Sowing Seeds of Mobility: the Gendered Impact of Land Reform *

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Abstract

We study how land rights insecurity slows women's transition out of agriculture more than men's by keeping women on the farm as guard labor. Exploiting two major Chinese land reforms, we construct a county-level reform index using large-language-model textual analysis of policy documents. Combining this index with a large panel dataset, we provide causal evidence that strengthening land rights shifts rural women out of agriculture significantly more than men. We then develop a multi-sector model incorporating couples' labor allocation, featuring reform and non-reform rural regions, and calibrate key elasticity parameters to match the predicted cross-regional differences with our empirical estimates. This allows us to quantify the contribution of land reform operating through the guard labor mechanism. The effects are gendered and quantitatively important: land reform accounts for 31% of rural women's transition out of agriculture, compared to 11% for men.

JEL classification: J16, J61, E24

Keywords: Gender, Land rights insecurity, Labor mobility, Structural transformation

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1 Introduction

In the early stages of development, agricultural work is the dominant form of employment, yet women transition out of agriculture into non-agricultural employment at lower rates than men (Boserup, 1970; ILO, 2019).[1] The existing literature has largely attributed these disparities to comparative advantages, norms, institutions, or technology that are specific to men and women. This paper highlights a novel mechanism: land rights insecurity, despite being gender-neutral, can generate a gendered path of structural transformation.

Land rights insecurity, prevalent in developing countries (World Bank, 2024), arises in settings where land ownership is communal and farmers hold only use rights, such that leaving land uncultivated risks its reallocation to others. This insecurity creates an incentive for households to retain labor on the farm as "guard labor" to protect their land-use rights.[2] We posit that women are more likely to serve this role due to their perceived comparative advantage in agriculture and home production (Boserup, 1970; Goldin, 1995). First, non-agricultural jobs, especially in manufacturing, often involve tasks and work environments considered more suitable for men. 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] Thus, when interacting with intra-household labor allocation, *de jure* gender-neutral land rights insecurity operate as *de facto* gender-specific mobility barriers.

Despite the prevalence of land rights insecurity, the underlying arrangements are often informal, making them difficult to measure or track over time. China offers a unique op-

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

[2] The role of guard labor under insecure property rights has been documented in other contexts, including urban titling programs in Peru (Field, 2007) and rural land certification programs in Mexico (de Janvry et al., 2015).

[3] Boserup (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) and Goldin (1995). For more recent evidence, see Dinkelman and Ngai (2022) for Africa and Ngai et al. (2024) for the 19th century United States.

portunity, as land rights insecurity is embedded in its land policy under the *hukou* system.[4] Consistent with our proposed guard labor mechanism, we find that villages with more frequent land re-allocation pre-reform have lower non-farm employment and out-migration, and that this disproportionately affects women. The two recent land reforms that strengthened farmers' rental rights allow us to construct a reform index with rich spatial and temporal variation. This index, combined with the *Rural Fixed Point Survey* (RFPS),[5] a large panel dataset tracking rural individuals and households, allows us to causally identify the impact of removing land rights insecurity on the employment outcomes across household members.

This paper makes two contributions. First, we construct a granular, county-level land reform index using a large-language-model (LLM)-based text analysis approach and provide *causal* evidence that land reform reducing land rights insecurity disproportionately increases rural women's transition into non-agricultural employment. Second, by incorporating couples' labor allocation decisions, we develop a multi-sector model with reform and non-reform rural regions, calibrating the key elasticity parameters to match the predicted cross-region differences with our empirical estimates. This allows us to quantify the contribution of land reform, operating through the guard labor mechanism, to the gender difference in the transition out of agriculture.

Our novel county-level reform index covers 2,813 counties from 2003 to 2020. We trace the rollout of two 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 using LLM-based validation, we extract location and timing information to construct the index.

Using the reform index as a direct measure of the removal of land rights insecurity and hence mobility barrier, we identify the causal impact of reform by gender across rural and urban populations. Our analysis shows that the reforms strengthened rental rights

[4]The *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.

[5]Renowned for its data quality, the RFPS has been used in influential studies such as Chari et al. (2021), Adamopoulos et al. (2024), and Gai et al. (2025).

by increasing rental activity and farm size. As a result, they increased the transition out of agriculture, especially for women, and increased joint spousal employment in non-farm work. We then examine the impact of the reform on urban populations through a shift-share design. We find that land reforms widen the gender gap in employment and wage income, particularly among individuals with lower education levels. This points to a substitution effect of rural female migrants on urban women, concentrated among less-educated workers.

While the empirical evidence establishes that land reforms disproportionately affect women, an important question remains: what is the aggregate impact of the reform on female employment? To address this question, we develop a multi-region, multi-sector model with household-level employment decisions. The model features an urban region and a continuum of rural regions consisting of reform and non-reform regions.

In rural regions without land reform, leaving agriculture means losing land income, so households have an incentive to retain a member on the farm as "guard labor" to protect the household land income. In the model, this role of "guard labor" interacts with sector-specific gender intensities to generate gender-differential outcomes from the reform. When women have a comparative advantage in agriculture and home production, the model predicts that they are more likely to serve as guard labor. As a result, reforms that eliminate the need for guard labor disproportionately increase women's transition into non-agriculture. This movement of workers out of agriculture increases the relative supply of female workers in non-agriculture, hence lowering the employment and wage rates of urban women relative to men.[6]

Our quantitative model is tightly disciplined by our empirical estimates. In particular, the empirical reform index from 2000 to 2020 is used to calibrate the share of reform regions, and the empirical regression results are used to pin down elasticities such that the difference in female and male non-agricultural employment shares across reform and non-reform regions reproduces the estimated reform effects from the regressions. The calibrated model allows us

[6]This provides a novel mechanism for the widening gender gap among urban workers in China (Feng et al., 2017; Brussevich et al., 2021; Qian, 2023).

to perform quantitative exercises to isolate the contribution of reform, productivity growth, and other labor market wedges to structural transformation. Reform is the only force that encourages women more than men to transition from agriculture to non-agriculture. Quantitatively, land reform accounts for 31% of the transition of women out of agriculture into non-agricultural sectors, and 11% for men.

Related Literature Boserup (1970) was among the first to highlight the differential pace at which men and women transition out of agriculture into non-farm employment. Goldin (1995) argues that non-farm work, unlike agricultural work on family farms, requires women to leave the home and is therefore difficult to combine with home production.[7] For women, leaving the home is further complicated by supply-side barriers, such as gender norms restricting mobility and concerns over transport safety (Jayachandra, 2021; Garlick et al., 2025).[8] On the demand side, women also face constraints, including men's comparative advantage in physically intensive manufacturing tasks (Rendall, 2024) and legal restrictions on female participation in non-farm sectors (Hyland et al., 2020). We contribute to this literature by identifying land rights insecurity as a barrier that keeps women on the farm as guard labor, and by providing causal evidence on how land reforms facilitate women's transition into non-farm employment.[9]

The gendered patterns of structural transformation carry quantitatively important implications for gender-specific labor market outcomes, as well as for aggregate and sectoral productivity (Ngai and Petrongolo, 2017; Hsieh et al., 2019; Lee, 2024; Chiplunkar and Kleineberg, 2024). The common approach in this literature models labor misallocation by

[7] The trade-off between women's time in home production and market work has been documented using time-use surveys across a broad set of countries (Bridgman et al., 2018; Dinkelman and Ngai, 2022; Gottlieb et al., 2024; Andreescu et al., 2025). Ngai et al. (2024) establish this pattern for the historical United States and develop a theory linking home production and structural transformation to the U-shaped pattern of female employment first emphasized by Goldin (1995).

[8] Alesina et al. (2013) trace back the historical origins of gender norms to agricultural technology, which has a persistent effect on female labor force participation.

[9] While our paper identifies gender-neutral land reforms that shift women into non-agriculture by releasing them from their role as guard labor, Ager et al. (2026) identify a gender-biased technological change within agriculture that displaced female-intensive tasks and thereby pushed women into non-farm employment.

summarizing the aforementioned factors as gender-specific labor market wedges, calibrated to match observed employment outcomes. While such outcome-based measures offer valuable insight into where barriers are severe, they do not identify their underlying causes. Our paper complements this approach by directly measuring a gender-neutral barrier, land rights insecurity, and quantifying its gendered impact.

Land rights insecurity and its economic consequences have been examined across a range of developing-country settings (Chernina et al., 2014; de Janvry et al., 2015; Chen, 2017; Gottlieb and Grobovšek, 2019; Manysheva, 2022). In the context of China, our empirical analysis advances upon Chari et al. (2021), Liu et al. (2023), and Shi et al. (2024) by constructing a comprehensive county-level reform index that captures both reforms and uncovers novel effects along gender and rural-urban dimensions.[10] Our quantitative analysis relates 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, rather than relying on survey-based reallocation frequencies. Second, we emphasize the joint employment decisions of couples in order to understand the impact by gender.[11]

The remainder of the paper is structured as follows. Section 2 provides background on China's land institutions and describes the construction of the 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.

[10]Specifically, Chari et al. (2021) and Shi et al. (2024) rely on a provincial-level index of land-contracting reform to study labor reallocation. We show that the county is the key margin of variation, confirming that provincial-level indices obscure substantial within-province variation across nearly 2,000 counties. Liu et al. (2023) collect the timing of land titling reform completion for a considerably smaller set of counties and do not study impacts by gender.

[11]See 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) examine migration decisions across age groups, we focus on gender and rural-urban interaction, driven by our empirical findings. Related results across different age groups are discussed throughout the paper.

2 Land Institutions and the Land Reform Index

This section outlines China's land institutions under the *hukou* system, posits the link between land rights 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 5.

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.[12] 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.[13] This insecurity discouraged land rental activities. Formal land rental markets were virtually non-existent, and when land was rented out, it was typically through informal, verbal agreements that yielded very low rents.[14]

2.1 Land Rights Insecurity and Gendered Patterns of Guard Labor

In the presence of land rights 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".

[12]The *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. More details on China's *hukou* system are provided in the Online Appendix B2.

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

[14]The 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).

First, we examine the employment patterns of married couples with agricultural *hukou*.[15] Figure 1 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.[16] Over time, more households are sending a member to work in non-agriculture, as shown in the decline of the type with both members working in agriculture. In the earlier years, the husband-only type dominated. Over time, the share of couples where both spouses work in non-agriculture has increased, coinciding with the timing of the two land reforms.[17] The idea that women are less likely than men to work in non-agriculture is also supported by Appendix Table A1, using individual 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.

Second, we show that historical land rights insecurity reduces non-farm employment and out-migration, mainly for women. The RFPS provides household-level information from 1995 and individual-level information from 2003. In 2006, it conducted a one-off supplementary survey on village governance, where village cadres were asked to recall instances of major land reallocation (*D*atiaozheng) from the 1970s through 2006.[18] Using this measure of historical land rights insecurity at the village level, Table 1 examines the impact on non-farm employment and migration. More historical land reallocations reduce both non-

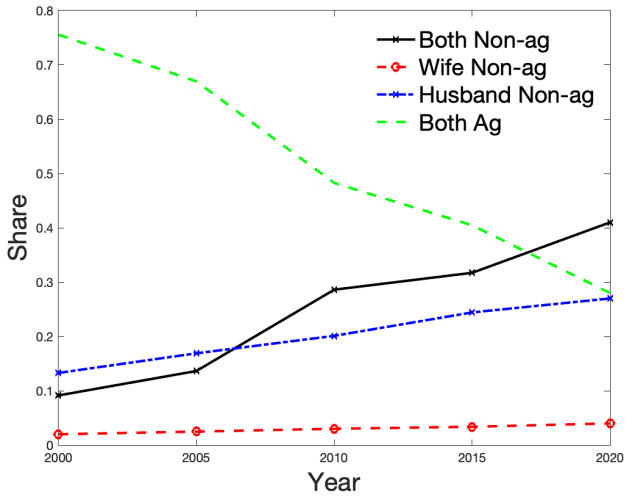
[15]Most couples share the same *hukou* type. Census data on married couples show that the share of mixed *hukou* couples among those aged 18–55 was about 5% from 1990 to 2010 and increased slightly to around 9% during the 2015–2020 period.

[16]To 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) were 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).

[17]The Online Appendix Figure B1 presents a similar breakdown by age, for both the young and old age groups.

[18]Major land reallocation refers to a village cadre's redistribution of more than 30% of village land. Despite this strict threshold, a substantial fraction of these villages experienced such reallocation, with considerable regional disparities in reallocation frequency. Since 1976, our data show that 26% of villages experienced reallocation once, 31% twice, 23% three times, 10% four times, and 10% five or more times, with some villages experiencing up to eight reallocations during this period.

Figure 1: Employment Patterns, Agricultural *Hukou* Married Couples



Notes: This figure plots the employment patterns of married couples with agricultural *hukou* using Census data. "Both Non-ag" denotes that both the husband and wife are employed in non-agricultural sectors. "Wife Non-ag" denotes that wives work in non-agriculture while husbands do not. "Husband Non-ag" denotes that husbands work in non-agriculture while wives do not. "Both Ag" denotes that both spouses work in agriculture. The four shares sum to 1.

farm employment and out-migration among households. More importantly, individual-level results reveal that this negative impact is primarily driven by women, consistent with our hypothesis that women are more likely to serve 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 only strengthened 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 (RLCL) 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, prefecture and county governments introduced

Table 1: Impact of Pre-Reform Land Rights Insecurity on Labor Reallocation

	Rural Households (1995–2006)		Rural Individuals (2003–2006)			
	Non-farm	Migration	Non-farm		Migration	
	(1)	(2)	(3)	(4)	(5)	(6)
Land Reallocation	-0.008**	-0.024**	-0.014**	0.006	-0.035**	0.010
	(0.004)	(0.010)	(0.006)	(0.004)	(0.017)	(0.018)
Land Reallocation*Female				-0.021**		-0.048**
				(0.010)		(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)–(6) use the individual sample of individuals aged 18–55 for 2003–2006. The dependent variable is a dummy equal to 1 if any household member spent any labor days in non-farm employment during the year in column (1), and if any household member spent any labor days in out-migration work during the year in column (2). In columns (3)–(4), the dependent variable is a dummy equal to 1 if the surveyed individual spent any labor days in non-farm employment during the year; in columns (5)–(6), it is a dummy equal to 1 if the surveyed individual spent any labor days in out-migration work during the year. Land reallocation is defined as the redistribution of more than 30% of village land by village cadres, as recalled for the period from the 1970s through 2006 and reported in a one-off supplementary survey on village governance conducted by the RFPS in 2006. Columns (1)–(2) include household and year fixed effects, while columns (3)–(6) include individual and year fixed effects. Standard errors clustered at the county level in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

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 2009,[19] addressed these gaps. Unlike the contracting reform's sequential provincial-to-county rollout, the titling reform was implemented directly at the county level, with counties given the 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, digitized national land registration system. This system records comprehensive information, including household ownership, demographic data, the identity of household heads, family relationships, and precise GIS-based parcel boundaries. The completion of the land

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

titling reform in 2019 marked a major milestone in improving land rights security in China's agricultural sector.

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.[20] 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.[21] 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 B2 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

[20] As 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.

[21] We 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 as (1), return 1; otherwise, return 0. If your answer is 1, please extract the most relevant sentences from (2) that support your judgment."

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 the 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 the contracting reform. In sum, our land reform index ranges from 0 to 3. Figure 2 illustrates the spatial diffusion of the phased reforms.

Our land reform index improves upon existing measures in two fundamental ways. First, on the contracting reform, it exploits granular county-level variation stemming from the staggered rollout. 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 more than three million government documents from a centralized database. This comprehensive coverage of rural counties makes it possible to analyze the impact of the reform not only on rural populations, but also on urban ones, as migrants from counties with varying reform intensities collectively shape 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 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).

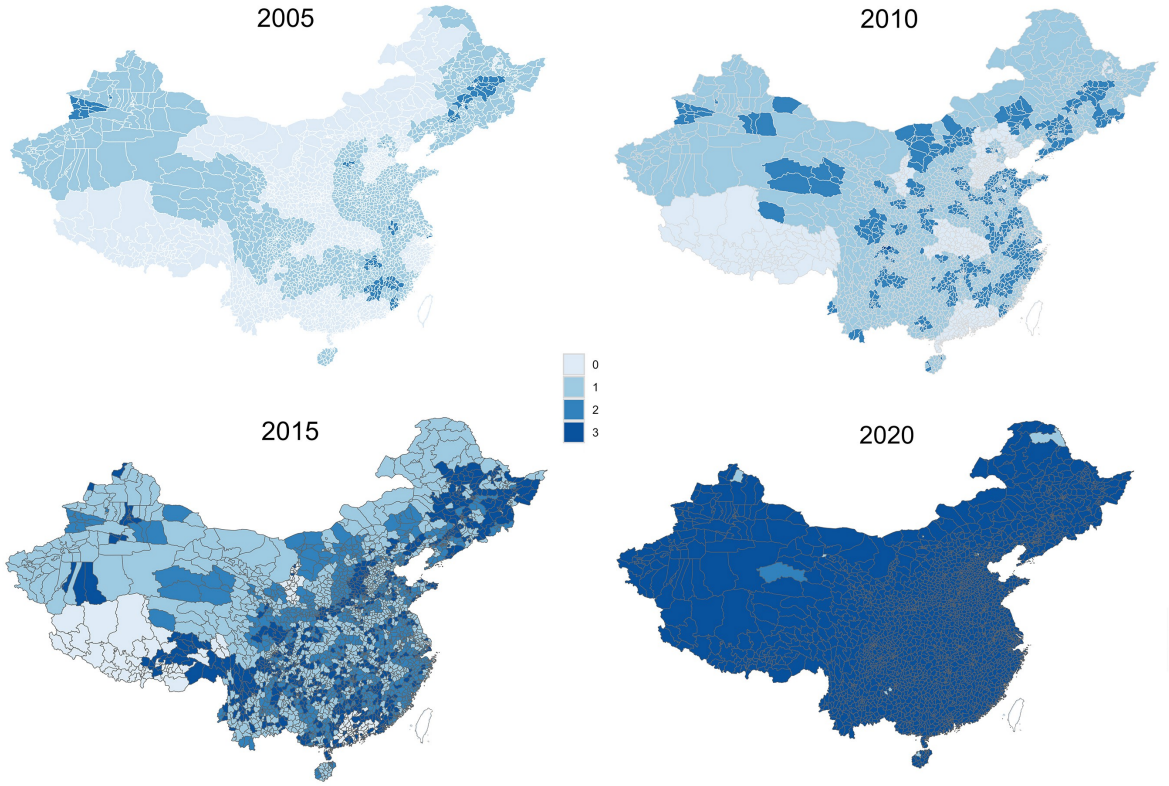
We formally test what predicts the timing of reform adoption by regressing the year of county-level reform completion on a set of predetermined county characteristics, including economic prosperity, economic structure, fiscal capacity, urbanization, farm size, geographic connectivity, and, crucially, 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 Appendix Table A2.

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. More importantly, neither the pre-reform level of total out-migration nor that of female out-migration in 2000 is significantly correlated with reform timing, which supports the identifying assumption that the rollout was not systematically related to baseline migration patterns. 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, as land disputes and encroachment by real estate development created urgency to formalize land rights. In our main specification (columns (4) and (8) of Table 2), we accordingly control for county-level real estate investment.

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 3 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

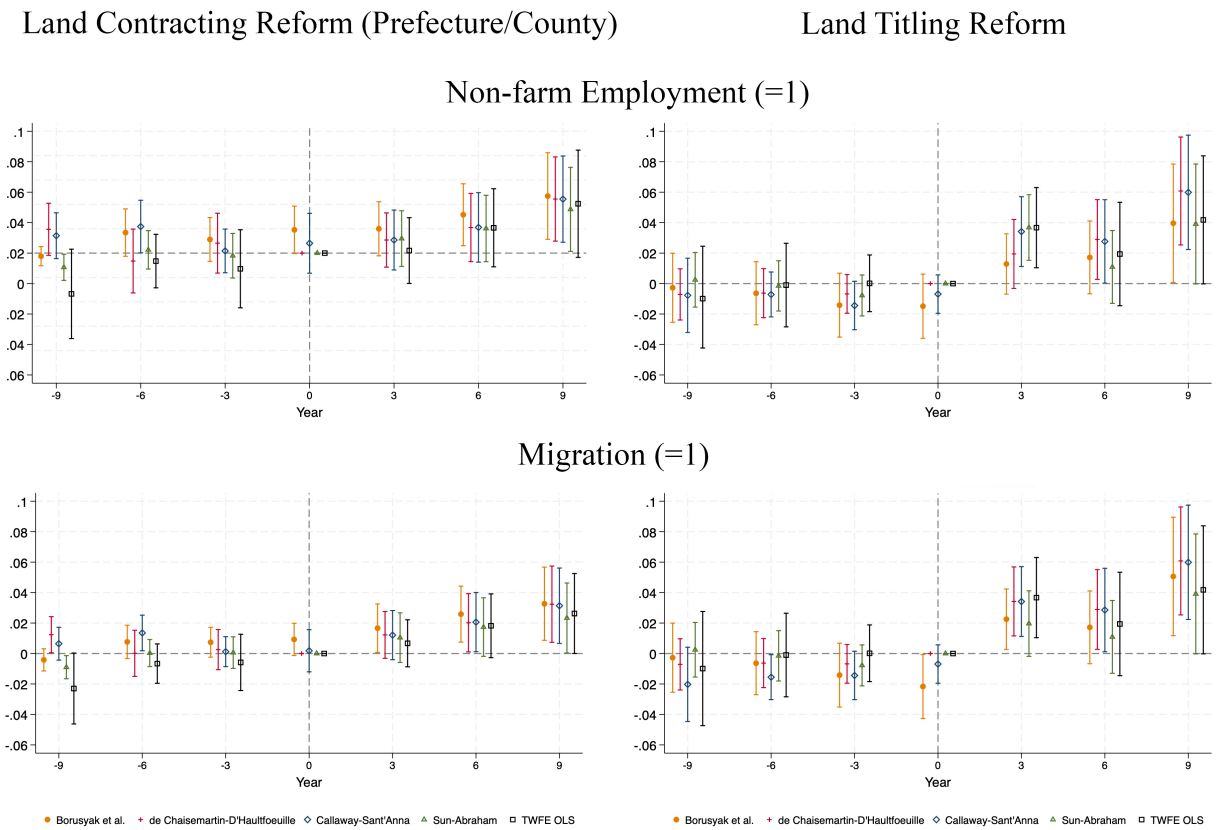
(Borusyak et al., 2024; de Chaisemartin and D'Haultfœuille, 2020; Callaway and Sant'Anna, 2021; Sun and Abraham, 2021), we confirm the absence of pre-reform trends in both non-farm employment and migration, formally validating the parallel trends assumption central to our identification strategy. Furthermore, 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 2: The Spatial Diffusion of Land Reform, 2005–2020



Notes: This figure maps the county-level land reform index onto the 2010 Chinese county-level GIS map. Darker colors indicate a higher land reform index, corresponding to deeper and broader reform progress within a county. The reform index takes four values: 0 (no reform), 1 (provincial regulations on land-contracting reform issued), 2 (prefecture- or county-level regulations on land-contracting reform issued), and 3 (land-titling reform completed, regardless of whether contracting reform had been implemented).

Figure 3: Event Studies of Land Reforms on Non-farm Employment and Migration



Notes: Rural Fixed Point Survey, 2003–2017; individuals aged 18–55. The figure reports event-study estimates for the land-contracting reform (left panels) and the land-titling reform (right panels) using multiple estimators, from nine years before to nine years after local reform implementation. The upper panels plot effects on non-farm employment, and the lower panels plot effects on migration. All specifications include individual and year fixed effects. Individual controls include 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 index and a dummy for adoption of the New Rural Pension Reform (Gai et al., 2025).

3 Empirical Analysis

This section presents causal evidence on the impact of land reforms across genders. 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.

3.1 Impact of Land Reforms on Land Rights Insecurity

Two pieces of survey evidence have documented improvements in land rights security for Chinese farming households following the 2003 reforms. Using a five-wave survey spanning 1999–2010, 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. More recently, Adamopoulos et al. (2024) report that reallocation events declined from 155 occurrences between 1991–2003 (a probability of 9.9%) to 16 occurrences between 2003–2018 (0.9%), using two retrospective surveys on the historical frequency of village-wide land reallocations among 120 sampled villages.

However, directly testing the impact of reform on land security remains challenging, as yearly 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, we combine our county-level land reform index with RFPS, and find that land reforms significantly increased household-level land rental activities. As shown in Appendix Table A3, reforms 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. We use RFPS data from 2003 to 2017 and 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 analyze their employment and migration outcomes using the following specification:

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

where Y_{iot} is a dummy capturing one of two 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, 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 it 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,[22] and a dummy for the adoption of the New Rural Pension Reform. Since RFPS tracks the same agricultural *hukou* households and individuals over time, we include individual fixed effects δ_i to control for unobserved selection factors. We also include 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 within-county correlation.

Individual Employment and Migration Outcomes—Table 2 presents the results on non-farm employment (columns (1) to (4)) and migration (columns (5) to (8)). Overall, land reform increases non-farm employment and out-migration, with the positive effects being stronger for women than for men. The gender-differential effects are stable as we add controls stepwise. In columns (1) and (5), we include two dummy variables proxying for caretaking burden: whether the household has a child under age 7 and whether it has an elderly member above age 65. Having a child significantly reduces both non-farm employment and migration, while the presence of an elderly member does not.[23] In columns (2) and (6), we add proxies for two concurrent reforms: the weighted non-land-related *hukou* reform index to control

[22]This 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.

[23]One concern is that elderly household members may help guard land, thereby attenuating the observed impact of land reform on female labor reallocation. We find no empirical support for this channel, which may reflect the fact that multi-generational households are relatively rare: they account for only about 20% of the RFPS sample, while the corresponding share among rural households in the 2000 Census is about 13.3%.

for the pull effect from urban-side *hukou* reforms, and the New Rural Pension Reform to capture its 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 accordingly add county-level real estate investment, given our earlier finding that real estate development pressure was one of the key drivers of reform adoption. 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 margin 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, both local-level contracting and titling reforms significantly increase non-farm employment, with a significantly larger effect 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's migration, both significantly increase women's migration relative to men's. 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 local-level implementations.

Our findings using the three reform-margin dummies help reconcile the 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 significantly affect non-farm employment, with the effects on women emerging at the local level.[24]

[24]Our results also speak to Shi et al. (2024), who use provincial-level contracting variation and find a larger effect for men. Our estimates are consistent with theirs at the provincial margin, where we find no

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

	Rural Individual Sample RFPS (2003-2017)							
DV	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; individuals aged 18–55. The dependent variable in columns (1)–(4) is a dummy equal to one if the surveyed individual spent any labor days in non-farm employment during the year; the dependent variable in columns (5)–(8) is a dummy equal to one if the surveyed individual spent any labor days in out-migration work during the year. Individual controls include age, a marital-status dummy, and education level. Economic controls include logged county- and village-level GDP per capita, manufacturing and tertiary industry shares, and logged population size. Columns (1) and (5) additionally control for indicators for whether the household has a child under age 7 and an elderly member above age 65. Columns (2) and (6) further control for the weighted non-land-related *hukou* reform index (Gai et al., 2025) and a dummy for adoption of the New Rural Pension Reform. Columns (3) and (7) additionally control for weighted employment growth and GDP growth in destination cities to absorb time-varying labor demand shocks. Columns (4) and (8) further control for county-level real estate investment. All specifications include individual and year fixed effects. Standard errors clustered at the county level in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

Results reported in Table 3 are used to calibrate our quantitative model. To calculate the full-reform impact, we weight each reform margin by the share of counties that have adopted it by 2020: 92.3% for provincial contracting reform, 39.8% for prefecture/county contracting, and 100% for titling reform. The weighted sum implies that a full reform raises male non-

significant female differential (Table 3). The female-favoring pattern in our setting emerges at the finer prefecture/county and titling reform margins, suggesting that the two sets of findings reflect different levels of reform variation.

Table 3: Impact of Individual 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)
Land Contracting Reform (Provincial)	0.007	0.014**
	(0.008)	(0.007)
Land Contracting Reform (Provincial) * Female	0.006	-0.005
	(0.007)	(0.005)
Land Contracting Reform (Prefecture/County)	0.031**	0.009
	(0.013)	(0.011)
Land Contracting Reform (Prefecture/County) * Female	0.024*	0.023***
	(0.012)	(0.008)
Land Titling Reform	0.028*	0.039**
	(0.016)	(0.015)
Land Titling Reform * Female	0.022*	0.017**
	(0.013)	(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; individuals aged 18–55. The dependent variable in column (1) is a dummy equal to one if the surveyed individual spent any labor days in non-farm employment during the year; the dependent variable in column (2) is a dummy equal to one if the surveyed individual spent any labor days in out-migration work during the year. Three separate reform dummies indicate provincial-level contracting reform, prefecture- or county-level contracting reform, and county-level titling reform, respectively. Individual controls include age, a marital-status dummy, and education level. Economic controls include logged county- and village-level GDP per capita, manufacturing and tertiary industry shares, and logged population size. All specifications additionally control for indicators for whether the household has a child under age 7 and an elderly member above age 65, the weighted non-land-related *hukou* reform index (Gai et al., 2025), a dummy for adoption of the New Rural Pension Reform, weighted employment growth and GDP growth in destination cities, county-level real estate investment, and individual and year fixed effects. Standard errors clustered at the county level in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

agricultural employment by approximately 4.7 percentage points ($0.923 \times 0.007 + 0.398 \times 0.031 + 0.028 \approx 0.047$). The corresponding gender-differential effect is 3.7 percentage points ($0.923 \times 0.006 + 0.398 \times 0.024 + 0.022 \approx 0.037$), bringing the total impact on female non-agricultural employment to 8.4 percentage points. These two numbers are used to calibrate the shape parameters of Fréchet distributions that govern labor mobility in the quantitative model.

