Impacts of Land Rental Markets on Rural Poverty in Kenya

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Selected Paper prepared for presentation at the Agricultural & Applied Economics Association's 2011 AAEA & NAREA Joint Annual Meeting, Pittsburgh, Pennsylvania July 24-26, 2011.

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Abstract: This study uses panel data from 1,142 Kenya smallholder households over four survey periods to examine the determinants of participation in land rental markets and to quantify the impact of renting land on households' crop income and total income. We find that land rental markets in Kenya enhance productivity and are equitable. The results are consistent across different estimation methods and model specifications. Dynamic panel models were used to assess the impact of rental participation on households' crop income and total income. After controlling for the endogeneity of rental market participation and the persistent effects of lagged income, we find that the decision to rent land increased tenant households' net crop (net total) income by 25.1 (6.6) percent. These percentage gains are inversely related to household landholding size. Hence, land rental markets in Kenya appear to play an important role in raising incomes and reducing poverty for land-constrained smallholder farmers.

Land is one of the most important productive assets of rural residents in developing countries. How land is owned, used, and exchanged has far reaching implications for productivity, equity, and overall economic growth. While the impacts of land tenure security on land related investments, credit access, and agricultural productivity have been widely studied in the literature (Feder and Feeny 1991; Besley 1995; Alston *et al.* 1996; Brasselle *et al.* 2002; Jacoby *et al.* 2002; Galiani and Schargrodsky 2005; Do and Iyer 2008) there have been relatively few studies of the performance and impact of land rental markets (Pender and Fafchamps 2006).

Within the literature on land rental markets, previous studies have focused mainly on Asia (Holden, Otsuka and Place 2009). There is little evidence from Africa on the reasons why rural households would want to rent or lease land in the first place and the consequences of participating in land rental markets, especially for poor and/or highly land-constrained households with limited off-farm employment opportunities. These knowledge gaps need to be addressed as issues of land access and international investment in land become increasingly prominent on the policy agenda (Commission for Africa Report 2005; U.N. Millennium project 2005; the World Bank's World Development Report 2008; and the joint FAO, IFAD and IIED report on international investment in agricultural land in Africa, 2009). In the meantime, land markets in Africa have recently been found to be more widespread than previously realized, though the types of exchange, contractual arrangements, and the extent of land market activity vary considerably from country to country (Holden, Otsuka and Place 2009). The rising importance of land markets and the need for empirical evidence to guide African governments and development partners in formulating actionable land policies are the main motivation for this paper.

This study uses panel data from 1,142 smallholder households over four surveys covering a 10-year period to examine the determinants of participation in land rental markets in Kenya. In addition, we quantify the impact of renting land on households' net crop and total income. To our knowledge, this is the first study that quantifies the impact of land rental market participation on farmers' income using a relatively long panel. The paper also adds to our understanding of the processes by which land rental markets affect resource reallocations and agricultural productivity within smallholder farm sectors.

We first develop a simple model of household participation in land rental markets under the assumption that households are endowed with heterogeneous farming ability and land-labor ratios, and that renting and leasing land incurs transaction costs, building on Carter and Yao (2002), and Deininger *et al.* (2008). The hypotheses derived from the conceptual model are then tested empirically based on estimation of an ordered probit model. Both the descriptive and econometric results strongly support the hypothesis that rental markets transfer land from less efficient to more efficient producers and also improve access to land for households with relatively small farms. The results from ordered probit models are also consistent with those based on simple probit or tobit models, as well as those based on panel fixed effect linear probability models that control for unobserved heterogeneity. The findings from the dynamic panel income regression further suggest that the overall income gains to the smallest farms from increased access to land through rental markets are quite remarkable. Renting in land would lead to an increase in per capita total net income and per capita net crop income by 6.6 and 25.1 percent, respectively. However, these gains are not sufficient to pull significant proportion of rural households out of poverty.

Background and Conceptual Framework

A growing number of studies in recent years have assessed the functioning of land rental markets and their impact of land rental markets on productivity in Southeast Asia and East and Central European countries (Deininger *et al.* 2004; Swinnen and Vranken 2006; Ciaian and Swinnen 2006; Deininger *et al.* 2008; Jin and Deininger 2009). These studies have generally found that land rental markets enhance productivity and equity even in the presence of transaction costs. Sources of transaction costs include restrictive regulations on the leasing of land (Deininger *et al.* 2008), restrictive local policies over leasing (Deininger and Jin 2005), negotiation or disputes with the renter or land owner, or imperfections in other factor markets (Pender and Fafchamps 2006). In spite of this emerging evidence, there remain quite entrenched perceptions that the existence of land rental markets may lead to land concentration and increased poverty, and therefore that close government control over land rental is necessary. To our knowledge, there is no empirical evidence to date using household survey data to rigorously assess the impacts of participation in land rental markets on farm household incomes. Evidence of impacts over time through relatively long-term panel data can help inform and guide these policy debates.

Land access and land rental markets in Kenya and Africa

Because the majority of land farmed by African smallholders is under customary tenure systems, the purchase and sale of land is often prohibited (Bassett and Crummy 1993; Holden, Otsuka, and Place 2009). In such environments, land rental markets become an important mechanism for readjusting land-labor ratios among farm households, at least in theory. Other motivations for land leasing may be at play. For example, leasing land may be one of the few ways by which poor and land constrained households can generate quick cash in response to emergency needs; land leasing in such cases may not represent movements toward optimal land-labor ratios. Moreover, poor households may not be able to make advance fixed-rent payments especially under imperfect credit markets, which would constrain their ability to acquire needed land through rental markets. These concerns give rise to nagging doubts in about the long-term impacts of land commercialization on rural equity and productivity (Yamano et al. 2009).

