# Do Farmers Adopt Fewer Conservation Practices on Rented Land? Evidence from Straw Retention in China

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# Do Farmers Adopt Fewer Conservation Practices on Rented Land? Evidence from Straw Retention in China

**Abstract**

We examine how land tenure arrangements affect Chinese crop farmers' adoption of straw retention, a key conservation practice promoted by the Chinese government in part to curb rising air pollution. Using data from a 2016 farmer household survey covering 1,659 crop plots in Henan Province in central China, we analyze whether farmers are less likely to adopt straw retention on rented plots compared to own-contracted plots. To address the potential endogeneity of the choice of renting from others, we use an instrument exploiting the role of remittance income from household members migrated to cities in a bivariate probit model and a control function approach, respectively. Our main results reveal that the Chinese crop farmers' likelihood of adopting straw retention were almost cut in half on rented plots compared to their owned plots, assuming the assumptions for biprobit or control functions hold. This suggests greater attention is needed to examine the spillovers across agricultural and environmental policies as China pushes for both a nationwide land rental market and more sustainable agricultural practices.

**Keywords**: land tenure; conservation practice; control function; bivariate probit; Chinese agriculture

**JEL Codes**: Q15, Q18, O13

## 1. Introduction

Land tenure security is crucial in promoting the adoption of various conservation practices, including conservation tillage (Lee and Stewart, 1983; Soule et al., 2000), contour farming (Soule et al., 2000), conservation crops (Fraser, 2004), and stone terraces or soil bunds (Gebremedhin and Swinton, 2003). Arguably, more secure land tenure, which often refers to complete, permanent, or durable ownership of farmland, leads to higher willingness to adopt conservation practices, especially those practices with long-term soil fertility benefits. This is likely because greater land tenure security increases the likelihood of farmers reaping the benefits of land investments, which are often long-term (Feder et al., 1988; Soule et al., 2000; Fraser, 2004; Kabubo-Mariara et al., 2010). Despite the perceived significance of land tenure, there is a lack of evidence of the role of land tenure security in conservation practice adoptions, especially in developing countries such as China. This, in part, results from varying definitions of land tenure security and heterogeneity in the ownership and tenure systems across different countries (Kabubo-Mariara et al., 2010).

In China, land tenure security has particular relevance because, under the current Household Responsibility System (HRS), agricultural land is owned by the collectives at the village level, and each eligible farmer household is granted a land contract right to farm a village-allocated land parcel with up to 30 years of tenure (Hu, 1997). The distinct nature of rural tenure systems confronts Chinese farmers with greater land tenure insecurity, which could potentially hinder farmers’ investments in production and conservation practices, especially those with a long time horizon. For instance, researchers have found that frequent land reallocation by the village collectives to accommodate a growing rural population often dampens the stability and security of land tenure, resulting in a very uncertain land tenure length with an effective length of much less than 30 years (Liu et al., 1998; Brandt et al., 2002; Tan et al., 2006).

In fact, since Jacoby et al. (2002), many researchers have examined the impacts of land tenure insecurity in Chinese farmers' production decisions, with a focus on input use such as organic fertilizer (e.g., Jacoby et al., 2002), land use efficiency (Zhang et al., 2011; Leight, 2016), and forest output efficiency (Salant and Yu, 2016). In contrast, evidence of the impacts of land tenure in conservation practice adoption in China is relatively scarce. Wang et al. (2010) investigate the determinants of adopting conservation tillage as well as residue retention; however, they do not consider land tenure as a driving factor. Liu and Huang (2013) were among the first to assess how land tenure security affects conservation practice adoption and show that the ownership of land is slightly positively associated with increased likelihood of using contour cultivation.[1]

However, to the best of our knowledge, no study has explicitly modeled the role of land tenure in the adoption of straw retention, an increasingly important component of conservation practices (Pittelkow et al., 2015). Straw retention (i.e., returning straw to the field) refers to a residue management strategy of covering the crop straws on the soil surface after harvest, which has been proven to improve long-term soil productivity (Lu, 2015; Wang et al., 2015), and boost

[1] Chinese farmers do not own the farmland. As will be discussed in detail later, the "ownership" of a plot by a farmer household in China is actually represented as the land contract right.

yield (Huang et al., 2013; Wang et al., 2015). However, burning straw, which generates large amounts of PM 2.5 (Tao et al., 2013; Shon, 2015; Chen et al., 2017), is still common in China and has become one of its biggest environmental problems despite increased public attention, fines, and penalties. In an attempt to reduce open burning of straw, thus curbing PM 2.5 pollution, the Chinese government has recently undertaken various measures supporting sustainable utilization of crop residues. For instance, in May 2015, the government announced a straw retention subsidy pilot project, which offered a cash payment to farmers to encourage straw retention adoption. The project was extended to all of China in 2016.[2] In addition, we analyze land tenure insecurity in the context of new policies in which China increasingly promotes rural land transfers among farmers through a land rental market. There is a lack of understanding on whether and how farmers make production and conservation decisions differently on rented land obtained through the rental market versus their own-contracted farmland allocated by collectives.

This study aims to examine whether and how land rental decisions, which are increasingly prevalent under the new rural land transfer market, affect Chinese crop farmers' adoption of straw retention. We hypothesize that Chinese crop farmers are less likely to adopt straw retention and other conservation practices on land rented from others due to their perceived less secure land tenure arrangement. We argue that farmers on rented fields face less secure land tenure due to the short-term nature of leasing contracts, and thus have lower willingness to undertake a

[2] The pilot project in 2015 was undertaken in five provinces—Anhui, Shandong, Hunan, Sichuan and Zhejiang. Our study analyzes crop and crop residue choices by farmers in Henan province for the 2015 growing season, which was not included in the pilot project. In addition, the progress on the subsidy project varied dramatically among different regions. Based on the experiences of our interviewers, most farmer households in Henan were not aware of this subsidy program at the time of the survey.

conservation practice compared to land formally allocated to them via long-term contracts from the collectives.

To test our hypothesis, we use a 2016 rural household survey covering 1,659 crop plots in Henan Province in central China and analyze whether farmers are less likely to adopt straw retention on rented plots compared to own-contracted plots. To address the potential endogeneity of the land tenure variable, we rely on an instrument that proxies the remittance income from the household members who migrated into cities. Arguably, a higher ratio of migrants' income over agricultural profits, conditional on available laborers and farmers' household income, would lead to a smaller likelihood of renting from others but not directly shift conservation practice choices. Following Wooldridge (2010), we employ both a bivariate probit and a control function approach using the above instrument to address the endogenous explanatory variable of land tenure.

Our main results confirmed that Chinese crop farmers are less likely to adopt straw retention on fields rented from others compared to own-contracted plots. In particular, the bivariate probit and control functions controlling for the endogeneity of the land tenure variable suggest that on average, the likelihood of Chinese crop farmers adopting straw retention on rented fields are almost only half compared to that for their own-contracted fields. In contrast, simple probit regressions with endogenous land tenure variable show that a rented plot is associated with an eight percent reduction in the probability of adopting straw retention after harvest throughout 2015. Overall, our results confirm our hypothesis of less conservation practice adoption given less secure land tenure arrangements, and are comparable to many studies in other countries.

This study contributes to the literature of conservation practice adoption by quantitatively examining the link between land tenure security and straw retention adoption in China for the first time. In particular, our analysis provides evidence that Chinese crop farmers are significantly less likely to adopt straw retention, a critical conservation practice, on rented plots compared to those own-contracted plots. More importantly, our research is of great policy relevance since it reveals the previously overlooked, potentially negative interconnection between two policies promoted by the Chinese government—encouraging the adoption of straw retention and expanding the rural land rental market—and offers insights into how the government can better promote and balance them. Finally, the significant larger average partial effect in the main specifications suggests that one needs to control for the endogeneity of the land tenure variable using the biprobit or control function approach.

The rest of this article is organized as follows: Section 2 provides a brief review of China's land system and the development of the land rental market. Section 3 introduces the conceptual framework. Section 4 introduces the empirical implementation strategy. Section 5 describes the data used in this study and empirical implementation. Section 6 discusses the empirical results. Section 7 provides concluding remarks.

## 2. Land Tenure and Land Rental Market in China

China prohibits private land ownership. The current HRS was introduced in the early 1980s and allocates a parcel of contracted farmland to each eligible rural household on the basis of household size, which is referred to as the land contract (and use) right. Nevertheless, the allocated land is owned by village collectives represented by villager committee or township government (Hu, 1997). Farmers are free to make their own agricultural production decisions, though they are not permitted to convert the land to non-agricultural use. In the early stages of HRS, land contracts only lasted for a one- or two-year period, which led to significant land tenure insecurity and discouraged farmers from making land improvements (Krusekopf, 2002). Realizing this limitation, the Chinese government lengthened the land contract terms to 15 years, further extending it to 30 years in 1993 (Zhang et al., 2011).

However, the increase in duration of land contracts did not necessarily improve the tenure security for rural households for two reasons: (*a*) village collectives periodically reallocate village land through administrative means to reach egalitarian goals in response to household demographic changes, even in the midst of land contract periods (Liu et al., 1998; Brandt et al., 2002; Tan et al., 2006); and, (*b*) collective allocations efficiency is negatively impacted by an increasing number of rural migrants going off-farm and working outside the village, which tends to lead to productive inefficiency (Benjamin and Brandt, 2002).

In response to the rising need of more secure land tenure, the central government has codified a framework for the protection of land rights and development of a land rental market, including the Land Management Law (1998), the Rural Land Contracting Law (2003), and the Property Law (2007). Subsequently, the land rental market took off rapidly, with a participation rate of about 10 percent in 2001 (Deininger and Jin, 2005), and stayed around that level for almost a decade. However, in many areas, farmland continued to be illegally reallocated by local officials (Zhang et al., 2011). With the steady increase in the number of rural workers migrating to urban regions, more and more rural households no longer have the need to keep all contracted land due to lack of laborers, which to a large extent stimulates the development of a rural rental market. Figure 1 shows the national growth of transferred rural land in China. The percentage of transferred land over the total contracted land to rural households has tripled from less than 10 percent in 2008 to about 33 percent in 2015.

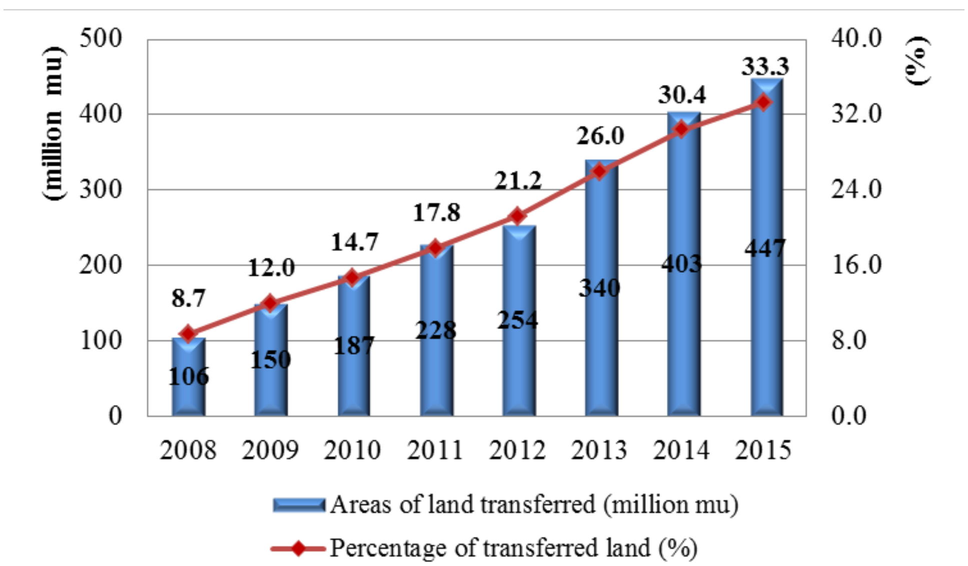


Figure 1. Scale of land Transfers in China, 2008–2015.
(Source: Author's calculation; Ye, 2015; Ministry of Agriculture, 2014, 2015, 2016)

More recently, the Chinese government has formally announced the intent to protect and split rural land rights into three parts: ownership, contract, and use. Ensuring permanent collective land ownership, the government allows rural households to lease out the land use right while maintaining the original land contract with the village government, which is largely to stimulate land transfer through the rental market.[3] However, the decomposition of rural land rights may raise a further question. The transferred land use right, which depends on how the leasing contract is made between rural households, may not be as secure as the land contract right.

While rural land transfer allows more flexible allocation of farmland across farmers, potentially moving from inefficient producers to more efficient producers, it remains uncertain how the tenants, who obtained the farmland via the rural land transfer market for a finite amount of time, would treat these land parcels differently compared to those owned and operated by its original contractees. Enlightened by previous literature on limited investments on rented land (Soule et al., 2000; Fraser, 2004; Kabubo-Mariara et al., 2010), we assume that tenants of rented land would have less incentive to make long-term investments, such as adopting conservation practices (i.e., straw retention) on these parcels.

## 3. Conceptual Framework

Following previous works by McConnell (1983) and Soule et al. (2000), we develop a three-stage model to analyze the adoption behavior of straw retention under different land tenure arrangements.

[3] See the 2014 No. 1 Policy Document available at http://www.moa.gov.cn/zwllm/zwdt/201401/t20140120_3742582.htm (in Chinese).

In the first stage, after the crop is harvested for the current growing season, the farmer household decides on the treatment of crop residues (straws), which involves a treatment cost denoted by $C_j$. In this context, we designate $j = s$ for adopting straw retention and $j = n$ for non-adoption. Straw retention requires straws to be smashed before they are covered and mixed with the soil surface, which takes additional machinery or labor costs and thus $C_s > 0$. If straws are not returned to the fields, farmers can either burn, discard, or use straws for other off-farm purposes such as feed, fuel, and to sell. Any of these treatments may also involve some costs including collecting, storing, and transporting of straws and thus $C_n > 0$. It is difficult to determine which of $C_s$ and $C_n$ is higher, as $C_j$ depends on specific straw treatment as well as the crop type.

Straw treatment in the first stage will also affect farming for the next season. Let $\pi_j$ be the second-season net returns under first-stage straw treatment, $j$. Although straws can improve soil quality, acting similar to fertilizer, it takes time for straw to decay, which may affect the next cropping given that farmers are often time-pressed between harvesting and sowing (Peng et al., 2016; Zhang et al., 2017), especially during autumn and in double-crop rotation areas like Henan Province. During the process of decaying, it may also lead to insect damage, which potentially increases the cost of pesticide use and hurts crop yields (Cai et al., 2011; Dinardo-Miranda and Fracasso, 2013; Carvalho et al., 2017). Therefore, there are many ways that straw retention may affect short-term profits, as a result it is common that $\pi_s < \pi_n$.

For the final stage, the farmer household is concerned with the terminal value of its farmland, denoted by $V_j$. Research has shown that straw retention can help reduce soil erosion and improve fertility and productivity over time, thus better retaining the long-term value of the land. Therefore, $V_s > V_n$.

Assume that a farmer household (either contractee or renter) selects a straw treatment option, $j$, to maximize the present value, $PV_j$, of the three terms previously discussed on a cultivated plot (either own-contracted or rented), as shown by the following equation:

$$\max_{(j)} PV_j = \pi_j - C_j + \lambda V_j/(1+r)^T, \tag{1}$$

where $r$ represents the discount rate and $T$ the number of periods. Following Soule et al. (2000), $\lambda$ is included as a tenure-security indicator weighting the third term that measures the farmer household's belief about its ability to use or sell the land in the future. Therefore, the more secure the land tenure, the higher $\lambda$ and greater importance of the long-term land value in the household's decision-making process.

Based on equation (1), it is optimal for a rational farmer household to adopt straw retention when

$$\pi_s - C_s + \lambda V_s/(1+r)^T > \pi_n - C_n + \lambda V_n/(1+r)^T \tag{2}$$

or

$$\lambda \Delta V/R > \Delta \Pi, \tag{3}$$

where $\Delta \Pi = (\pi_n - C_n) - (\pi_s - C_s)$, $\Delta V = V_s - V_n$ and $R = (1+r)^T$. Since $V_s > V_n$ and $R$

is positive, condition (3) actually suggests that the adoption of straw retention hinges on whether the potential short-term profit loss (i.e., the right-hand side) can be offset by the perceived gains in long-term land value (i.e., the left-hand side). That is, with higher $\lambda$, or more secure land tenure, the farmer household is more likely to undertake long-term improvement activities such as straw retention, as stated in our hypothesis.

An empirical estimate of the value of $\lambda$ is lacking. In countries where private land ownership is well-established, it is plausible to assume that $\lambda = 1$ for a land owner. While in the context of China, the corresponding "owner" of farmland may be the contractee who bears both the land contract and use rights. The value of $\lambda_{contractee}$ is possibly lower than, but close to 1, since land contractees are also confronted with tenure insecurity such as land reallocation. For renters, $\lambda$ should be much lower since they only possess the land use right and are mainly concerned with short-term profits. However, $\lambda_{renter}$ could be higher than zero if renters can continually use the land, which depends on the duration of the lease. In this case, more than 90 percent of the existing leases in the study region are oral and informal on a one-year basis. Thus, we hypothesize that $\lambda_{renter} < \lambda_{contractee}$ or the probability of adopting straw retention is higher for contractees than for renters, which we test in our empirical model.

## 4. Empirical Implementation

### *4.1. Base probit model*

To test the above hypothesis, we first employ a binary discrete choice model derived from the latent variable approach. Let $y$ denote a farmer household's decision to adopt straw retention or not, which is generated from a latent variable, $y^*$, equal to $\lambda \Delta V/R - \Delta \Pi$ from equation (3). The difference between short-term profits and long-term land values for the farmer household is unobserved, but one can observe the household's adoption decision. If $y^*$ is positive, straw retention is adopted and $y = 1$ is observed; otherwise, $y = 0$ is observed if $y^*$ is negative.

For each household, $i$, the latent variable $y_i^*$ is assumed to be a linear function of the vector of observable household, plot, and regional characteristics ($\mathbf{X}_i$) as follows:

$$y_i^* = \boldsymbol{\beta}\mathbf{X}_i + \varepsilon_i, \quad (4)$$

where $\boldsymbol{\beta}$ is the coefficient vector and $\varepsilon_i$ a random error term. The linkage between $y_i^*$ and $y_i$ is as follows:

$$y_i = \begin{cases} 1, & if\ y_i^* > 0 \\ 0, & if\ y_i^* < 0 \end{cases}, \quad (5)$$

then the probability that household $i$ adopts straw retention ($y_i = 1$) is given by

$$\begin{aligned} Pr[y_i = 1] &= Pr[y_i^* > 0] \quad (6) \\ &= Pr[\boldsymbol{\beta}\mathbf{X}_i + \varepsilon_i > 0] \\ &= 1 - Pr[\varepsilon_i \leq -\boldsymbol{\beta}\mathbf{X}_i] \\ &= F(\boldsymbol{\beta}\mathbf{X}_i), \end{aligned}$$

where $F(\cdot)$ is the cumulative distribution function of the error term $\varepsilon_i$. We assume that $\varepsilon_i$ follows the standard normal distribution and equation (6) is estimated by probit regression.

Empirically, the straw retention adoption equation, for each household $i$ and cultivated plot $k$, is assumed to be written as:

$$Pr[adopt_{ik} = 1] = \theta rent_{ik} + \boldsymbol{\beta}\mathbf{X}_{ik} + u_{ik}, \qquad (7)$$

where $adopt_{ik}$ represents the adoption dummy, which equals unity if the household $i$ adopts straw retention on plot $k$; $rent_{ik}$ is a land tenure dummy, which equals unity if the plot is rented from others, and is designed to capture the land tenure insecurity denoted earlier as $\lambda_{renter}$; $\mathbf{X}_{ik}$ is the vector of other plot-level and household-level control variables; and, $u_{ik}$ is the error term. $\hat{\theta}$ is of primary interest because it measures the marginal effect of renting from others on the adoption of straw retention.

The land tenure dummy may be endogenous to adoption decisions for several reasons. First, growers on rented plots may self-select into non-conservation activities. For instance, renters could be large-scale producers who are keener to maximize profits on all grounds (including rented plots) and thus are more cognizant of all possible cost savings and prone to select the profit-maximizing practice.[4] Given that air quality impacts from straw burning is currently external to a producer's decision process, the renters are more likely to adopt the least costly option as opposed to more costly straw retention. In this case, the naïve probit estimate of the rental dummy may be underestimated. In addition, there may be unobserved characteristics that

[4] As indirect evidence to support the claim that renters are probably large-scale producers, we make a simple statistic for our surveyed sample. Of all 670 valid households, 588 households are purely own-contractees, which means they only cultivate on their own contracted plots throughout 2015; the remaining 82 households have rented plots to cultivate besides their own-contracted plots. The average area of cultivable land in 2015 for 82 renters is 16.06 Mu, about four times larger than that of 588 purely own-contractees, which is only 4.86 Mu.

are correlated with both adoption choices and rental decisions, including on-farm characteristics such as land shapes, soil quality and machinery use, and off-farm factors such as government-supported agricultural policies.

Technically, the endogeneity arises due to the presence of correlation between $rent_{ik}$ and $u_{ik}$, which may bias the estimate of $\hat{\theta}$. In order to obtain a consistent $\hat{\theta}$, we further consider two approaches separately dealing with the endogeneity issue, a bivariate probit model and a control function method.

### *4.2. Bivariate probit model*

The bivariate probit model is frequently used in estimating the effect of an endogenous binary treatment on a dichotomous outcome, while accounting for unobserved confounders (Wooldridge, 2010). For expositional purposes, we rewrite the adoption equation and construct the two-equation latent variable framework as follows:

$$y^* = x'\beta + \theta d^* + u,\ y = 1\ if\ y^* > 0; \qquad (8)$$

$$d^* = z'\gamma + \varepsilon,\ d = 1\ if\ d^* > 0; \qquad (9)$$

where $y^*$ and $d^*$ represent the latent outcome (i.e., adoption) and treatment (i.e., renting) variables, respectively; $x$ and z are covariates in each equation; and, $u$ and $\varepsilon$ are corresponding errors. We assume that households' rental decisions are endogenous to straw retention adoption, thus $\varepsilon$ is correlated with $u$. The bivariate probit builds on the assumption that $(u, \varepsilon)$ follows a bivariate standard normal joint distribution with zero means, unit variances, and correlation coefficient $\rho$ (i.e., $\rho = corr(u, \varepsilon)$):

$$(u, \varepsilon) \sim N\left(\begin{bmatrix} 0 \\ 0 \end{bmatrix}, \begin{bmatrix} \rho & 1 \\ 1 & \rho \end{bmatrix}\right). \quad (10)$$

To identify the parameters $\theta$ and $\beta$ in the adoption equation, the covariates $z$ in equation (9) must contain at least one variable as an instrument for $d^*$ (or exclusion restriction) not included in covariates $x$ in equation (8). Otherwise, simply relying on nonlinearities without excluded instruments will lead to ill specification and the estimated partial effects may also be biased (Altonji et al., 2005; Wooldridge, 2010).

In particular, we construct a ratio of annual income for a family's migrant workers to annual agricultural profits for the farmer household as the instrument for the rental variable. The higher the ratio, the higher the remittances the family receives from migration than from farming, and thus a smaller likelihood that the household would rent more farmland from other contractees. The idea is that the ratio serves as a proxy of a household's ability to earn additional income from urban employment and thus is directly associated with household's rental decisions. It is possible that the migrants' income ratio is still endogenous to adoption because higher remittances means a lack of labor supply available for agricultural production, or more money to potentially buy straw retention machinery. In other words, the ratio may affect the adoption choice through laborer numbers or gross household income. Empirically, this can be overcome by including these two variables as controls.[5] That way, the ratio variable is exogenous to

[5] In our empirical models, as will be noted later, the number of laborers that engaged in farming activities and annual family income will be included as control variables in all regressions.

adoption conditional on available laborers and household income.

### *4.3. Control function method*

The control function method addresses the endogeneity problem by directly augmenting the generalized residuals in the outcome equation, which accounts for any correlation between $d^*$ and $u$. Specifically, we can rewrite equation (8) as:

$$y^* = x'\beta + \theta d^* + \mu\tilde{\varepsilon} + \nu, \quad (11)$$

where $\tilde{\varepsilon}$ is the fitted residuals from equation (9) and identified as the inverse Mills ratio for the whole sample, which takes the form (Gourieroux et al., 1987; Vella, 1998):

$$\tilde{\varepsilon} = d * \left[\frac{\phi(z'\hat{\gamma})}{\Phi(z'\hat{\gamma})}\right] + (1 - d) * \left[\frac{-\phi(z'\hat{\gamma})}{1-\Phi(z'\hat{\gamma})}\right] \quad (12)$$

where $d$ is the rental dummy; $\hat{\gamma}$ denotes the estimated parameter vector for the probit estimation of equation (9); and, $\phi(\cdot)$ and $\Phi(\cdot)$ are normal density and normal distribution functions.

Following Vella (1998), a consistent estimate of $\hat{\theta}$ can be obtained from regressing equation (11) by linear probability model (LPM), assuming the generalized residuals, $\tilde{\varepsilon}$, has fully removed any correlation between $d^*$ and the new error $\nu$.[6] Although technically one could rely on the nonlinearity of the probit model in the first stage of the control function approach, Wooldridge has forcefully argued that it is better to have more independent variation in the generalized residuals with inclusion of instrumental variables (Wooldridge, 2010, p.599).

[6] For detailed discussions about why equation (11) is required being estimated by LPM and characteristics of the generalized residuals, see Vella (1998).

As a result, we rely on the migrants' income ratio in the first stage of the control function approach when generating the generalized residuals. It is important to note that the control function method with binary endogenous explanatory variables in nonlinear models is more controversial, because it requires nonstandard assumptions including that the generalized residuals act as a kind of sufficient statistic for capturing the endogeneity of the land tenure variable (Wooldridge, 2015). However, these assumptions are no more or less general than the bivariate probit assumptions (Wooldridge, 2015). In addition, equation (11) itself can serve as an endogeneity test of the binary treatment variable, the rental dummy in our model, and as an analogue of the Rivers-Vuong approach, which deals with a continuous endogenous variable in a binary response model (Rivers and Vuong, 1988).[7] Specifically, rejection of the null hypothesis that $\mu = 0$ in equation (11) implies that $d$ is endogenous.

Empirically, we provide the estimation results of the two approaches—a bivariate probit and a control function method—to quantify the marginal effect of rental decision on straw retention adoption after controlling for the endogeneity of the land tenure variable.

## 5. Data Description

The data used in this study is drawn from a rural household survey conducted by Henan Agricultural University in 2016 in Henan Province, a major grain production province in central

[7] In Rivers-Vuong's two-step approach, first run an OLS of the endogenous continuous variable, say $y_2$, on covariates $z_2$ and obtain the residuals $\hat{v}$; then run the probit of the binary outcome variable, say $y_1$, on covariates $z_1$, $y_2$ and $\hat{v}$. While in our model, the first-stage regression is probit and the second-stage is OLS but with generalized residuals.

China. Henan is also a big crop straw producer, accompanied by severe air pollution resulting from straw burning in open fields (Fu et al., 2017).

We have followed a stratified sampling approach in collecting our sample: for each of the 17 prefecture cities in Henan province, we randomly selected two to four counties, and within each county, we randomly selected one or two villages among the agriculturally significant counties based on historical planting area. About 50 trained undergraduate research assistants went to these sampled villages for face-to-face interviews of farmers. With the help of local village and county officials, we randomly selected 10 percent of rural households within each village based on the household registry roster. In total, our sample consists of 670 valid rural household responses spanning across 175 villages in 47 counties (see Figure 2).

The household-level data covers detailed information about agricultural production and operation throughout 2015, as well as household and personal characteristics (see Appendix 1 for the full version of the questionnaire). We also supplement the survey data with climatic information obtained from the National Meteorological Information Center (NMIC), which reports county-level historical averages for climate variables from 1981 to 2000 in Henan Province.[8]

In our study area, many farmers grow at least two crops every year, typically corn during the autumn season (from June to October) and winter wheat during the winter season (from October to June). As a result, farmers face two annual straw retention choices. To better identify

[8] As mentioned earlier, since Henan province is not included in the pilot project for a straw retention subsidy, in our study area for the 2015 crop year, there is no subsidy available for farmers specifically targeted to incentivize straw retention.

the impact of tenure, we treat straw retention choices for these crops separately and our unit of observation is at the plot-level for each season, with each representing a particular crop for a particular growing season.

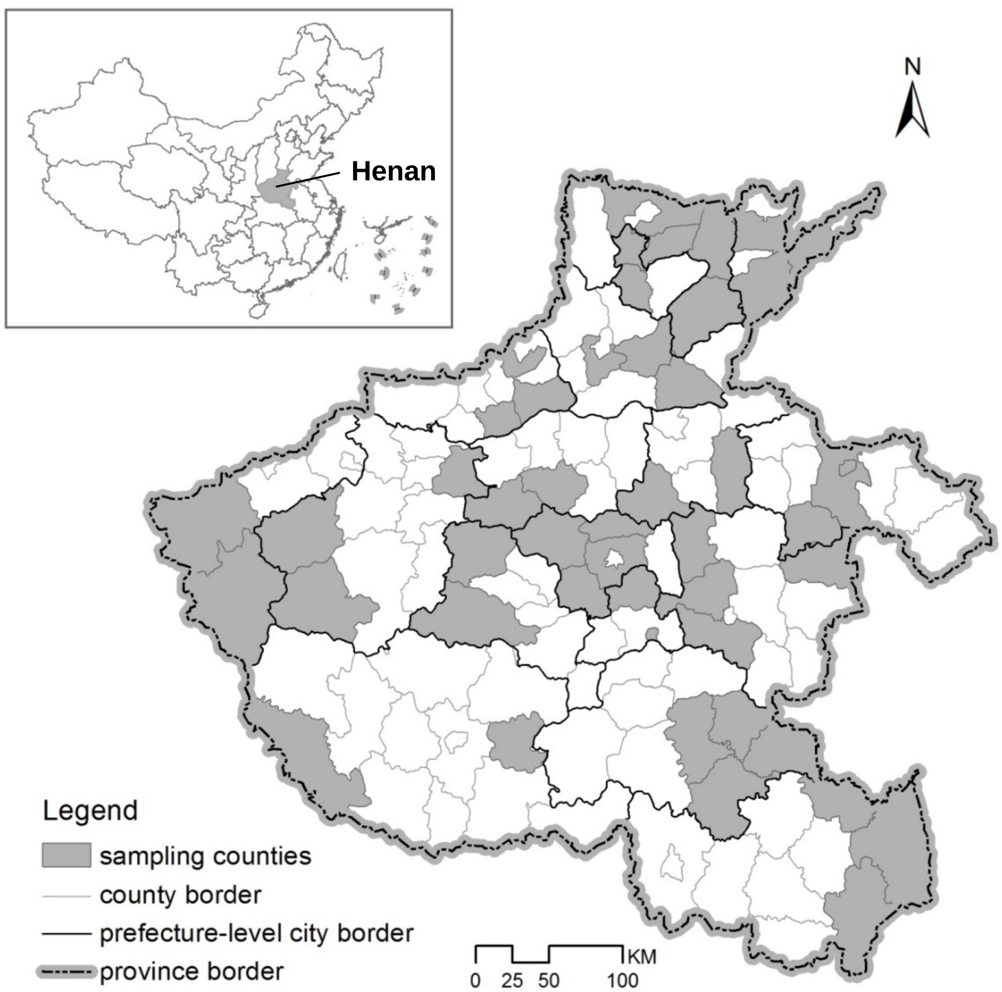


Figure 2. Map of Henan Province in China, the sampled region.

For each specified plot, the farmer is asked how he or she dealt with the straws after harvest and asked to choose among the following options: straw retention, burn, use for fuel, use for feed, sell, and discard. We assume that straw retention is adopted for a particular plot if straw retention is chosen, no matter whether there are other selected options as well and regardless of the specific share of straws being returned to the field (retention).[9] Of all observations, 9.3 percent of plots are rented and 90.7 percent are own-contracted, but the areas of rented plots account for 20.3 percent of total cultivated land areas in our sample. In addition, 70.8 percent of plots adopted straw retention, 44.1 percent are wheat plots, 38.4 percent are corn plots, and 17.5 percent are other crops including rice, soybean, peanuts, cotton, etc.

All plots are classified by a binary land tenure variable: own-contracted versus rented from others. Own-contracted refers to plots with current growers that are the original contractees of land allocated directly from the village collectives and who hold both the land contract and use rights. Rented plots are those which current growers rented from original contractees and thus only hold the land use rights. After eliminating (*a*) all plots that are not cultivated in 2015; and, (*b*) plots where the planted crops do not generate straws, and thus do not involve any treatment of straw, we have 1,659 plot-level observations from 670 farmer households for analysis.[10]

As mentioned above, we form our instrument by constructing a ratio of annual income for family's migrant workers to annual agricultural profits for the farmer household.[11] More

[9] The survey does not collect information about the shares of straws for different straw treatments.

[10] Due to widely used informal contracts between farmer households in most cases, it is not clear how the decision is made on a rented plot in this context. The decision-making process has been shown crucial in affecting adoption of conservation practices (Soule et al., 2000; Kurkalova et al., 2006). For instance, in the United States, there are two types of land renter: share-renters' decisions may be affected by land owners because both the owners and renters share the revenues and costs of production; cash-renters, however, may behave more independently since they only pay a fixed rent to the landlords while the owners do not participate in any activities. Failure to consider the differences in decision-making process may obscure the effect of land tenure on adoption.

[11] Besides migrant income and agricultural profits, the annual household income also includes income from off-farm business or employment, transfer income such as pensions and government subsidies, etc. Statistics from our survey reveals that migrants'

specifically, the annual migrant income is calculated by the number of family migrant members working outside the village, times the average working months, times the average monthly wages. The annual agricultural profits are simply the profits of planting crops for the whole family. All this information has been collected through the survey.

We follow the extant literature of soil conservation practices in selecting other control variables of farmer and field characteristics (Feder et al., 1988; Fox et al., 1991; Knowler and Bradshaw, 2007; Baumgart-Gets et al., 2012; Zhang et al. 2016). Table 1 shows the descriptions of selected variables. First, we include three farmer characteristics of the household head (who is also the main operator): age, years of farming, and risk preference.[12] Farmers with more experience and more risk-taking attitudes are expected to have a higher likelihood of adopting conservation practices. A second set of variables control for household-level characteristics, which include highest years of education of all household members, number of laborers, family income, participation in agricultural economic organization, and purchase of agricultural insurance. We hypothesize that a richer, well-connected farmer household with better education, more laborers, and insurance coverage is more likely to adopt straw retention. In addition, July temperature and July precipitation are two weather variables included as controls of local climate effect on farmer's adoption choice. Finally, at the plot level, we consider plot size, a season dummy, and crop dummies to control for possible scale and seasonal effects, and heterogeneous

income and agricultural profits jointly account for about 75 percent of gross household income on average.

[12] The household head's attitude towards risk is measured through response to a binary question "whether you will utilize a newly developed fertilizer that may increase yields but has not been used yet" and is assigned "risk-seeking" if answering "yes" and "risk averse" otherwise.

impacts of crop types. We do not consider other topographic features, such as plot slope, because most of our study area is flat and thus we do not expect significant within-province variation.

Table 1. Descriptions of Selected Variables

| **Variable** | **Description** |
|---|---|
| Straw retention | Straw retention is adopted in the plot (1=yes, 0=no) |
| Rented plot | The plot is rented from other contractees (1=yes, 0=no) |
| Age | Age of the household head |
| Farming experience | Number of farming years of the household head |
| Risk preference | Risk preference of the household head (1=risk seeking, 0=risk averseness) |
| Education | Highest years of education of all household members |
| Number of laborers | Number of laborers that engaged in farming activities in 2015 |
| Income | Household annual income in 2015 (10,000 Yuan) |
| Organization | The household has participated in rural economic organization (1=yes, 0=no) |
| Insurance | The household has purchased agricultural insurance (1=yes, 0=no) |
| Plot size | Plot area (mu=0.0667 hectares) |
| Mig-ag ratio | Ratio of annual income for family's migrant workers to annual agricultural profits for the farmer household in 2015 |
| July temperature | Average daily temperature of July days from 1981 to 2000 of the county that the household locates (℃) |
| July precipitation | Average annual daily precipitation of July days from 1981 to 2000 (mm) |
| Winter season | The plot is cultivated in the winter season (1=winter season, 0=autumn season) |
| Wheat | The plot is planted wheat (1=yes, 0=no) |
| Corn | The plot is planted corn (1=yes, 0=no) |
| Other crop | The plot is planted other crops (1=yes, 0=no) |

Table 2 shows summary statistics for the selected variables used in our model by tenure type. The last column reports the sample t-test results of mean differences (rented minus own-contracted) of each of the variables. Basically, the raw descriptive adoption rate for straw retention on rented plots is about five percent lower, though insignificantly, compared to own-contracted plots, confirming our general expectation. It is worthwhile to note that the instrumental variable—the migrants' income ratio on rented plots—is significantly less than that on own-contracted plots, which is in line with our assumption that higher remittances from migration decrease the probability of renting. Another notable fact is that rented plots are generally (3.56 mu) larger in size, which provides partial evidence that renters are probably large-scale producers. Other factors with significant differences include age, education, household income, etc.

Table 2. Summary Statistics of Selected Variables by Tenure Type

| **Variable** | **All plots** | | **Own contracted plots** | | **Rented plots** | | **T-test of** |
|---|---|---|---|---|---|---|---|
| | **Mean** | **Std. Dev.** | **Mean** | **Std. Dev.** | **Mean** | **Std. Dev.** | **mean diff** |
| Straw retention | 0.708 | 0.45 | 0.712 | 0.45 | 0.665 | 0.47 | -0.048 |
| Rented plot | 0.093 | 0.29 | - | - | - | - | - |
| Age | 44.28 | 11.76 | 44.08 | 11.93 | 46.28 | 9.71 | **2.20**** |
| Farming experience | 28.28 | 13.59 | 28.32 | 13.55 | 27.92 | 14.04 | -0.40 |
| Risk preference | 0.679 | 0.47 | 0.676 | 0.47 | 0.710 | 0.46 | 0.034 |
| Education | 12.40 | 3.58 | 12.32 | 3.59 | 13.20 | 3.47 | **0.88***** |
| Number of laborers | 2.611 | 1.26 | 2.620 | 1.27 | 2.516 | 1.21 | -0.104 |
| Income | 4.201 | 3.15 | 4.087 | 2.96 | 5.351 | 4.48 | **1.26***** |
| Organization | 0.041 | 0.20 | 0.041 | 0.20 | 0.039 | 0.19 | -0.003 |
| Insurance | 0.129 | 0.34 | 0.116 | 0.32 | 0.258 | 0.44 | **0.14***** |
| Plot size | 4.135 | 4.25 | 3.802 | 2.61 | 7.363 | 10.80 | **3.56***** |
| Mig-ag ratio | 6.703 | 11.54 | 6.896 | 11.83 | 4.829 | 8.02 | **-2.07**** |
| July temperature | 26.98 | 0.36 | 26.97 | 0.36 | 27.05 | 0.34 | **0.08***** |
| July precipitation | 337.6 | 70.4 | 336.5 | 67.7 | 348.2 | 92.0 | **-11.74**** |
| Winter season | 0.461 | 0.50 | 0.463 | 0.50 | 0.439 | 0.50 | -0.024 |
| Wheat | 0.441 | 0.50 | 0.443 | 0.50 | 0.413 | 0.49 | -0.031 |
| Corn | 0.384 | 0.49 | 0.387 | 0.49 | 0.355 | 0.48 | -0.032 |
| Other crop | 0.175 | 0.38 | 0.170 | 0.38 | 0.232 | 0.42 | **0.06*** |
| Observations | 1659 | | 1504 | | 155 | | |

Note: ***, **, and * represent significance at 1%, 5%, and 10% level, respectively.

## 6. Empirical Results

### *6.1. Base probit regressions with an endogenous regressor*

We first report the results of several simple probit regressions in which we test whether the probability of adopting straw retention is higher in own-contracted plots than rented plots. That is, we test the null hypothesis that $\beta = 0$, where $\beta = \lambda_{renter} - \lambda_{contractee}$ is the coefficient on the tenure dummy and represents the difference of tenure-security indicator between own-contracted and rented plots. Table 3 presents the resulting estimated average partial effects.

Table 3. Results of the Binary Probit Regressions (Dependent: Adoption=1)

| Variable | (1) Full sample<br>Marg. eff. | (2) Own contracted<br>Marg. eff. | (3) Rented<br>Marg. eff. |
|---|---|---|---|
| Rented plot | -0.0819** | | |
| Age | -0.0021 | -0.0036* | 0.0068 |
| Farming experience | 0.0002 | 0.0013 | -0.0035 |
| Risk preference | 0.0349 | 0.024 | 0.0903 |
| Education | 0.0166*** | 0.0147*** | 0.0274** |
| Number of laborers | 0.0659*** | 0.0635*** | 0.1041*** |
| Income | 0.0054 | 0.0056 | -0.0013 |
| Organization | 0.1853** | 0.1986** | 0.1909 |
| Insurance | 0.1239*** | 0.1380*** | 0.0842 |
| Plot size | 0.0093** | 0.0122** | 0.0027 |
| July temperature | -0.0792* | -0.1153 | 0.1766 |
| July precipitation | 0.0013*** | 0.0013*** | 0.0011** |
| Winter season | 0.1284** | 0.1600** | 0.0176 |
| Crop dummies | Yes | Yes | Yes |
| Observations | 1659 | 1659 | 1659 |
| % correctly predicted | 0.8005 | 0.8005 | 0.8005 |
| LR test (p-value) | | $\chi^2(14)$=30.06 (0.0075) | |

Note: ***, **, and * represent significance at 1%, 5%, and 10% level, respectively. Standard errors are clustered at the household level. The marg. eff. columns report the average partial effects across all individuals in the sample. The null of the likelihood ratio test is no systematic difference in coefficients across tenure types.

Column (1) in Table 3 displays the results using the full sample. The marginal effect of the "rented plot" dummy is negative and significant at the five percent level, and shows that a rented plot is associated with about eight percent lower probability of adopting straw retention than an own-contracted plot, ceteris paribus. The result confirms our hypothesis and is consistent with previous studies regarding the impacts of land tenure on conservation practice adoption (Soule et al., 2000; Fraser, 2004; Kabubo-Mariara et al., 2010). The household-level characteristics appear to be more important in the adoption decision—higher educational attainment, greater number of laborers, participation in an agricultural economic organization, and purchase of agricultural insurance are significantly and positively related to increased likelihood of adoption. More specifically, each extra year of education increases the probability of adoption by about two percent and one more laborer leads to an increase of the likelihood of adoption by about seven percent. Participation in agricultural economic organization and purchase of agricultural insurance have a larger effect on adoption, with values of about 19 percent and 12 percent, respectively. Two climate variables are significant, suggesting that straw retention occurs more frequently where the climate is warmer and wetter (more rain). The estimated coefficient of plot size indicates that one more unit increase of plot size results in an about one percent increase of the likelihood of adoption, which reflects that larger farmland in fact lowers the average cost of adoption. The positive and significant winter dummy verifies the existence of seasonal effect on residue management.

We also examine whether the estimated coefficients systematically differ by tenure type. We conduct separate probit regressions for each of the two tenure categories, own-contracted and rented plots, with all other independent variables the same as those in the full-sample regression. A Chi-square statistic of 30.06 in the likelihood ratio test rejects the null hypothesis that the coefficients are identical across tenure types, implying that the impact of all other characteristics on adoption of straw retention relies on the specific tenure status of the plot.[13] The resulting

[13] In conducting the likelihood ratio test, we do not specify clustered standard errors at any level in the regressions.

coefficients and average partial effects by tenure type are reported in columns (2) and (3) in Table 3, which shows that despite similarity in other variables, the age of household head, organization participation, purchase of insurance, plot size, and seasonal effect only matter for own-contracted plots, while they have no effect on adoption for rented plots. The sharp disparity reflects a significantly different decision mechanism for a renter in adopting straw retention. Although the simple probit models shown in Table 3 provide suggestive evidence of the negative impacts of land rental decisions on straw retention adoptions, it is important to note that the key variable of interest, the land rental variable, is potentially endogenous and thus the estimates might be biased.

### *6.2. Results of bivariate probit model and control function method*

To control for or mitigate the endogeneity of the rental variable, we also employ two alternative approaches—a bivariate probit and a control function method with a linear probability model. In both models, the ratio of migrants' income to agricultural profits for the farmer household is used as the instrument.[14] Results are reported in Tables 4 and 5, respectively.

[14] We have also tried "migrants' income in levels" and "relative share of migrants' income" as exclusion restrictions. Results are very similar, but using the ratio variable as the restriction has the best fits. Regression results for bivariate probit and control function method using the other two restrictions are available from authors upon request.

Table 4. Results of Bivariate Probit Model

| Variable | Adoption Equation | | Renting Equation | |
|---|---|---|---|---|
| | Marg. eff. | p-value | Coef. | p-value |
| Rented plot | -0.5113*** | 0.000 | | |
| Age | -0.0008 | 0.547 | 0.0187*** | 0.004 |
| Farming experience | -0.0006 | 0.603 | -0.0107** | 0.034 |
| Risk preference | 0.0353* | 0.070 | -0.0377 | 0.699 |
| Education | 0.0165*** | 0.000 | 0.0276** | 0.037 |
| Number of laborers | 0.0585*** | 0.000 | -0.0597 | 0.119 |
| Income | 0.0064** | 0.043 | 0.0301** | 0.021 |
| Organization | 0.1650*** | 0.001 | 0.0321 | 0.888 |
| Insurance | 0.1425*** | 0.000 | 0.5818*** | 0.000 |
| Plot size | 0.0149*** | 0.000 | 0.0804*** | 0.000 |
| July temperature | -0.0610* | 0.064 | | |
| July precipitation | 0.0011*** | 0.000 | | |
| Winter season | 0.1058* | 0.071 | | |
| Mig-ag ratio | | | -0.0110** | 0.042 |
| ρ (p-value) | 0.7100** (0.0108) | | | |
| Crop dummies | Yes | | No | |
| Observations | 1659 | | | |

Note: ***, **, and * represent significance at 1%, 5%, and 10% level, respectively. Standard errors are clustered at the household level. The marg. eff. column reports the average partial effects across all individuals in the sample.

As seen in the bivariate probit model, the rental dummy is still negatively significant at the one percent level, but with a much higher average partial effect of 0.5113, meaning that being on a rented plot would lower the probability of adoption of straw retention by half. The signs and significance of most of other explanatory variables remain nearly unchanged when compared with the base model, except that household head's risk preference and household income become significant in the bivariate probit model.

Table 5. Results of Control Function with LPM[15]

| Variable | LPM | | Probit | |
|---|---|---|---|---|
| | Coef. (Marg. eff.) | p-value | Coef. | p-value |
| Rented plot | -0.5251* | 0.054 | | |
| Age | -0.0004 | 0.863 | 0.0218*** | 0.005 |
| Farming experience | -0.0012 | 0.512 | -0.0122* | 0.081 |
| Risk preference | 0.0386 | 0.241 | 0.0188 | 0.893 |
| Education | 0.0210*** | 0.000 | 0.0310* | 0.091 |
| Number of laborers | 0.0607*** | 0.000 | -0.0508 | 0.400 |
| Income | 0.0082 | 0.170 | 0.0274 | 0.152 |
| Organization | 0.1427** | 0.043 | -0.0117 | 0.974 |
| Insurance | 0.1718*** | 0.000 | 0.5829*** | 0.000 |
| Plot size | 0.0137*** | 0.003 | 0.0772*** | 0.000 |
| July temperature | -0.0821 | 0.121 | | |
| July precipitation | 0.0015*** | 0.000 | | |
| Winter season | 0.1696** | 0.030 | | |
| Generalized residual | 0.2293 | 0.109 | | |
| Mig-ag ratio | | | -0.0048 | 0.517 |
| Intercept | 1.9038 | 0.165 | -2.7877*** | 0.000 |
| Crop dummies | Yes | | No | |
| Observations | 1659 | | | |

Note: ***, **, and * represent significance at 1%, 5%, and 10% level, respectively. Standard errors are clustered at the household level.

As for the control function method, the marginal effect of the rental dummy is -0.5251 and significant at the ten percent level. It is noteworthy that the generalized residual term appears insignificant at any significance level but with a very close p-value of 0.109, which fails to reject the null hypothesis that renting is exogenous, suggesting that at least the endogeneity may not be that apparent in our case. However, it is important to note that one could not fully rely on a

[15] In an identical model but without exclusion restriction variable (i.e. "mig-ag ratio" is excluded in the probit regression), whose identification of the endogenous rental variable simply relies on the nonlinearity nature of the generalized residual, the coefficient (marginal effect) of the rental dummy is -0.3707 but insignificant with a p-value of 0.16.

marginally insignificant test statistic to claim full exogeneity.[16] We argue that our control function approach at least mitigated, if not fully addressed, the endogeneity concerns. We also want to caution readers that alternative specifications of our control function approach, using alternative set of instruments such as migrants' income in levels or without any instrumental variables in the first stage, lead to statistically insignificant coefficient for the land tenure variable. This suggests further research is needed to examine the robustness of these results but even in these alternative specifications the coefficient for land tenure on straw retention is still negative.

Contrary to the results in the adoption regression, characteristics of the household head, including age and farming experience, appear to play an important role in driving land rental decisions. Higher family education level, higher family income, and purchase of insurance are positively related to the probability of renting more farmland. In addition, the positive and significant coefficient of plot size confirms our conjecture that renters are more likely to rent larger fields or more farmland. The exclusion restriction variable (migrants' income ratio) is negatively significant in the bivariate probit model, which confirms our expectation that households with higher remittances from migrant members are less willing to rent from other contractees. This variable has negative coefficient in the control function approach as well but it is statistically insignificant.

### *6.3. Discussion*

[16] We thank an anonymous reviewer for pointing this out.

Our empirical results provide evidence that land tenure insecurity, as proxied by rented plots, tend to negatively impact farmers' decisions to adopt straw retention in China. Combining results of all three regression models, we discover that after dealing with the potential endogeneity issue, there is a dramatic increase in magnitude of the average marginal effect of renting on conservation practice adoption, from about -0.08 in the base model to around -0.50, which is much higher than that in the existing literature (Soule et al., 2000; Gebremedhin and Swinton, 2003; Kabubo-Mariara et al., 2010; Baumgart-Gets et al., 2012). This increase is intuitive in the sense that the base probit model may be underestimated, either due to renters' self-selection behaviors or omission of some important covariates. More specifically in our data analysis, the omitted unobservables could be positively related to both farmers' rental decisions and adoption choices. Evidence includes the positive coefficient on the residual term (although insignificant) from the second-stage regressions in two control function models and the positive and significant $\rho$ in the bivariate probit that measures correlation between adoption and renting equations.

Availability of machinery and agricultural subsidies from local government may be two omitted factors and unfortunately they were not covered in our survey data.[17] Use of machinery can facilitate straw retention and save laborers to a large extent, so households who own relevant machines, or are able to rent machines from outside supplies, are more likely to adopt straw retention. Meanwhile, machinery availability may further lead to an increased likelihood of

[17] We thank an anonymous review for pointing out that these two factors may be correlated, though slightly, with the instrument variable, the migrants' income ratio. We acknowledge that the constructed ratio variable may not fully solve the endogeneity of the rental dummy. While in this sense, the use of control function method could mitigate the endogeneity problem to a large extent.

renting more farmland to gain economies of scale and enhance productivity. China's central government has granted rural farmers four categories of agricultural subsidies for many years to boost grain production.[18] Although these subsidies do not directly target conservation, it is reasonable to infer that households receiving more subsidies are less vulnerable to uncertain short-term profit loss from adopting straw retention. Moreover, households residing in areas where the subsidies are available are more willing to grow crops and thus more likely to rent more farmland simply to increase income. Recent studies in China also support the claim that machinery use and subsidies play an important and positive role in inducing straw retention behaviors, and their effects warrant further research.[19]

Furthermore, the descriptive data from our sample seems to suggest that current land rental arrangements may be significantly insecure. Of all 82 surveyed households (out of a total of 690) who rent land from others, 75 households (about 95 percent) are based on oral land leases without legal validity; only four households (about five percent) have signed formal and written contracts, and 65 households' leases (about 82 percent) are on a one-year-tenure basis. The figures reflect that the tenure of a rented plot, or the transferred land use right, may be far less secure than land contract rights in Henan province.[20]

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[18] Four categories of subsidies include direct grain subsidy, superior grain varieties subsidy, agricultural machinery and tools subsidy, and general agricultural suppliers subsidy. Unlike the straw retention subsidy that is about to take effect in recent years, these subsidies have been available since 2004 in most areas of China.

[19] See Qi et al. (2016) and Tong and Liu (2017) as examples.

[20] To test the potential effects of contract type and lease term on adoption, we have also run several probit regressions for only the 155 renters in our sample. Results are reported in Table A.1 in Appendix 2, where we list four specifications with column (1) the base model and columns (2)–(4) with one or two of the contract variables augmented. However, results reveal that neither of the two contract variables is significant in any specification. Therefore, we are unable to identify the effect of contract type and renewal frequency in this study.

## 7. Concluding Remarks

Based on data from a 2016 farmer household survey in Henan Province in central China, we apply several probit regressions, including a bivariate probit model and a control function method, to examine how land tenure arrangements especially land rental decisions affect the probability of adopting crop straw retention. We employ an instrument that relies on the remittance income from household members who migrate to cities to address the potential endogeneity of the land tenure variable. Empirical results show that farmer households are more likely to adopt straw retention after harvest on own-contracted than rented plots, which is consistent with previous literature that shows insecure land tenure often hinders adoption of conservation practices. This study provides some of the first evidence in assessing the negative impact of land tenure insecurity on the adoption of a long-term land improvement practice, straw retention.

Our findings have important and valuable policy implications. The Chinese government is currently encouraging participation of market-based rural land transfer as an alternative to committee-intervening land reallocation to improve agricultural production efficiency. However, our results imply that more careful deliberation or monitoring is warranted regarding the potential impediments to less-sustainable farming practices on the growing acres of rented grounds. Our results also suggest that the air quality and other environmental benefits from encouraged straw retention could be impacted without further efforts to boost adoption on rented land. Due to less secure land tenure, farmers on rented plots may have lower incentives to adopt conservation practices, causing environmental problems such as more PM 2.5 emissions from straw burning and more severe soil degradation. Improving the land tenure security, especially for rented plots, may be one effective measure. For instance, the government could regulate the rental market by enforcing the use of a more stable and formal written contract, which incorporates specific requirements on conservation practices and residue management. On the other hand, it may also be helpful to extend the current range of land contract rights transfer from within-village to cross-village or within-township, in order to allow farmer households to own more contracted land and stimulate long-run land improvements.

The importance of our findings notwithstanding, they are limited by the fact that some of important factors such as availability of machinery and agricultural subsidies were not covered in our survey. Moreover, the effects of different aspects of tenure security need further investigation. For instance, the duration of tenure, another important dimension of tenure security either for own-contracted or rented land, may also play a role in affecting the adoption rate. In addition, with a panel dataset, one can examine whether the adoption rate of straw retention varies over time under different policy scenarios, such as varying strictness on restricting the burning of straws and the forthcoming straw retention subsidy. It is possible that the higher adoption rate is mainly due to the prohibition of straw burning.

Another caveat is that our empirical results may not always be robust in terms of the significance and magnitude of the land rental variable when we rely on alternative instruments in both the biprobit and control function models. For instance, using migrant income in levels or relative share of migrants' income could lead to differing estimated marginal effects of rental ranging from 38 percent to 51 percent. However, it is important to note that the rental variable has been shown consistently to negatively impact the adoption of straw retention, even in cases of statistical insignificance. This could result from our data limitation of only relying on one year's data, or lack of information about some important omitted factors, which suggests additional research to examine the role of land tenure in conservation practice adoptions.


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## Appendix 1: Translated Version of Survey Questionnaire

### Survey on Farmers' Planting Behavior in Henan Province 2015

__________City__________County__________Town__________Village Date______________

Name of interviewee__________Contact number__________ Name of interviewer__________

### A. On-farm Operation of Tillage and Straw Return in 2015

**A1 In 2015, total areas of family cultivable land is__________Mu; # of plots__________; the area of the biggest plot is__________Mu.**

| A2 Land area for each property type (Mu) | Own-contracted_______ | | | | Rented-in_______ | | | | Trusteeship_ | |
|---|---|---|---|---|---|---|---|---|---|---|
| Plot Index | X1 | X2 | X3 | X4 | Y1 | Y2 | Y3 | Y4 | Z1 | Z2 |
| A3 Plot area (Mu) | | | | | | | | | | |
| A4 Contract type | NA | NA | NA | NA | | | | | NA | NA |
| A5 Lease term (ys) | NA | NA | NA | NA | | | | | NA | NA |
| Season 1: Autumn | | | | | | | | | | |
| A6a Crop | | | | | | | | | | |
| A7a Straw treatment | | | | | | | | | | |
| A8a Reasons for SR adoption | | | | | | | | | | |
| A9a Reasons for SR non-adoption | | | | | | | | | | |
| Season 2: Winter-Summer | | | | | | | | | | |
| A6b Crop | | | | | | | | | | |
| A7b Straw treatment | | | | | | | | | | |
| A8b Reasons for SR adoption | | | | | | | | | | |
| A9b Reasons for SR non-adoption | | | | | | | | | | |

Note:

1. **Plot division**: Plots with the same property type and identical planted crops for both seasons is treated as ONE plot (no matter whether the plot(s) is(are) connected).
2. **A4 Contract type**: 1=oral; 2=written.
3. **A6 Crop**: 0=not planted; 1=wheat; 2=corn; 3=rice; 4=soybean; 5=peanut; 6=cotton; 7=inter-panting; 8=others.
4. **A7 Straw treatment (multiple choices)**:1=straw return (SR); 2=burn; 3=use as feed; 4=use as fuel; 5=for

sale; 6=discard; 7=no straw；8=others.

5. **A8 Reasons for SR adoption (multiple choices)**: 1=soil improvement; 2=fertilizer saving; 3=policy factors; 4= increasing production/income; 5=unable to handle; 6=environmental protection; 7=others.
6. **A9 Reasons for SR non-adoption (multiple choices)**: 1=affecting planting; 2=cost saving; 3=policy factors; 4=lack of laborers; 5=lack of machinery; 6=feeling troublesome; 7= avoiding infected by pests and diseases; 8=others.

## B. Production and Operation of Crop Planting in 2015

**B01 Purpose of farming（　　）**

1=to increase income; 2=to hold land property; 3=to eat；4=without other job opportunities if not farming; 5=scale management; 6=others.

**B02 What do you think is the most currently necessary technology for farming（　　）**

1=new varieties; 2=new fertilizers; 3=new pesticides; 4=new machinery; 5=new irrigation facilities; 6=others.

**B03 What do you think is the harm of excessive use of chemical fertilizer and pesticide（　　）**

1=causing residues harmful to human; 2=soil compaction; 3=waste of resources; 4=pollution；5=have to gradually increase the input of chemical fertilizer and pesticide; 6=reducing soil fertility; 7=no harm; 8=others.

| Crop index | W1 Wheat | W2 Corn | W3 Rice | W4 Soybean | W5 Other (__) |
|---|---|---|---|---|---|
| B1 Sown area (Mu) | | | | | |
| B2 Yield (kg/Mu) | | | | | |
| B3 Sale price (￥/kg) | | | | | |
| B4 Seed price (￥/Mu) | | | | | |
| B5a Base fertilizer cost (￥/Mu) | | | | | |
| B5b Top-dressing cost (￥/Mu) | | | | | |
| B6 Pesticide cost (￥/Mu) | | | | | |
| B7 Irritation cost (￥/Mu) | | | | | |
| B8 Machinery cost (￥/Mu) | | | | | |
| B9 Labor cost (￥/Mu) | | | | | |
| B10 Other cost (￥/Mu) | | | | | |

Note:

1. In this table, only record crops that involve straw treatment; for those without straws (e.g. watermelon, pepper, vegetables, etc.), no need to fill in.
2. B9 Labor cost: If the household hires workers at any part of the cost link (B5 – B8), count the cost as labor cost and fill it in B9, and leave the corresponding blank (B5 – B8) empty.

## C. Characteristics of Farmer Household and Members

| | | |
|---|---|---|
| Risk Awareness | **D1 Have you bought fake fertilizers, fake seeds or fake pesticides（　　）** | 0=No; 1=Yes |
| | **D2 Do you care about new agricultural techniques or products（　　）** | 0=No; 1=Yes |
| | **D3 If there is a newly developed pesticide or fertilizer, but no one has ever used it. Will you be willing to try it（　　）**<br>1=Absolutely not; 2=Try in some fields; 3=Use it in all fields | |
| | **D4 Do you get relevant information of straw return through the following media?**<br>a. neighbor/villager（　　）<br>b. relatives/friends（　　）<br>c. village committee（　　）<br>d. experts in agriculture（　　）<br>e. radio/television media（　　）<br>f. internet（　　）<br>g. other（please note____________________） | 0=No; 1=Yes |
| Household Information | **D5 Total # of family members_________** | |
| | **D6 Total # of family laborers_________** | |
| | **D7 # of family migrant members working outside the village in 2015_________** | |
| | **D8 Average # of working (outside) months in 2015_________** | |
| | **D9 Average wages for migrant members in 2015_________￥/month** | |
| | **D10 Total income of migrant members in 2015______________** | |
| | **D11 Does your family have a village cadre（　　）** | 0=No; 1=Yes |
| | **D12 Does your family have a party member（　　）** | 0=No; 1=Yes |
| | **D13 Whether a basic-living-allowance family（　　）** | 0=No; 1=Yes |
| | **D14 Does your family participate in agricultural economic organization（ ）** | 0=No; 1=Yes |
| | **D15 Does your family purchase agricultural insurance（　　）** | 0=No; 1=Yes |
| | **D16 Does your family participate in land trusteeship（　　）** | 0=No; 1=Yes |
| | **D17 Is there a family member who participated in technical training（　　）** | 0=No; 1=Yes |
| | **D18 # of households with which your family often contact________** | |
| | **D19 Are there any relatives who have settled down or worked in city（　）** | 0=No; 1=Yes |
| | **D20 In 2015, the annual household income was______________￥, among which the crop profit was______________￥** | |
| Household Head Information | **D21 Gender（　　）** | 0=Female; 1=Male |
| | **D22 Age________** | |
| | **D23 Health condition（　　）** | 1=Great; 2=good; 3=medium; 4=bad; 5=disabled |
| | **D24 Years of education________** | |
| | **D25 The highest # of years of education of all your family members is________** | |
| | **D26 Farming experience________years** | |

## Appendix 2: Additional Regression Results

Table A.1 Probit Regression Results for Renters Only (Dependent: adoption=1)

| Variable | (1) Base | (2) | (3) | (4) |
|---|---|---|---|---|
| Age | 0.0297 | 0.0289 | 0.0280 | 0.0276 |
| Farming experience | -0.0155 | -0.0150 | -0.0126 | -0.0124 |
| Risk preference | 0.3962 | 0.4074 | 0.4172 | 0.4215 |
| Education | 0.1200** | 0.1215** | 0.1156** | 0.1165** |
| Number of laborers | 0.4566*** | 0.4500*** | 0.4702*** | 0.4654*** |
| Income | -0.0058 | -0.0039 | -0.0294 | -0.0269 |
| Organization | 0.8375 | 0.8453 | 0.8886 | 0.8883 |
| Insurance | 0.3692 | 0.3212 | 0.3770 | 0.3508 |
| Plot size | 0.0118 | 0.0093 | 0.0071 | 0.0054 |
| July temperature | 0.7747 | 0.7950 | 0.7542 | 0.7644 |
| July precipitation | 0.0049** | 0.0050** | 0.0044** | 0.0045** |
| Winter season | 0.0772 | 0.0609 | 0.0886 | 0.0784 |
| Contract type | | 0.3711 | | 0.2126 |
| Term of lease | | | 0.1109 | 0.1058 |
| Intercept | -25.6801** | -26.2282** | -24.9656* | -25.2419** |
| Crop dummies | Yes | Yes | Yes | Yes |
| Observations | 155 | 155 | 155 | 155 |

Note: ***, **, and * represent significance at 1%, 5%, and 10% level, respectively. Standard errors are clustered at the household level. “Contract type” is a dummy equal to unity if the lease is written (oral otherwise). “Term of lease” is a variable indicating how many years the lease covers.