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.

Results by marital status also help distinguish the guard-labor channel from an alternative interpretation that land reform raises migration by relaxing household financial constraints. By activating rental markets, reform generates rental income that can finance the upfront costs of migration, an income channel emphasized in the credit-constraint literature on rural-urban mobility (Cai, 2020). This channel, however, cannot explain why the migration effect is larger for married women than for unmarried women, unmarried men, and married men. Since rental income merely relaxes liquidity constraints, the associated income effect should not differ markedly across these groups without an additional group-specific mechanism. Guard labor, by contrast, is inherently a feature of joint employment decisions within couples: land reforms that remove the need to leave a member behind should disproportionately free married women, who are most likely to serve as guard labor. The concentration of the effect among married women in Table 4 therefore fits the guard-labor mechanism more naturally than a pure income effect.

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-differential impacts 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 in non-agriculture, as discussed in Section 2—we extract the RFPS subsample in which married couples can be reliably identified.[25] We consider four outcomes for households'

[25]RFPS 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.

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

	Rural Individual Sample RFPS (2003-2017)	
DV	Non-farm Employment (=1)	Migration (=1)
	(1)	(2)
Land Reform	0.011**	0.010**
	(0.005)	(0.005)
Reference Group: Married Male		
Land Reform * Unmarried Male	0.005	0.002
	(0.006)	(0.005)
Land Reform * Unmarried Female	0.013	0.005
	(0.008)	(0.005)
Land Reform * Married Female	0.009**	0.008***
	(0.004)	(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; individuals aged 18–55. The dependent variable in column (1) is a dummy equal to one if the surveyed individual spent any labor days in non-farm employment during the year; the dependent variable in column (2) is a dummy equal to one if the surveyed individual spent any labor days in out-migration work during the year. Individual controls include age, a marital-status dummy, and education level. Economic controls include logged county- and village-level GDP per capita, manufacturing and tertiary industry shares, and logged population size. All specifications control for indicators for whether the household has a child under age 7 and an elderly member above age 65, the weighted non-land-related *hukou* reform index (Gai et al., 2025), a dummy for adoption of the New Rural Pension Reform, weighted employment growth and GDP growth in destination cities, county-level real estate investment, and individual and year fixed effects. Standard errors clustered at the county level in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

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 (neither 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' Employment and Migration

	Multinomial Logistic Regression		Logit Regression	
	Non-Farm	Migration	Non-Farm	Migration
	(1)	(2)	(3)	(4)
	vs. No Non-Farm/Migration			
Wife only Non-Farm/Migration	0.067	0.085		
	(0.067)	(0.065)		
	vs. No Non-Farm/Migration			
Husband only Non-Farm/Migration	0.077**	0.121***		
	(0.033)	(0.036)		
	vs. No Non-Farm/Migration		vs. Husband only Non-Farm/Migration	
Both 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; married couples aged 18–55 with agricultural *hukou*. Columns (1) and (2) use multinomial logit models to compare wife-only non-farm employment/migration, husband-only non-farm employment/migration, and both spouses in non-farm employment/migration with the baseline category of both spouses in farming/neither spouse migrating. Columns (3) and (4) use logit models to compare both spouses in non-farm employment/migration with husband-only non-farm employment/migration. Individual controls include the husband's age, the wife's age, the husband's education level, and the wife's education level. Economic controls include logged county- and village-level GDP per capita, manufacturing and tertiary industry shares, and logged population size. All specifications control for indicators for whether the household has a child under age 7 and an elderly member above age 65, the weighted non-land-related *hukou* reform index (Gai et al., 2025), a dummy for adoption of the New Rural Pension Reform, weighted employment growth and GDP growth in destination cities, county-level real estate investment, and couple 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,[26] 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 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:

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

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

where Emp_{jdt} indicates whether non-agricultural *hukou* holder j in prefecture d in year t is currently employed in a paid, full-time 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 the regression in 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.

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	0.037	-0.017	0.096	-0.073	0.024
	(0.021)	(0.046)	(0.024)	(0.130)	(0.105)	(0.042)
Weighted Land Reform * Female	-0.028***	-0.051**	-0.042***	-0.071**	0.070	-0.033
	(0.008)	(0.026)	(0.009)	(0.036)	(0.046)	(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 uses urban samples from CHIP, 1995–2018, including individuals aged 18–55 with non-agricultural *hukou*. The dependent variables are employment status, defined as an indicator equal to one if the individual is employed in a full- or 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, an indicator for cohabitation with a partner, education level, work experience, ethnicity, city GDP per capita, secondary and tertiary industry shares, logged population size, and the interaction between total migrants received at destination d and a yearly trend. All specifications include prefecture 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 on urban female employment may reflect a substitution effect from the influx of rural female migrants. If so, urban women with less 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 mi-

grants have no more than a high school education.[27] Taken together, the results support the interpretation of a substitution effect.

4 Model

This section develops a multi-sector model with household-level employment decisions. The model has two sets of regions: an urban region hosting the non-agricultural sector 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 is urban households that only work in the non-agricultural sector. A fraction $1 - \alpha_u$ is rural households that have land use rights in their own rural region. This captures the land institutional setting in many developing countries and those with agricultural *hukou* in the context of China. To focus on the effect of land policy, we assume goods markets are frictionless and all households face the same prices for agricultural and non-agricultural goods.

Each rural household decides the sector of employment for both members. The sectoral choices imply that there will be four types of rural 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 their equilibrium shares:

$$\alpha^o + \alpha^f + \alpha^m + \alpha^b = 1 - \alpha^u \tag{3}$$

Land policy plays a crucial role in the sectoral choices of rural 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

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

reform. We next formalize this potential rental income through a parameter λ^{ir}, which varies by type of rural 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:

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

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:

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

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

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 receives $\lambda^{ir} I_k^r$:

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

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

$$Y_a = \phi Y_a^1 + (1-\phi) Y_a^0 \tag{8}$$

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

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

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

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:

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

4.2 Households

Rural Households Depending on their sectoral employment choices i, rural households in region r have the following joint utility function:

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

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 υ^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}:

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

The non-market hours include both home production hours and leisure hours. Depending on the type of household, the non-market good H^{ir} is produced as follows:

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

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

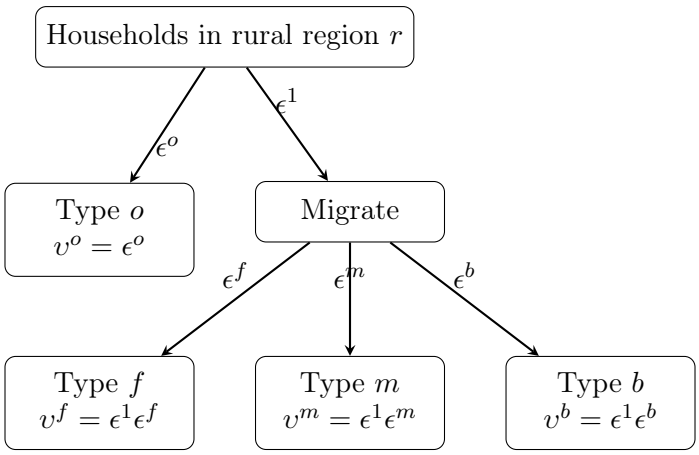
where $\delta_f \leq 1$ implies a lower efficiency in home sector when a female member moves to non-agriculture.[28] It captures the idea of Goldin (1995) that transitioning into non-farm work requires leaving the home, making it difficult to combine with home production. The parameter $\delta_s \leq 1$ captures the disutility when couples split into different sectors. Together, these parameters reflect social norms and attitudes toward women's roles in home production (δ_f) and perceptions of split couples (δ_s).

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^0), 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 decide which member(s) to migrate. Figure 4 illustrates the decision tree and specifies the corresponding values of υ^i.[29]

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

[29] Imbert et al. (2025), while not focusing on gender, also adopt a nested migration decision framework where households first choose whether to migrate, then whether to bring dependent family members, and finally the migration destination.

Figure 4: Agricultural Household Employment Decision Tree



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

Given the employment choice, the budget constraint for the consumption decisions of each type of households is:

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

Members remaining in the agricultural sector choose their hours of work. 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. The wedge $(1 - \mu)$ is a reduced-form measure that captures all mobility barriers, including travel costs, information frictions, labor market discrimination, lack of social subsidy when moving to an urban area.[30]

The total household income (I^{ir}) for household type i in region r is composed of both

[30] In the context of China, see Meng and Zhang (2001) on the empirically observed wage discount or Ngai et al. (2019) who show that migrants receive lower effective wages due to another aspect of the *hukou* system, whereby migrants do not receive local public services or subsidies such as healthcare, education, and housing.

labor and land income, the latter of which depends crucially on the land policy:

$$I^{or} = w^r_{fa} n^{or}_f + w^r_{ma} n^{or}_m + \lambda^{or} I^r_k, \tag{15}$$

$$I^{fr} = \mu w_{fu} n^{fr}_f + w^r_{ma} n^{fr}_m + \lambda^{fr} I^r_k, \tag{16}$$

$$I^{mr} = w^r_{fa} n^{mr}_f + \mu w_{mu} n^{mr}_m + \lambda^{mr} I^r_k, \tag{17}$$

$$I^{br} = \mu w_{fu} n^{br}_f + \mu w_{mu} n^{br}_m + \lambda^{br} I^r_k. \tag{18}$$

Urban Households The utility is given by

$$U^u = \ln \left[(c^u_a - \bar{c})^{\omega} \left(c^u_u\right)^{(1-\omega)} \right] + \theta \ln H^u; \quad H^u = (h^u_f)^{\xi_h} (h^u_m)^{1-\xi_h}. \tag{19}$$

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:

$$h^u_f = 1 - E_f \bar{n}, \quad h^u_m = 1 - E_m \bar{n}, \tag{20}$$

and the budget constraint for the urban household is:

$$p_a c^u_a + p_u c^u_u \leq I^u = w_{fu} \bar{n} E_f + w_{mu} \bar{n} E_m. \tag{21}$$

4.3 Equilibrium

The labor market clearing conditions by gender and sector satisfy:

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

$$N^{r}_{fa} = \alpha^{or} n^{or}_{f} + \alpha^{mr} n^{mr}_{f}, \tag{23}$$

$$N^{r}_{ma} = \alpha^{or} n^{or}_{m} + \alpha^{fr} n^{fr}_{m}, \tag{24}$$

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

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

The full derivation of the model is presented in the 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 examines the quantitative importance of the reform on structural transformation by gender. The model is calibrated using the land reform index constructed in Section 2, matching the aggregate household employment patterns in China. The cross-region empirical results on rural households guide the calibration of the key parameters for labor mobility, κ_1, κ_2, which then discipline the quantitative predictions for the aggregate effects of the reform.

5.1 Baseline Calibration

We first discuss parameters related to land reform and land policy. The share of rural reform regions ϕ is set to the national average of the land reform index divided by its maximum value of 3. As shown in Figure 2, 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 the reform. The values of ϕ are summarized in Table 7a. 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 entire household's land income, i.e., $\lambda^{f0} = \lambda^{m0} = 1$.

Table 7: Land Policy Parameters

(a) Share of reform regions ϕ

	2000	2005	2010	2015	2020
ϕ	0	0.212	0.365	0.591	0.997

(b) Land policy parameters in each region

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

Notes: ϕ is calculated as the national average of the reform index divided by 3. The land policy parameters are set such that that leaving one partner in agriculture can guard the full land income share.

We now turn to other directly set parameters. We normalize initial A_a and A_u in 2000, and the total endowment of land, K, to 1. The elasticity of substitution between female and male workers, η, is set to 3, which is around the average of existing estimates in the literature (Acemoglu et al., 2004). The land share, β, is set to 0.5, the reported land income share in China (Cao and Birchenall, 2013). Following Adamopoulos et al. (2024), we set $\omega = 0.02$ to capture the long run share of agriculture. The number of hours worked required in non-agriculture is set at 49 hours per week, i.e., $\bar{n} = 0.49$, based on the average hours in the non-agricultural sector calculated from the Census 2005. Finally, we set the share of urban households α^u using Census data on the share of non-agricultural *hukou*, reported in Appendix Figure A1a.

The rest of the parameters are internally calibrated. The Fréchet distribution parameters, κ_1 and κ_2, are chosen such that the cross-region differences in female and male non-agriculture employment share match the impact of reform identified from regressions estimates from Table 3: 8.4 percentage points and 4.7 percentage points. Since the regressions use data

spanning 2003 to 2017, we calculate the average of the model outcomes across all five model periods, weighted by the reform region share.[31] The implied values are $\kappa_1 = 0.90, \kappa_2 = 2.43$. 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.[32] We set θ to match the average urban employment rate (0.70). The subsistence consumption, $\bar{c}$, is set to match the value-added share of agriculture at 14.7% in 2000.

The remaining internally calibrated parameters are allowed to change over time. The growth rates of productivity A_a and A_u, denoted by g_a and g_u, are set to match the growth rate of labor productivity per worker in agriculture and non-agriculture respectively, calculated from the China Statistical Yearbook (CSY). The gender parameter ξ_u is calibrated to match the relative female-to-male wage ratio in non-agriculture;[33] ξ^h is calibrated to match the employment rate of urban females relative to urban males during 2000–2020. These two data moments are reported in Appendix Figure A1a.

Finally, we calibrate μ, δ_f, and δ_s to match the share of different types of rural households reported in Figure 1. Since the four types sum to the total share of rural households (see equation (3)), there are three independent targets: the share of households with both members working in agriculture ($\alpha^o/(1-\alpha^u)$), the share of households with female working in non-agriculture ($(\alpha^f+\alpha^b)/(1-\alpha^u)$), and the share of split households ($(\alpha^f+\alpha^m)/(1-\alpha^u)$). The calibrated value of μ increases substantially after 2015, reflecting the urban *hukou* reform since 2014 that significantly reduces migration barriers. δ^f is stable over time, while δ^s increases, reflecting changes in communication technology and social norms towards couples living separately.

To summarize, the baseline calibration matches the cross-region empirical results in Ta-

[31] We calculate $\frac{\sum_t x_t \phi_t}{\sum_t \phi_t}$, where x_t denotes our variable of interest and ϕ_t denotes the share of the reform regions at time t. This captures the fact that later periods have more reform regions in the empirical sample, thus carrying larger weights in the estimated coefficient.

[32] Our model allows workers to choose agricultural hours, but the RFPS only reports days. We use the relative days to proxy for relative hours in the model.

[33] We use the gender wage ratio from the Urban Household Survey from 2000 to 2009, and the gender wage ratio from the Chinese Household Finance Survey in 2011, 2013, 2015, 2017, 2019, and 2021.

ble 2, and the observed share of reform regions and the aggregate employment patterns from 2000 to 2020. Table 7, Appendix Figure A1b, and Table A4 summarize all parameter values.

5.2 The Guard Labor Mechanism

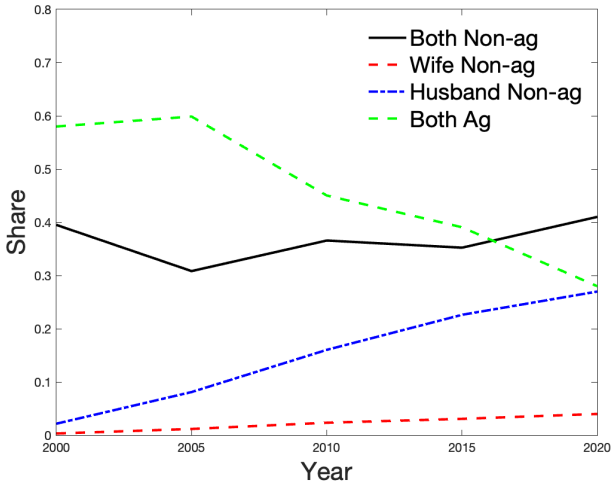
Our baseline model matches the employment trends of rural households in Figure 1. In the model, these aggregate trends reflect the underlying patterns in the reform and non-reform regions, together with the rising share of reform regions. Figure 5 presents the predicted employment patterns in each region type. In both regions, higher productivity and lower mobility barriers (higher μ) move workers from agriculture to non-agriculture over time. However, the reform regions have a much smaller share of households with both members remaining in agriculture. They also show a much larger gap between the share with both members in non-agriculture and the share with only the husband in non-agriculture. Weighted by the share of reform regions across the calibrated years, these patterns imply that reform increases the female non-agricultural employment share by 8.4 percentage points and the male by 4.7 percentage points. These figures are derived from empirical regression coefficients that the model matches by varying κ_1 and κ_2.

The "guard labor" mechanism, households keeping a member behind to protect household land income under insecure tenure, is central to generating the gender-differential impacts. To isolate its role, we consider a counterfactual in which we set the land policy parameters to $\lambda^{f0} = \lambda^{m0} = 0.5$ in the non-reform regions, meaning that the member staying behind can keep her/his own land share, but cannot protect the land income of the member who moves to non-agriculture. Keeping all other parameters at their baseline values, including the mobility parameters (κ_1 and κ_2) and the time path of reform region shares, reform now affects both genders nearly symmetrically: the female non-agricultural employment share increases by 7.7 percentage points relative to non-reform regions and the male share by 7.5 percentage points.

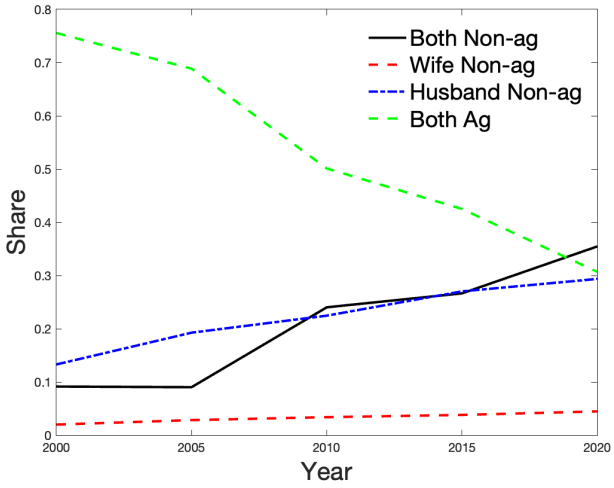
The intuition is as follows. In both the baseline and the "no-guarding" cases, women

Figure 5: Employment Patterns of Rural Households by Regions

(a) Reform Regions



(b) Non-reform Regions



Notes: This figure plots the shares of the four household types in reform and non-reform regions, α^{i1} and α^{i0}, respectively.

remain disproportionately in agriculture because it is more female-intensive ($\xi_a > \xi_u$), and when women leave agriculture, the production of non-market goods becomes less efficient ($\delta_f < 1$). In the baseline, the sectoral gender intensities and women's role in home production imply that women are more likely to take on the role of guard labor. Reform that eliminates the need for guarding thus has a stronger effect on women's transition out of agriculture. In the "no-guarding" case, this mechanism is absent and the reform is essentially gender-neutral.

5.3 Aggregate Effects of Land Reform

We then use our model to quantify the aggregate impact of the reform. From 2000 to 2020, the share of rural married females working in non-agriculture increases by 33.9 percentage points, from 11.1% to 45.0%. The share for rural married males increases by 45.6 percentage points, from 22.4% to 68.0%. We summarize these changes in Table 8.

Several time-varying parameters contribute to these patterns, and our goal is to isolate the impact of reform. We organize the parameters into four groups: (1) the share of the reform regions (ϕ), (2) productivity (A_a and A_u) and the share of urban households (α_u),

which govern the speed of structural transformation, (3) gender-neutral parameters (μ and δ_s), and (4) gender-specific parameters (ξ_u, ξ_h, and δ_f). Since our model is non-linear, there are interactions across these forces. To address this, we apply the Shapley–Owen decomposition, which evaluates the marginal contribution of each parameter group across all possible orderings in which the other parameter groups could be switched from their beginning of period to their end of period values, and averages over these orderings.[34] By construction, the resulting contributions sum exactly to the total observed change.

Table 8: Aggregate Changes in Employment Patterns of Rural Households

	Non-agricultural Employment Share	
	(1)	(2)
Change 2000–2020	Female	Male
	+33.9 pp	+45.6 pp
Contribution of		
Share of reform regions ϕ	+10.6 pp	+5.1 pp
A_a, A_u, and α_u	+23.5 pp	+23.5 pp
μ and δ_s	−1.8 pp	+16.8 pp
ξ_u, ξ_h, and δ_f	+1.5 pp	+0.1 pp

Notes: This table reports the results of the Shapley–Owen decomposition. Columns (1) and (2) report the changes in female and male non-agricultural employment shares, respectively.

Table 8 decomposes the change in the non-agriculture employment share of rural females and males. A few observations stand out. First, land reform (ϕ) is the only force that generates a substantially larger increase in female than male non-agricultural employment in both decades. Quantitatively, reform accounts for about 31% (=(10.6)/(33.9)) of the increase for females and 11% (=(5.1)/(45.6)) of the increase for males. Second, productivity growth and urbanization are the largest contributors to the rise in non-agricultural employment, but the impact is similar across genders.

Third, the gender-neutral parameters (μ and δ_s) together disproportionately encourage male migration out of agriculture, primarily driven by the increase in δ_s implied by our

[34] See Hubmer et al. (2023) for one recent example that uses this approach to decompose the factors contributing to the excess wealth of the top 0.1%.

calibration. A higher δ_s makes split migration more socially acceptable, which interacts with women's role in home production, inducing husband-only non-agricultural employment. Finally, the contributions of the gender-specific parameters (ξ_u, ξ_h, and δ_f) are quantitatively small.

Turning to urban households, the reform-induced relative increase in female labor supply to non-agriculture widens both the gender employment gap and the gender wage gap. Applying the same Shapley–Owen decomposition, Appendix Table A5 reports that the reform induces a 4.7 pp decline in the female employment rate relative to males and a 3.8 log-point decline in the female-male wage ratio. These contribute substantially to the observed increase in urban gender gaps: 1.9 pp in employment and 7.0 log points in wages.

In summary, the quantitative exercise illustrates that land rights insecurity, though gender-neutral in design, plays an important role in shaping the gender dimension of structural transformation through women's role as guard labor. Reform that strengthens land rights security contributes substantially to rural women's transition out of agriculture, and in turn to widening gender gaps among the urban population.

6 Conclusion

Exploiting quasi-natural experiments from two major Chinese land reforms, we show that improving land rights security induces more women than men to transition into non-agriculture, with the shift occurring through joint household employment in non-agriculture rather than husband-only transition. We develop a multi-sector, multi-region model with couples' labor allocation decisions that rationalizes these findings through women's role as guard labor. Using our reform index, we show that land reform is quantitatively important in closing the gap between men's and women's transition out of agriculture. These findings speak to the broader consensus that land institutions are critical for fostering resilient and equitable growth (World Bank, 2024).

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Appendix

A1 Data

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, the 2015 and 2020 censuses did not provide *hukou* type. We used whether an individual had rural land contracting rights to proxy for their *de facto hukou* type.

For individuals reporting their 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 including ethnic minority status, age, marital status, and high skill (education at college or above). Summary statistics for these variables are reported in the 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 18–55 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 or more (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 with respect to 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 the 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 regions, income, cropping patterns, population, and non-farm activities. Within each village, households were drawn by random sampling. Our study uses RFPS data from 2003 to 2017 to examine how land reform affects rural residents' migration, sectoral employment, and intra-household migration patterns. The Online Appendix Table B3 presents 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 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 the Online Appendix Table B4.

A2 Appendix Figures and Tables

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 an individual-level regression examining the gender difference in non-agricultural employment. Controls include minority status, age, age squared, marital status, and a high-skill indicator (college education or above). Source: the Census and One Percent Population Survey for 1990, 2000, 2005, 2010, 2015, and 2020. All specifications control for province fixed effects and year fixed effects. The sample includes agricultural-*hukou* holders aged 18–55. Standard errors are reported in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

Table A2: Determinants of Reform Timing

	Year of Reform	
	Contracting	Titling
	(1)	(2)
GDP (Logged)	-0.017	-0.023
	(0.016)	(0.015)
Population size (Logged)	-0.180	-0.162
	(0.295)	(0.276)
Agricultural share	-0.614	-0.476
	(0.913)	(0.840)
Manufacturing share	-0.013	0.004
	(0.017)	(0.016)
Fiscal Revenue (Logged)	0.038	0.041
	(0.074)	(0.088)
Real Estate Investment (Logged)	-0.304*	-0.419***
	(0.183)	(0.104)
Land Dispute	-0.225**	-0.278***
	(0.093)	(0.096)
Average farm size	0.092	0.075
	(0.096)	(0.097)
Urbanization Rate	-0.017	-0.023
	(0.016)	(0.015)
Distance to First Tier Cities	-0.768	-0.555
	(0.876)	(0.805)
Distance to Railway	-0.180	-0.162
	(0.295)	(0.276)
Total out-migration (Logged)	0.031	0.024
	(0.051)	(0.078)
Female out-migration (Logged)	0.008	0.011
	(0.029)	(0.041)
Observations	2813	2813

Notes: This table reports county-level cross-sectional regressions examining the determinants of the timing of adoption of the two land reforms: the contracting reform and the titling reform. All explanatory variables are measured in either 2000 or 2003, depending on data availability, to ensure that they are predetermined relative to reform implementation. Data on GDP, population, agricultural share, manufacturing share, fiscal revenue, real estate investment (defined as total investment in the real estate sector), and the urbanization rate are drawn from the *China Statistical Yearbook for Regional Economy* and the *China County (City) Socio-Economic Statistical Yearbook*. Land disputes, defined as the number of land dispute cases filed with local courts, are drawn from *China Land and Resources Statistical Yearbook*. Total out-migration and female out-migration are constructed from the 2000 Population Census. Robust standard errors are reported in parentheses; * $p<0.1$, ** $p<0.05$, *** $p<0.01$.

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

	Rural Household Sample, RFPS (1995–2017)		
Dependent Variables	Rental Out (=1)	Rental Area (Logged)	Rental Income (Logged)
	(1)	(2)	(3)
Land Reform Index	0.100***	0.129**	0.097**
	(0.028)	(0.050)	(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 Survey, 1995–2017. The dependent variables are household-level measures of land rental activity: (1) an indicator for whether the household rented out land in a given year; (2) the amount of land rented out; and (3) income from renting out land. All specifications control for household fixed effects and year fixed effects. Standard errors clustered at the county level are reported in parentheses; * p<0.1, ** p<0.05, *** p<0.01.

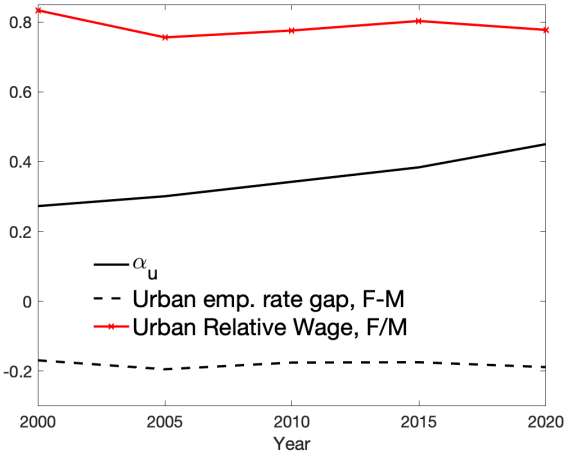
Table A4: Baseline Parameters

Para.	Values	Moments
Directly calibrated		
η	3	Acemoglu et al. (2004)
$\bar{n}$	0.49	49 hours per week (Census 2005)
β	0.50	Land share from Cao and Birchenall (2013)
ω	0.02	Adamopoulos et al. (2024)
α_u	varying	Non-agricultural hukou share (Census)
Internally calibrated, non-time-varying		
$\bar{c}$	0.44	Value added share of agriculture in 2000 (CSY)
κ_1	0.90	Regression coefficients
κ_2	2.43	Regression coefficients
θ	2.05	Employment rate; non-agricultural hukou holders (UHS 2000)
ξ_a	0.48	Agricultural days per worker, female over male (RFPS 2003)
g_a	6.7%	Labor productivity growth based on employment and real GDP (CSY)
g_u	6.5%	Labor productivity growth based on employment and real GDP (CSY)
Internally calibrated, time varying to match 2000, 2005, 2010, 2015, and 2020		
ξ_h	varying	Female over male employment rate; non-ag hukou (Census)
ξ_u	varying	Female over male wage rate in non-agriculture (UHS and CHNS)
μ	varying	Share of α_o (Census)
δ_f	varying	Share of $\alpha_f + \alpha_b$ (Census)
δ_s	varying	Share of split households $(\alpha_f + \alpha_m)$(Census)

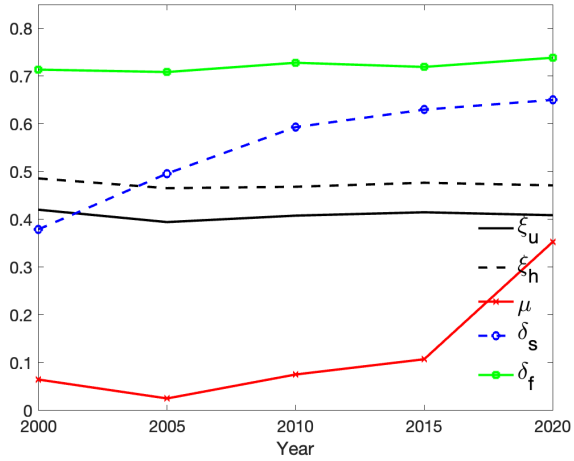
Note: Values are rounded to two decimal places. See Figure A1b for values of the time-varying parameters.

Figure A1: Time Varying Data Moments and Parameters

(a) Time Varying Data Moments



(b) Baseline Time-varying Parameters



Notes: Panel (a) of this figure plots three time-varying data moments: the share of urban households, the female-to-male wage ratio in non-agricultural employment, and the employment rate of urban women relative to urban men over 2000–2020. Panel (b) plots the following time-varying parameters: the weights of female labor in non-agricultural and home production (ξ_u and ξ_h), the exogenous labor market wedge μ, the efficiency of non-market production when women leave agriculture, δ_f, and the efficiency of non-market production when couples split across sectors, δ_s.

Table A5: Changes in Gender Gaps Among Urban Households, 2000-2020

	Female relative to male	
	(1) Employment	(2) Wage ratio
	−1.9 pp	−7.0 log pt
Contribution of		
Share of reform regions ϕ	−4.7 pp	−3.8 log pt
A_a, A_u, and α_u	−5.0 pp	−2.5 log pt
μ and δ_s	+7.1 pp	+4.6 log pt
ξ_u, ξ_h, and δ_f	+0.6 pp	−5.3 log pt

Notes: The first column reports the urban female employment rate minus the urban male employment rate. The second column reports the change in the log wage ratio between urban females and males, $\log(w_{fu}/w_{mu})$.

Online Appendix–Not For Publication

B1 Model Derivations

B1.1 Consumption Choices

The utility function implies the relative consumption:

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

Together with the budget constraint, the consumption decisions are:

$$c_a^i - \bar{c} = \frac{\omega\left(I^i - p_a\bar{c}\right)}{p_a}; \qquad c_u^i = \frac{(1-\omega)\left(I^i - p_a\bar{c}\right)}{p_u}. \tag{28}$$

The utility achieved from consumption and non-market goods becomes:

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

where $B = \left(\frac{\omega}{p_a}\right)^{\omega}\left(\frac{1-\omega}{p_u}\right)^{1-\omega}$ is identical for all households.

B1.2 Rural Households

We solve the labor allocation decisions in two steps. First, we derive the optimal market and non-market work for each of the four types of rural households $i = o, f, m, b$, taking wage rates (w_{fa}^r, w_{ma}^r, w_{fu}, and w_{mu}), land income (I_k^r), and the corresponding allocation rules (λ^{ir}) as given. Second, we solve for the share of each type of households.

Type o Households Given the time constraint (Equation 12), a type o household solves:

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

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

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

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

They imply the following relative home hours:

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

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

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

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

Type f Households Given the time constraint (Equation 12), a type f household 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 male member's choice of hours is:

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

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

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

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

Type m Households Given the time constraint (Equation 12), a type m household solves:

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

where $n_m^{mr} = \bar{n}, h_m^{mr} = 1 - \bar{n}$. The first-order condition of the female member's hours choice:

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

Together with Equation 12, we obtain:

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

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, $n_f^{br} = n_m^{br} = \bar{n}, h_f^{br} = h_m^{br} = 1 - \bar{n}$. The utility is:

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

Choice Among the Four Types We assume a nested structure in which households first decide whether at least one member migrates and then choose the which member/members migrate, as illustrated in Figure 4. Let V^{ir} denote the indirect utility derived from the

optimal employment and time allocations of a type i household. We have:

$$\ln(V^{ir}v^i) \equiv \max U^{ir}(v^i), \quad i = o, f, m, b. \tag{38}$$

$$v^o = \epsilon^o, \quad v^f = \epsilon^1\epsilon^f, \quad v^m = \epsilon^1\epsilon^m, \quad v^b = \epsilon^1\epsilon^b. \tag{39}$$

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

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

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

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

In general, we have

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

To solve the first stage, we require an expression for the expected utility conditional on migration, $\mathbb{E}\left[\max(V^f\epsilon^f,\ V^m\epsilon^m,\ V^b\epsilon^b)\right]$. Under the assumption that ϵ^i follows a Fréchet

distribution with shape parameter κ_2, it follows that:

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

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

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

which is the CDF of a Fréchet distribution with shape parameter κ_2 and scale parameter S:

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

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

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

With this expression, we can solve the first-stage choice. They choose the option that maximizes their expected utility. Since ϵ^1 and ϵ^o follow independent Fréchet distributions with parameter κ_1, we can calculate $\alpha^{or}, r = 0, 1$ as:

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

Together with equation (40), we have:

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

where ϕ denotes the share of reformed rural regions ($r = 1$). This determines the share of different types of rural households in the economy.

B1.3 Urban Households

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

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

where E_f^u and E_m^u denote urban female and male employment rates, respectively. The household head chooses the optimal E_f^u and E_m^u and the first-order conditions are:

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

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

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

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

Together with the household time constraint, this implies:

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

Given h_f^u, we can solve for h_m^u, and hence for 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 of giving 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, resulting in 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* statuses into a unified *hukou* type—reforms that several localities had already begun experimenting with as early as 2007 (e.g., Chongqing Municipality in 2007 and 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 were able to 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.

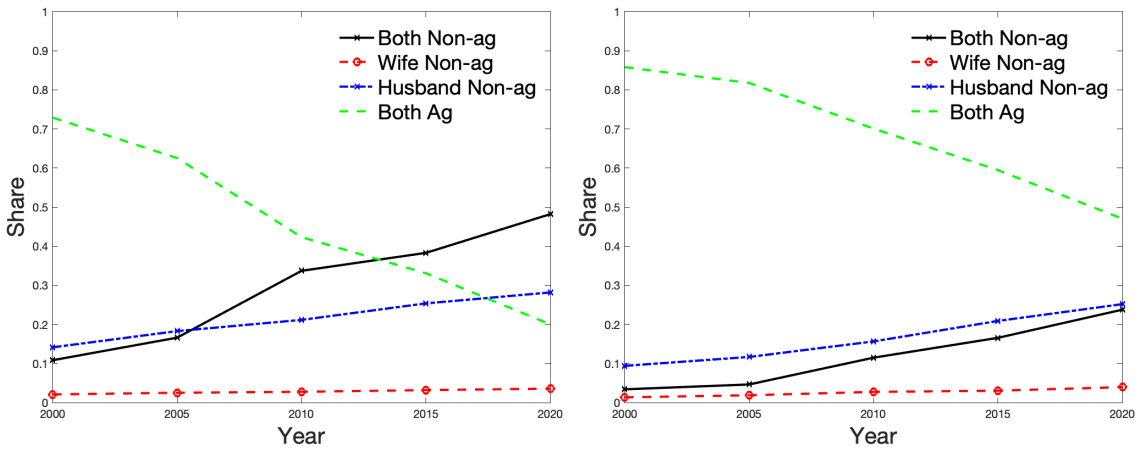
In summary, while these reforms represented some institutional adjustments, they did not fundamentally alter the *hukou* system or eliminate the rural–urban divide, especially with respect to rural land property rights. Following the 2014 reform, the 2015 and 2020 censuses ceased reporting individuals' *hukou* type. We therefore use rural land tenure rights

as a proxy for *de facto hukou* type. In the 2015 data, 62% of respondents held land rights, indicating that *hukou* unification did not substantially change the underlying land use rights of the former agricultural-*hukou* population.

B3 Additional Figures and Tables

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

(a) Young (aged 17–44) (b) Old (aged 45–64)



Notes: Census and One-Percent Population Survey. Pairs of rural husbands and wives aged 17–44 (young cohort) in Panel A and 45–64 (old cohort) in Panel B. "Both Non-ag" denotes that both the husband and wife are employed in non-agricultural sectors. "Wife Non-ag" denotes that the wife works in non-agriculture while the husband does not. "Husband Non-ag" denotes that the husband works in non-agriculture while the wife does not.

Figure B2: 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8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）

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第一章　总　　则
第二章　家庭承包
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第二节承包的原则和程序
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第四节土地承包经营权的保护
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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　国家保护集体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。

(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：2019-09-16　　访问次数:3285　　信息来源：省林业局　　字号：大 中 小　打印本页

(2006年9月30日浙江省第十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
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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。

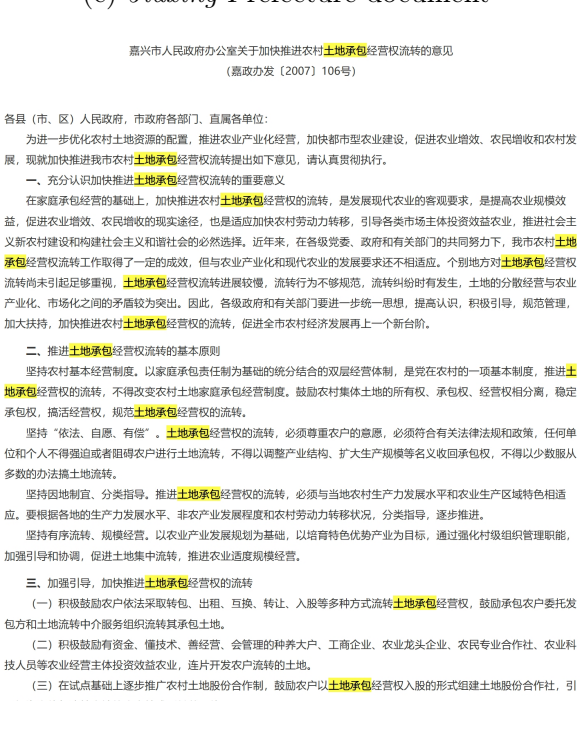
第五条　各级人民政府应当加强对农村土地承包工作的领导，维护农村土地承包当事人的合法权益，促进农业、农村经济发展和农村社会稳定。农村土地承包管理日常工作经费应当列入各级财政预算。

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。

(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，各级政府和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一个新台阶。

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。

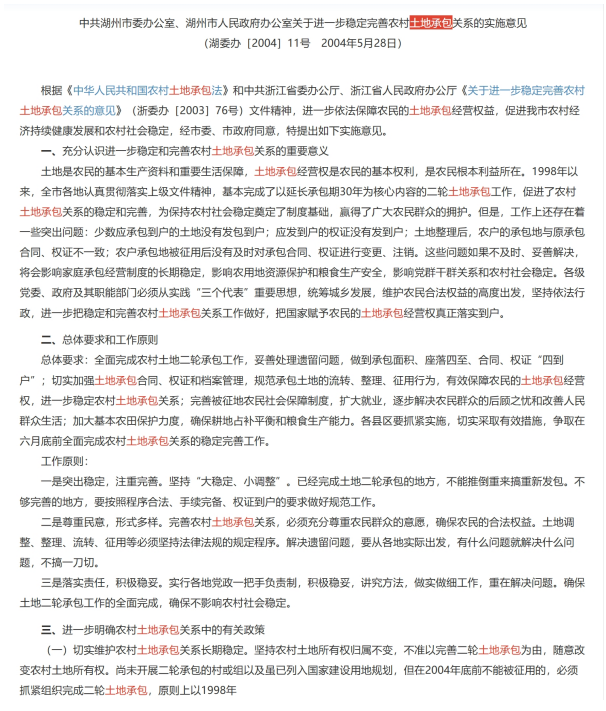
三、加强引导，加快推进土地承包经营权的流转

（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、会管理的种养大户、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

(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和重要生活保障，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，赢得了广大农民群众的拥护。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包到户；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户。

二、总体要求和工作原则

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。

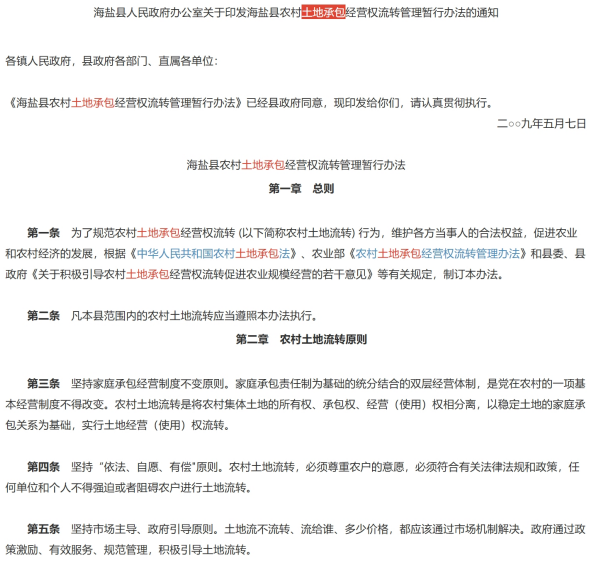
二是尊重民意，形式多样。完善农村土地承包关系，必须充分尊重农民群众的意愿，确保农民的合法权益。土地调整、整理、流转、征用等必须坚持法律法规的规定程序。解决遗留问题，要从各地实际出发，有什么问题就解决什么问题，不搞一刀切。

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○○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。

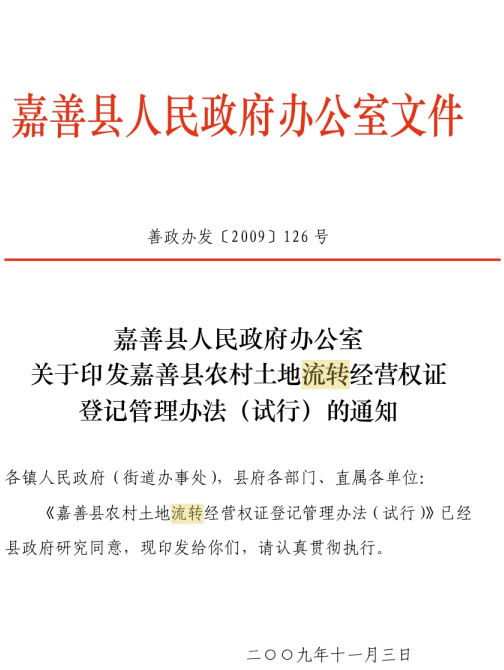
第二章　农村土地流转原则

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○○九年十一月三日

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Notes: Source: pkulaw.com. Panels (a)–(f) show examples of policy documents on land contracting reform (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 from the Census and the One Percent Population Survey for 1990, 2000, 2005, 2010, 2015, and 2020.

Table B2: CHIP (Urban): Summary Statistics

Variable	N	Mean	SD	Min	Max
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), urban sample, non-agricultural *hukou* individuals aged 18–55.

Table B3: RFPS: Summary Statistics

Variable	N	Mean	SD	Min	Max
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
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
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), agricultural *hukou* individuals aged 18–55.

Table B5: Keywords: Search Documents from pkulaw.com

Keywords	Total	Province	City	County
	Land Contracting Reform			
Land Contracting ("土地承包")	1390	182	646	562
Land Transfer ("土地流转" *or* "土地转让")	574	142	289	143
Land Contracted Management ("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			
Land Titling ("土地确权")	303	27	199	77
Land Certification ("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

	Rural Individual Sample RFPS (2003–2017)			
	Aged between 18 and 44		Aged between 45 and 55	
Dependent Variables	Non-farm (=1)	Migration (=1)	Non-farm (=1)	Migration (=1)
	(1)	(2)	(3)	(4)
Land Reform	0.004	0.006	0.020***	0.015***
	(0.005)	(0.004)	(0.006)	(0.004)
Land Reform * Female	0.014***	0.008***	-0.008	-0.004
	(0.004)	(0.003)	(0.005)	(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 equal to one if the surveyed individual devoted any labor days to non-farm employment during the year, while the dependent variable in columns (2) and (4) is a dummy variable equal to one if the surveyed individual devoted any labor days to out-migration work during the year. Control variables include age, a marital-status dummy, and education level. All specifications include individual fixed effects and year fixed effects. Standard errors clustered at the county level are reported 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

	Rural Individual Sample					
	All	High School and Below	College and Above	All	High School and Below	College and Above
Dependent Variables	Non-farm (=1)			Migration (=1)		
Panel A	CFPS (2010–2020)					
	(1)	(2)	(3)	(4)	(5)	(6)
Land Reform Index	0.016**	0.016**	0.029	0.013**	0.013**	0.024
	(0.007)	(0.007)	(0.020)	(0.007)	(0.007)	(0.042)
Land Reform Index * Female	0.009**	0.009**	0.012	0.008***	0.008***	0.013
	(0.004)	(0.004)	(0.023)	(0.003)	(0.003)	(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	0.024	-0.139*	0.009	0.008*	-0.008
	(0.027)	(0.028)	(0.080)	(0.010)	(0.005)	(0.012)
Land Reform Index * Female	0.017***	0.019***	-0.01	0.014*	0.015***	0.007
	(0.004)	(0.004)	(0.013)	(0.009)	(0.004)	(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 uses the CFPS Rural Sample (2010–2020), while Panel B uses the CHIP Rural Sample (1995–2018). The dependent variables are dummy variables equal to one if the surveyed individual devoted any labor days to non-farm employment during the year or devoted any labor days to out-migration work during the year. Columns (1), (4), (7), and (10) use the full sample, while columns (2)–(3), (5)–(6), (8)–(9), and (11)–(12) use subsamples categorized by education level. Control variables include age, a dummy variable for 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	Male	Female	Male
	(1)	(2)	(3)	(4)
Age	-0.008***	-0.004***	-0.013***	-0.011***
	(0.001)	(0.000)	(0.001)	(0.001)
With Child under 7 (=1)	-0.081***	0.014**	-0.068***	0.013
	(0.008)	(0.007)	(0.007)	(0.008)
With Elderly above 65 (=1)	0.029***	0.002	0.010	0.001
	(0.009)	(0.007)	(0.008)	(0.010)
Married (=1)	-0.045***	-0.021***	-0.046***	-0.051***
	(0.010)	(0.008)	(0.009)	(0.010)
Education Level	0.038***	0.016***	0.038***	0.017**
	(0.008)	(0.006)	(0.007)	(0.008)
Health	0.016	0.020**	-0.001	-0.001
	(0.012)	(0.010)	(0.010)	(0.011)
Year FE	Yes	Yes	Yes	Yes
Observations	256875	275275	256875	275275

Notes: Rural Fixed Point Survey, 2003–2017, individuals aged 18–55. The dependent variable in columns (1)–(2) is a dummy equal to one if the surveyed individual devoted any labor days to non-farm employment during the year, while the dependent variable in columns (3)–(4) is a dummy equal to one if the surveyed individual devoted any labor days to out-migration work during the year. Individual control variables include age, household presence of children under age 7, household presence of elderly individuals above age 65, marital status, education level, and self-reported health; individuals reporting poor health are excluded. All specifications control for year fixed effects. Standard errors clustered at the county level are reported 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	0.044	-0.038	0.068	-0.084	0.047
	(0.031)	(0.042)	(0.047)	(0.074)	(0.102)	(0.052)
Weighted Land Reform * Female	-0.034***	-0.068***	-0.048***	-0.087**	0.013	-0.028
	(0.009)	(0.022)	(0.008)	(0.042)	(0.036)	(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 the urban samples from CHIP (1995–2018), including individuals aged 18–55 with non-agricultural *hukou*. The dependent variable is an employment indicator in columns (1), (3), and (5), equal to one if the respondent is employed in a full- or part-time paid job at the time of the survey, and logged annual wage income in columns (2), (4), and (6). Control variables include gender, age, a cohabitation dummy, education, work experience, ethnicity, city GDP per capita, the shares of secondary and tertiary industries, the logged population size, 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.