The proportion of agricultural land under rent is considered to be lower in Africa than in most areas of Asia (Holden, Otsuka, and Place 2009). However, Kenya is unique in Africa in that it undertook a large-scale land registration and titling program beginning in the 1960s. According to the nationwide survey data to be described below, 58.4% of smallholder households in 2007 owned a land title deed for at least part of the land they controlled. Other factors constant, this might lead to less active land rental markets than elsewhere in Africa, because the transaction costs associated with the sale and purchase of land would presumably be relatively low. Cross-country African estimates of the prevalence of land leasing are spotty, but a review by Holden, Otsuka, and Place (2009) suggests that land rental markets appear to be most active in densely populated areas where land is highly fragmented, as in some areas of Burundi and Rwanda. Evidence from several districts of Kenya in the 1990s suggests that less than 10 percent of households rented in land (Wangila 1999) but more recent evidence from 15 districts in 2004 reports that 17.9 percent of households rented in land (Yamano et al. 2009). In our survey drawn from 22 districts, the proportion of households renting in land rose from 18 percent in 1997 to 20% in 2007, an expansion of roughly 1% per year. Fixed rental rates paid in cash are by far the most common form of informally arranged land rental contracts, followed by share cropping.[1]

The Government of Kenya's National Land Policy (2007) takes a decidedly positive stance toward land leasing, stating that it has "the potential to provide access to land to those who are productive but own little or no land" and that government policy is to "encourage the development of land rental markets while protecting the rights of smallholders by providing better information about transactions to enhance their bargaining power (Government of Kenya, 2007, paras 162 and 163). Given the explicitly promotional position of the Kenyan government toward land rental markets and the fact that an increasing proportion of farmers are participating in land rental markets over time, it seems important to better understand the productivity and equity effects of land rental markets within the smallholder farming sector.

Hypotheses on relationship between farming ability, land endowment and land transfer

To motivate our empirical analysis, we first build a simple model to explain a household's rental participation decision and derive the key testable hypotheses. Following Carter and Yao (2002), and Deininger et al. (2008), we assume that a rural household's decision to participate in land and labor markets is to maximize total household income by optimally reallocating its endowed productive resources (i.e., land, labor and productive assets) through land and labor markets according to its relative production ability in different sectors. We also follow previous studies that assume (1) farmers face no binding credit constraints associated with renting in land, and (2) land market imperfections create transaction costs that, when aggregated, vary proportionally with the area rented in or rented out. The first assumption is reasonable because one of the main advantages of rental markets over the sale markets is that amount of cash required to pay landowners up-front is much smaller when renting land as opposed to buying land (Deininger 2003). Credit constraints are greatly reduced when farmers make land rental payments after the crop is harvested.[2] The second assumption of variable transaction costs follows previous studies and is mainly for ease of modeling (Deininger et al. 2008; Kimura et al. 2010).[3] While certain kinds of transaction costs associated with land rental transactions are fixed (Bell and Sussangkarn 1988; Skoufias 1995; Pender and Fafchamps 2006) others vary with the amount of land rented, resulting in aggregated transaction costs that increase with plot size albeit with several discontinuous jumps. The main way in which fixed transaction costs would change our hypotheses is that there would be no rental transactions for very small plots of land because the gains from such transactions would be less than the fixed transaction costs involved. However, we note that in the survey data, 32 percent of the transactions by renters involved less than one acre of land, and 16 percent of them involved less than 0.5 acres. This might indicate fairly low fixed transaction costs. In the empirical model, we make the specification flexible enough to allow us to test for the validity of these assumptions.

Given these assumptions, households allocate their labor endowment between farming their own land (l_a) and off-farm employment (l_o) at an exogenous wage (w_i). Let T^{in} and T^{out} be the aggregated transaction costs associated with renting in and renting out respectively[4]. Household i's production decision can be modeled as (for notational brevity, subscript i is suppressed in the model):

$$\max_{l_a, l_o, A} pf(\alpha, l_a, A) + wl_o - I^{in}[(A - \overline{A})(r + T^{in})] + I^{out}[(\overline{A} - A)(r - T^{out})] \quad (1)^5$$

s.t. $l_a + l_o \leq \overline{L}$ (1a)

$l_a, l_o, A \geq 0$ (1b)

where p is the farm output price, r is the land rental rate, I^{in} is an indicator variable for renting in (I^{in} =1 if household i participated in renting in land; = 0 otherwise). Similarly, I^{out} is an indicator for renting out land (=1 if renting out land; =0 otherwise). A is household i's operational land size. Regarding the two constraint equations, (1a) means household's total labor use in farming and off-farm activities cannot exceed its total endowment, and (1b) does not allow negative labor use and operational land size.

The following propositions can be derived from this simple household model (see Deininger et al. (2008) for detailed derivation of the propositions):

Proposition 1. The operational land size is strictly increasing in farming ability, α. In other words, households with greater agricultural ability are more likely to rent in land; and households with lower agricultural ability are more likely to rent out land. Therefore land rental markets enhance efficiency by transferring land from less productive to more productive households regardless of the presence of transaction costs.

Proposition 2. The area rented in (out) is strictly decreasing (increasing) in their land endowment $\overline{A}$. