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Does Women's Land Ownership Promote Their Empowerment? Empirical Evidence from Nepal

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Summary. — Land rights equity is seen as an important tool for increasing empowerment and economic welfare for women in developing countries. Accordingly, the objective of this paper is to empirically examine the role of women's land ownership, either alone or jointly, as a means of improving their intra-household bargaining power in the areas of own healthcare, major household purchases, and visiting family or relatives. Using the 2001 and 2011 Nepal Demographic and Health Surveys and relevant econometric techniques, we find that land ownership has a positive and significant impact on women's empowerment. In particular, we find two important patterns of results. First, accounting for the endogeneity of land ownership with inverse probability weighting, coarsened exact matching and instrumental variable methods makes its impact on empowerment higher. Previous research in this area had largely ignored the potential endogeneity of land ownership. Second, the impact is generally stronger in 2011 than in 2001. As evidenced in a number of empirical studies, the increase in women's bargaining power can in turn translate into a redirection of resources toward women's preferences, including higher investment in human capital of the household such as education, health, and nutrition. Therefore, our study indicates that in places where agriculture is the main source of economy for women, policies enhancing land rights equity have the potential to increase women's empowerment and associated beneficial welfare effects.

Key words — gender, land ownership, empowerment, household decision making, Nepal, South Asia

1. INTRODUCTION

Women in many developing countries have long faced gender discrimination. This discrimination is fundamentally driven by societal views of women as economically less productive (due to their limited involvement in direct income-generating activities) and of lesser value to parents for the purpose of long-term asset accumulation (Anukriti, 2014). This problem is more acute in regions where the dowry system inflicts considerable costs on the girls' parents. [1] Sex-based discrimination has resulted in numerous negative outcomes for women. One prevalent consequence of this phenomenon is manifested in the "missing women" problem; for example, male-biased sex ratios are increasing in India and China, where parents selectively abort female fetuses to gain perceived and real benefits from sons and avoid losses from daughters (Sen, 1990).

In poor rural areas where agriculture is the primary source of income, women are wrongly perceived as even less valuable, mostly engaged in household work and less so in direct income-generating activities. Exacerbating this perception in rural areas is the centrality of land ownership, because women generally have restricted access to land. Although women constitute the majority of the agricultural workforce in developing countries (SOFA Team & Doss, 2011), they only control about 19% of agricultural land holdings (FAO., 2010).

Economic theories have predicted that access to assets, such as land, gives financial security to women and improves their household bargaining power (Agarwal, 1994, 1997; Anderson & Eswaran, 2009; Haddad, Hoddinott, & Alderman, 1997; Kabeer, 1999). The improvement in bargaining power in turn reduces gender discrimination by giving women more control over decisions that affect their lives (such as child bearing) and by a reallocation of resources toward women's preferences (Ashraf, Karlan, & Yin, 2010; Aslam & Kingdon, 2012; Doss, 2013; Malhotra & Schuler, 2005; Thomas, Contreras, & Frankenberg, 2002). In the same realm, Janssens (2010) finds that more empowered women are more likely to participate in community development projects such as construction and maintenance of schools, roads, and bridges. Furthermore, related studies have found that an increase in women's access to resources, including property rights, results in a higher investment in human capital such as education, health, and nutrition (Anderson & Eswaran, 2009; Doss, 2006; Duflo, 2003; Mason, 1996; Menon, van der Meulen Rodgers, & Nguyen, 2014; Pandey, 2010; Thomas, 1990; Wiig, 2013).

In light of this discussion, the objective of this paper is to empirically investigate the role of land ownership as a means of improving Nepali women's intra-household bargaining power (hereon referred to as household bargaining power). In doing so, we make a number of contributions to the literature on the determinants and effects of women's empowerment. First, our study adds to a limited number of studies that directly explore the relationship between women's land ownership and empowerment (Allendorf, 2007; Mason, 1996; Pandey, 2010; Wiig, 2013). More importantly, unlike many of these studies, we estimate the impact of women's land ownership on their empowerment using econometric methods that allow us to control for endogeneity of land ownership. [2] The endogeneity could arise in the usual sense with omitted factors that affect both women's land ownership and their decision-making power. For example, progressive households may have both more legal ownership of land by women and higher empowerment of women (Menon *et al.*, 2014). The

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endogeneity could also stem from reverse causality if more empowered women are more likely to own land. Estimates of land ownership on women's empowerment that fail to account for this endogeneity are prone to both bias and inconsistency. In order to deal with the endogeneity issue, we employ a propensity score method with inverse probability weight and an instrumental variable approach. We discuss the details of these two approaches in Sections 4 and 5.

Second, we take advantage of the most recent 2011 Nepal Demographic and Health Survey (NDHS) that was collected after the 2007 amendment of the Interim Constitution of Nepal, which for the first time granted equal rights to sons and daughters to ancestral property with no restrictions on age and marital status. We also supplement these data with the 2001 NDHS in order to draw a time wise comparison of the estimates. The 2011 data provide identification of women's land ownership in a household context by specifically capturing land ownership of other household members unlike in previous studies. This allows us to focus our study on women owners and non-owners of land in landed households only, unlike Pandey (2010) for example. We discuss the merit of restricting the sample to landed households in Section 4 below. Although Allendorf (2007) accounts for other family members' land ownership, families in Nepal (and in many developing countries) are often large, consisting of many households. Therefore, estimating a woman's land ownership in a family context might underestimate its impact on her bargaining power in a household context.

Third, we consider more relevant empowerment indicators: own healthcare decision, major household purchases, and visits to family or relatives, which better reflect the evolution of women's bargaining power. [3] We exclude decisions on what food to cook (Kishor, 1997) and purchase of daily needs (Allendorf, 2007) from our study because these decisions, although important, do not necessarily reflect empowerment since they are traditionally made by women anyway (Basu & Koolwal, 2005; Kabeer, 1999; Ministry of Health & Population (MOHP) [Nepal], New ERA, & ICF International Inc., 2012).

Our empirical results indicate two major findings. First, endogeneity-corrected logit and probit estimates are significantly higher than their uncorrected counterparts. For example, the instrumental variable (IV) estimate of women's land ownership on the probability of having the final-say in healthcare decisions is 2.76 times higher than that of the regular Probit estimate (Table 4, Panel B). Second, whether corrected or uncorrected for endogeneity, we find that the estimated effects of land ownership in 2011 are quantitatively higher than those in 2001. For example, women's land ownership increases the probability of having the final-say in major household purchases by 19% in 2011 but only by 11% in 2001 (Table 4, Model 3). Together, our results demonstrate robustly that land ownership plays an important role in combating gender discrimination by enhancing bargaining power of women. The higher bargaining power in turn is expected to result in increased household and societal welfare as discussed above. This link calls for policies that enhance and facilitate land rights equity among men and women.

The remainder of the paper is structured as follows. Section 2 discusses the relevance of women's land ownership in Nepal and provides a background on women's land rights. Section 3 discusses the empirical framework. Section 4 describes the data and construction of the variables of interest. The discussion of empirical methodology and results are presented in Section 5. Section 6 concludes and discusses policy implications.

2. NEPALI CONTEXT

Nepal is an excellent case for studying the impact of land ownership on empowerment for a number for reasons. First, Nepal has adopted a string of progressive laws of land rights equality with the latest passed in 2007, which states that sons and daughters have equal rights to inheritance regardless of their age and marital status (more on this below). Second, land ownership is vital in Nepal given the preponderant role of agriculture in Nepal's economy, as it is for many developing countries. According to the Ministry of Agricultural Development [Nepal], 2014, over 66% of the total population is employed in the agricultural sector, which contributes to 36% of Nepal's GDP. In a society where agriculture is the major source of income, land ownership (including size and quality) is critical to social status and economic participation (Bhandari, 2001; Sharma, 1999).

Third, the noted importance of the agriculture sector notwithstanding, there is a significant asymmetry between women's agricultural labor force participation and their land ownership. Only 19% of women own land even though they are predominantly engaged in agriculture (Ministry of Health and Population (MOHP) [Nepal], New ERA, and ICF International Inc., 2012) and are responsible for most of the agricultural activities such as fertilizing, transplanting, and harvesting (Acharya & Bennet, 1983; Pun, 2000). The involvement of women in agriculture has increased over time because more men are migrating into non-agricultural work to urban areas or abroad, creating "feminization of agriculture" (Asian Development Bank, 1999; Crowley, 1998) in labor participation and in some cases in decision-making. The latter is higher in households where women are the de facto household heads as opposed to those living in patrilineal households of in-laws (Gartaula, Niehof, & Visser, 2010). [4] As for labor force participation, over 90% of women workers were employed in the agricultural sector, compared to 64% of male workers in 2001 (Ministry of Health [Nepal], New ERA, & & ORC Macro, 2002). Fourth, the progressive changes to the Nepali constitution, with very limited land rights for women in 1977 to equal rights in 2007, can serve as a legislative springboard for advancing gender equality for asset ownership rights in countries that are struggling to achieve it.

The main way of gaining land in Nepal is through inheritance, which is largely patrilineal. Otherwise, women gain access to land or property through kinship or marital relationships to men. Nepali property law has its roots in Manusmriti; [5] influenced by this book, Nepal's first legislation (the 1853 National Code) restricted women's property rights to gifts and bequests. As long as the father, mother, brothers, brothers' sons, or other male relatives on the father's side were alive, a daughter could not inherit paternal property. Divorced women did not have any property rights, and if they instigated a divorce, they lost potential alimony. Over one hundred years later, the 1963 amendment of the National Code addressed some of the gender discriminatory issues, but did not address property rights (Pandey, 2010).

It was not until 1977, that a constitutional fix (sixth amendment of the 1963 National Code) brought changes to women's land ownership rights. An unmarried daughter of 35 years or older was equally entitled to parental property as her brothers. However, she had to return the property after she was married unless her mother, father, brothers and brothers' sons were dead. A married woman of 35 years or older (30 years for widows) was entitled to a portion of her husband's property if she had been married for 15 years. Although ancestral property was to be equally divided among the mother, father and sons, sharing the property with anyone other than the rightful owners required consent from the husband and sons. Women only enjoyed absolute legal rights to property such as gifts, bequests, wills, and personal earnings (Malla & Shrestha, 2000).

More relevant to our research are the two constitutional amendments passed in 2002 and 2007, which significantly improved land ownership rights of Nepali women. Specifically, the 2002 amendment expanded women's rights by guaranteeing equal inheritance of property at birth by sons and unmarried daughters, providing married women rights to a share of their husband's property immediately after marriage, and lifting the age limit on widows. However, daughters had to still renounce their share of the inherited property upon marriage (Shrestha, 2008). The latest amendment of the Interim Constitution in 2007 further forbade gender-based discrimination regarding land ownership. It guaranteed joint land ownership by both husband and wife of the land provided by the state, and implemented policies to facilitate a wife's joint ownership of her husband's land. The amendment also removed the "unmarried" requirement from inheritance of property by daughters, meaning both sons and daughters have equal rights to ancestral property regardless of their marital status (Nepal Law Commission, 2007). On the one hand, these fundamental changes in the later amendments in 2002 and especially 2007 give a sense of security over land ownership to women, which may translate into higher household bargaining power. On the other hand, lack of awareness of these legislative developments among eligible women (Pandey, 2010) and social norms of patrilineal inheritance may prevent them from exercising their constitutional rights. Therefore, it may take some time for the expected social changes arising from these legislative developments to unfold, making this a relevant topic to explore and draw policy conclusions from.

3. EMPIRICAL FRAMEWORK

Before delving into the empirical framework, we provide a brief discussion of the theoretical work underpinning our econometric specification. Various theoretical models have been developed to understand household bargaining mechanisms. Manser and Brown (1980) were the first to analyze household decision-making as a bargaining problem between husband and wife with separate utility functions. This model assumes that bargaining is conducted in the shadow of divorce such that the better off one partner would be upon divorce (relative to the other partner), the better compensation is provided for him or her to stay in marriage. Hence the threat of divorce (when supplemented by legal rights) increases the wife's bargaining power because by law she is entitled to a share of the household assets (e.g., land) in case the marriage dissolves. However, the threat of divorce may not be common practice or practical in many marriages. For societies where divorce is not culturally popular – such as in South Asia, another model proposed by Lundberg and Pollak (1993) seems more plausible. This model assumes that bargaining is conducted under the threat of noncooperation whereby a spouse will refuse to share his or her individually owned resources while still remaining in marriage. Consequently, this model implies that a change of the ownership of assets from husband to wife will increase the utility of the wife and decrease the utility of the husband if cooperation fails. Therefore, an increase in a woman's land ownership and thus an increase in her contribution to the household gives her more bargaining power.

Even more closely related to our paper is Agarwal's (1997) work which draws from Agarwal (1994) and conceptualizes ownership and control of arable land as the most important pathway for improving household bargaining power for women in agrarian economies.[6] Agarwal (1994, 1997) offers several explanations and anecdotal evidence supporting her theory. For example, she argues that in times of an economic crisis, households first sell their liquid assets which tend to be the women's only assets (e.g., jewelry, small animals) while the land, the main productive resource, is kept. This pecking order of asset disposal in times of crisis leaves women in a more vulnerable position (than men), but also with a more reduced ability to contribute to household income. This puts them at high risk of little to no bargaining power. These arguments are also echoed by Haddad *et al.* (1997), who claim that joint ownership of assets upon marriage and equal rights upon divorce (if guaranteed by the social norms and legal environment) affect households' allocation decisions through changes in the relative bargaining positions of household members.

Based on these theories, we hypothesize that women's relative increase of land ownership (alone or jointly) compared to men, will increase their relative bargaining power. We therefore model the empirical relationship between empowerment (outcome variable), land ownership and other determinants of empowerment as follows:

$$Y_{ijk} = \alpha + \theta L_{ijk} + \boldsymbol{V_{ij}}\beta + \boldsymbol{X_{jk}}\lambda + \boldsymbol{G_k}\gamma + \varepsilon_{\boldsymbol{ijk}} \quad (1)$$

where i, j, and k index individual, household, and ecological zone, respectively. θ is the parameter of interest that measures the impact of women's land ownership (L_i) on empowerment variables defined by household decision-making in own healthcare, major household purchases, and visits to family or relatives, and represented by Y_{ijk}.[7]

We control for several other covariates that have been hypothesized to impact empowerment. V_i is a vector of the respondent's individual characteristics that may impact empowerment measures. Following Allendorf (2007) and Kabeer (1999), we control for the level of women's education, employment remuneration, age, and whether the respondent is the household head's wife. In South Asian countries, age is a proxy for authority, increasing the likelihood of both land ownership and empowerment (Basu & Koolwal, 2005; Kabeer, 1999; Kishor & Gupta, 2004; Mahmud *et al.*, 2012). Similarly, being the wife of the household's prime decision maker can give an upper hand in influencing decision-making. Education and employment can empower women socially and economically by giving them economic independence and a sense of self-worth (Kabeer, 1999; Kishor & Gupta, 2004; Samarakoon & Parinduri, 2015; Trommlerová, Klasen, & Leßmann, 2015). We also control for spousal age differences, total number of children ever born, media exposure, and the husband's education level based on a study by Kishor and Subaiya (2008), which found these variables to be relevant covariates of women's empowerment in Nepal. Media exposure (defined by access to radio, television, and newspaper at least once a week) is an important source of information for women, exposing them to the changes happening elsewhere since many of these women are culturally walled off from the outside world (Aslam & Kingdon, 2012; Basu & Koolwal, 2005; Kishor & Gupta, 2004; Mahmud *et al.*, 2012). Although the expected effect of the husband's education on empowerment is ambiguous ex-ante, several studies have found it significant for women's household bargaining power (Basu & Koolwal, 2005; Kishor & Gupta, 2004). Therefore, we include this variable in our model.

X_j is a vector of household characteristics that are hypothesized to impact empowerment measures. For example, religion and caste play important roles in societal and legal systems (such as women's land ownership and decision-making roles) in South Asia. Following previous studies (Basu & Koolwal, 2005; Kabeer, 1999; Trommlerová *et al.*, 2015), we control for caste and religion in our model. Likewise, a woman's participation in household decision-making is likely associated with wealth because wealthier households have more access to information and resources that can affect women's empowerment (Basu & Koolwal, 2005; Kabeer, 1999; Kishor & Subaiya, 2008). Wealth is controlled for by an index, which is a composite measure constructed using household assets, construction materials of the house, and water access and sanitation facilities (Rutstein & Johnson, 2004). The traditional barriers to women's empowerment are likely stronger and harder to challenge in rural areas where women are often relegated to subordinate roles than in urban areas (Kabeer, 1999; Kishor & Gupta, 2004; Trommlerová *et al.*, 2015). We therefore control for whether the respondent resides in a rural or urban area.

G_k is a set of ecological zone (mountains, hills and Tarai [8]) dummies, which are important to control for in this model because these regions are significantly different both culturally and in terms of soil fertility. This dissimilarity in culture and soil quality may create variation in land ownership and inheritance patterns across the regions. Finally, ε_i is a vector of individual level unobserved characteristics.

As noted earlier, an empirical challenge in identifying the causal effect of land ownership on bargaining power is the endogeneity of land ownership. First, there could be reverse causality due to empowered women being more likely to inherit land from parents (e.g., those that are liked by their parents more). Conversely, parents may be more inclined to give land to children that lag behind (less empowered) since more empowered women may draw less utility from their inheritance (Wiig, 2013). Second, there could be omitted factors that affect both women's land ownership and their empowerment. For example, women advocacy groups may lobby to increase women's land ownership as well as their autonomy in the household. A preliminary analysis of a mean comparison *t*-test between land owning and non-owning women suggests that these two groups are significantly different. Therefore, to deal with the potential bias caused by endogeneity, we use propensity score matching and instrumental variable (IV) approaches, which we discuss below.

4. DATA

(a) *Data source*

This study uses the 2001 and 2011 Nepal Demographic and Health Surveys (NDHS), conducted by the Nepal Ministry of Health and funded by the United States Agency for International Development (USAID). The NDHS is a nationally representative cross-sectional household survey, and its objective is to provide reliable estimates for population characteristics such as fertility, contraceptive prevalence, health indicators, infant mortality, and women's empowerment (Ministry of Health [Nepal], New ERA, and & ORC Macro, 2002; Ministry of Health and Population (MOHP) [Nepal], New ERA, and ICF International Inc., 2012). The empowerment section collects information on women's participation in three types of decision-making: own health care, major household purchases, and visits to family or relatives. In addition, the 2001 NDHS collected information on women's decision-making in household purchases for daily needs and what food to cook; however, it was found that these two decisions were predominantly made by women anyway (Ministry of Health [Nepal], New ERA, and & ORC Macro, 2002) and were excluded from the 2011 NDHS. The survey uses a stratified two stage cluster method to sample "ever married" individuals of ages 15–49. The 2001 NDHS surveyed 8,726 women from a total of 8,633 households with a response rate of 98%. Similarly, the 2011 NDHS surveyed a total of 12,674 women from 10,826 households with a response rate of 98%.

Our study sample is based on household structure and occupation. Following Allendorf (2007), we select women who are involved in agriculture and whose households own land. This is because the information on other household members' land ownership is only available for women who are engaged in agriculture. It is important to correctly identify the independent variable and limit the study to landed households, else we would be comparing land owning women in landed households to those in landless households. This could bias the estimates since households with land are generally wealthier and have more access to resources that may impact both women's land ownership and empowerment (Allendorf, 2007). A simple mean comparison *t*-test of women across landed and landless households shows that the characteristics of women in households with and without land are significantly different, so comparing across them could produce biased estimates. [9] As discussed in the model specification above, we also include household wealth index in our model because wealthier households may have more access to information and resources that can affect women's empowerment. We further limit the sample to those women who are currently married and residing with their husbands because women who are unmarried or live without their husbands could be the primary decision makers in their households. Including such women could result in upward biased estimates of land ownership on empowerment. After eliminating observations with incomplete information, we end up with 4,066 and 3,047 observations in 2001 and 2011, respectively. [10]

(b) *Construction of primary variables*

(i) *Land ownership dummy variable*

For the 2011 NHDS, we construct the land ownership dummy variable as one if the woman owns land (alone or jointly) and her household members own land, and zero if the woman does not own land herself but her household members own land. For the 2001 NDHS, however, information on other household members' land ownership is not available. Instead, women are asked if they work on family land or someone else's land. Following Allendorf (2007), we utilize this information as a proxy to identify whether other household members own land. Specifically, the land ownership variable equals one if the woman owns land (either alone or jointly) and she works on family land, and zero if the woman does not own land and she works on family land. This way of constructing the land ownership variable results in about 10% women who own land either jointly or alone in both 2001 and 2011 (Table 1). [11]

(ii) *Instrumental variable*

We employ an instrumental variable (IV) approach as one way to deal with the endogeneity of land ownership discussed earlier. For the IV approach to provide consistent estimates,

Table 1. *Descriptive statistics for variables analyzed in the study*

	2001 (observations = 4,067)		2011 (observations = 3,047)	
	Sample median	Std. Dev.	Sample median	Std. Dev.
Age	31 years	9.08	34	9.07
Age at first marriage	16 years	2.83	17	3.08
Spousal age difference	4 years	4.73	4	4.85
# Children ever born	3	2.49	3	2.14
	Sample Prop. (%)	Std. Err.	Sample Prop. (%)	Std. Err.
Own healthcare	20.43	0.006	57.76	0.009
Major household purchases	24.24	0.007	47.29	0.009
Visits to family or relatives	32.28	0.007	57.76	0.009
Aggregate decision scale				
0	57.51	0.008	27.01	0.008
1	20.29	0.006	18.64	0.007
2	9.93	0.005	18.87	0.007
3	12.27	0.005	35.48	0.009
Final-say-alone	13.15	0.005	29.01	0.008
Women own land	9.74	0.005	9.58	0.005
Urban residence	5.56	0.004	11.94	0.006
Hindu	85.41	0.006	86.84	0.006
High caste	37.57	0.008	46.31	0.009
Household head's wife	71.60	0.008	72.79	0.008
Women's education				
None	77.40	0.007	24.06	0.009
Primary	14.33	0.005	17.16	0.007
Secondary or higher	8.26	0.004	19.75	0.007
Husband's education				
None	36.96	0.008	24.06	0.008
Primary	27.81	0.007	28.49	0.008
Secondary or higher	35.23	0.007	47.46	0.009
Media exposure	1.42	0.008	58.39	0.009
Owns livestock	27.83	0.007	96.68	0.003
Employment remuneration				
Unpaid	89.03	0.005	88.22	0.006
Paid in cash plus kind	2.21	0.002	6.20	0.004
Paid in kind only	8.75	0.002	5.58	0.004
Household wealth				
Poorest	29.16	0.007	32.88	0.009
Poorer	20.68	0.006	24.71	0.008
Middle	16.62	0.006	21.76	0.007
Richer	20.75	0.006	14.15	0.006
Richest	12.79	0.005	6.50	0.004
Ecological region				
Mountain	19.77	0.006	27.99	0.008
Hill	41.65	0.008	45.29	0.009
Tarai	38.58	0.007	26.71	0.008

we need a variable (instrument) that is highly correlated with land ownership but does not directly explain empowerment, i.e., the instrument should only affect empowerment indirectly via land ownership. In the absence of a panel dataset, we opt for a matching procedure to obtain a valid instrument for the 2011 wave of the survey.[12] Specifically, we match the 2001 sample to the 2011 sample using the coarsened exact matching (CEM) method based on the following individual and household characteristics: urban location, women's education, caste, exposure to media, household wealth index, number of children ever born, age at first marriage, and husband's education (see Table 5 in Appendix A for matching). The idea is to obtain a subset of women in the two waves that are similar in observables so that "lagged" land ownership status based on the matched data mimics that of unobserved true lagged land ownership status. In other words, we seek to create a counterfactual land ownership status in 2001 for women in the latest wave and use that as an instrument (see endnote 12).

Procedurally, the CEM method divides the sample into several strata, each of which has identical values for all the coarsened preoperative covariates. Then, the CEM drops all individuals from any stratum that does not have at least one 2001 observation for each unique observation from 2011 (David, Rawley, & Polsky, 2013).[13] This process results in a total of 462 matched strata. Next, we take means of the land ownership variable per stratum for the 2001 sample. We then code this average as one (indicating land ownership) for values greater than 0.5, and zero (indicating non-ownership of land) otherwise. This newly created variable based on 2001 land ownership is our instrument for the 2011 land ownership of women. Furthermore, we control for relevant socioeconomic variables as mentioned earlier in order to isolate all channels via which the instrument could directly impact empowerment. We argue that the 2011 women who are similar in socioeconomic conditions to the 2001 women are equally or even more likely (due to the 2002 and 2007 constitution amendments) to have land ownership. The partial F-test from the first stage regression (of the instrument and remaining covariates on land ownership) is 10.80, corroborating the relevance of the instrument in the sense of Bound, Jaeger, and Baker (1995). While we cannot directly test the exogeneity assumption since we only have one instrument, we conduct a falsification test for a subsample of women who are heads of their households in 2011 (therefore excluded from our sample). To do so, we regress the empowerment variables on the instrument and other controls; if indeed the instrument directly causes women's empowerment, its coefficient estimate would be statistically significant. However, our results indicate that the 2001-based instrument is not a significant predictor of the 2011 empowerment measurements at any conventional level of significance, suggesting that it does not directly affect empowerment. These results are available upon request.

(iii) *Empowerment variables*

We draw from the extant literature as well as the NDHS' measures of empowerment to construct various measures of the outcome variable. Specifically, we construct three binary indicators of household decision-making processes: own healthcare, household purchases, and visits to family or relatives. These variables equal one if the women have the final-say (alone or with husband) for healthcare decisions, major household purchases, and visits to family or relatives, respectively, and zero otherwise. Following Allendorf (2007), we also create an aggregate measure of empowerment, *aggregate decision scale*, by summing the three binary variables defined above. This results in an ordered scale from zero to three, where three indicates the highest level of empowerment. Moreover, we create another binary indicator, *final-say-alone*, with an aim of better identifying women's sole autonomy in decision-making. We assign the value one to *final-say-alone* if the women are the sole final decision makers in any of the three household decisions, else zero; this results in a more strict definition of empowerment. The proportion of women with the *final-say-alone* in each of the three household decisions is very low. Thus, we cannot construct separate categories for this measure.

5. DISCUSSION OF RESULTS

Table 1 presents the descriptive statistics of our data, categorized by wave. The descriptive statistics indicate that women's bargaining power has increased significantly from 2001 to 2011. For example, in 2001, 20% of women reported having final-say in their healthcare, 24% in major household purchases and 32% in visits to family or relatives; these percentages increase to 58, 47 and 58%, respectively in 2011. Similarly, the proportion of women having *final-say-alone* in one or more household decisions more than doubles from 13% in 2001 to 29% in 2011. Other factors that may impact empowerment, such as women's average age, age at first marriage, and education level have also increased during 2001–11. For example, the average age of women interviewed in the 2011 wave is 34 years, compared to 31 years in the 2001 wave, and about 20% of women have secondary or higher education in 2011, compared to only about 8% in 2001.

We estimate several variants of our empirical model, which we group in three phases, in order to obtain results that are robust to alternative measures of empowerment, functional form specification, endogeneity of land ownership, and to explore if and how the impact of land ownership on empowerment changes over time. The estimation phases are designed to progressively build more robust results.

In the first estimation phase, we employ ordered logit and ordinary logit models. The ordered logit model is used for our *aggregate decision scale* and requires a proportionality of odds assumption, which implies that the relationship between pairs of the empowerment scale (i.e., zero to three) is the same. To account for the possibility of sample selection with respect to land ownership based on observable characteristics, we utilize a propensity score method with inverse probability weights (IPW) (Rosenbaum & Rubin, 1983). The IPW approach yields an estimate of the causal effect of land ownership on empowerment if land ownership status is explained solely by observed data.

In the second estimation phase, our goal is to more accurately explore if the impact of land ownership on empowerment has changed over time. We predict that the impact estimates might be higher for 2011 compared to 2001 for two possible reasons: (i) the definition of land ownership variable in a household context in 2011 (as opposed to a family context in 2001) and (ii) the 2002 and 2007 constitutional amendments in favor of land rights equity for women. In order to construct a fair comparison across the two waves, we match the 2001 and 2011 waves by CEM method over covariates that are significantly different across the waves (see instrumental variable construction in Section 4 above for a list of the covariates). Hence, we build a more stringent matched sample in order to obtain comparable estimates of land ownership over the waves.

In the third and final phase, we drop the ordered logit model and conduct analyses on disaggregate measures of empowerment. This allows us to obtain differential impacts of land ownership on the three indicators of empowerment separately. Furthermore, we allow for selection based on unobservables and reverse causality with respect to land ownership by employing an instrumental variable approach.

The empirical results for phases one, two and three are presented in Tables 2–4, respectively and discussed below. For space consideration and expositional simplicity, the tables omit the coefficients on other controls and focus on the coefficients for land ownership alone.

(a) *Comparing the logit, ordered logit and IPW models for 2001 and 2011*

Table 2 reports ordered logit (Models 1 and 2) and logit (Models 3 and 4) estimations. Models 1 and 3 are estimated assuming exogeneity of land ownership as in Allendorf (2007), Pandey (2010), and Wiig (2013) while Models 2 and 4 are estimated using the IPW propensity score approach to eliminate differences in observable characteristics between land owning and non-land owning women.[14] All of our models include the control variables discussed in the model specification above.[15] The IPW method is considered an efficient propensity score method for small samples (no observations are lost) and does not assume any functional form (Hirano & Imbens, 2001; Hirano, Imbens, & Ridder, 2003).

Table 2. *Odds ratios from the ordered logit model of the aggregate decision scale and the logit model of the final-say-alone dummy for 2001 and 2011*

	Aggregate decision scale		Final-say-alone	
	Model 1	Model 2	Model 3	Model 4
Panel A: 2001 wave				
Owns land	1.60***	1.71***	1.51***	1.35*
	(0.16)	(0.24)	(0.21)	(0.22)
Number of observations	4,066	4,066	4,066	4,066
- 2 Log likelihood	8,656	18,775	3,014	6,534
Model chi-squared	475***	14***	151***	3***
Panel B: 2011 wave				
Owns land	1.97***	2.38***	1.42**	2.43***
	(0.26)	(0.47)	(0.20)	(0.48)
Number of observations	3,047	3,047	3,047	3,047
- 2 Log likelihood	7,708	15,535	3,554	7,861
Model chi-squared	495***	19***	116***	20***
Additional covariates	Yes	Yes	Yes	Yes
Ecological region control	No	Yes	No	Yes

Notes: (i) The standard errors are robust for all the models and are given in parenthesis. (ii) $^{*}p < 0.1$, $^{**}p < 0.05$ and $^{***}p < 0.01$. (iii) Models 1 and 3 are based on control variables in Allendorf (2007): urban, Hindu and caste dummies, household wealth, age, women's education level, employment remuneration, and whether the respondent is household head's wife and owns livestock. Models 1 and 3 are ordinary ordered and binary logits, respectively. (iv) Models 2 and 4 contain additional controls for the propensity score: spousal age difference, total number of children ever born, media exposure, and ecological region dummy. These models employ inverse probability weighing (IPW) with propensity scores.

Table 3. *Odds ratios from the ordered logit model of the aggregate decision scale and the logit model of the final-say-alone dummy for 2001 and 2011 matched dataset*

	Aggregate decision scale		Final-say-alone	
	Model 1	Model 2	Model 3	Model 4
Panel A: 2001 wave				
Owns land	2.01***	2.11***	1.72***	1.74***
	(0.27)	(0.39)	(0.31)	(0.37)
Number of observations	2,709	2,709	2,709	2,709
- 2 Log likelihood	5,626	12,486	1,896	4,326
Model chi-squared	321***	16***	87***	7***
Panel B: 2011 wave				
Owns land	2.23***	2.56***	2.22***	3.63***
	(0.47)	(0.69)	(0.49)	(1.04)
Number of observations	1,316	1,316	1,316	1,316
- 2 Log likelihood	3,264	6,654	1,470	3,272
Model chi-squared	268***	12***	80***	20***
Additional covariates	Yes	Yes	Yes	Yes
Ecological region control	Yes	Yes	Yes	Yes

Notes: (i) The standard errors are robust for all the models and are given in parenthesis. (ii) $^{***}p < 0.01$. (iii) Controls included are: urban, Hindu and caste dummies, household wealth, age, women's education level, husband's education level, employment remuneration, whether the respondent is household head's wife, spousal age difference, number of children ever born, media exposure, and ecological regional dummy. (iv) Models 1 and 3 are ordinary ordered and binary logits, respectively. Models 2 and 4 employ inverse probability weighing (IPW) with propensity scores.

The results in Table 2 indicate that the odds of high *aggregate decision scale* versus the combined middle and low *decision scale* are 1.60 times higher for land owning women than for non-owners for 2001 (Panel A, Model 1). Using the IPW approach results in slightly higher odds of 1.71 (Panel A, Model 2). Similarly, the logit model for binary variable predicts that the odds of having the *final-say-alone* in one or more of the decisions are 1.51 times higher for land owning women than for non-land owners (Panel A, Model 3). The results for 2011 follow a similar trend. For example, the odds of high *aggregate decision scale* versus the combined middle and low *decision scale* are 1.97 times higher for land owning women than for non-owners (Panel B, Model 1). With the IPW, the same odds increased to 2.38 (Panel B, Model 2). The odds of *final-say-alone* are 1.42 and 2.43 times higher using the binary logit and IPW approaches, respectively for 2011 (Panel B, Models 3–4). One possible reason for this pattern within years (with higher IPW estimates) can be attributed to selection based on observables. A mean comparison *t*-test between the land owning and non-owning samples within the 2001 and 2011 waves shows that these samples are significantly different on most of the observables with an exception of caste and higher level education (results are available upon request). We note, however, that differences between land owning and

Table 4. *Marginal effects from the Probit, IPW-PS and CMP-IV models for the disaggregated household decisions*

	Own healthcare		Major household purchases		Visits to family or relatives	
	Model 1	Model 2	Model 3	Model 4	Model 5	Model 6
Panel A: 2001 wave						
Owns land	.0884***	.1021**	.1092***	.1659***	.1436***	.1553***
	(0.03)	(0.04)	(0.06)	(0.04)	(0.03)	(0.05)
Number of observations	2,709	2,709	2,709	2,709	2,709	2,709
-2 Log likelihood	2,536	5,676	2,632	6,273	3,104	6,834
Model Chi-Squared	121***	8***	250***	16***	243***	12***
Panel B: 2011 wave						
Owns land	.1391***	.3836**	.1888***	.2347	.1885***	.2880*
	(0.05)	(0.18)	(0.05)	(0.17)	(0.05)	(0.16)
Number of observations	1,316	1,316	1,316	1,316	1,316	1,316
- 2 Log likelihood	1,686	2,338	1,574	2,224	1,548	2,200
Model Chi-Squared	111***	272***	220***	336***	217***	351***
Hausman Endo. test	0.03	0.20				
Additional covariates	Yes	Yes	Yes	Yes	Yes	Yes
Eco. region control	Yes	Yes	Yes	Yes	Yes	Yes

Notes: (i) The standard errors are robust for all the models and are given in parenthesis. (ii) $^{*}p < 0.1$, $^{**}p < 0.05$ and $^{***}p < 0.01$. (iii) Controls included are: urban, Hindu and caste dummies, household wealth, age, women's education level, husband's education level, employment remuneration, whether the respondent is household head's wife, spousal age difference, number of children ever born, media exposure, and ecological regional dummy. (iv) Models 1, 3 and 5 are binary probits. Models 2, 4, and 6 employ inverse probability weighing (IPW) with propensity scores for the 2001 wave and instrumental variable (CMP-IV) for the 2011 wave.

non-land owning women may be attributable to unobservable factors making it necessary to use an instrumental variable approach to identify a causal effect. We turn to this task later.

When comparing estimates over the 2001 and 2011 waves, we find that the pattern is consistent with the sample statistics; i.e., the odds of empowerment are generally higher for 2011 (with an exception of Model 3). For example, the odds of empowerment for *aggregate decision scale* are 1.71 and 2.38 times higher for 2001 and 2011, respectively, for land owning women (Model 2). Likewise, the odds of having *final-say-alone* are 1.35 times higher for land owning women for the 2001 wave, compared to 2.43 for the 2011 wave (Model 4).

(b) *Constructing a better comparison over the 2001 and 2011 waves*

In order to draw a parallel comparison between 2001 and 2011, we match the two waves on several covariates using the CEM method (Blackwell *et al.*, 2009).[16] This method produces a more stringent sample with 2,709 and 1,316 matched observations for 2001 and 2011, respectively (Appendix A, Table 5). Then, we employ logit, ordered logit and IPW methods to estimate our models as done in the first phase using the control variables discussed in the model specification.

The results from the matched sample are presented in Table 3. We find that the impacts of land ownership on empowerment measures are statistically significant and positive in all models. When comparing within the years, again, the estimated impact is higher when controlling for selection based on observables for both 2001 and 2011. Across years, we find that the odds of empowerment are higher for 2011, compared to 2001 for all four models. For example, the odds of being in the higher *aggregate decision scale* for land owning women are 2.11 times higher for 2001 and 2.56 times higher for 2011, compared to women without land (Model 2). Likewise, the odds of having *final-say-alone* for land-owning women are 1.74 and 3.63 times higher for 2001 and 2011, respectively (Model 4). These results suggest that land ownership has a higher impact on empowerment in the 2011 wave. This trend could have been spurred by two factors. As mentioned earlier, our definition of land ownership for 2011 is limited to a household unit, as opposed to the one used in 2001, which is based on a family unit (due to data limitations). Since a household is a smaller unit, both the proportion of women with land and their participation in household decision-making may be higher in a household context, which could produce a higher impact estimate. Therefore, estimating the woman's land ownership in a family context might underestimate its impact on her bargaining power in a household context. Another possible explanation is the major change in the constitutional rights of women pertaining to land ownership with the passage of the 2002 and 2007 amendments, which occurred after the 2001 wave. These constitutional changes could have brought a sense of security over the use and ownership of land, which in turn could have led to greater confidence and participation in household decision-making. Unfortunately, we cannot directly test for these conjectures and hence only speculate if one versus the other or both caused increased women's empowerment over the years.

(c) *Exploring the impact of land ownership using disaggregate data and an IV method*

As discussed above, the models in the previous phases make two simplifying assumptions. First, the proportionality of odds assumption built into the ordered logit means that the distance between each category of empowerment scale is equivalent. We employ the Brant test to test this proportional odds assumption (Brant, 1990). The test gives a p-value of 0.01, suggesting rejection of the assumption. This signals that aggregating indicators of empowerment may not be the best approach to analyze empowerment. Hence in this phase, we examine the impact of land ownership on disaggregate measures of empowerment, which also has the virtue of allowing differential impacts of land ownership on the three indicators of empowerment separately. Second, the assumption of sample selection based on observables is inadequate to address endogeneity of land ownership in the presence of reverse causality or selection based on unobservables. In order to deal with the endogeneity issue, we employ an IV approach using the instrumental variable described above. To implement our IV estimator, we opt for the probit model, since it lends itself more readily to the conditional mixed process (CMP) model, which incorporates the IV approach (Chyi & Mao, 2012; Ruppert, Stancanelli, & Wasmer, 2009).[17]

The marginal effects of women's land ownership on disaggregate binary household decisions are presented in Table 4. We find that the significant positive impact of women's land ownership remains even after relaxing the proportional odds assumption and accounting for endogeneity issues (with an exception of Panel B, Model 4). Parallel to the findings in previous phases, we get consistently higher marginal effects for the 2011 wave relative to the 2001 wave and when we control for endogeneity.[18] For example, per Models 1 and 2, women who own land are 13.91 and 38.36% more likely to have the final-say in their healthcare decision using the probit and CMP-IV models, respectively, for 2011 (Panel B). Similarly, the probability of having the final-say in decisions regarding visits to family or relatives is 9.95 percentage points higher when using the CMP-IV, compared to the probit model (Panel B, Models 5–6). Although the IV estimate of the effect of land ownership on the final-say in major household purchases is higher than the probit estimate and positive, it is insignificant (Panel B, Models 3–4). This pattern of obtaining higher coefficient estimates when accounting for sample selection and endogeneity suggest that the impacts of land ownership on empowerment in Nepal reported in Allendorf (2007) and Pandey (2010) are downward-biased.

Overall, we find empirical support for the hypothesis that women's land rights are important for increasing women's bargaining power (Agarwal, 1997; Anderson & Eswaran, 2009; Haddad *et al.*, 1997; Kabeer, 1999). We therefore contribute to the extant literature on the causes and consequences of women's empowerment and specifically to the limited literature that directly investigate the impact of women's land ownership on their empowerment as mentioned earlier.

6. CONCLUSION AND POLICY IMPLICATIONS

This paper empirically examines the impact of land ownership on empowerment for cross-sections of Nepali women interviewed in two rounds of the NDHS survey. By employing several econometric techniques and robustness checks, we find that women's land ownership in Nepal significantly increases their empowerment, defined by household decision-making in areas of own healthcare, major household purchases, and visits to family or relatives. Furthermore, we find that the estimated impacts of land-ownership are higher over time and when an instrumental variable approach is used. The fact that the estimated impacts of women's land ownership on empowerment have increased over time could be the result of two factors. First, the data for the 2011 wave allow us to create a more precise definition of land ownership using the household as the basic social unit for the determination of land ownership rather than the family, which in Nepal can be comprised of several households. It is likely that both the proportion of women with land as well as women's participation in decision-making is higher in a household context than in a family context. Hence, a focus on households (as we do for the 2011 wave) could produce a higher impact estimate of land ownership. Second, the higher estimated impact of land ownership over time could be the outcome of stronger land ownership rights afforded to women by the constitutional amendments of 2002 and 2007. Moreover, our finding of higher estimated impacts with the IV approach highlights that it is crucial to address the sample selection and endogeneity caused by the non-random assignment of land ownership (Anderson & Eswaran, 2009; Wiig, 2013).

These results have important policy implications. First, they imply that land ownership can play a significant role in combating gender discrimination by enhancing the bargaining power of women. As evidenced in a number of empirical studies, this increase in women's autonomy is expected to translate into a redirection of resources toward women's preferences, including higher investment in human capital of the household such as education, health and nutrition (Aslam & Kingdon, 2012; Doss, 2006, 2013; Menon *et al.*, 2014). Therefore, in places where agriculture is the main source of economy for women, policies enhancing land rights equity have the potential to increase women's empowerment and associated beneficial welfare effects. While we cannot empirically test the impact of the land rights reforms of 2002 and 2007 in Nepal, it is important to note the impact of land ownership on empowerment seems to have risen following their enactment, suggesting that staggered implementation of similar progressive laws could result in more empowerment for women. Second, lack of education and awareness prevents poor rural women from claiming their constitutional rights. Therefore, administrative capacity building at the local level that can advance gender equity in land titling and disseminate information on its importance and procedures to local households can prove useful to women's empowerment. We caution, however, that legislative and local administrative reforms alone may not be enough to generate significant gains in women's empowerment without further institutional changes (such as access to credit markets and social safety nets) and changes in cultural attitudes that disfavor women (Menon *et al.*, 2014).

NOTES

1. The dowry system is most prevalent in South Asian countries such as India and Nepal.

2. Wiig (2013) is an exception; he assumes that land reform in rural Peru was an exogenous process, independent of both household and community characteristics.

3. Drawing from the extensive literature on empowerment, we consider a woman to have "bargaining power" within her household if she has the ability to influence decision-making processes pertaining to her own and household welfare. For example, if a woman is able to make decisions that affect her health without other household members' input then she is deemed to have all the bargaining power over those decisions.

4. Gartaula *et al.* (2010) find that male out-migration is causing increased agricultural labor force participation of women in Eastern Nepal but not necessarily managerial control of agricultural decisions. They find a higher level of "feminization" of managerial control in agriculture in households where women are the de facto autonomous heads-of-households but a lower level of "feminization" in cases where women stay within the patrilineal household of their parents-in-law who assume the decision-

making position of their husbands. Therefore, the authors caution against policy formulations that assume that women have primary control over agriculture (due to out-migration of men) and target them for long-term development of the sector.

5. Manusmriti (200 BC) describes the rules of life, including those for women. According to the rulebook, a woman should be under her father's control in childhood, husband's control in youth, and sons' control after her husband's death (translated by Doniger & Smith, 1991).

6. Agarwal (1997) states that a rural woman's intra-household bargaining power depends on eight factors, which can be conceptualized as:

$$\begin{gathered}\textit{Bargaining Power} = f(\textit{access to employment}, \textit{ access to} \\ \textit{communal resources}, \textit{ access to traditional social support system}, \\ \textit{NGO support}, \textit{ government support}, \textit{ social perception}, \textit{ socialnorms})\end{gathered}$$

7. See, for example, Allendorf (2007), Anderson and Eswaran (2009), Ashraf *et al.* (2010), Basu and Koolwal (2005), Kishor and Gupta (2004), Kishor and Subaiya (2008), Mahmud, Shah, and Becker (2012), Samarakoon and Parinduri (2015), and Weber and Ahmad (2014) for recent studies that have used household decision-making variables as measures of empowerment.

8. Tarai is the southern plain belt of Nepal with the most fertile soil, which is responsible for producing majority of the food supply in the country.

9. A mean comparison t-test between these two groups is available upon request.

10. Although the raw data for 2001 is smaller than that of 2011, the pattern is reversed for the subsamples in our empirical analysis. This is because for the 2011 dataset, we only select women whose households own land while for the 2001 dataset, we select women whose families own land (due to data limitations), a less strict screen.

11. Our 10% ownership rate in 2011 is at odds with the official figure in the National Population Census Report, which is reported as 19% in 2011. However, the definition of land ownership in the National Population Census includes the ownership of a house as well. Therefore, the land ownership for our study sample differs from the National Report and ranges around 10% for the 2011 sample.

12. If the NDHS surveys of 2001 and 2011 followed the same women (resulting in a panel), a candidate instrument for land-ownership in the cross-sectional regression based on 2011 data would have been land ownership status for the same women 10 years earlier as captured in the 2001 survey. The reason is that land ownership status in 2001 should still be highly correlated with land ownership 10 years later given the near-static nature of land ownership in Nepal and (in part because of the depth of the lag) should not *directly* affect empowerment measures in 2011 once other observable covariates (such as education, wealth, and caste, to name a few) are controlled for.

13. The CEM temporarily coarsens (categorizes) the data based on the specified covariates for matching. For example, years of education might be coarsened into primary, secondary, college, and university degrees. Then, observations with the same values for all the coarsened variables are grouped into a single stratum. Those strata that do not have at least one of the treatment and control observations are dropped out of the dataset. The CEM method does not require any functional form assumption and is based on ex-ante user choice (Blackwell, Iacus, King, & Porro, 2009; Iacus, King, & Porro, 2012).

14. Through the IPW method, we control for selection bias due to covariates in a two-step process. First, we estimate the propensity score ($p(X_i)$ by a logit model. Second, we employ logit or probit regressions by weighting the treatment (land owners) unit by $1/p(X_i)$ and the control unit (non-land owners) by $1/(1-p(X_i))$.

15. Models 1 and 3 incorporate control variables used in Allendorf's (2007). These are: urban, Hindu (religion) and caste dummies, household wealth, age, women's education level, employment remuneration, whether the respondent is household head's wife and owns livestock. Models 2 and 4 contain additional controls for the propensity score which are: spousal age difference, total number of children ever born, media exposure, and ecological region dummies based on other related studies (Aslam & Kingdon, 2012; Basu & Koolwal, 2005; Jejeebhoy & Sathar, 2001; Kabeer, 1999; Kishor & Gupta, 2004; Kishor & Subaiya, 2008; Mahmud *et al.*, 2012; Malhotra & Mather, 1997; Roy & Niranjan, 2004; Samarakoon & Parinduri, 2015; Trommlerová *et al.*, 2015).

16. The variables used for matching (urban location, women's education, caste, exposure to media, household wealth, number of children ever born, age at first marriage, and husband's education) are the key socio-economic characteristics of the Nepali population that were found to be significantly different across the 2001 and 2011 wave samples, based on a mean comparison t-test. Moreover, the survey methodology is the same across 2001 and 2011. These reasons justify the matching and trimming of the sample to make the waves as similar as possible.

17. The CMP is the first general Stata tool for discrete choice models, appropriate for instrumental variable approaches (Roodman, 2009).

18. We do not have an instrument for the 2001 wave and therefore the instrumental variable method is only done for the most recent 2011 wave.

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APPENDIX

Table 5. *Summary results from coarsened exact matching (CEM) for 2001 and 2011*

Variables	$L1$ stat.	Mean Diff.
Urban	4.80E−16	−1.90E−16
Women's education	1.10E−15	−2.20E−15
Hindu	7.20E−16	7.80E−16
Caste	2.10E−15	−5.60E−16
Media exposure	1.70E−15	−9.20E−16
Wealth	1.70E−15	−1.90E−15
#Children ever born	0.03501	0.03501
Age at first marriage	0.06992	−0.00691
Husband's education	1.90E−15	4.00E−15
	2001	2011
All sample	4,067	3,047
Matched	2,709	1,316
Unmatched	1,358	1,731

Notes: (i) The $L1$ statistic gives overall imbalance of the covariates, defined as the difference between the multidimensional histogram of all covariates in the 2001 and 2011 waves. A perfect global balance is $L1 = 0$ and complete separation is $L1 = 1$. (ii) We checked the difference in empirical quantiles of the distributions of the 2001 and 2011 groups for each variable above. Except for *age at first marriage* variable at means, all other variables are balanced in all quantiles of the distributions.