation, exploiting two major Chinese land reforms that strengthened farmers' rental rights. Using large-language-model text analysis, we construct a county-level reform index and combine it with large panel data to identify causal effects. The reforms shift rural women out of agriculture more than men and reduce urban women's employment and wages relative to men. A multisector model with intra-household labor allocation rationalizes these findings: gender-neutral land market frictions act as gender-specific mobility barriers, and their removal disproportionately accelerates women's transition into non-agricultural employment.

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gender, land, labor mobility, structural transformation

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1 Introduction

Structural transformation, the reallocation of labor across sectors, is a central feature of economic development. As Kuznets (1966) emphasized, mobility barriers that impede this process hinder economic growth. Although a substantial literature has examined the causes and consequences of the slow transition out of agriculture in developing countries, the gender dynamics of this process remain underexplored. Agricultural work has long been the dominant form of female employment in developing economies (Boserup, 1970), and women transition into non-agricultural employment at significantly lower rates than men (ILO, 2019).¹ If barriers to structural transformation hinder growth in general, those that fall disproportionately on women compound these costs and raise distinct questions regarding their origins and potential remedies. A natural explanation points to well-documented *gender-specific* factors: social norms, comparative advantage, or discriminatory institutions. This paper draws attention to a distinct and novel mechanism: *gender-neutral* institutions can interact with intra-household labor allocation to function as gender-specific mobility barriers.

We focus on land market institutions prevalent across developing countries, where ownership is typically communal and farmers hold only use rights (Bank, 2024) – leaving land uncultivated often risks its reallocation. This tenure insecurity creates a strong incentive for households to retain a member on the farm as “guard labor” (de Janvry et al., 2015) to protect their land-use rights.² Women are the natural candidates for this role for two reasons. First, non-agricultural jobs, especially in manufacturing, often involve tasks and environments where men hold a comparative advantage. Second, agricultural work, particularly on family farms, a traditional form of “working from home”, is more compatible with women’s greater responsibilities in home production such as childcare, eldercare, and domestic chores.³ Our objective is to establish, both empirically and theoretically, that these *de jure* gender-neutral land market institutions can operate as *de facto* gender-specific mobility barriers through intra-household labor allocation, generating a gendered path of structural transformation.

¹In 2018, 38% of women in Northern Africa were employed in agriculture, 15 percentage points more than men; in Southern Asia, the gap was 21 percentage points, as men migrated to urban areas while women remained in the rural sector (ILO, 2019). See also Lee (2024) and Chiplunkar and Kleineberg (2024) for evidence across a large panel of developing countries.

²Guard labor refers to the amount of labor households must allocate to protecting informal property claims. Field (2007) shows in her study of Peru’s urban titling program that providing formal titles eliminates this need, freeing adults for the labor market, reducing child labor, and correcting labor misallocations.

³Boserup (1970) highlights unpaid family farm work as the dominant form of female employment at early stages of economic development. This is supported by Durand (1975), who draws on census data from 100 countries over the period 1946–66. For more recent evidence, see Dinkelman and Ngai (2022) for Africa and Ngai et al. (2024) for the 19th century United States.

This paper makes two main contributions. Empirically, we exploit China’s land institutions under the *hukou* system and subsequent land reforms that strengthen land-rental rights as a natural laboratory to study the causal impact of labor mobility barriers arising from land market frictions.⁴ This setting allows us to construct a granular county-level land reform index using a large language model (LLM)-based textual analysis, providing a direct measure of the implicit land market frictions common in developing countries. By combining this reform index with the *Rural Fixed Point Survey* (RFPS, a large panel dataset tracking rural individuals and households over two decades) and the Chinese Household Income Project (CHIP, a nationally representative household survey), we identify the *causal* impact of reform on employment outcomes by gender and by rural and urban status, highlighting the importance of examining different demographic groups and their interactions. Theoretically, we develop a multisector model incorporating intra-household labor allocation to explain why gender-neutral land reforms generate gender-specific outcomes and to quantify their impact on the gendered patterns of structural transformation.

The county-level land reform index is constructed by tracing the rollout of two land reforms aimed at strengthening farmers’ land rental rights: the land contracting reform (2003–2014) and the land titling reform (2009–2019). Applying keyword filtering to over three million government policy documents and then using LLM-based validation, we identify about 4,400 reform-related documents and construct a novel index capturing the rich temporal and spatial diffusion of reforms across more than 2,000 counties. Using this index as a direct measure of mobility barriers, our analysis based on RFPS shows that the reforms have significantly strengthened rental rights, reflected in higher rental activity and larger farm size. As a result, they have increased the transition out of agriculture, especially for women, and have increased joint spousal employment in non-farm work. We then use the CHIP data to estimate the impact on urban populations. Using a shift-share design, we find that land reforms widen the gender gap in employment and wage income, particularly among individuals with a lower education level. This points to a substitution effect from rural female migrants on urban women, concentrated among lower-educated workers.

Our empirical analysis identifies the causal effects of land reforms by gender by comparing counties with different reform statuses. To understand the aggregate effects and the mechanism through which gender-neutral land reforms generate gender-specific impacts, we develop a multi-region and multi-sector model with household-level employment decisions. There is an urban region and a continuum of rural regions. Motivated by our empirical findings, we allow for two types of rural regions: reform and non-reform regions.

⁴The *hukou* system classifies citizens as either agricultural or non-agricultural. Agricultural *hukou* holders can cultivate government-allocated land rent-free but may lose it if they do not farm it.

Households in rural regions have agricultural *hukou* and are entitled to farm communal land free of rent. They choose whether the female member, the male member, or both should leave agriculture to work in non-agriculture in the urban region. In rural regions without land reform, leaving agriculture means forgoing land income, so households have an incentive to retain a member as “guard labor” to protect the land income of the household. Since female labor is more intensively used in agriculture than in non-agriculture, and women’s migration to the urban region reduces home sector productivity, women are more likely to serve as guard labor. Reforms that eliminate the need for guard labor therefore disproportionately increase women’s transition into non-agriculture. This increases the relative supply of female workers in non-agriculture, hence lowering the employment and wage rate of urban women relative to men.

To discipline the model, we calibrate the share of reform regions using our empirical reform index from 2000 to 2020, and the remaining parameters to match household employment patterns and sector productivity in 2000 and 2020. The model’s predicted differences between reform and non-reform regions on the untargeted moments – employment patterns by gender, agricultural labor productivity, and hours per worker – are very close to the empirical findings. With these empirical validations, we use the model to conduct two counterfactual exercises. First, to study the aggregate impact of reform, we consider a no-reform counterfactual by restricting the share of reform regions to zero from 2000 to 2020. We find that land reform increases rural women’s non-agricultural employment relative to men by 2.1 percentage points, while decreasing urban employment of women relative to men by 2.8 percentage points and the female-to-male wage ratio by 1.9%.⁵ The reform also improves agricultural productivity, accounting for about 31% of the decline in the productivity gap across non-agriculture and agriculture. Second, to isolate the role of guard labor, we consider a counterfactual exercise in which the remaining member can only protect her/his own share of land. In this ‘no-guarding case’, the land reform no longer implies more women than men transitioning into non-agriculture.

Related Literature Women’s employment has historically been intertwined with agriculture, and the nature of agricultural technology can play a decisive role in shaping gender norms (Boserup, 1970; Alesina et al., 2013). Recent work has examined two distinct channels through which technological change generates gendered labor market outcomes: Ager et al. (2026) identify the effects of gender-biased technological change within agriculture, while Ngai et al. (2024) show that even gender-neutral technological change across sectors

⁵This provides a novel mechanism for the widening gender gaps among urban workers in China (Feng et al., 2017; Brussevich et al., 2021; Qian, 2023). See Dustmann et al. (2016) for a review on local labor market effects of immigration.

can interact with comparative advantage to produce gendered transition patterns out of agriculture. We contribute to this literature by establishing how gender-neutral land market institutions can function as gender-specific mobility barriers during structural transformation. Our approach complements both studies. Whereas Ngai et al. (2024) document and explain the U-shaped pattern of female employment using historical US data, Ager et al. (2026) exploit a quasi-natural experiment to identify the role of gender-biased technology. We instead employ a quasi-natural experiment to establish a causal link between gender-neutral land reforms and the gendered pattern of structural transformation.

Our paper is also related to a growing body of work on the misallocation of female talent across sectors and occupations and its implication for aggregate productivity (Hsieh et al., 2019; Lee, 2024; Chiplunkar and Kleineberg, 2024; Gottlieb et al., 2024). A common approach in this literature is to assume and infer gender-specific mobility barriers from matching data moments.⁶ We take a different approach by empirically identifying and directly measuring a form of gender-neutral barriers, land market frictions, that can operate as gender-specific barriers. Since land market frictions of the kind we study are prevalent across developing countries (Bank, 2024), the mechanism we identify has implications that extend well beyond China.

Land market frictions and their economic implications have been examined in developing countries across different continents (Chernina et al., 2014; de Janvry et al., 2015; Chen, 2017; Gottlieb and Grobovšek, 2019; Manyшева, 2022). Our empirical analysis on the two land reforms in China relates to Chari et al. (2021), who construct a provincial-level index for the land-contracting reform and find improvements in land allocation and agricultural productivity but no effect on labor reallocation. Building on this provincial-level index, Shi et al. (2024) find positive effects on non-farm employment but weaker effects for women. However, a provincial-level index is likely to obscure substantial within-province variation across nearly 2,000 counties. Liu et al. (2023) collect the timing of completing land titling reform for a selected set of counties and find a positive effect on migration. Our county-level index captures both reforms, covers nearly 2000 counties in China, and exploits the within-province heterogeneity for a much longer time span, allowing us to detect novel effects by gender and by rural and urban populations. Our quantitative model for studying the role of land insecurity in China is closely related to Ngai et al. (2019) and Adamopoulos et al. (2024), but differs in two key ways. First, we use our land reform index to directly inform the share of reform regions in the model instead of relying on survey-based reallocation frequencies. Second, we emphasize the employment decisions of couples to understand the

⁶Reduction of gender-specific barriers is often related to policies that expand women’s legal and economic rights (Bailey et al., 2024; Doepke et al., 2012; Heath and Tan, 2020).

impact by gender.⁷

Finally, Lagakos (2020) highlights a literature gap, noting the need for more quasi-natural experimental evidence to identify the effects of mobility barriers. Recent contributions include Sarvimäki et al. (2022), which exploits forced migrations in historical Finland, and Gai et al. (2025), which leverages China’s Rural Pension Reform. We contribute by identifying the effects of reducing mobility costs through land reforms in China on structural transformation by gender, a setting of broad relevance for other developing economies.

The remainder of the paper is structured as follows. Section 2 provides background on China’s land institutions and describes the construction of our land reform index. Section 3 presents the main empirical findings for rural and urban populations. Section 4 introduces the model, and Section 5 conducts the quantitative exercise. Finally, Section 6 concludes.

2 Land Institutions and the Land Reform Index

This section outlines China’s land institutions under the *hukou* system, posits the link between land insecurity and women’s role as guard labor, and employs an LLM-based approach to build a novel county-level reform index for the empirical analysis in Section 3 and the quantitative analysis in Section 4.

The household registration system, known as *hukou*, was established in 1958 and acts as an internal passport to regulate migration. At birth, an individual is assigned one of two *hukou* types: “agricultural” or “non-agricultural”. We focus on the land policy aspect of the *hukou* system.⁸ The 1986 Land Management Law explicitly recognized agricultural *hukou* households’ rights to farm their allocated agricultural land, while the ownership of land remained collective. However, land tenure was insecure – land could be reallocated if household members did not engage in farming.⁹ This insecurity discouraged land rental activities. Formal land rental markets were virtually non-existent, and when land was rented

⁷See Cao et al. (2024), Imbert et al. (2025) and Matsuda et al. (2026) for the importance of employment and migration decisions at the household level. While Adamopoulos et al. (2024) examines migration decisions across age groups, we focus on gender and rural-urban interaction, driven by our empirical findings. Related results on different age groups will be discussed throughout the paper.

⁸The *hukou* system classifies individuals by agricultural/non-agricultural status, as well as by registration place – the administrative region that issued their *hukou*. While our paper focuses on the land aspect, a second important aspect is that entitlements to social services, such as education and healthcare, are often tied to one’s registration place. Migrant workers who leave their *hukou* location typically do not have access to these services at their destination. Although changing *hukou* status has gradually become easier, it remains difficult for most. As of 2020, over 50% of the population still holds an agricultural *hukou*. More details on China’s *hukou* system are provided in the Online Appendix B2.

⁹This practice has been extensively documented by the literature, see, for example, Kung and Liu (1997); Brandt et al. (2002).

out, it was typically through informal, verbal agreements that yielded very low rents.¹⁰

2.1 Land Insecurity and Gendered Patterns of Guard Labor

In the presence of land insecurity, the transition from agriculture to non-agriculture could result in a loss of land, incentivizing households to leave one member behind to farm and safeguard their land rights. We present two sets of suggestive evidence that women are more likely to take on this role of “guard labor”.

First, women are less likely than men to work in the non-agricultural sector. Appendix Table A1 documents this pattern for individuals with agricultural *hukou*, using data from the Chinese Population Census and the One Percent Population Survey (1990-2020), controlling for ethnic minorities, age, age squared, marital status, skill (education), province fixed effects, and year fixed effects. More importantly, this employment pattern also holds for married couples.¹¹ Appendix Figure A1 displays the shares of four types of couples: (1) both spouses employed in non-agriculture, (2) husband-only employed in non-agriculture, (3) wife-only employed in non-agriculture, and (4) neither employed in non-agriculture.¹² In the earlier years, households where only the husband worked in the non-agricultural sector dominated. Over time, the share of couples where both spouses work in non-agriculture has increased, especially after 2005.¹³

Second, we use data from Rural Fixed Point Survey (RFPS) to show that historical land insecurity reduces non-farm employment and out-migration, mainly for women. The RFPS provides household-level information between 1995 and 2017 and individual-level information from 2003 to 2017. In 2006, RFPS conducted a one-off supplementary survey on village governance, where village cadres were asked to recall instances of major land redistribution (*Datiao Zheng*) from the 1970s through 2006.¹⁴ In our regressions, we use this to measure

¹⁰The 1998 revision of the law aimed to strengthen land-use rights by formalizing 30-year contracts between village collectives and individual households. However, the transfer or leasing of these land-use rights remained heavily restricted. By 2000, fewer than 10% of farmers rented out land (Deininger and Jin, 2005).

¹¹Most couples share the same *hukou* type. Census data on married couples show that the share of mixed *hukou* couples among those aged 18–55 was about 5% from 1990 to 2010 and increased slightly to around 9% during the 2015–2020 period.

¹²To emphasize the “use-it-or-lose-it” aspect of land policy, we focus on total non-agricultural employment, which includes both local non-farm work and migration outside the county for non-farm jobs. In the 1980s and early 1990s, Township and Village Enterprises (TVEs) are the main source of local non-farm work. Since 2000, however, growth in non-agricultural employment has been driven mainly by the urban non-agricultural sector. See Figure 1 of Ngai et al. (2019), based on data from Brandt and Zhu (2010).

¹³The online Appendix Figure B1 shows that the pattern outlined in Figure A1 holds for both the young and old age groups. Using CHIP data for 1995–2018, Online Appendix Figure B2 shows a similar pattern when focusing on migration instead of employment.

¹⁴Major land redistribution refers to the reallocation of more than 30% of village land by village cadres. Despite this strict threshold, a substantial fraction of these villages experienced such reallocation, with

historical land insecurity at the village level. Appendix Table A2 presents the regression results using household data (1995–2006) and individual data (2003–2006). More frequent historical land redistribution reduced both households’ and individuals’ participation in non-farm employment and out-migration. More importantly, individual-level results reveal that this negative impact is primarily driven by women. This corroborates our hypothesis that female household members are disproportionately retained on the farm to act as guard labor.

2.2 Two Land Policy Reforms

Since 2003, China has implemented two major top-down land reforms: the land contracting reform (2003–2014) and the land titling reform (2009–2019). Crucially, both reforms strengthened only farmers’ rental rights rather than conferring any right to sell. The collective ownership of agricultural land remained intact throughout.

The land contracting reform, initiated by the Rural Land Contracting Law (RCLC) in 2003, granted agricultural *hukou* households the legal right to lease and rent farmland—rights that had previously been largely based on informal agreements. Following the national legislation, provincial governments issued detailed regulations on leasing, transfers, and dispute resolution between 2003 and 2010. In turn, the prefecture and county governments introduced specific administrative procedures and legal directives.

While the contracting reform established the legal framework, individual farmers had no certified documentation of their land rights and there was no national land registration system. Land titling reform, piloted in 2009¹⁵, addressed both gaps. Unlike the contracting reform’s sequential provincial-to-county rollout, the titling reform was implemented directly at the county level, with counties given autonomy to set their own start dates and implementation strategies.

Under the titling reform, households received a standardized certificate recording plot location, area, and ownership information, directly facilitating land transactions and judicial dispute resolution. Beyond individual certificates, the titling reform established a unified, digitalized national land registration system. The system records comprehensive information, including household ownership, demographic data, identity of household heads, family relationships, and precise GIS-based parcel boundaries. The completion of the land titling reform in 2019 marked a major milestone in improving land tenure security in China’s agricultural sector.

considerable regional disparities in reallocation frequency. Since 1976, our data show that 26% of villages experienced redistribution once, 31% twice, 23% three times, 10% four times, and 10% five or more times, with some villages experiencing up to eight redistributions during this period.

¹⁵The central government launched the nation-wide land titling reform in 2013 by issuing its No. 1 Document, but small-scale pilot programs began in eight villages in 2009.

2.3 Land Reform Index

To construct a numerical measure that captures the intensity of these two reforms, we begin by performing keyword searches in the `pkulaw.com` database to identify legal and government documents related to land contracting and land titling reform.¹⁶ Table B5 in the Online Appendix lists the keywords used and the number of documents identified by administrative level. In total, we located 4,522 potentially relevant documents: 3,451 for contracting reform and 1,071 for titling reform.

Next, we employ *Doubao*, a leading Chinese large language model developed by ByteDance, to verify the substantive relevance of each document. For each reform, we first provide *Doubao* with the central government policy document to establish a benchmark. Then, using its API service, we prompt the model to assess whether each local document pertains to the same reform and, if so, to extract the most relevant sentences supporting its classification.¹⁷ Through this validation process, we identify 4,413 documents as reform-related. Finally, we extract information on the administrative level, location, and year from these verified documents. Figure B3 in the Online Appendix provides examples of policy documents from various administrative levels.

Given the two-stage rollout of the land contracting reform, a county scores 1 point if its provincial government has issued a relevant document, plus an additional 1 point if either the county itself or its prefecture has released a local regulation. Thus, a county can achieve a maximum of 2 points for land contracting reform. By 2014, about 92% of counties were governed by either provincial or local regulations, with 39% covered by both. Since the land titling reform began in 2009 before the completion of the contracting reform, some counties completed titling without finishing contracting. Because titling formally certifies individual land rights through official registration, it effectively subsumes the protections that contracting reform sought to establish. We therefore assign a score of 3 to counties that have completed the titling reform, regardless of their status on contracting reform. In sum, our land reform index ranges from 0 to 3. Figure 1 illustrates the spatial diffusion of the phased reforms.

Our land reform index improves existing measures in two fundamental ways. First, on the contracting reform, it exploits granular county-level variation stemming from the staggered

¹⁶As of September 2024, this database – maintained by the Legal Information Center of Peking University and one of the most comprehensive of its kind – contains 4,049,703 documents issued by local governments (provincial, prefectural, and county levels) and 454,035 documents issued by central authorities.

¹⁷We use the following prompt: “Read (1) the {central policy document} on land contracting/titling reform, compare it with (2) the {local policy document}, and if (2) concerns the implementation of the same reform in (1), return 1; otherwise, return 0. If your answer is 1, please extract the most relevant sentences from (2) that support your judgement.”

roll-out. Second, it incorporates both contracting and titling reforms, providing a more comprehensive measure of changes in land property rights from 2003 to 2019.

Liu et al. (2023) collect the timing of county-level land titling reform for the counties corresponding to 229 villages (a subsample of all villages in the RFPS) by visiting the official websites of local agricultural bureaus county by county. Our index, by contrast, covers 2,813 counties, and is constructed by sourcing over three million government documents from a centralized database. As our index covers all counties in China and incorporates both the contracting and titling reforms, we can analyze the impact of land reform on the urban population. This is important because migrants originate from rural counties with varying reform intensity, and reform across all rural counties shapes the overall impact on urban labor markets.

2.4 Timing and Exogeneity of Reforms

Both reforms followed a top-down implementation process, initiated by the central government’s “No. 1 Document” (Chari et al., 2021; Liu et al., 2023), but there are significant spatial variations in the timing of county-level contracting and titling reforms. Several potential factors that can affect the timing of reforms have been highlighted in the literature, including practical constraints on program rollout (Liu et al., 2022), pending land ownership disputes, and insufficient funding in fiscally weak areas (Cao et al., 2022).

We formally test what predicts the timing of adopting reforms by regressing the year of county-level reform completion on a set of predetermined county characteristics, including economic prosperity, economic structure, fiscal capacity, urbanization, farm size, geographic connectivity, and, crucially, land pressure arising from real estate development and existing land disputes. We conduct this exercise separately for the contracting and titling reforms, and the results are reported in Table 1.

Two patterns emerge. First, most conventional county characteristics, including GDP, population size, agricultural and manufacturing shares, fiscal revenue, average farm size, urbanization rate, and distance to first-tier cities or railways, do not predict the timing of either reform. This is consistent with the findings of Bu and Liao (2022) and Cheng et al. (2019), who argue that the rollout was primarily driven by administrative and logistic considerations rather than local economic conditions. Second, two variables, real estate investment and land disputes, do predict earlier adoption. Counties with greater real estate development pressure and more pre-existing land disputes adopted both reforms significantly earlier. This is intuitive, since land disputes and encroachment by real estate development created an urgency to formalize land rights. In our main specification (columns (4) and (8)

of Table 2), we will control for county-level real estate investment.

Table 1: Determinants of Reform Timing

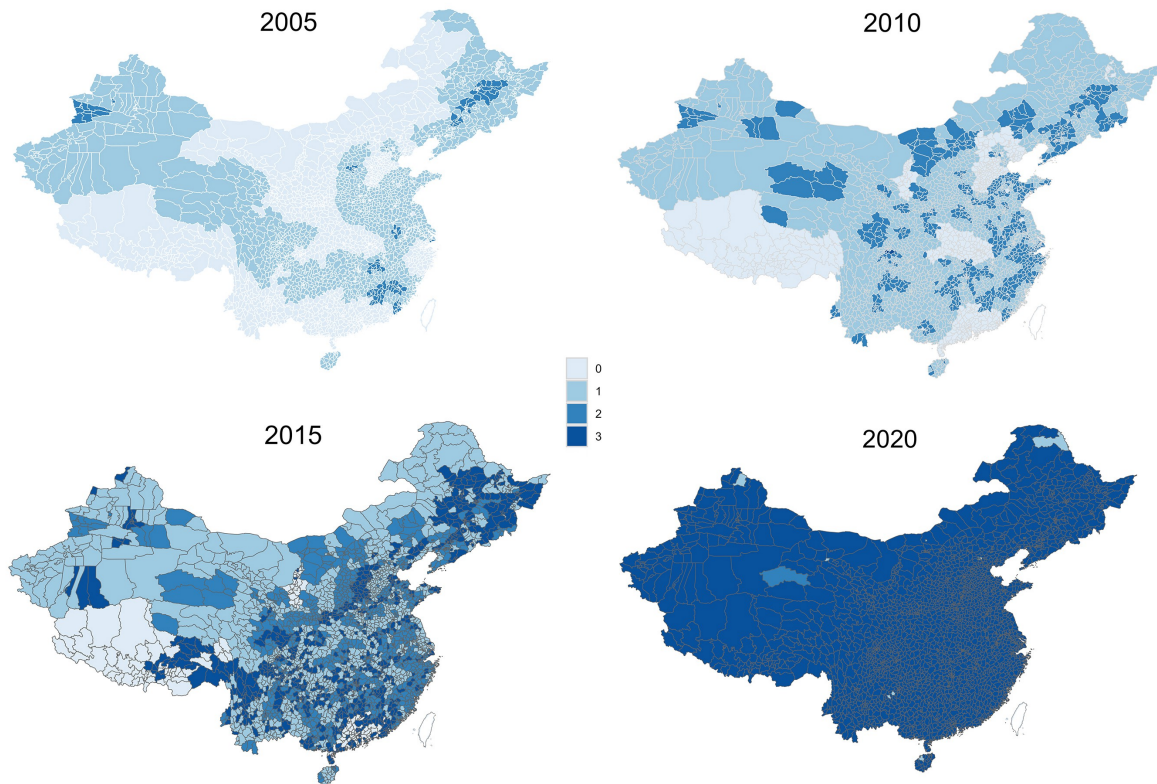
	Year of Reform	
	Contracting (1)	Titling (2)
GDP (Logged)	-0.033 (0.029)	-0.013 (0.038)
Population size (Logged)	-0.081 (0.154)	-0.113 (0.182)
Agricultural share	-0.614 (0.913)	-0.476 (0.840)
Manufacture share	-0.013 (0.017)	0.004 (0.016)
Fiscal Revenue (Logged)	0.038 (0.074)	0.041 (0.088)
Real Estate Investment (Logged)	-0.304* (0.183)	-0.419*** (0.104)
Land Dispute	-0.225** (0.093)	-0.278*** (0.096)
Average farm size	0.092 (0.096)	0.075 (0.097)
Urbanization Rate	-0.017 (0.016)	-0.023 (0.015)
Distance to First Tier Cities	-0.768 (0.876)	-0.555 (0.805)
Distance to Railway	-0.180 (0.295)	-0.162 (0.276)
Observations	2813	2813

Notes: This table reports county-level, cross-sectional regressions examining the determinants of adoption timing for both land reforms: the contracting reform and the titling reform. All explanatory variables are measured either in 2003 or 2000, depending on data availability, to ensure they are predetermined relative to reform implementation. Robust standard errors are reported in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

To further establish the exogeneity of land reforms, we treat the completion of each reform at the county level as a separate policy shock and employ a staggered difference-in-differences (DiD) design to examine its impact on employment and migration patterns. Figure 2 illustrates the dynamic effects of land contracting and titling reforms (left and right panels, respectively) on non-farm employment (upper panels) and migration decisions (lower panels). We merge individual-level data (2003–2017) from the RFPS with our county level reform index, controlling for a rich set of individual, regional economic, and policy variables. Applying multiple estimators from recent methodological advances in the DiD literature, we confirm the absence of pre-reform trends in both non-farm employment and migration, formally validating the parallel trends assumption central to our identification strategy. Furthermore, the event-study estimates reveal that approximately three years

after local implementation, both reforms produced statistically significant positive effects on agricultural households' non-farm employment and out-migration.

Figure 1: The Spatial Diffusion of Land Reform, 2005–2020

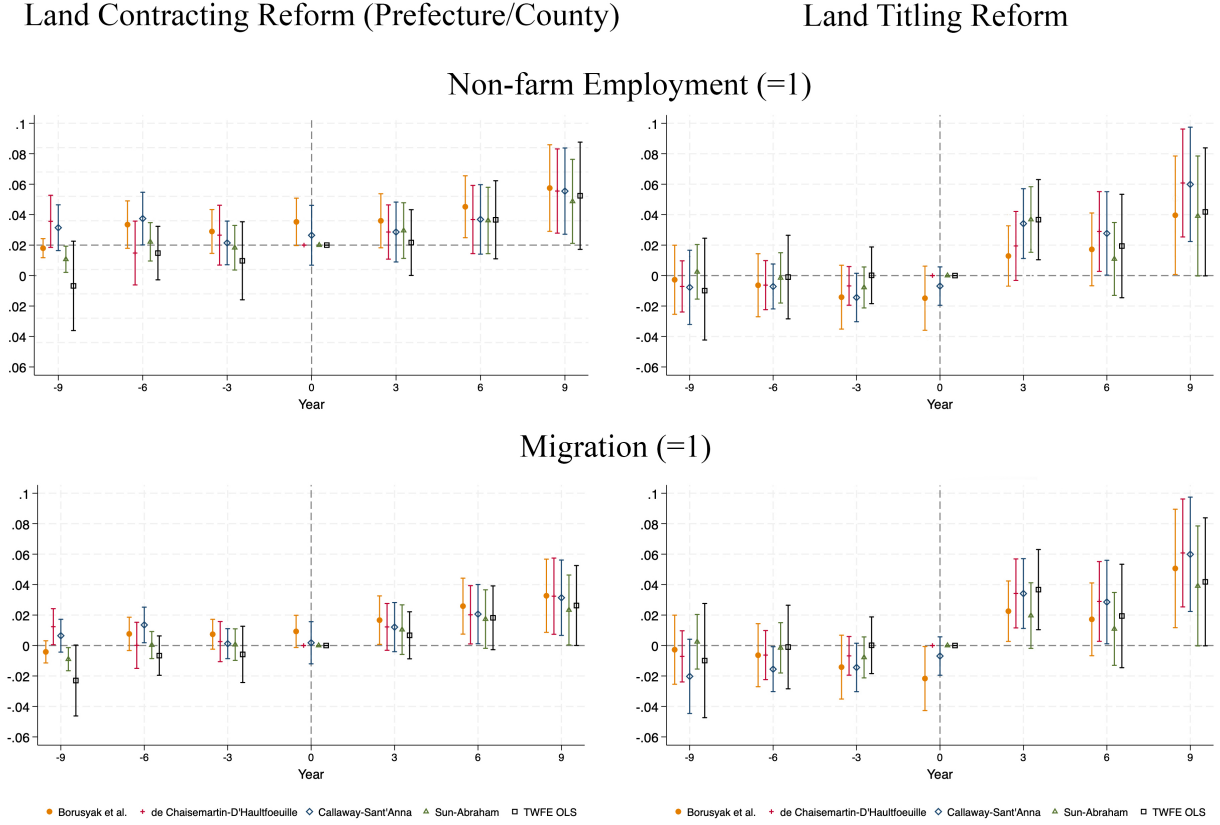


Note: This figure matches the county-level land reform index with the 2010 Chinese county-level GIS map. Darker colors represent a higher land reform Index, reflecting greater depth and broader reform scope within each county. Our reform index takes four potential values: 0 (No reform), 1 (Provincial regulations on land contracting reform issued), 2 (Prefecture/county regulations on land contracting reform issued), and 3 (Land titling reform completed, regardless of contracting reform).

3 Empirical Analysis

This section presents empirical evidence on the impact of land reforms across gender and between rural and urban populations since 2003. We analyze the employment and migration outcomes of rural agricultural *hukou* holders, as well as the employment and wage income of non-agricultural *hukou* holders. Details on the data sources and the main variables used in this section can be found in Appendix A1.

Figure 2: Event Studies of the Land Reforms on Non-farm Employment and Migration



Note: Rural Fixed Point Survey, 2003–2017; individuals aged 18–55. The figure presents event-study estimates of the land contracting reform (left panels) and the land titling reform (right panels), spanning nine years before and nine years after each reform’s local implementation. Upper panels show effects on non-farm employment; lower panels show effects on migration. All specifications include individual and year fixed effects. Individual controls comprise age, a cohabitation dummy, and education level. Economic controls include logged county- and village-level GDP per capita, manufacturing and tertiary industry shares, and logged population size. Policy controls include the weighted non-land-related *hukou* reform indices (Gai et al., 2025) and a dummy for adoption of the New Rural Pension Reform.

3.1 Impact of Land Reforms on Land Security

Two pieces of survey evidence have documented improvements in land security for Chinese farming households following the 2003 reforms. Using a five-wave survey spanning 1999–2010 and covering 17 provinces and 77% of the rural population, Feng et al. (2014) find that while 82% of villages experienced land reallocation by 2001, this figure had dropped to 40% by 2010. A more recent study by Adamopoulos et al. (2024), using two retrospective surveys on the historical frequency of village-wide land reallocations among 120 sampled villages, finds that reallocation events declined from 155 occurrences between 1991–2003 (a probability of

9.9%) to 16 occurrences between 2003–2018 (0.9%).

However, directly testing the impact of reform on land security remains challenging, as high-quality longitudinal data on village-level land reallocations are not widely available. Chari et al. (2021) provide an indirect validation by showing that provincial-level contracting reforms led to more active land rental markets among affected households. In a similar vein, by combining our county-level land reform index, which captures both contracting and titling reforms, with household-level data from RFPS, we find that land reforms significantly increased household-level land rental activities. Specifically, as shown in Appendix Table A3, they raise the probability and area of land rented out, as well as rental income.

3.2 Impact on Rural Households

We now turn to the impact of reform on rural agricultural *hukou* holders, using RFPS data covering the years 2003 to 2017. We restrict the sample to individuals aged 18–55 who self-report as physically healthy to ensure demographic comparability between the rural and urban samples. We then analyze their employment and migration outcomes using the following specification:

$$Y_{iot} = \alpha + \beta_1 LandRef_{ot} + \beta_2 LandRef_{ot} * Female_i + X_{it}\Phi + \delta_i + \kappa_t + X_{ot}\Gamma + \epsilon_{iot} \quad (1)$$

where Y_{iot} denotes a dummy variable capturing one of two related outcomes for individual i in county o and year t : (1) non-agricultural employment, indicating whether the individual devoted any labor days in the non-farm sector—either locally or non-locally, or (2) migration, indicating whether the individual spent any labor days working outside county o in the non-agricultural sector.

$LandRef_{ot}$ is the land reform index in county o in year t . We interact this index with the dummy variable $Female_i$ to assess the impact by gender. We include a rich set of regional economic and policy control variables (X_{ot}), including the logged county and village GDP per capita, the GDP shares of manufacturing and tertiary industries, logged population size, the intensity of non-land-related *hukou* reforms in destination cities,¹⁸ and a dummy variable indicating whether a county adopted the New Rural Pension Reform. Moreover, the RFPS tracks the same agricultural *hukou* households and individuals over time, allowing

¹⁸This non-land-related *hukou* reform index, constructed by Gai et al. (2025), measures the difficulty migrants face in obtaining urban *hukou* status in their work locations. While land reforms push agricultural *hukou* holders out of agriculture, these *hukou* reforms in destination cities serve as pull factors for rural out-migration. In our regression, we weight Gai et al. (2025)’s index by the share of migration from origin county o to destination city d relative to the total out-migration from county o (based on the 2000 Census). This weighted index captures the exposure of county o to *hukou* reforms in all potential destination cities.

us to control for unobserved selection factors at the individual level. Thus, we control for individual fixed effects δ_i , year fixed effects κ_t , and a rich set of individual-level variables X_{it} (such as age, education level, and marital status). Standard errors are clustered at the county level to safeguard against potential heteroskedasticity.

Individual Employment and Migration Outcomes—Table 2 presents the results on non-farm employment (columns (1) to (4)) and migration (columns (5) to (8)). Land reform has a stronger positive effect on rural women’s non-farm employment and out-migration than on men’s. Specifically, the probability of women transitioning to non-farm employment increases by about 0.9 percentage points more than men’s when the reform index increases by 1 point. With the index ranging from 0 to 3, a full reform would correspond to an additional 2.7 percentage-point increase in women’s non-farm employment (column (4)). This serves as our empirical reference point when we later compare the model with the data. For migration, rural women’s probability increases by about 0.7 percentage points more than that of men for each unit increase in the land reform index (column (8)).

These gender-differential effects remain stable as we progressively add controls to shut down alternative channels. In columns (1) and (5), we include two measures proxying for families’ care-taking burden: two dummies indicating whether the household has a child under age 7 or an elderly member above age 65. Having a child significantly reduces both non-farm employment and migration, while the presence of elderly does not. In columns (2) and (6), we add proxies for two concurrent reforms relevant to rural employment and migration: the weighted non-land-related hukou reform index to control for the pull effect from urban-side hukou reforms, and the New Rural Pension reform to capture the push effect (Gai et al., 2025). Only the hukou reforms have a significant effect, and only on migration. In columns (3) and (7), we further control for weighted employment growth and GDP growth in destination cities to absorb time-varying labor demand shocks. Finally, in columns (4) and (8), we add county-level real estate investment, because we find that one of the driving forces of land reform adoption was land pressure from the real estate investment boom, especially in the periphery of urban areas connected to rural areas. Controlling for this channel has no discernible effect on either outcome, suggesting that our estimates are not confounded by local real estate development pressures.

To compare the impact of each reform separately, we replace the comprehensive land reform index with three dummy variables, provincial-level contracting reform, prefecture/county-level contracting reform, and county-level titling reform. Table 3 reports the results, using the same specification as columns (4) and (8) of Table 2. For non-farm employment, provincial-level contracting reform has no significant effect for either gender. By contrast,

Table 2: Impact of Reforms on Rural Individuals' Employment and Migration Outcomes, by Gender

DV	Rural Individual Sample RFPS (2003-2017)							
	Non-farm Employment (=1)				Migration (=1)			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Land Reform	0.009*	0.012**	0.012**	0.012**	0.008**	0.010**	0.011**	0.011**
	(0.005)	(0.005)	(0.005)	(0.005)	(0.004)	(0.005)	(0.005)	(0.005)
Land Reform * Female	0.008**	0.009**	0.009**	0.009**	0.008***	0.007***	0.007***	0.007***
	(0.004)	(0.004)	(0.004)	(0.004)	(0.002)	(0.003)	(0.003)	(0.003)
With Child under 7	-0.020***	-0.020***	-0.020***	-0.020***	-0.007**	-0.006**	-0.006**	-0.006**
	(0.005)	(0.005)	(0.005)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)
With Elderly above 65	-0.001	-0.001	-0.001	-0.001	-0.007	-0.007	-0.007	-0.007
	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)	(0.005)
Weighted Hukou Reform		0.004	0.004	0.004		0.014**	0.014**	0.014**
		(0.005)	(0.005)	(0.005)		(0.006)	(0.006)	(0.006)
New Rural Pension Reform		0.010	0.010	0.010		-0.000	-0.000	-0.000
		(0.010)	(0.010)	(0.010)		(0.005)	(0.005)	(0.005)
Weighted Employment Growth			0.054	0.054			0.120**	0.129**
			(0.040)	(0.040)			(0.054)	(0.055)
Weighted GDP Growth			0.071	0.071			0.234**	0.235**
			(0.070)	(0.070)			(0.103)	(0.102)
County Real Estate Investment				0.000				0.000
				(0.001)				(0.001)
Control Variables	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Individual Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year Fixed Effect	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	537896	537896	537896	537896	537896	537896	537896	537896

Notes: Rural Fixed Point Survey 2003–2017, aged 18–55. The dependent variable in columns 1 to 4: A dummy variable equal to one if the surveyed individual devoted any labor days to non-farm employment during a year, while that in columns 5 to 8: A dummy variable equal to one if the surveyed individual devoted any labor days to out-migration work during a year. Individual control variables include age, a dummy variable of marital status, and education level. Economic control variables include county and village GDP per capita (logged), shares of manufacturing and tertiary industries, and population size (logged). Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

both local-level contracting and titling reforms significantly increase non-farm employment for men, with significantly larger effects for women. For migration, provincial contracting reform significantly increases men's migration, but has no significant effect on women. At the local level, while contracting and titling reforms differ in their significance for men, both significantly increase women's migration relative to men. Taken together, these results are consistent with those obtained using the comprehensive reform index and indicate that the gender-differential effects are primarily driven by the local-level implementations.

Our findings from including three dummies for each reform also align with existing evidence: Chari et al. (2021) report no significant impact on off-farm employment or migration when examining only the provincial-level contracting reform, whereas Liu et al. (2023) find a significant effect of the county-level titling reform. We confirm that both reforms exert significant effects on migration, especially when they progress to the local level.

Table 3: Impact of Individual Reform on Rural Individuals' Employment and Migration Outcomes, by Gender

DV	Rural Individual Sample RFPS (2003-2017)	
	Non-farm Employment (=1) (1)	Migration (=1) (2)
Land Contracting Reform (Provincial)	0.007 (0.008)	0.014** (0.007)
Land Contracting Reform (Provincial) * Female	0.006 (0.007)	-0.005 (0.005)
Land Contracting Reform (Prefecture/County)	0.031** (0.013)	0.009 (0.011)
Land Contracting Reform (Prefecture/County) * Female	0.024* (0.012)	0.023*** (0.008)
Land Titling Reform	0.028* (0.016)	0.039** (0.015)
Land Titling Reform * Female	0.022* (0.013)	0.017** (0.008)
Control Variables	Yes	Yes
Individual Fixed Effect	Yes	Yes
Year Fixed Effect	Yes	Yes
Observations	537896	537896

Notes: Rural Fixed Point Survey 2003–2017, aged 18–55. The dependent variable in column 1 : A dummy variable equal to one if the surveyed individual devoted any labor days to non-farm employment during a year, while that in column 2: A dummy variable equal to one if the surveyed individual devoted any labor days to out-migration work during a year. Individual control variables include age, a dummy variable of marital status, and education level. Economic control variables include county and village GDP per capita (logged), shares of manufacturing and tertiary industries, and population size (logged). All specifications control for individual fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

To test whether the positive effects on women vary by marital status, we replace the female dummy in Equation 1 with three group indicators: married female (41.21% of individual-year observations), unmarried female (7.06%), and unmarried male (10.31%), using married male (41.42%) as the reference group. Using the same specification as in columns (4) and (8) of Table 2, Table 4 shows that the stronger positive effect on rural women's migration is driven entirely by married women.

Finally, to assess robustness across age groups, as emphasized by Adamopoulos et al. (2024), we re-run the baseline regression (Equation 1) on two sub-samples: individuals aged 18–44 and those aged 45–55. The Online Appendix Table B6 suggests that gender differences in the impact of land reform are more pronounced in the younger group (aged 18–44).

Household Employment and Migration Outcomes—To assess the impact of land reforms at the household level—specifically, the rise of households where both spouses work

Table 4: Impact of Reforms on Rural Individuals' Employment and Migration Outcomes, by Gender and Marital Status

DV	Rural Individual Sample RFPS (2003-2017)	
	Non-farm Employment (=1) (1)	Migration (=1) (2)
Land Reform	0.011** (0.005)	0.010** (0.005)
<i>Reference Group: Married Male</i>		
Land Reform * Unmarried Male	0.005 (0.006)	0.002 (0.005)
Land Reform * Unmarried Female	0.013 (0.008)	0.005 (0.005)
Land Reform * Married Female	0.009** (0.004)	0.008*** (0.003)
Control Variables	Yes	Yes
Individual Fixed Effect	Yes	Yes
Year Fixed Effect	Yes	Yes
Observations	537896	537896

Notes: Rural Fixed Point Survey 2003–2017, aged 18–55. The dependent variable in column 1: A dummy variable equal to one if the surveyed individual devoted any labor days to non-farm employment during a year, while that in columns 2: A dummy variable equal to one if the surveyed individual devoted any labor days to out-migration work during a year. Individual control variables include age, a dummy variable of marital status, and education level. Economic control variables include county and village GDP per capita (logged), shares of manufacturing and tertiary industries, and population size (logged). All specifications control for individual fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

in non-agriculture, as discussed in Section 2—we extract the RFPS subsample in which married couples can be reliably identified.¹⁹ We consider four outcomes for households' employment: joint non-farm employment, husband-only non-farm employment, wife-only non-farm employment, and no non-farm employment, along with the analogous migration outcomes: joint migration, husband-only migration, wife-only migration, and no migration.

We use both multinomial and binary logistic regressions to examine the effect of reform on the likelihood of different non-farm employment and migration patterns, relative to no non-farm employment or no migration. In all specifications, we add the same regional control variables as in Table 2, and the micro-level control variables including both spouses' education and age, as well as couple and year fixed effects.

Table 5 reports the estimation results. Columns (1) and (2) report outcomes for non-farm employment and migration, respectively. Each of the three main rows compares one type of non-farm employment (migration) outcome relative to both spouses staying in farming (nei-

¹⁹RFPS does not provide identifiers for all married couples within each household, but we can still identify household heads and their spouses using information on the relationship with the household head.

ther migrating). The results indicate that land reform significantly increases the likelihood of “husband only” and “both” engaging in non-farm employment (migration) compared to the base category. Notably, the coefficient is larger for the “both” category, suggesting a stronger reform effect on joint couple mobility. This finding is further corroborated by the binary logit estimations in columns (3) and (4), which directly compare “both” against “husband only” cases, confirming that reforms increase the relative likelihood of joint participation versus husband-only participation in non-farm employment and migration.

Table 5: Impact of Reforms on Rural Married Couples

	Multinomial Logistic Regression		Logit Regression	
	Non-Farm	Migration	Non-Farm	Migration
	(1)	(2)	(3)	(4)
Wife only Non-Farm/Migration	vs. No Non-Farm/Migration			
	0.067	0.085		
	(0.067)	(0.065)		
Husband only Non-Farm/Migration	vs. No Non-Farm/Migration			
	0.077**	0.121***		
	(0.033)	(0.036)		
Both Non-Farm/Migration	vs. No Non-Farm/Migration		vs. Husband only Non-Farm/Migration	
	0.106***	0.151***	0.086***	0.192***
	(0.041)	(0.050)	(0.029)	(0.029)
Control Variables	Yes	Yes	Yes	Yes
Couple FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	107296	105016	37985	64820

Notes: RFPS 2003–2017, pairs of married couples aged 18–55 with agricultural *hukou*. Columns (1) and (2) use multinomial logistic regression to compare Wife only Non-Farm/Migration, Husband only Non-Farm/Migration, and Both Non-Farm/Migration versus both in farming/neither migrating. Columns (3) and (4) use logit regression to compare Both non-farm/migration versus the husband only non-farm/migration. Individual control variables include the husband’s age, the wife’s age, the husband’s education level, and the wife’s education level. Economic control variables include county and village GDP per capita (logged), shares of manufacturing and tertiary industries, and population size (logged). Policy control variables include the weighted non-land-related *hukou* reforms (Gai et al., 2025), along with a dummy variable indicating whether a county adopted the New Rural Pension Reform. All specifications control for couple fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Robustness Checks on Individual Outcomes—The results discussed so far rely on RFPS data from 2003 to 2017. Next, we assess robustness using two additional data sources: the rural agricultural *hukou* samples from CFPS and CHIP. The CFPS, a panel dataset covering 2010 to 2020, offers more balanced sub-samples across various education levels compared to the RFPS,²⁰ whereas the CHIP is a nationally representative survey available for the years 1995, 2002, 2007, 2013, and 2018. Given the differences in data structure, we

²⁰In RFPS, the sample of rural individuals with a college education and above accounts for only 2.65% of the total observations, as opposed to the 7.20% in CFPS. Further details about CFPS can be found in Appendix A1 D.

include year and individual fixed effects when using the CFPS, and year and county fixed effects when using the CHIP, while maintaining the same set of controls as in Table 2.

We first apply the specification in Equation 1 to the full rural samples from CFPS and CHIP. The results are reported in Panel A and Panel B of the online Appendix Table B7, respectively. Consistent with the findings in Table 2, land reforms encourage rural *hukou* individuals to transition to non-farm employment, with a larger impact on women than men (column (1) in Panel A and column (7) in Panel B). Similarly, different impacts by gender are also observed for migration (column (4) in Panel A and column (10) in Panel B). In the remaining columns, we divide the samples by education levels and run the same regression on lower- and higher-educated sub-samples. Reassuringly, these heterogeneity tests confirm the earlier pattern: land reforms disproportionately benefit less educated rural women.

3.3 Impact on Urban Households

To examine the spillover effects of reform on urban non-agricultural *hukou* holders, we use the urban sample from six CHIP waves (1995, 1999, 2002, 2007, 2013, and 2018). As with the rural sample, we restrict to individuals aged 18 to 55 who self-report as physically healthy. The specification is as follows:

$$Emp_{jdt} = \theta + \gamma_1 Index_{dt} + \gamma_2 Index_{dt} * Female_j + \gamma_3 Female_j + X_{jt}\Phi + \sigma_d + \pi_t + X_{dt}\Gamma + v_{jdt}, \quad (2)$$

where Emp_{jdt} indicates whether non-agricultural *hukou* holder j in prefecture d in year t is currently employed in a paid, full- or part-time job. It is important to note that agricultural *hukou* holders who migrate to work and reside in prefecture d are excluded from the urban sample. We also estimate the same regression using the reported log annual income as the outcome, in place of Emp_{jdt} .

$Index_{dt} = \sum_{o=1}^O \frac{MigInflow_{od} * LandReform_{ot}}{MigInflow_d}$ represents the weighted land reform index for destination prefecture d in year t . This shift-share-like variable weights each origin county o 's land reform index, $LandReform_{ot}$, by a weight variable $\frac{MigInflow_{od}}{MigInflow_d}$, where $MigInflow_{od}$ denotes the migrant flow from origin o to destination d , and $MigInflow_d$ is the total migrants received at destination d . The migration flow variables are constructed using the 2000 Population Census, which predates the implementation of land reforms and thus helps to mitigate endogeneity concerns. Analogous to a shift-share instrumental variable, this measure captures the heterogeneous exposure of each destination prefecture to land reforms, with the share component serving as a pre-reform "intent-to-treat" measure. For ease of interpretation, we rescale the weighted land reform index to range from 0 to 1.

Similar to Equation 1, we interact the weighted land reform index with a female dummy.

To capture potentially different growth trajectories by initial inflow level, we control for the interaction between total migrants received at destination d and a yearly trend. Regional controls include prefecture-level log GDP per capita, the GDP share of secondary and tertiary industries, and log population size. Individual controls X_{dt} include age, cohabitation with partner, education level, work experience, and ethnicity. We also include prefecture and year fixed effects. Standard errors are clustered at the prefecture level to safeguard against potential heteroskedasticity.

Empirical Results— Columns (1) and (2) of Table 6 report results for employment status (=1 if employed in a paid, full or part-time job at the time of the survey) and annual logged wage income. The results show that when the weighted index rises from 0 to 1, land reforms lower the employment of urban non-agricultural hukou females by 2.8 percentage points and reduce their annual wage income by 5.1 percent. In contrast, no such effects are observed for urban males. Alternatively, we use total employment from the 2000 population census as the denominator to construct the weighted reform index. In this specification, we control for the interaction between the initial level of total employment in a prefecture and a yearly time trend to capture potentially different growth trajectories by total employment. The results reported in Table B9 in the Online Appendix are consistent with those of Table 6.

Table 6: Impacts of Reforms on Urban Employment

Dependent Variable	Urban Individual Sample CHIP (1995–2018)					
	Employed (=1)	Income (Logged)	Employed (=1)	Income (Logged)	Employed (=1)	Income (Logged)
	All		Low Education		High Education	
	(1)	(2)	(3)	(4)	(5)	(6)
Weighted Land Reform	-0.011 (0.021)	0.037 (0.046)	-0.017 (0.024)	0.096 (0.130)	-0.073 (0.105)	0.024 (0.042)
Weighted Land Reform * Female	-0.028*** (0.008)	-0.051** (0.026)	-0.042*** (0.009)	-0.071** (0.036)	0.070 (0.046)	-0.033 (0.118)
Control Variables	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	104175	74175	79370	51678	24805	22497

Notes: This table is based on urban samples from CHIP (1995–2018), including individuals aged 18–55 with non-agricultural *hukou*. The dependent variables are employment status, where 1 denotes employment in a full/part-time paid job at the time of the survey (columns 1, 3, and 5) and logged annual wage income (columns 2, 4, and 6). Control variables include gender, age, a dummy indicator for cohabitation with a partner, education level, work experience, ethnicity, city GDP per capita, shares of secondary and tertiary industries, and population size (logged), and the interaction between total migrants received at destination d and a yearly trend. All specifications control for prefecture fixed effects and year fixed effects. Standard errors clustered at the prefecture level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

The negative impact of land reforms on urban female employment may reflect a substitution effect from the influx of rural female migrants. If so, urban women with lower education would be expected to bear a larger impact. Table 6 further reports regression results for lower-education (columns (3) and (4)) and higher-education (columns (5) and (6)) subsamples. Rural land reforms have a significant negative effect on urban women with a high school education or below, consistent with the fact that the vast majority of the rural population and migrants have no more than a high school education.²¹ Taken together, the results support the interpretation of a substitution effect.

4 Model

This section develops a multisector model with household-level employment decisions to understand how a gender-neutral land reform can generate gender-specific outcomes. The model has two sets of regions: an urban region hosting the non-agricultural sector (denoted by subscript u) and a continuum of rural regions hosting the agricultural sector. There is a continuum of households with measure 1, each consisting of one female and one male member. A fraction α^u of households that originated in the urban region have a non-agricultural *hukou*, and a fraction $(1 - \alpha^u)$ that originated in the rural regions have an agricultural *hukou*.²² To focus on the mobility barriers created by land policy for agricultural *hukou* holders, we make two assumptions. The goods markets are frictionless and all consumers face the same price. A representative household head makes consumption and employment decisions on behalf of all non-agricultural *hukou* households, who only work in the non-agricultural sector.²³

Households with agricultural *hukou* are allocated land in their *hukou* region by the government rent-free. They can either work in the agricultural sector by cultivating this land or move to work in the non-agricultural sector. Each household decides the sector of employment for both members. The sectoral choice implies that there will be four types of agricultural *hukou* households: both members remain in the agricultural sector (type o), the female member works in the non-agricultural sector (type f), the male member works in the non-agricultural sector (type m), and both members work in the non-agricultural sector (type b). Let α^i , where $i = o, f, m, b$, represent the equilibrium share of these four types of

²¹Among rural individuals aged 18 to 55 in the RFPS, only 2.65% have a college education or above.

²²Both household members are assumed to hold the same type of *hukou* as mixed-*hukou* couples are rare according to census data.

²³This is consistent with the fact that the non-agriculture *hukou* holders do not have land-use right, and in line with the empirical pattern observed in China.

households:

$$\alpha^o + \alpha^f + \alpha^m + \alpha^b = 1 - \alpha^u \quad (3)$$

Land policy plays a crucial role in the sectoral choices of agricultural *hukou* households and hence, affects the speed of structural transformation. It determines whether they are able to receive rental income, and how much, when they move away from farming. Let superscript $r = 1$ denote rural regions that have carried out land reform, and $r = 0$ denote those without reform. We next formalize this potential rental income through a parameter λ^{ir} , which varies by type of agricultural household ($i = o, f, m, b$) and by two types of rural region r ($r = 1, 0$).

4.1 Land Policy and Production

The production of agricultural goods is region specific. Rural regions are ex-ante identical except in terms of reform status. The production of agricultural goods in a rural region r is:

$$Y_a^r = A_a (N_a^r)^{1-\beta} K^\beta; \quad N_a^r \equiv \left[\xi_a (N_{fa}^r)^{\frac{\eta-1}{\eta}} + (1 - \xi_a) (N_{ma}^r)^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}}. \quad (4)$$

where K denotes land, while N_{fa}^r and N_{ma}^r denote the total female and male work hours in the agricultural sector in region r . In equilibrium, the agricultural hours will depend on whether the region has implemented reform. The wage is equal to the value of the marginal product of the worker:

$$w_{fa}^r = (1 - \beta) p_a A_a \xi_a \left(\frac{K}{N_a^r} \right)^\beta \left(\frac{N_a^r}{N_{fa}^r} \right)^{\frac{1}{\eta}} \quad (5)$$

$$w_{ma}^r = (1 - \beta) p_a A_a (1 - \xi_a) \left(\frac{K}{N_a^r} \right)^\beta \left(\frac{N_a^r}{N_{ma}^r} \right)^{\frac{1}{\eta}} \quad (6)$$

The total land income $\beta p_a Y_a^r$ is allocated to households. The land policy parameter λ^{ir} determines the land income received by household type i in region r . Specifically, type i household receive $\lambda^{ir} I_k^r$:

$$0 \leq \lambda^{br} \leq \lambda^{fr} = \lambda^{mr} \leq \lambda^{or} = 1; \quad I_k^r \equiv \frac{\beta p_a Y_a^r}{\sum_{i=b,f,m,o} \alpha^{ir} \lambda^{ir}}, \quad (7)$$

where I_k^r is the land income received by household with both members staying in agriculture (type o household) and $\lambda^{ir} < 1$ captures the case that land is reallocated away from household

i if one or both member leave agriculture. We will discuss how we set the values of λ^{ir} in subsection 5.1.

Since rural regions only differ by reform status, the total agricultural output can be written as:

$$Y_a = \phi Y_a^1 + (1 - \phi) Y_a^0 \quad (8)$$

where ϕ denotes the share of regions that have experienced reform.

The production of the non-agricultural goods uses only labor:

$$Y_u = A_u N_u; \quad N_u \equiv \left[\xi_u (N_{fu})^{\frac{\eta-1}{\eta}} + (1 - \xi_u) (N_{mu})^{\frac{\eta-1}{\eta}} \right]^{\frac{\eta}{\eta-1}}. \quad (9)$$

where N_{fu} and N_{mu} denote the quantity of female and male labor used in the non-agricultural sector, respectively. The wage is equal to their corresponding value of the marginal product of labor for each gender:

$$w_{fu} = p_u A_u \xi_u \left(\frac{N_u}{N_{fu}} \right)^{\frac{1}{\eta}}, \quad w_{mu} = p_u A_u (1 - \xi_u) \left(\frac{N_u}{N_{mu}} \right)^{\frac{1}{\eta}}. \quad (10)$$

4.2 Households

Agricultural *Hukou* Households Depending on their sectoral employment choices i , agricultural *hukou* households in region r have the following joint utility function:

$$U^{ir} = \ln \left[(c_a^{ir} - \bar{c})^\omega (c_u^{ir})^{(1-\omega)} \right] + \theta \ln H^{ir} + \ln v^i; \quad i = o, f, m, b, \quad (11)$$

where $\bar{c}$ is the subsistence level of agricultural goods, c_a^{ir} and c_u^{ir} represent the consumption of agricultural and non-agricultural goods purchased from the market, H^{ir} is the non-market goods produced at home, and v^i represents household preferences for becoming different household types.

Each individual member has one unit of time to allocate between market work n_g^{ir} and non-market activities h_g^{ir} :

$$n_g^{ir} + h_g^{ir} = 1; \quad i = o, f, m, b; \quad g = f, m. \quad (12)$$

The non-market hours include both home production hours and leisure hours. Depending

on the type of household, the non-market good H^i is produced as follows:

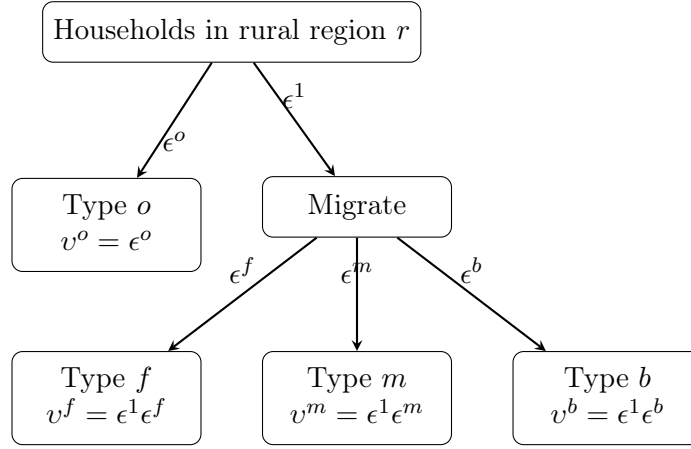
$$H^{ir} = \delta_f^i \delta_s^i (h_f^{ir})^{\xi_h} (h_m^{ir})^{1-\xi_h}; \quad i = o, f, m, b. \quad (13)$$

$$\delta_f^i = \begin{cases} \delta_f \leq 1 & \text{if } i = f, b \\ 1 & \text{if } i = o, m \end{cases}; \quad \delta_s^i = \begin{cases} \delta_s \leq 1 & \text{if } i = f, m \\ 1 & \text{if } i = o, b. \end{cases}$$

where $\delta_f \leq 1$ implies a lower efficiency in home sector when a female member move to non-agriculture,²⁴ and $\delta_s \leq 1$ captures the disutility when couples split into different sectors. These parameters reflect social norms and attitudes toward women's roles in home production and perceptions of split couples.

Households make sectoral employment decisions based on idiosyncratic preferences in two stages. First, each household draws preferences for migration (ϵ^1) and non-migration (ϵ^o), and decides if at least one household member will migrate to work in the non-agricultural sector. In the second stage, conditional on migration, they draw ϵ^f , ϵ^m , and ϵ^b , and to decide which member(s) to migrate. Figure 3 illustrates the decision tree and specifies the corresponding values of v^i .

Figure 3: Agricultural Household Employment Decision Tree



Note: ϵ^1 and ϵ^o follow independent Frechet distributions with shape parameter κ_1 . ϵ^f , ϵ^m , and ϵ^b follow independent Frechet distributions with shape parameter κ_2 .

Given the employment choice, the budget constraint for the consumption decisions of

²⁴CHIP survey of rural married aged 18–55 shows that women cite “looking after seniors/children” as a key reasons for not migrating in 2008, 2013, and 2018 (about 28% of women vs. 10% of men). Online Appendix Table B8 reports regression evidence on the determinants of non-farm employment and migration for females and males, respectively.

each type of households are:

$$p_a c_a^{ir} + p_u c_u^{ir} \leq I^{ir}; \quad i = o, f, m, b. \quad (14)$$

Members remaining in the agriculture sector choose the hours of work above a minimum number of hours to demonstrate active land use for protecting against land reallocation: $n_g^{ir} \geq n_{min}, i \neq b$. Those who move to the non-agricultural sector work a fixed number of hours ($\bar{n}$), and earn a fraction μ of the wage received by the urban worker. This wage discount reflects evidence from the literature that migrants often face discrimination or possess skills that are not well suited to the requirements of urban sectors (Meng and Zhang, 2001). Moreover, as shown by Ngai et al. (2019), another important aspect of *hukou* system is that migrants are not qualified to receive local public services or subsidies such as health, education and housing, which reduces their effective wages. In other words, the parameter μ captures other labor market barriers that are beyond the scope of this paper.

The total household income (I^{ir}) for household type i in region r is composed of both labor and land income, which depends crucially on the land policy:

$$I^{or} = w_{fa}^r n_f^{or} + w_{ma}^r n_m^{or} + \lambda^{or} I_k^r, \quad (15)$$

$$I^{fr} = \mu w_{fu} n_f^{fr} + w_{ma}^r n_m^{fr} + \lambda^{fr} I_k^r, \quad (16)$$

$$I^{mr} = w_{fa}^r n_f^{mr} + \mu w_{mu} n_m^{mr} + \lambda^{mr} I_k^r, \quad (17)$$

$$I^{br} = \mu w_{fu} n_f^{br} + \mu w_{mu} n_m^{br} + \lambda^{br} I_k^r. \quad (18)$$

Non-Agricultural *Hukou* Households The utility is given by

$$U^u = \ln \left[(c_a^u - \bar{c})^\omega (c_u^u)^{(1-\omega)} \right] + \theta \ln H^u; \quad H^u = (h_f^u)^{\xi_h} (h_m^u)^{1-\xi_h}. \quad (19)$$

The representative household head chooses the employment rates for the female and male members, E_f and E_m , respectively. Given that non-agricultural work involves fixed hours $\bar{n}$, the time constraint implies:

$$h_f^u = 1 - E_f \bar{n}, \quad h_m^u = 1 - E_m \bar{n}, \quad (20)$$

and the budget constraint for the urban household is:

$$p_a c_a^u + p_u c_u^u \leq I^u = w_{fu} \bar{n} E_f + w_{mu} \bar{n} E_m. \quad (21)$$

4.3 Equilibrium

The labor market clearing conditions by gender and sector satisfy:

$$\alpha^i = \phi \alpha^{i1} + (1 - \phi) \alpha^{i0}, \quad i = o, f, m, b \quad (22)$$

$$N_{fa}^r = \alpha^{or} n_f^{or} + \alpha^{mr} n_f^{mr}, \quad (23)$$

$$N_{ma}^r = \alpha^{or} n_m^{or} + \alpha^{fr} n_m^{fr}, \quad (24)$$

$$N_{fu} = (\alpha^b + \alpha^f) \bar{n} + \alpha^u \bar{n} E_f, \quad (25)$$

$$N_{mu} = (\alpha^b + \alpha^m) \bar{n} + \alpha^u \bar{n} E_m. \quad (26)$$

The full derivation of the model is presented in Online Appendix B1, where we solve for α^{ir} , assuming that ϵ^1 and ϵ^o have independent Fréchet distributions with parameter κ_1 , and that ϵ^f , ϵ^m and ϵ^b have independent Fréchet distributions with parameter κ_2 .

5 Quantitative Results

This section calibrates the model using the land reform index constructed in Section 2 and targeting observed household employment patterns in 2000 and 2020. We then simulate a no-reform counterfactual scenario to quantify its gender-specific effects.

5.1 Calibration

We begin by setting the land policy parameters for reform and non-reform regions. In reform regions, households receive the same land income irrespective of type, i.e. $\lambda^{i1} = 1, \forall i = o, f, m, b$. In non-reform regions, for households with both members working in agriculture, λ^{o0} is set to 1. For households with both members working in non-agriculture, we set $\lambda^{b0} = 0$. For households with one member in agriculture, the remaining member can protect the land income of the household, i.e., $\lambda^{f0} = \lambda^{m0} = 1$. This is our baseline case. We will also explore the alternative case of “no-guarding” where the remaining member can only keep his/her share of land.

The share of rural regions that have experienced reform ϕ is set to the national average of the land reform index divided by its maximum value of 3. As shown in Figure 1, the diffusion of land reform across counties implies an increasing ϕ over time, reaching close to 1 in 2020 when almost all counties have completed reform. The values of ϕ are summarized in Table 8.

The growth rate of A_u , denoted as g_u , is directly set using sectoral employment and real

Table 7: Baseline Land Policy Parameters

	Non-reform	Reform
	$r = 0$	$r = 1$
$\lambda^{or} = \lambda^{fr} = \lambda^{mr}$	1	1
λ^{br}	0	1

Notes: This is the case where leaving one partner in agriculture can guard full land income share.

Table 8: Share of Reform Regions

	2000	2005	2010	2015	2020
ϕ	0	0.212	0.365	0.591	0.997

Notes: ϕ is calculated as the national average of the reform index divided by 3.

output data from the China Statistical Yearbook (CSY). The growth rate of A_a , denoted as g_a , is endogenously calibrated so that the model-implied growth rate of labor productivity in agriculture matches its empirical counterpart calculated from the CSY. We normalize A_a and A_u in 2000 to be 1.

We now turn to other parameters that are directly set and do not vary over time. The share of non-agricultural *hukou* α^u is set to be 0.27, calculated from the 2000 Census. The elasticity of substitution between female and male workers, η , is set to 2, a common value in the literature (Ngai and Petrongolo, 2017). We normalize the total endowment of land, K , to 1. The land share, β , is set to 0.5 to match the reported land income share in China (Cao and Birchenall, 2013). The number of hours worked per employed individual in the non-agricultural sector is set at 48 hours per week, i.e., $\bar{n} = 0.48$, based on the average hours in the urban sector reported by the National Bureau of Statistics (NBS) for 2022 and 2023.²⁵ The minimum hours required in agriculture is set to five hour per week, i.e. $n_{min} = 0.05$.²⁶

We set κ_1 , the migration elasticity, to 1.4. In a setting with three layers of decisions, Imbert et al. (2025) estimate a value close to 1.4 for the decision to migrate at the household level. It is close to the common value (1.5) seen in the literature that studies migration in China.²⁷ We set κ_2 , capturing the elasticity between different modes of migration, to be 2.5 in the baseline, and experiment with values from 2 to 4 in Table A5.²⁸

²⁵The NBS began reporting hours for non-agricultural workers from January 2022 onwards.

²⁶We report results for $n_{min} = 0.01$ and $n_{min} = 0.10$ in Table A7.

²⁷For example, Tombe and Zhu (2019) uses a migration elasticity of 1.5 at the individual level.

²⁸Imbert et al. (2025) provide the closest structural analog in the literature, estimating, in a layered setting, an elasticity of 2.72 to 4.2 between different modes of migration (whether to bring dependent family members such as children or elderly parents), conditional on migration. Our baseline value of 2.5 is slightly below their estimated range, making our main results conservative with respect to mode-switching sensitivity. While a

The rest of the parameters are internally calibrated. The subsistence consumption, $\bar{c}$, governs the speed of decline in agricultural employment. We start by assuming a value for $\bar{c}$ for 2000, calibrate the remaining parameters, then simulate the model forward, and finally adjust the value of $\bar{c}$ to match the 2020 share of households with both members in agriculture ($\alpha^o/(1 - \alpha^u)$).

We now explain the calibration of parameters to match the data moments for the year 2000. The gender parameter ξ_u is calibrated to match the relative female-to-male wage ratio in the non-agricultural sector (0.84), ξ^h to match the employment rate of urban females relative to urban males (0.78), and θ to match the average urban employment rate (0.70). These moments on urban populations are calculated from Urban Household Survey (UHS) 2000. The parameter ξ_a is set to match the ratio of days worked in agriculture by female workers relative to male workers (0.93), based on data from the 2003 RFPS.

Using the observed employment pattern in 2000, we calibrate ω , δ_f , and δ_s to match the share of households with both members working in agriculture ($\alpha^o/(1 - \alpha^u) = 0.76$), the relative share of households with female working in non-agriculture ($(\alpha^f + \alpha^b)/(\alpha^f + \alpha^m + \alpha^b) = 0.46$), and the relative share of split households ($(\alpha^f + \alpha^m)/(\alpha^f + \alpha^m + \alpha^b) = 0.63$).²⁹ The exogenous labor market wedge μ is set to match the agricultural productivity gap (APG) in 2000 (5.81), which is calculated using CSY data on real output and employment by sector.³⁰

With the above parameters calibrated to match 2000, we then simulate the model at five-year intervals from 2005 to 2020, allowing δ_s , δ_f and μ to vary.³¹ Together, they allow the model to match APG, the relative share of households with the female working in non-agriculture ($(\alpha^f + \alpha^b)/(\alpha^f + \alpha^m + \alpha^b)$), and the relative share of split households ($(\alpha^f + \alpha^m)/(\alpha^f + \alpha^m + \alpha^b)$). Since $\bar{c}$ matches the share of non-migrating households ($\alpha^o/(1 - \alpha^u)$) in 2020, the model reproduces the exact employment shares in 2020 and the relative employment shares in 2005, 2010, and 2015.

The baseline is referred as the *Reform* scenario where the share of reform regions gradually increases from 0 in 2000 to close to 1 in 2020 as in Table 8, with the other parameter values

direct estimate of the elasticity corresponding to κ_2 is not available, we experiment with a range of values in Appendix Table A5.

²⁹Since the shares of different *hukou* types must satisfy Equation 3, and α^u is given exogenously, the three targeted moments pin down the shares of the four agricultural *hukou* household types.

³⁰APG is defined as the value added per worker in non-agriculture relative to that of in agriculture (Gollin et al., 2014). In the model, it is calculated as $(p_u y_u)/(p_a y_a)$, where y_u and y_a denote output per worker in non-agriculture and agriculture.

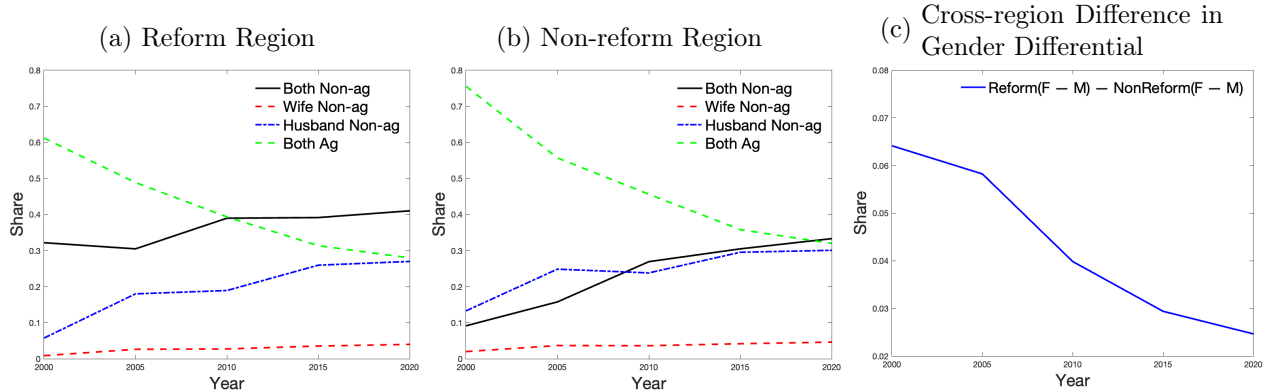
³¹Changes in δ_s and δ_f reflect potential changes in social norms, as well as technological advancements that facilitate long-distance communication and reduce the dependence on female home production, for example, through household durables. Changes in μ capture other labor market wedges outside the scope of our model.

summarized in Table 7, Appendix Figure A2, and Table A4. The employment pattern of households with agricultural *hukou* in the data and the baseline reform scenario are reported in Figure A3 Panel (a) and Panel (b). Although the model only targets the employment pattern in 2000 and 2020 and the relative employment shares in other years, the overall employment pattern also matches data in other years quite well.

5.2 Comparing Reform and Non-reform Regions

We begin by examining the predictions of the model against the empirical results across reform and non-reform regions, which are untargeted moments since the model is calibrated to aggregate employment patterns. Figure 4 plots the predicted employment patterns for reform and non-reform regions hidden behind the aggregate pattern in Figure A3 Panel (b). Panel (c) of Figure 4 turns to comparing the predicted gender difference in non-agriculture employment across reform and non-reform regions. It shows that female non-agriculture employment share relative to male is 6.4 percentage points higher in reform regions relative to non-reform regions in 2000 and the gap declines to 2.5 percentage points in 2020 as more regions undergo land reforms.

Figure 4: Employment Pattern by Region, Agricultural *Hukou* Households



Notes: Panels (a) and (b) plot the employment patterns of agricultural *hukou* households in reform and non-reform regions in the baseline reform case, represented by $\frac{\alpha^{ir}}{1-\alpha^u}$, $i = o, f, m, b$, where the four shares sum to one. Panel (c) plots the cross-region difference in gender differential (female minus male non-agricultural employment share), i.e., $\frac{\alpha^{f1}-\alpha^{m1}}{1-\alpha^u} - \frac{\alpha^{f0}-\alpha^{m0}}{1-\alpha^u}$.

The results reported in Panel (c) are related to the regression results in Table 2, which pool the impact of reforms across years 2003 to 2017. To compare the model's predictions with the empirical results, we calculate an average of the model's predictions across all periods, weighted by the reform region share.³² This captures the fact that later periods have

³²We calculate $\frac{\sum_t x_t \phi_t}{\sum_t \phi_t}$, where x_t denotes our variable of interest and ϕ_t denotes share of reform regions

more reform regions in the empirical sample, thus carrying larger weights in the estimated coefficient. With this approach, the weighted average of the cross-region difference in female non-agricultural employment relative to male is 3.2 percentage points in our baseline, which lies slightly above the empirical estimate of 2.7 percentage points (0.009×3). The elasticity κ_2 plays an important role. Our baseline value of $\kappa_2 = 2.5$ is close to the lower end of the range estimated by Imbert et al. (2025). As no direct estimate of this parameter is available for our setting, we examine sensitivity to alternative values of κ_2 in Appendix Table A5. The model-implied cross-region difference in gender-differential effect is increasing in κ_2 : at $\kappa_2 = 2$ the model produces an effect of 2.4 percentage points, slightly below our empirical reference point of 2.7 percentage points. The value of κ_2 that exactly matches the empirical moment lies between 2 and 2.5, suggesting that our baseline choice is broadly consistent with the data.

Beyond the gender-differential effect, we further check the model’s predictions on three additional untargeted moments: the reform-induced increase in labor productivity, land per worker, and hours per worker, presented in Table 9. Column (1) summarizes the empirical estimates from Table A8, with the caveat that RFPS reports farm days rather than hours. Column (2) presents the weighted average of the model predictions across years. The model’s prediction on hours (11.1%) closely aligns with the empirics on farm days (11.4%). For land per worker, the model overpredicts the increase: land freed by departing workers is absorbed by remaining household farmers in the model, but in practice, it is also transferred to the non-household sector, such as rural cooperatives and agricultural firms, so the model’s prediction exceeds what household data can capture. For labor productivity, the overprediction of land per worker tends to overstate productivity, while omitting endogenous productivity gains, such as land reallocation toward more productive farmers and increases in investment due to more secure land rights, tends to understate it.³³ These forces work in opposite directions, which likely explains why the model’s prediction of a 12.6% increase in labor productivity sits close to the empirical estimate of 14.4%.

5.3 Aggregate Effects of Land Reform

With the model’s cross-region predictions validated against the empirical results, we turn to the aggregate effects of land reform. We conduct a counterfactual no-reform scenario that sets the share of reform regions to zero throughout 2000–2020. Figure A3 Panel (c) plots the no-reform scenario against the baseline reform scenario in Panel (b). Comparing

at time t .

³³See Chari et al. (2021) and Adamopoulos et al. (2022) for the effects on agricultural productivity due to improved land and labor allocation across farmers.

Table 9: Reform Regions vs Non-reform Regions

	Empirics	Model Baseline
Agricultural <i>hukou</i>		
Non-ag emp. share, female – male	+2.7 pp	+3.2 pp
Agricultural labor productivity	+14.4 %	+12.6 %
Land per worker	+6 %	+13.9 %
Hours per worker	+11.4 %	+11.1 %

Note: This table presents differences between reform and non-reform regions, averaged across years and weighted by the share of reform regions in the corresponding year.

the two panels, land reforms that improve land security have a stronger effect on women’s transition out of agriculture by increasing joint non-agricultural employment relative to male-only non-agricultural employment. Zooming in on 2020, reform increases rural female non-agricultural employment by 5.8 percentage points and male non-agricultural employment by 3.7 percentage points, with a gender differential of 2.1 percentage points. These numbers are summarized in the upper panel of Table 10 under column (1).³⁴

Table 10: The Uneven Impact of Reform vs No-reform

	(1) Baseline	(2) No-guarding
Reform Scenario vs No-reform Scenario, 2020		
Agricultural <i>hukou</i>		
Non-ag emp. share, male	+3.7 pp	+6.8 pp
Non-ag emp. share, female – male	+2.1 pp	–0.6 pp
Non-agricultural <i>hukou</i>		
Emp. rate, female – male	–2.8 pp	–1.4 pp
Wage ratio, female/male	–1.9 %	–1.0 %

Note: Column (1) reports the model predictions in the baseline case. Column (2) reports the case where the remaining member can only keep his/her own share of land.

Role of Guard Labor In our baseline, the member staying behind can protect the land income of the household. To examine the role of “guard labor” in generating the gender-specific impact, we consider a “no-guarding” alternative, where the remaining member can only keep her/his own land share. To reflect this, we set $\lambda^{fr} = \lambda^{mr}$ to 0.5 (see Table 11). The

³⁴Our baseline sets $\kappa_2 = 2.5$. Appendix Table A5 reports the 2020 difference for three alternative cases, $\kappa_2 = 2$ and $\kappa_2 = 3$, and $\kappa_2 = 4$. Overall, a larger κ_2 generates more migration and enlarges the differential impact by gender.

other parameters in the “no-guarding” case are calibrated using the same strategy outlined in subsection 5.1 to match the observed employment patterns. We then repeat the same no-reform counterfactual exercise for this “no-guarding” case. The results are reported in column (2) of Table 10.

Table 11: Land Policy Parameters: No-guarding Case

	Non-reform	Reform
	$r = 0$	$r = 1$
λ^{or}	1	1
$\lambda^{fr} = \lambda^{mr}$	0.5	1
λ^{br}	0	1

Notes: This is the case where leaving one partner in agriculture does not help guard the land income share of the migrating member.

Relative to the baseline, the reform in the “no-guarding” case generates more employment reallocation from agriculture to non-agriculture, because it affects all migrating households, while in the baseline it only affects type b . In contrast to the baseline, however, the reform in this case shifts fewer women out of agriculture compared to men (by -0.6 percentage points), because land frictions do not hold women back in the first place. Intuitively, in both the baseline and the “no-guarding” cases, women remain disproportionately in agriculture because the non-agricultural sector is less female-intensive ($\xi_u < \xi_a$), and when women exit agriculture, the production of non-market goods becomes less efficient ($\delta_f < 1$). The difference between the two cases is that having one partner in agriculture helps guard land income in the baseline case. The sectoral gender intensities and women’s role in home production imply that women are more likely to take on this role. Reform freeing women from this role thus has a stronger effect on women’s transition out of agriculture. In the “no-guarding” case, this mechanism is absent and the reform turns out to increase slightly more men to leave agriculture.

Urban Population The reform-induced labor re-allocation increases the relative supply of female workers in the urban non-agricultural sector, which lowers the female wage rate relative to male in non-agriculture. Urban households respond by decreasing the female employment rate relative to males and by increasing the share of females specializing in non-market production. The lower panel of Table 10 summarizes the model’s predictions. Column (1) shows that the urban female employment rate decreases by 2.8 percentage points relative to men in 2020, and the gender wage ratio decreases by 1.9%.³⁵

³⁵The elasticity of substitution between women and men (η) is relevant for the predictions on urban household. Our baseline sets $\eta = 2$, close to the empirical estimate of 1.92 based on firm-level data from

It is worth noting that, under the no-guarding case, the large increase in female migrants ($6.8 - 0.6 = 6.2$ pp) still generates a decrease in the employment rate of urban women relative to men and a lower gender wage ratio, because the wage ratio depends on the relative supply of total female workers to total male workers in the non-agricultural sector. However, the magnitude of the gender-differential impact is much smaller compared to the baseline.

Agricultural Productivity Gap We conclude the quantitative exercise by reporting the effect on APG. With reform, the APG decreases from 5.81 in 2000 to 3.73 in 2020. Without reform, it goes to 4.37 in 2020. In other words, reform explains about 31% of the APG decline from 2000 to 2020 ($= \frac{4.37-3.73}{5.81-3.73}$). One main reason is that the reform scenario has higher labor productivity in agriculture. The channels that contribute to higher labor productivity in the cross-region comparison, higher land-per-worker ratio and higher hours worked, apply here as well.

6 Conclusion

Exploiting quasi-natural experiments from two major land reforms, we show that *de jure* gender-neutral land market institutions can operate as *de facto* gender-specific mobility barriers during structural transformation. A multisector model focusing on intra-household labor allocation reveals that dismantling these frictions is quantitatively important: land reform accounts for a significant share of women’s transition out of agriculture and the observed rise in joint household migration relative to husband-only migration. These findings speak to the broader consensus that land institutions are critical for fostering resilient and equitable growth (Bank, 2024).

China (Ostry et al., 2018). At the macro-level, the degree of substitution could differ. Appendix Table A6 presents the alternative cases where $\eta = 1.5$ and $\eta = 2.5$. A higher η implies that female and male workers are more substitutable in non-agriculture, hence the relative increase in female migrant workers generates a smaller differential impact by gender.

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Appendix

A1 Data

This section provides details on the data used in this paper.

A. Censuses and One Percent Population Surveys Our data on employment by sector and *hukou* status were obtained from the Chinese Population Censuses (1990, 2000, 2010, and 2020) and the One Percent Population Surveys (2005 and 2015). The microdata are nationally representative and offer detailed information on ethnicity, gender, age, employment status, occupation, industry, *hukou* type (rural agricultural or non-agricultural urban), current residence, marital status, and education across survey waves. We focus on the age group 18-55. As discussed in the text, Censuses 2015 and 2020 did not provide *hukou* type. We used whether an individual has rural land contracting rights to proxy for the de-facto *hukou* type.

For individuals reporting work status as employed, we group them into agricultural and non-agricultural sectors following the National Statistical Bureau’s sectoral classification system. For the individual-level regression on sectoral employment of agricultural-*hukou* holders (Appendix Table A1), we compile several control variables for ethnic minority status, age, marital status, and high skill (education at college and above). Summary statistics for these variables are reported in Online Appendix Table B1.

B. Chinese Household Income Project (CHIP) We use both rural and urban samples from six waves of the nationally representative CHIP surveys conducted between 1995 and 2018. Married households are defined as couples aged 18 to 55 and include various family structures, such as intergenerational arrangements. In the rural household questionnaires, respondents are also asked about time spent away from the household or hometown. Those who indicate “ ≥ 6 months (180 days)” are classified as migrants. Although the wording varies slightly across survey waves, the core objective remains the same. This classification allows for meaningful comparisons across households and over time.

The urban resident sample follows the same age and employment criteria as the rural sample and is used to assess how external land reform in migration-origin counties affects the urban labor market, particularly employment and wage outcomes for the urban population. Specifically, we select urban residents aged 18–55 who were not in school at the time of the survey and were healthy enough to work. Summary statistics for the relevant variables are presented in Online Appendix Table B2.

C. Rural Fixed Point Survey (RFPS) This is a longitudinal panel survey conducted by the Research Center of Rural Economy of the Chinese Ministry of Agriculture, providing micro-level data on rural household agricultural production. Initiated in 1986, the RFPS tracks a nationally representative sample of approximately 20,000 rural households across 300 villages in all 31 Chinese provinces.

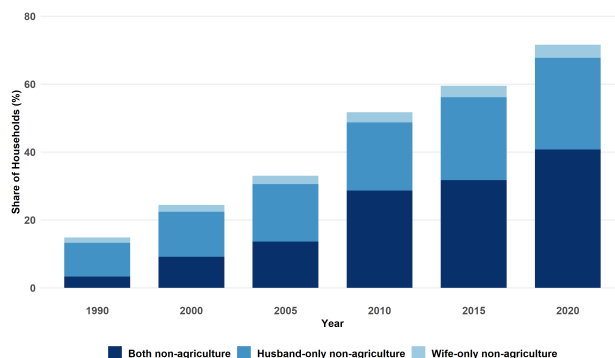
These villages were carefully selected to ensure representativeness across region, income, cropping patterns, population, and non-farm activities. Within each village, households were drawn by random sampling. Renowned for its data quality, RFPS has been used in influential studies such as Chari et al. (2021) and Adamopoulos et al. (2024). Our study uses RFPS data from 2003 to 2017 to examine how land reform affects rural residents' migration, sectoral employment, and intra-household migration pattern. Online Appendix Table B3 presents the summary statistics for the variables used in our analysis.

D. China Family Panel Studies (CFPS) A biennial longitudinal survey conducted by the Institute of Social Science Survey (ISSS) at Peking University. Launched in 2010, CFPS tracks a nationally representative sample of about 19,237 individuals from 8,582 households across 1,004 counties. The surveys span from 2010 to 2020, covering both rural and urban populations. In our analysis, we use CFPS data to test the robustness of our empirical results for the rural resident sample.

Using six waves of CFPS, we construct an individual panel for rural agricultural residents, detailing migration, employment, and income. The panel structure allows for two-way fixed effects estimation, controlling for unobserved, time-invariant individual heterogeneity and time-specific shocks. Summary statistics for the rural CFPS sample are reported in Online Appendix Table B4.

A2 Appendix Figures and Tables

Figure A1: Employment Pattern of Rural Married Couples



Notes: Census and one percent population survey. Pairs of rural husbands and wives aged 18–55 by sectoral employment. “Both non-agriculture” denotes that both the husband and wife are employed in non-agricultural sectors. “Husband-only non-agriculture” denotes that husbands work in non-agriculture and wives do not. “Wife-only non-agriculture” denotes that wives work in non-agriculture while husbands do not.

Table A1: Non-agricultural Employment of Agricultural-hukou Holders

Dependent Variable	Non-Agricultural Employment
Female	-0.101*** (0.006)
Individual Controls	Yes
Province Fixed Effects	Yes
Year Fixed Effects	Yes
# of Obs	13,140,026
adj. R-sq	0.201

Notes: This table is based on individual-level regression to examine the gender difference in non-agricultural employment. Controls include minority status, age, age squared, marital status, and a high skill indicator (education at the college level and above). Source: Census and One Percent Population Survey for 1990, 2000, 2005, 2010, 2015, and 2020. Sample: agricultural-*hukou* holders (aged 18–55). Standard errors in parentheses * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A2: Impact of Pre-Reform Land Insecurity on Labor Reallocation

	Rural Household (1995-2006)		Individual (2003-2006)			
	Non-farm (1)	Migration (2)	Non-farm (3)	Non-farm (4)	Migration (5)	Migration (6)
Land Redistribution	-0.008** (0.004)	-0.024** (0.010)	-0.014** (0.006)	0.006 (0.004)	-0.035** (0.017)	0.010 (0.018)
Land Redistribution*Female				-0.021** (0.010)		-0.048** (0.023)
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	308345	308345	125720	125720	125720	125720

Notes: Rural Fixed Point Survey 1995–2006. Columns (1) and (2) use the household sample for 1995–2006, while columns (3) to (6) use the individual sample (aged 18–55) for 2003–2006. Dependent variables: Column (1) is a dummy equal to 1 if any household member spent any labor days in non-farm employment during a year; Column (2) is a dummy equal to 1 if any household member spent any labor days in out-migration work during a year. Columns (3)–(4) are dummies equal to 1 if the surveyed individual spent any labor days in non-farm employment during a year; Columns (5)–(6) are dummies equal to 1 if the surveyed individual spent any labor days in out-migration work during a year. Land redistribution is defined as the reallocation of more than 30% of village land by village cadres, as recalled for the period from the 1970s through 2006, and sourced from a special one-off supplementary survey on village governance conducted by RFPS in 2006. Columns 1 to 2 control for household individual fixed effects and year fixed effects. Columns 3 to 6 control for individual fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A3: Impact of Reforms on Rural Households' Rental Activities

Dependent Variables	Rural Household Sample, RFPS (1995–2017)		
	Rental Out (=1) (1)	Rental Area (Logged) (2)	Rental Income (Logged) (3)
Land Reform Index	0.100*** (0.028)	0.129** (0.050)	0.097** (0.047)
Control Variables	Yes	Yes	Yes
Household FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Observations	409455	409455	409455

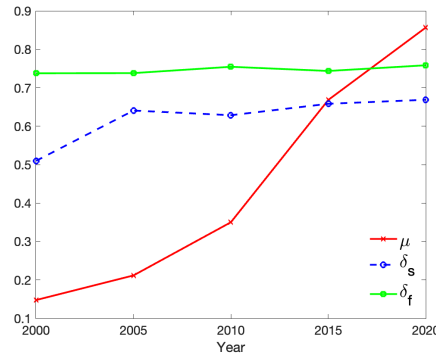
Notes: Rural Fixed Point Surveys 1995–2017. The dependent variables are at the household-level and include: 1) a dummy variable for renting out land that year; 2) the size of land rented out; and 3) the income from renting out land. All specifications control for household fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table A4: Baseline Parameters

Para.	Values	Moments
<i>Directly calibrated parameters</i>		
α_u	0.27	Non-agricultural hukou share (Census)
g_u	6.61%	Labor productivity growth based on employment and real GDP (CSY)
η	2	Ngai and Petrongolo (2017)
$\bar{n}$	0.48	48 hours per week (NBS)
n_{min}	0.05	5 hours per week required in agriculture
β	0.50	Land share from Cao and Birchenall (2013)
$\kappa_1; \kappa_2$	1.4; 2.5	Imbert et al. (2025)
<i>Internally calibrated parameters to match 2000</i>		
ξ_u	0.39	Female over male wage rate in non-agriculture (UHS)
ξ_h	0.47	Female over male employment rate; non-ag hukou (UHS)
θ	2.22	Employment rate; non-agricultural hukou holders (UHS)
ξ_a	0.47	Agricultural days per worker, female over male (RFPS 2003)
ω	0.09	Share of households with both members in agriculture (Census)
μ	0.15	Agricultural productivity gap (CSY)
δ_f	0.74	Relative share of households with female in non-agriculture (Census)
δ_s	0.51	Relative share of split households (Census)
<i>Time varying to match 2005, 2010, 2015, and 2020</i>		
μ	varying	Agricultural productivity gap (Census)
δ_f	varying	Share of households with female in non-agriculture (Census)
δ_s	varying	Share of split households (Census)
<i>Internally calibrated parameters to match 2020</i>		
g_a	6.8%	Labor productivity growth based on employment and real GDP (CSY)
$\bar{c}$	0.35	Share of households with both members in agriculture (Census)

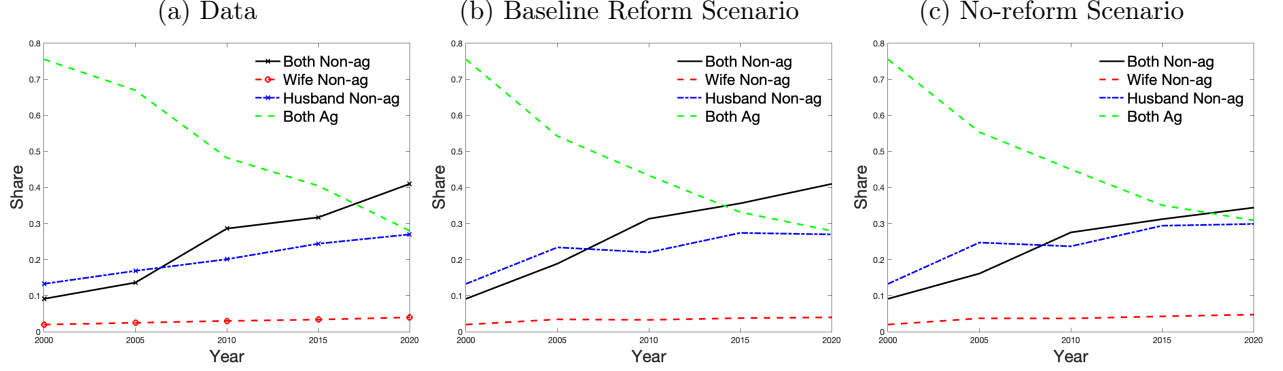
Note: This table reports the baseline parameters. The implied value of $\frac{\bar{c}}{Y_a}$ is 0.80 in 2020. Parameter values are rounded to two digits. See Figure A2 for the values of time-varying parameters.

Figure A2: Baseline Time-varying Parameters



Note: This figure displays time-varying parameters for the exogenous labor market wedge μ , the efficiency of non-market production when women work in the non-agricultural sector δ_f , as well as the efficiency of non-market production when couples split into different sectors δ_s .

Figure A3: Employment Pattern of Agricultural *Hukou* Households



Notes: This figure plots the employment pattern of households with agricultural *hukou*, represented by $\frac{\alpha^i}{1-\alpha^u}$, $i = o, f, m, b$, where the four shares sum to one. The reform case is calibrated to match the data pattern in the 2000 and 2020 Census. The counterfactual exercise “No-reform” refers to the case where the share of reform regions is fixed at 0 throughout 2000-2020.

Table A5: The Role of Migration Elasticity Between Household Type

	(1) $\kappa_2 = 2$	(2) Baseline $\kappa_2 = 2.5$	(3) $\kappa_2 = 3$	(4) $\kappa_2 = 4$
Reform Scenario vs No-reform Scenario, 2020				
Agricultural <i>hukou</i>				
Non-ag emp. share, male	+3.7 pp	+3.7 pp	+3.7 pp	+3.8 pp
Non-ag emp. share, female – male	+1.7 pp	+2.1 pp	+2.6 pp	+3.5 pp
Non-agricultural <i>hukou</i>				
Emp. rate, female – male	–2.4 pp	–2.8 pp	–3.2 pp	–4.0 pp
Wage ratio, female/male	–1.7 %	–1.9 %	–2.2 %	–2.8 %
Reform Region vs Non-reform Regions, weighted average				
Agricultural <i>hukou</i>				
Non-ag emp. share, female – male	+2.4 pp	+3.2 pp	+4.1 pp	+5.9 pp

Note: The baseline is column (1) of Table 10. For each κ_2 value, we recalibrate the model following the same calibration strategy described in subsection 5.1.

Table A6: The Role of the Elasticity of Substitution Between Women and Men

	(1) $\eta = 1.5$	(2) Baseline $\eta = 2$	(3) $\eta = 2.5$
Reform Scenario vs No-reform Scenario, 2020			
Agricultural <i>hukou</i>			
Non-ag emp. share, male	+3.7 pp	+3.7 pp	+3.6 pp
Non-ag emp. share, female – male	+2.0 pp	+2.1 pp	+2.2 pp
Non-agricultural <i>hukou</i>			
Employment rate, female – male	–3.3 pp	–2.8 pp	–2.4 pp
Wage ratio, female/male	–2.3 %	–1.9 %	–1.7 %
Reform Region vs Non-reform Regions, weighted average			
Agricultural <i>hukou</i>			
Non-ag emp. share, female – male	+3.2 pp	+3.2 pp	+3.2 pp

Note: The baseline is column (1) of Table 10. For each η value, we recalibrate the model following the same calibration strategy described in subsection 5.1.

Table A7: The Role of Minimum Hours in Agriculture

	(1) $n_{min} = 0.01$	(2) Baseline $n_{min} = 0.05$	(3) $n_{min} = 0.10$
Reform Scenario vs No-reform Scenario, 2020			
Agricultural <i>hukou</i>			
Non-ag emp. share, male	+3.6 pp	+3.7 pp	+3.8 pp
Non-ag emp. share, female – male	+2.1 pp	+2.1 pp	+2.2 pp
Non-agricultural <i>hukou</i>			
Employment rate, female – male	–2.7 pp	–2.8pp	–2.9 pp
Wage ratio, female/male	–1.9 %	–1.9 %	–2.0 %
Reform Region vs Non-reform Regions, weighted average			
Agricultural <i>hukou</i>			
Non-ag emp. share, female – male	+3.1 pp	+3.2 pp	+3.2 pp

Note: The baseline is column (1) of Table 10. For each n_{min} value, we recalibrate the model following the same calibration strategy described in subsection 5.1.

Table A8: Impact of Reforms on Agricultural Households' Labor Productivity, Farm Size, and Labor Days

	Labor Productivity (1)	Land per worker (2)	Labor Days per Worker (3)
Land Reform	0.048** (0.021)	0.020** (0.010)	0.038** (0.018)
Control Variables	Yes	Yes	Yes
Household Fixed Effects	Yes	Yes	Yes
Year Fixed Effects	Yes	Yes	Yes
Observations	139322	139322	139322

Notes: RFPS 2003–2017, households with agricultural *hukou*. The dependent variables in columns 1–3 are: 1) labor productivity measured as total value of agricultural output divided by the number of agricultural workers in the household (logged), 2) farming area per worker (logged) and 3) total labor days per worker in a year (logged). Household control variables include the share of male agricultural labor and dependency rate. Economic control variables include county and village GDP per capita (logged), shares of manufacturing and tertiary industries, and population size (logged). Policy control variables include the weighted non-land-related *hukou* reform indices (Gai et al., 2025), along with a dummy variable indicating whether a county adopted the New Rural Pension Reform. All specifications control for household fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Online Appendix–Not For Publication

B1 Model Derivations

This section derives the household consumption and employment decisions, taking wage rates and prices as given.

B1.1 Consumption Choices

The utility function implies the relative consumption:

$$\frac{\omega}{1-\omega} = \frac{p_a}{p_u} \frac{c_a^i - \bar{c}}{c_u^i}, \quad i = f, m, b, o, u \quad (27)$$

Together with the budget constraint, the consumption decisions are:

$$c_a^i - \bar{c} = \frac{\omega (I^i - p_a \bar{c})}{p_a}; \quad c_u^i = \frac{(1-\omega) (I^i - p_a \bar{c})}{p_u}, \quad (28)$$

which are independent of hours allocation. The utility achieved from consumption and non-market goods becomes:

$$\ln \left[(c_a^i - \bar{c})^\omega (c_u^i)^{(1-\omega)} \right] + \theta \ln H^i = \ln B + \ln (I^i - p_a \bar{c}) + \theta \ln H^i \quad (29)$$

where the first term, $B = \left(\frac{\omega}{p_a} \right)^\omega \left(\frac{1-\omega}{p_u} \right)^{1-\omega}$ is the same for all households.

B1.2 Agricultural *Hukou* Households

Rural households in region r solve the utility maximization problem for each of the four choices $i = o, f, m, b$ and select the one that yields the highest utility. They take wage rates (w_{fa}^r , w_{ma}^r , w_{fu} , and w_{mu}), land income (I_k^r), and its allocation rules (λ^{ir}) as given.

Type o Households Subject to the time constraint (Equation 12), a type o household in region r solves:

$$\max_{n_g^{or}, h_g^{or}} U^{or} = \ln B + \ln [w_{fa}^r n_f^{or} + w_{ma}^r n_m^{or} + \lambda^{or} I_k^r - p_a \bar{c}] + \theta \ln \left[(h_f^{or})^{\xi_h} (h_m^{or})^{1-\xi_h} \right] + \ln \epsilon^o$$

The first order conditions for interior solutions h_f^o and h_m^o are given by:

$$\frac{w_{fa}^r}{w_{fa}^r n_f^{or} + w_{ma}^r n_m^{or} + \lambda^{or} I_k^r - p_a \bar{c}} = \frac{\theta \xi_h}{h_f^{or}} \quad (30)$$

$$\frac{w_{ma}^r}{w_{fa}^r n_f^{or} + w_{ma}^r n_m^{or} + \lambda^{or} I_k^r - p_a \bar{c}} = \frac{\theta (1 - \xi_h)}{h_m^{or}} \quad (31)$$

They imply the following relative home hours:

$$\frac{w_{fa}^r}{w_{ma}^r} \frac{1 - \xi_h}{\xi_h} = \frac{h_m^{or}}{h_f^{or}} \quad (32)$$

Using Equation 12, Equation 30 and Equation 32, we obtain a closed-form solution for h_f^o :

$$h_f^{or} = \frac{\theta \xi_h}{1 + \theta} \frac{w_{fa}^r + w_{ma}^r + \lambda^{or} I_k^r - p_a \bar{c}}{w_{fa}^r} \quad (33)$$

Knowing h_f^{or} , Equation 32 solves h_m^{or} .

Type f Households Subject to the time constraint (Equation 12), a type f household in region r solves:

$$\max_{n_g^{fr}, h_g^{fr}} U^{fr} = \ln B + \ln \left[\mu w_{fu} n_f^{fr} + w_{ma}^r n_m^{fr} + \lambda^{fr} I_k^r - p_a \bar{c} \right] + \theta \ln \left[\delta_f \delta_s \left(h_f^{fr} \right)^{\xi_h} \left(h_m^{fr} \right)^{1 - \xi_h} \right] + \ln \epsilon^1 \epsilon^f$$

where the female member works fixed hours in non-agriculture: $n_f^{fr} = \bar{n}$, $h_f^{fr} = 1 - \bar{n}$.

The first order condition for the hours choice of the male member is:

$$\frac{w_{ma}^r}{\mu w_{fu} \bar{n} + w_{ma}^r n_m^{fr} + \lambda^{fr} I_k^r - p_a \bar{c}} = \frac{\theta (1 - \xi_h)}{h_m^{fr}} \quad (34)$$

Together with $n_m^{fr} + h_m^{fr} = 1$, we get an explicit solution for h_m^{fr} :

$$h_m^{fr} = \frac{\theta (1 - \xi_h)}{\theta (1 - \xi_h) + 1} \frac{\mu w_{fu} \bar{n} + w_{ma}^r + \lambda^{fr} I_k^r - p_a \bar{c}}{w_{ma}^r} \quad (35)$$

If $h_m^{fr} > 1 - n_{min}$, we set $h_m^{fr} = 1 - n_{min}$.

Type m Households Subject to the time constraint (Equation 12), a type m household in region r solves:

$$\max_{n_f^{mr}, h_f^{mr}} U^{mr} = \ln B + \ln [w_{fa}^r n_f^{mr} + \mu w_{mu} n_m^{mr} + \lambda^{mr} I_k^r - p_a \bar{c}] + \theta \ln [\delta_s (h_f^{mr})^{\xi_h} (h_m^{mr})^{1-\xi_h}] + \ln \epsilon^1 \epsilon^m,$$

where the male member works fixed hours in non-agriculture: $n_m^{mr} = \bar{n}$, $h_m^{mr} = 1 - \bar{n}$.

The first order condition of female hours is given by:

$$\frac{w_{fa}^r}{w_{fa}^r n_f^{mr} + \mu w_{mu} \bar{n} + \lambda^{mr} I_k^r - p_a \bar{c}} = \frac{\theta \xi_h}{h_f^{mr}} \quad (36)$$

Together with Equation 12, we obtain:

$$h_f^{mr} = \frac{\theta \xi_h}{\theta \xi_h + 1} \frac{w_{fa}^r + \mu w_{mu} \bar{n} + \lambda^{mr} I_k^r - p_a \bar{c}}{w_{fa}^r} \quad (37)$$

If $h_f^{mr} > 1 - n_{min}$, we set $h_f^{mr} = 1 - n_{min}$.

Type b Households When both members are working in non-agriculture, they do not choose hours. A type b household in region r has utility:

$$U^{br} = \ln B + \ln [\mu w_{fu} n_f^{br} + \mu w_{mu} n_m^{br} + \lambda^{br} I_k^r - p_a \bar{c}] + \theta \ln [(h_f^{br})^{\xi_h} (h_m^{br})^{1-\xi_h}] + \ln \epsilon^1 \epsilon^b$$

where

$$n_f^{br} = n_m^{br} = \bar{n}; \quad h_f^{br} = h_m^{br} = 1 - \bar{n}. \quad (38)$$

Choice Among the Four Types We assume a nested structure where households first decide if at least one member will migrate, and then decide on which type of migration, as illustrated in Figure 3. Define V^{ir} as the indirect utility derived from optimal employment and time allocations by household type i in region r . Given the preference shifter, we have:

$$\ln(V^{ir} v^i) \equiv \max U^{ir}(v^i), \quad i = o, f, m, b. \quad (39)$$

where

$$v^i = \begin{cases} \epsilon^o & i = o \\ \epsilon^1 \epsilon^f & i = f \\ \epsilon^1 \epsilon^m & i = m \\ \epsilon^1 \epsilon^b & i = b \end{cases}$$

We solve the problem backwards. In the second stage, conditional on having at least one migrant, households choose which among type f , type m , and type b . Let ϵ^f , ϵ^m , ϵ^b be independent random variables each following a standard one-parameter Fréchet distribution with shape parameter $\kappa_2 > 1$:

$$F(\epsilon) = \Pr(\epsilon^i \leq \epsilon) = \exp(-\epsilon^{-\kappa_2}), \quad \epsilon > 0$$

Hence, the share of households with both spouses migration is given by:

$$\begin{aligned} \alpha^{br} &= (1 - \alpha^u - \alpha^{or}) \int_0^\infty \Pr(V^{fr}\epsilon^f < V^{br}\epsilon^b) \Pr(V^{mr}\epsilon^m < V^{br}\epsilon^b) f(\epsilon^b) d\epsilon^b \\ &= (1 - \alpha^u - \alpha^{or}) \int_0^\infty \Pr\left(\epsilon^f < \frac{V^{br}}{V^{fr}}\epsilon^b\right) \Pr\left(\epsilon^m < \frac{V^{br}}{V^{mr}}\epsilon^b\right) f(\epsilon^b) d\epsilon^b \\ &= (1 - \alpha^u - \alpha^{or}) \int_0^\infty e^{-\left(\frac{V^{br}}{V^{fr}}\epsilon\right)^{-\kappa_2}} e^{-\left(\frac{V^{br}}{V^{mr}}\epsilon\right)^{-\kappa_2}} \kappa_2 \epsilon^{-1-\kappa_2} e^{-\epsilon^{-\kappa_2}} d\epsilon \\ &= (1 - \alpha^u - \alpha^{or}) \int_0^\infty \kappa_2 \epsilon^{-1-\kappa_2} e^{-\left[\left(\frac{V^{br}}{V^{fr}}\right)^{-\kappa_2} + \left(\frac{V^{br}}{V^{mr}}\right)^{-\kappa_2} + 1\right] \epsilon^{-\kappa_2}} d\epsilon \\ &= (1 - \alpha^u - \alpha^{or}) \frac{e^{-\left[\left(\frac{V^{br}}{V^{fr}}\right)^{-\kappa_2} + \left(\frac{V^{br}}{V^{mr}}\right)^{-\kappa_2} + 1\right] \epsilon^{-\kappa_2}}}{\left[\left(\frac{V^{br}}{V^{fr}}\right)^{-\kappa_2} + \left(\frac{V^{br}}{V^{mr}}\right)^{-\kappa_2} + 1\right]} \Bigg|_0^\infty \\ &= (1 - \alpha^u - \alpha^{or}) \frac{(V^{br})^{\kappa_2}}{(V^{fr})^{\kappa_2} + (V^{mr})^{\kappa_2} + (V^{br})^{\kappa_2}} \end{aligned}$$

In general, we have

$$\alpha^{ir} = (1 - \alpha^u - \alpha^{or}) \frac{(V^{ir})^{\kappa_2}}{(V^{fr})^{\kappa_2} + (V^{mr})^{\kappa_2} + (V^{br})^{\kappa_2}} \quad \forall i = b, f, m. \quad (40)$$

To solve the first stage, we need an expression for the expected utility conditional on migration, $\mathbb{E}[\max(V^f\epsilon^f, V^m\epsilon^m, V^b\epsilon^b)]$. The assumption that ϵ^i follows a Fréchet distribution with shape parameter κ_2 implies:

$$\Pr(V\epsilon \leq x) = \Pr\left(\epsilon \leq \frac{x}{V}\right) = \exp\left(-\left(\frac{x}{V}\right)^{-\kappa_2}\right) = \exp(-V^{\kappa_2} \cdot x^{-\kappa_2})$$

Since ϵ^f , ϵ^m and ϵ^b are independent:

$$\begin{aligned}
& \Pr(\max(V^{fr}\epsilon^f, V^{mr}\epsilon^m, V^{br}\epsilon^b) \leq x) \\
&= \Pr(V^{fr}\epsilon^f \leq x) \cdot \Pr(V^{mr}\epsilon^m \leq x) \cdot \Pr(V^{br}\epsilon^b \leq x) \\
&= \exp(-(V^{fr})^{\kappa_2} x^{-\kappa_2}) \cdot \exp(-(V^{mr})^{\kappa_2} x^{-\kappa_2}) \cdot \exp(-(V^{br})^{\kappa_2} x^{-\kappa_2}) \\
&= \exp(-[(V^{fr})^{\kappa_2} + (V^{mr})^{\kappa_2} + (V^{br})^{\kappa_2}] x^{-\kappa_2})
\end{aligned} \tag{41}$$

Notice that Equation 41 is a CDF of Fréchet distribution with shape parameter κ_2 and scale parameter S where

$$S \equiv [(V^{fr})^{\kappa_2} + (V^{mr})^{\kappa_2} + (V^{br})^{\kappa_2}]^{1/\kappa_2}$$

The mean of a Fréchet distribution with shape κ_2 and scale S is $S \cdot \Gamma(1 - 1/\kappa_2)$, provided $\kappa_2 > 1$. Therefore:

$$\mathbb{E}[\max(V^{fr}\epsilon^f, V^{mr}\epsilon^m, V^{br}\epsilon^b)] = [(V^{fr})^{\kappa_2} + (V^{mr})^{\kappa_2} + (V^{br})^{\kappa_2}]^{1/\kappa_2} \cdot \Gamma\left(1 - \frac{1}{\kappa_2}\right) \equiv C^r$$

With this expression, we can solve the first stage choice, assuming that households are risk neutral, and they choose the option that maximize their expected utility. Since ϵ^1 and ϵ^o are independent Frechet distributions with paramter κ_1 , we can calculate α^o as:

$$\begin{aligned}
\alpha^{or} &= (1 - \alpha^u) \int_0^\infty \Pr(\mathbb{E}[\max(V^{fr}\epsilon^f, V^{mr}\epsilon^m, V^{br}\epsilon^b)] \epsilon^1 < V^{or}\epsilon^o) f(\epsilon^o) d\epsilon^o \\
&= (1 - \alpha^u) \int_0^\infty \Pr(C^r \epsilon^1 < V^{or}\epsilon^o) f(\epsilon^o) d\epsilon^o \\
&= (1 - \alpha^u) \int_0^\infty \Pr\left(\epsilon^1 < \frac{V^{or}}{C^r} \epsilon^o\right) f(\epsilon^o) d\epsilon^o \\
&= (1 - \alpha^u) \int_0^\infty e^{-(\frac{V^{or}}{C^r} \epsilon)^{-\kappa_1}} \kappa_1 \epsilon^{-1-\kappa_1} e^{-\epsilon^{-\kappa_1}} d\epsilon \\
&= (1 - \alpha^u) \frac{(V^{or})^{\kappa_1}}{(C^r)^{\kappa_1} + (V^{or})^{\kappa_1}}
\end{aligned} \tag{42}$$

Given 40 and 42, we can then calculate:

$$\alpha^i = \phi \alpha^{i1} + (1 - \phi) \alpha^{i0}, \quad i = o, f, m, b \tag{43}$$

where ϕ is the share of rural regions that have reformed ($r = 1$). This gives the share of different types of rural household in the economy.

B1.3 Non-agricultural Hukou Households

Given the time constraint (Equation 12), the representative household head solves

$$\max_{E_f^u, E_m^u} U^u = \ln B + \ln [w_{fu}\bar{n}E_f^u + w_{mu}\bar{n}E_m^u - p_a\bar{c}] + \theta \ln [(h_f^u)^{\xi_h} (h_m^u)^{1-\xi_h}] \quad (44)$$

where E_f^u and E_m^u denote urban female and male employment rates. The household head chooses the optimal E_f^u and E_m^u . The corresponding FOCs are given by:

$$\frac{w_{fu}}{w_{fu}\bar{n}E_f^u + w_{mu}\bar{n}E_m^u - p_a\bar{c}} = \frac{\theta\xi_h}{h_f^u} \quad (45)$$

$$\frac{w_{mu}}{w_{fu}\bar{n}E_f^u + w_{mu}\bar{n}E_m^u - p_a\bar{c}} = \frac{\theta(1-\xi_h)}{h_m^u} \quad (46)$$

This implies a relative relationship between h_f^u and h_m^u :

$$\frac{w_{fu}}{w_{mu}} = \frac{\xi_h}{1-\xi_h} \frac{h_m^u}{h_f^u} \quad (47)$$

Together with the household time constraint, we get:

$$h_f^u = \frac{\theta\xi_h}{1+\theta} \frac{w_{fu} + w_{mu} - p_a\bar{c}}{w_{fu}} \quad (48)$$

Knowing h_f^u , we can easily solve for h_m^u , and hence E_f and E_m .

B2 Recent Changes in China's *Hukou* System

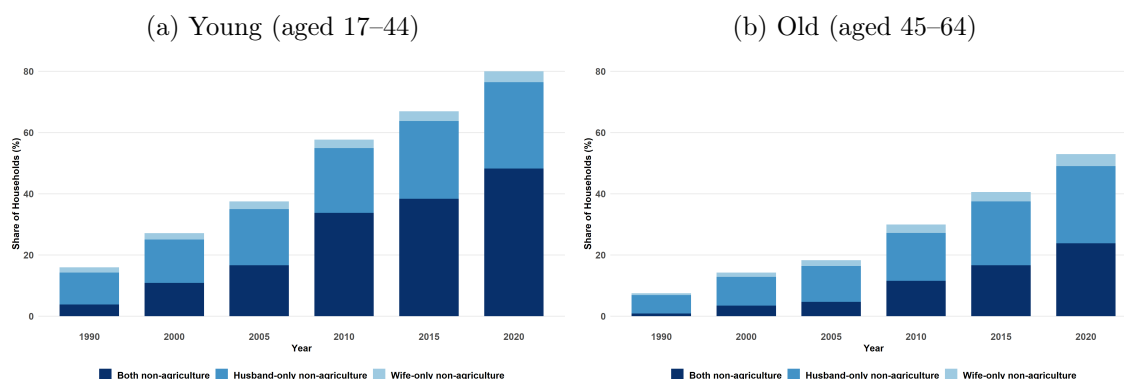
This section summarizes recent changes in the *hukou* system based on the comprehensive reviews by Chan (2019). Since the early 2000s, several cities across 11 provinces have eliminated the agricultural and non-agricultural *hukou* classifications. Residents in these areas now typically hold a unified *hukou* type, labeled “jumin hukou.” In some cases, former agricultural *hukou* holders were given the option to give up their land in exchange for urban social security benefits, such as old-age pensions. For those who retained their land, the *hukou* relabeling is largely symbolic, yielding few meaningful changes in the benefits they receive.

The National New-Type Urbanization Plan, announced in 2014, set a target of granting urban registration to 100 million people by 2020. While priority was given to skilled workers and long-term migrants, many current and former farmers were also included. In July 2014, the State Council issued the *Opinions on Hukou Reform*, detailing measures to merge agricultural and non-agricultural *hukou* into a unified hukou type—reforms that several localities had already begun experimenting with as early as 2007 (e.g., Chongqing Municipality in 2007, Chengdu Prefecture in 2010). Following the Opinions, each province issued its own implementation guidelines over the next two years. However, because hukou registration remains tied to one's registered place of residence at birth, this reform had limited impact on migrants' access to urban public services unless they can change their registration address. More importantly, the reform did not affect the land tenure system: rural farmland remains collectively owned and is allocated to pre-2014 agricultural *hukou* holders with no substantial changes to their land rights.

In summary, while these reforms represented some institutional adjustments, they did not fundamentally alter the *hukou* system or eliminate the rural–urban divide, especially with respect to rural land property rights. Note that following the 2014 reform, the 2015 and 2020 censuses ceased reporting individuals' *hukou* type. We instead use whether an individual holds rural land tenure rights as a proxy for de-facto *hukou* type. In the 2015 data, 62% of the respondents had land rights, indicating that *hukou* unification did not substantially change the underlying land use rights for the former agricultural-*hukou* population.

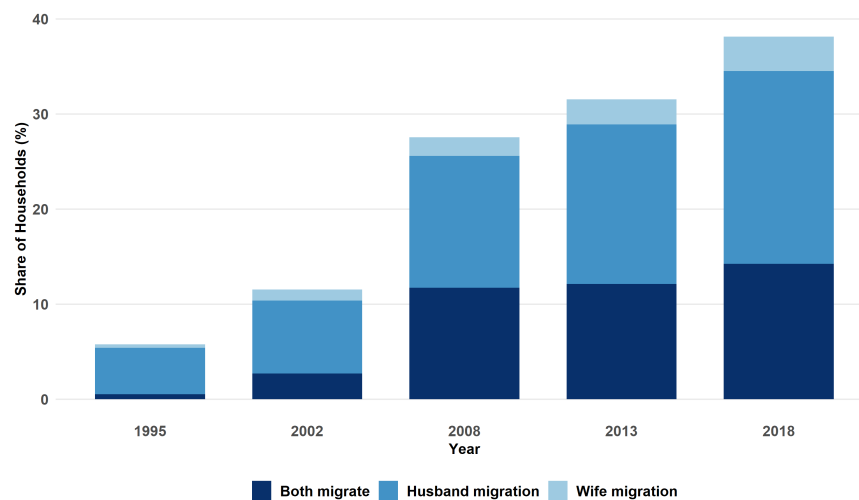
B3 Additional Figures and Tables

Figure B1: Employment Pattern of Rural Married Households by Age Groups



Notes: Census and one percent population survey. Pairs of rural husbands and wives aged 17–44 (young) in Panel A and 45–64 (old) in Panel B. “Both non-agriculture” denotes that both the husband and wife are employed in non-agricultural sectors. “Husband-only non-agriculture” denotes that husbands work in non-agriculture and wives do not. “Wife-only non-agriculture” denotes that wives work in non-agriculture while husbands do not.

Figure B2: Migration Pattern of Rural Married Couples



Notes: CHIP. Pairs of rural husbands and wives aged 18–55 by migration mode: both spouses migrate, only the husband migrates, and only the wife migrates. Migrants: if an individual answers ≥ 6 months (180 days) to “How many days did you work away (live/work outside) from the household (your hometown)?”

Figure B3: Examples of Land Reform Policy Documents at Different Levels

(a) Central government document

中华人民共和国农村 土地承包法 (2002年9月29日九届全国人民代表大会 常务委员会第二十九次会议通过)	
目 录	
第一章 总 则	
第二章 家庭承包	
第一节 发包方和承包方的权利和义务	
第二节 承包的原则和程序	
第三节 承包期限和承包合同	
第四节 土地承包经营权 的保护	
第五节 土地承包经营权 的流转	
第三章 其他方式的承包	
第四章 争议的解决和法律责任	
第五章 附 则	
第一章 总 则	
第一条 为稳定和完善以家庭承包经营为基础、统分结合的双层经营体制，赋予农民长期而有保障的土地使用权，维护农村 土地 承包 当事人的合法权益，促进农业、农村经济发展和农村社会稳定，根据宪法，制定本法。	
第二条 本法所称农村土地，是指农民集体所有和国家所有依法由农民集体使用的耕地、林地、草地，以及其他依法用于农业的土地。	
第三条 国家实行农村 土地 承包 经营制度。	
农村 土地 承包 采取农村集体经济组织内部的家庭承包方式，不宜采取家庭承包方式的荒山、荒沟、荒丘、荒滩等农村土地，可以采取招标、拍卖、公开协商等方式承包。	
第四条 国家依法保护农村 土地 承包 关系的长期稳定。	
农村 土地 承包 后，土地的所有权性质不变。承包地不得买卖。	
第五条 农村集体经济组织成员有权依法承包由本集体经济组织发包的农村土地。	
任何组织和个人不得剥夺和非法限制农村集体经济组织成员承包土地的权利。	
第六条 农村 土地 承包 ，妇女与男子享有平等的权利。承包中应当保护妇女的合法权益，任何组织和个人不得剥夺、侵害妇女应当享有的 土地 承包 经营权。	
第七条 农村 土地 承包 应当坚持公开、公平、公正的原则，正确处理国家、集体、个人三者的利益关系。	
第八条 农村 土地 承包 应当遵守法律、法规，保护土地资源的合理开发和可持续利用。未经依法批准不得将承包地用于非农建设。	
国家鼓励农民和农村集体经济组织增加对土地的投入，培肥地力，提高农业生产能力。	
第九条 国家保护承包土地所有者的合法权益，保护承包方的 土地 承包 经营权，任何组织和个人不得侵犯。	
第十条 国家保护承包方依法、自愿、有偿地进行 土地 承包 经营权流转。	
第十一条 国务院农业、林业行政主管部门分别依照国务院规定的职责负责全国农村 土地 承包 及承包合同管理的指导。县级以上地方人民政府农业、林业等行政主管部门分别依照各自职责，负责本行政区域内农村 土地 承包 及承包合同管理。乡(镇)人民政府负责本行政区域内农村 土地 承包 及承包合同管理。	

(c) *Jiaxing* Prefecture document

嘉兴市人民政府办公室关于加快推进农村 土地承包 经营权流转的意见 (嘉政办发〔2007〕106号)
各县(市、区)人民政府，市政府各部门，直属各单位：
为进一步优化农村土地资源的配置，推进农业产业化经营，加快都市型农业建设，促进农业增效、农民增收和农村发展，现就加快推进我市农村 土地 承包 经营权流转提出如下意见，请认真贯彻执行。
一、充分认识加快推进 土地 承包 经营权流转的重要意义
在家庭承包经营的基础上，加快推进农村 土地 承包 经营权的流转，是发展现代农业的客观要求，是提高农业规模效益，促进农业增效、农民增收的现实途径，也是适应加快农村劳动力转移，引导各类市场主体投资效益农业，推进社会主义新农村建设和构建社会主义和谐社会的必然选择。近年来，在各级党委、政府和有关部门的共同努力下，我市农村 土地 承包 经营权流转工作取得了一定的成效，但与农业产业化和现代农业的发展要求还不相适应。个别地方对 土地 承包 经营权流转尚未引起足够重视， 土地 承包 经营权流转进展较慢，流转行为不够规范，流转纠纷时有发生，土地的分散经营与农业产业化、市场化之间的矛盾较为突出。因此，各级政府 and 有关部门要进一步统一思想，提高认识，积极引导，规范管理，加大扶持，加快推进农村 土地 承包 经营权的流转，促进全市农村经济发展再上新台阶。
二、推进 土地 承包 经营权流转的基本原则
坚持农村基本经营制度。以家庭承包责任制为基础的统分结合的双层经营体制，是党在农村的一项基本制度，推进 土地 承包 经营权的流转，不得改变农村土地家庭承包经营制度。鼓励农村集体土地的所有权、承包权、经营权相分离，稳定承包权，搞活经营权，规范 土地 承包 经营权的经营流转。
坚持“依法、自愿、有偿”。 土地 承包 经营权的流转，必须尊重农户的意愿，必须符合有关法律法规和政策，任何单位和个人不得强迫或者阻碍农户进行土地流转，不得以调整产业结构、扩大生产规模等名义收回承包权，不得以少数服从多数的办法搞土地流转。
坚持因地制宜、分类指导。推进 土地 承包 经营权的流转，必须与当地农村生产力发展水平和农业生产区域特色相适应。要根据各地的生产力发展水平、非农产业发展程度和农村劳动力转移状况，分类指导，逐步推进。
坚持有序流转、规模经营。以农业产业发展规划为基础，以培育特色优势产业为目标，通过强化村级组织管理职能，加强引导和协调，促进土地集中流转，推进农业适度规模经营。
三、加强引导，加快推进 土地 承包 经营权的流转
(一) 积极鼓励农户依法采取转包、出租、互换、转让、入股等多种方式流转 土地 承包 经营权。鼓励承包农户委托发包方和土地流转中介服务组织流转其承包土地。
(二) 积极鼓励有资金、懂技术、善经营、会管理的种粮大户、工商企业、农业龙头企业、农民专业合作社、农业科技人员等农业经营主体投资效益农业，连片开发农户流转的土地。
(三) 在试点基础上逐步推广农村土地股份合作制，鼓励农户以 土地 承包 经营权入股的形式组建土地股份合作社，引

(b) *Zhejiang* province document

浙江省实施《中华人民共和国农村 土地承包法 》办法
发布日期：2010-09-16 访问次数:3285 信息来源：省林业局 字号：大 中 小 打印/转载
(2006年9月30日浙江省第十一届人民代表大会常务委员会第二十七次会议通过 根据2009年12月30日浙江省第十一届人民代表大会常务委员会第十五次会议《关于修改
目录
第一章 总则
第二章 土地发包和承包
第三章 土地承包经营权 的保护
第四章 土地承包经营权 的流转
第五章 土地承包经营权 证管理
第六章 争议的解决
第七章 法律责任
第八章 附则
第一章 总 则
第一条 根据《中华人民共和国农村 土地 承包法 》等法律、法规，结合本省实际，制定本办法。
第二条 本省行政区域内农村 土地 承包 及承包合同的管理，适用本办法。
本办法所称农村土地，是指农民集体所有和国家所有依法由农民集体使用的耕地、林地，以及其他依法用于农业的土地。
第三条 国家依法保护集体土地所有者的合法权益，保障承包方的 土地承包 经营权，维护农村 土地 承包 关系的长期稳定。
第四条 农村集体经济组织成员有权依法承包由本集体经济组织发包的农村土地。
农村 土地 承包 ，妇女与男子享有平等的权利。
任何组织和个人不得剥夺和非法限制农村集体经济组织成员承包土地的权利。
第五条 各级人民政府应当加强对农村 土地 承包 工作的领导，维护农村 土地 承包 当事人的合法权益，促进农业、农村经济发展和农村社会稳定。农村 土地 承包 管理日常工作经费应当列入各级财政预算。
省人民政府农业、林业行政主管部门按照各自职责，负责本省农村 土地 承包 及承包合同的管理。
设区的市、县(市、区)人民政府农业行政主管部门或者农村工作部门(以下统称农业行政主管部门)、林业行政主管部门按照各自职责，负责本行政区域内农村 土地 承包 及承包合同的管理。
乡(镇)人民政府负责本行政区域内农村 土地 承包 及承包合同的管理。

(d) *Huzhou* Prefecture document

中共湖州市委办公室、湖州市人民政府办公室关于进一步稳定完善农村 土地 承包 关系的实施意见 (湖委办〔2004〕11号 2004年5月28日)
根据《中华人民共和国农村 土地 承包法 》和中共浙江省委办公厅、浙江省人民政府办公厅《关于进一步稳定完善农村 土地 承包 关系的意见》(浙委办〔2003〕76号)文件精神，进一步依法保障农民的土地 承包 经营权益，促进我市农村经济持续健康发展和农村社会稳定，经市委、市政府同意，特提出如下实施意见。
一、充分认识进一步稳定和完善的农村 土地 承包 关系的重要意义
土地是农民的基本生产资料 and 重要生活保障， 土地 承包 经营权是农民的基本权利，是农民根本利益所在。1998年以来，全市各地认真贯彻落实上级文件精神，基本完成了以延长承包期30年为核心内容的二轮 土地 承包 工作，促进了农村 土地 承包 关系的稳定和完善，为保持农村社会稳定奠定了制度基础，赢得了广大人民群众的拥护。但是，工作上还存在着一些突出问题：少数应承包到户的土地没有发达到户；应发到户的权证没有发到户；土地整理后，农户的承包地与原承包合同、权证不一致；农户承包地被征用后没有及时对承包合同、权证进行变更、注销。这些问题如果不及及时、妥善解决，将会影响家庭承包经营制度的长期稳定，影响农用地资源保护和粮食生产安全，影响党群干群关系和农村社会稳定。各级党委、政府及其职能部门必须从实践“三个代表”重要思想，统筹城乡发展，维护农民合法权益的高度出发，坚持依法行政，进一步把稳定和完善的农村 土地 承包 关系工作做好，把国家赋予农民的 土地 承包 经营权真正落实到位。
二、总体要求和基本原则
总体要求：全面完成农村土地二轮承包工作，妥善处理遗留问题，做到承包面积、座落四至、合同、权证“四到户”；切实加强 土地 承包 合同、权证和档案管理，规范承包土地的流转、整理、征用行为，有效保障农民的土地 承包 经营权，进一步稳定农村 土地 承包 关系；完善被征地农民社会保障制度，扩大就业，逐步解决农民群众的后顾之忧和改善人民群众生活；加大基本农田保护力度，确保耕地占补平衡和粮食生产能力。各县区要抓紧实施，切实采取有效措施，争取在六月底前全面完成农村 土地 承包 关系的稳定完善工作。
工作原则：
一是突出稳定，注重完善。坚持“大稳定、小调整”，已经完成土地二轮承包的地方，不能推倒重来搞重新发包。不够完善的地方，要按照程序合法、手续完备、权证到户的要求做好规范工作。
二是尊重民意，形式多样。完善农村 土地 承包 关系，必须充分尊重农民群众的意愿，确保农民的合法权益。土地调整、整理、流转、征用等必须坚持法律法规的规定程序，解决遗留问题，要从各地实际出发，有什么问题就解决什么问题，不搞一刀切。
三是落实责任，积极稳妥。实行各地党政一把手负责制，积极稳妥，讲究方法，做实做细工作，重在解决问题。确保土地二轮承包工作的全面完成，确保不影响农村社会稳定。
三、进一步明确农村 土地 承包 关系中的有关政策
(一) 切实维护农村 土地 承包 关系长期稳定。坚持农村土地所有权归属不变，不准以完善二轮 土地 承包 为由，随意改变农村土地所有权。尚未开展二轮承包的村或组以及虽已列入国家建设用地规划，但在2004年底前不能被征用的，必须抓紧组织完成二轮 土地 承包 ，原则上以1998年

(e) Haiyan County

海盐县人民政府办公室关于印发海盐县农村土地承包经营权流转管理暂行办法的通知

各乡镇人民政府，县政府各部门、直属各单位：

《海盐县农村土地承包经营权流转管理暂行办法》已经县政府同意，现印发给你们，请认真贯彻执行。

二〇〇九年五月七日

海盐县农村土地承包经营权流转管理暂行办法

第一章 总则

第一条 为了规范农村土地承包经营权流转（以下简称农村土地流转）行为，维护各方当事人的合法权益，促进农业和农村经济的发展，根据《中华人民共和国农村土地承包法》、农业部《农村土地承包经营权流转管理办法》和县委、县政府《关于积极引导农村土地承包经营权流转促进农业规模经营的若干意见》等有关规定，制订本办法。

第二条 凡本县范围内的农村土地流转应当遵照本办法执行。

第二章 农村土地流转原则

第三条 坚持家庭承包经营制度不变原则。家庭承包责任制为基础的统分结合的双层经营体制，是党在农村的一项基本经营制度不得改变。农村土地流转是将农村集体土地的所有权、承包权、经营（使用）权相分离，以稳定土地的家庭承包关系为基础，实行土地经营（使用）权流转。

第四条 坚持“依法、自愿、有偿”原则。农村土地流转，必须尊重农户的意愿，必须符合有关法律法规和政策，任何单位和个人不得强迫或者阻碍农户进行土地流转。

第五条 坚持市场主导、政府引导原则。土地流不流转、流给谁、多少价格，都应该通过市场机制解决。政府通过政策激励、有效服务、规范管理，积极引导土地流转。

(f) Jiashan County

嘉善县人民政府办公室文件

善政办发〔2009〕126号

嘉善县人民政府办公室
关于印发嘉善县农村土地流转经营权证
登记管理办法（试行）的通知

各乡镇人民政府（街道办事处），县府各部门、直属各单位：

《嘉善县农村土地流转经营权证登记管理办法（试行）》已经县政府研究同意，现印发给你们，请认真贯彻执行。

二〇〇九年十一月三日

Notes: Source: pkulaw.com. Panels (a)–(f) show examples of policy documents on land contracting reform⁻¹⁻ (the keywords used in the search) from the central level to the province, prefecture, and county levels.

Table B1: Summary Statistics: Census Agricultural *hukou* Sample

Variables	Obs.	Mean	SD	Min	Max
Non-ag Employment	13,140,026	0.22	0.42	0	1
Gender (1=female)	13,140,026	0.49	0.50	0	1
Minority	13,140,026	0.09	0.29	0	1
Age	13,140,026	34.4	10.5	18	55
Marital Status	13,140,026	0.78	0.41	0	1
Education (high skill)	13,140,026	0.004	0.07	0	1

Notes: This table displays summary statistics for the variables used in Table A1. Source: Individual-level data of Census and One Percent Population Survey from 1990, 2000, 2005, 2010, 2015, and 2020.

Table B2: CHIP (Urban): Summary Statistics

Variable	N	Mean	SD	Min	Max
Weighted Land Reform Index	104175	0.45	1.13	0.92	
Hukou Reform Index	104175	4.01	1.71	1	6
Employment (=1)	104175	0.73	0.40	0	1
Income (Logged)	74175	6.83	3.86	0.65	15.84
Age	104175	36.9	9.70	18	55
Female	104175	0.50	0.50	0	1
Married (=1)	104175	0.81	0.39	0	1
Primary School and Below	19835	19.04%			
Middle School	23453	22.51%			
High School	36082	34.64%			
College or Above	24805	23.81%			

Notes: China Household Income Project (CHIP), aged 18–55, urban, non-agricultural *hukou* individuals.

Table B3: RFPS: Summary Statistics

Variable	N	Mean	SD	Min	Max
Land Reform	537896	1.089	0.834	0	3
Non-farm Employment (=1)	545835	0.613	0.487	0	1
Migration (=1)	545835	0.389	0.488	0	1
Female	545835	0.483	0.500	0	1
Age	545835	37.214	10.908	18	55
Education Level	545835				
- Primary School	169491	31.05%			
- Middle School	304354	55.76%			
- High School	57507	10.54%			
- College	14483	2.65%			
Married (=1)	545835	0.826	0.379	0	1
Weighted Hukou Reform	537896	2.941	0.872	1.031746	5.927757
New Rural Pension Reform	545835	0.393	0.488	0	1
With Child under 7 (=1)	545835	0.300	0.458	0	1
With Elderly above 65 (=1)	545835	0.212	0.409	0	1

Notes: Rural Fixed Point Survey (RFPS), aged 18–55, agricultural *hukou* individuals.

Table B4: CFPS (Rural): Summary Statistics

Variable	N	Mean	SD	Min	Max
Land Reform Index	120734	1.78	0.83	0	3
Weighted Hukou Reform Index	120734	3.52	0.88	1.10	5.92
Migration (=1)	120734	0.11	0.32	0	1
Female	120733	0.50	0.50	0	1
Age	120734	36.77	10.86	18	55
Married (=1)	120734	0.74	0.44	0	1
Primary School and Below	48778	40.48%			
Middle School	45484	37.75%		0	1
High School	17556	14.57%		0	1
College or Above	8682	7.20%		0	1

Notes: China Family Panel Studies (CFPS), aged 18–55, agricultural *hukou* individuals.

Table B5: Keywords: Search Documents from pkulaw.com

Keywords	Total	Province	City	County
Land Contracting Reform				
<i>Land Contracting</i> (“土地承包”)	1390	182	646	562
<i>Land Transfer</i> (“土地流转” or “土地转让”)	574	142	289	143
<i>Land Contracted Management</i> (“土地经营”)	1487	126	739	622
Total Number of Documents	3451	450	1674	1327
Unique Administrative Unit Issuing Policy		(27)	(256)	(348)
Land Titling Reform				
<i>Land Titling</i> (“土地确权”)	303	27	199	77
<i>Land Certification</i> (“土地登记”)	768	31	375	362
Total Number of Documents	1071	58	574	439
Unique Administrative Unit Issuing Policy		(31)	(286)	(431)

Notes: This table presents the keywords used in the search for land contracting and titling reforms, along with the corresponding document counts at various administrative levels sourced from pkulaw.com.

Table B6: Impact of Reforms on Rural Individuals’ Migration and Employment, by Gender and Age

Dependent Variables	Rural Individual Sample RFPS (2003–2017)			
	Aged between 18 and 44		Aged between 45 and 55	
	Non-farm	Migration	Non-farm	Migration
	(=1)	(=1)	(=1)	(=1)
	(1)	(2)	(3)	(4)
Land Reform	0.004 (0.005)	0.006 (0.004)	0.020*** (0.006)	0.015*** (0.004)
Land Reform * Female	0.014*** (0.004)	0.008*** (0.003)	-0.008 (0.005)	-0.004 (0.004)
Control Variables	Yes	Yes	Yes	Yes
Individual FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Observations	228861	228861	120024	120024

Notes: Rural Fixed Point Survey 2003–2017. The dependent variable in columns 1 and 3 is a dummy variable that equals one if the surveyed individual devoted any labor days to non-farm employment during a year, while that in columns 2 and 4 is a dummy variable that equals one if the surveyed individual devoted any labor days to out-migration work during a year. Control variables include age, a dummy variable of marital status, and education level. All specifications control for individual fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table B7: Impact of Reforms on Rural Individuals' Migration by Education Level, Alternative Data

Dependent Variables	Rural Individual Sample					
	All	High School and Below	College and Above	All	High School and Below	College and Above
	Non-farm (=1)			Migration (=1)		
Panel A	CFPS (2010–2020)					
	(1)	(2)	(3)	(4)	(5)	(6)
Land Reform Index	0.016** (0.007)	0.016** (0.007)	0.029 (0.020)	0.013** (0.007)	0.013** (0.007)	0.024 (0.042)
Land Reform Index * Female	0.009** (0.004)	0.009** (0.004)	0.012 (0.023)	0.008*** (0.003)	0.008*** (0.003)	0.013 (0.025)
	114647	105962	7250	114647	105962	7250
Control Variables	Yes	Yes	Yes	Yes	Yes	Yes
Individual FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Panel B	CHIP (1995–2018)					
	(7)	(8)	(9)	(10)	(11)	(12)
Land Reform Index	0.022 (0.027)	0.024 (0.028)	-0.139* (0.080)	0.009 (0.010)	0.008* (0.005)	-0.008 (0.012)
Land Reform Index * Female	0.017*** (0.004)	0.019*** (0.004)	-0.01 (0.013)	0.014* (0.009)	0.015*** (0.004)	0.007 (0.010)
Observations	83546	78282	5237	83546	78282	5237
Control Variables	Yes	Yes	Yes	Yes	Yes	Yes
County FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes

Notes: Panel A use CFPS Rural Sample (2010–2020), while Panel B use CHIP Rural Sample (1995–2018). The dependent variables are a dummy variable that equals one if the surveyed individual devoted any labor days to non-farm employment during a year and a dummy variable that equals one if the surveyed individual devoted any labor days to out-migration work during a year. Columns 1, 4, 7 and 10 use the full sample, while columns 2–3, 5–6, 8–9, and 11–12 use sub-samples categorized by education level. Control variables include age, a dummy variable of marital status, and education level. All specifications control for county fixed effects and year fixed effects. Standard errors clustered at the county level in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

Table B8: Determinants of Employment and Migration, by Gender

	Non-farm (=1)		Migration (=1)	
	Female (1)	Male (2)	Female (3)	Male (4)
Age	-0.008*** (0.001)	-0.004*** (0.000)	-0.013*** (0.001)	-0.011*** (0.001)
With Child under 7 (=1)	-0.081*** (0.008)	0.014** (0.007)	-0.068*** (0.007)	0.013 (0.008)
With Elderly above 65 (=1)	0.029*** (0.009)	0.002 (0.007)	0.010 (0.008)	0.001 (0.010)
Married (=1)	-0.045*** (0.010)	-0.021*** (0.008)	-0.046*** (0.009)	-0.051*** (0.010)
Education Level	0.038*** (0.008)	0.016*** (0.006)	0.038*** (0.007)	0.017** (0.008)
Health	0.016 (0.012)	0.020** (0.010)	-0.001 (0.010)	-0.001 (0.011)
Year Fixed Effects	Yes	Yes	Yes	Yes
Observations	256875	275275	256875	275275

Notes: Rural Fixed Point Survey 2003–2017, aged 18–55. The dependent variable in columns 1 to 2: a dummy equal to one if the surveyed individual devoted any labor days to non-farm employment during a year, while that in columns 3 to 4: a dummy equal to one if the surveyed individual devoted any labor days to out-migration work during a year. Individual control variables include age, household presence of child under 7, household presence of an elderly person above 65, marital status, education level, and self-reported health (exclude those reporting unhealthy). All specifications control for year fixed effects. Standard errors clustered at the county level in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

Table B9: Impacts of Reforms on Urban Employment, Alternative Measure

Dependent Variable	Urban Individual Sample CHIP (1995–2018)					
	Employed (=1)	Income (Logged)	Employed (=1)	Income (Logged)	Employed (=1)	Income (Logged)
	All		Low Education		High Education	
	(1)	(2)	(3)	(4)	(5)	(6)
Weighted Land Reform	-0.024 (0.031)	0.044 (0.042)	-0.038 (0.047)	0.068 (0.074)	-0.084 (0.102)	0.047 (0.052)
Weighted Land Reform * Female	-0.034*** (0.009)	-0.068*** (0.022)	-0.048*** (0.008)	-0.087** (0.042)	0.013 (0.036)	-0.028 (0.112)
Control Variables	Yes	Yes	Yes	Yes	Yes	Yes
Prefecture FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Observations	104175	74175	79370	51678	24805	22497

Notes: This table is based on urban samples from CHIP (1995–2018), including individuals aged 18–55 with non-agricultural *hukou*. The dependent variables are employment status, where 1 denotes employment in a full/part-time paid job at the time of the survey (columns 1, 3, and 5) and logged annual wage income (columns 2, 4, and 6). Control variables include gender, age, a dummy indicator for cohabitation with a partner, education level, work experience, ethnicity, city GDP per capita, shares of secondary and tertiary industries, and population size (logged), and the interaction between total employment population at destination d and a yearly trend. All specifications control for prefecture fixed effects and year fixed effects. Standard errors clustered at the prefecture level in parentheses; * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.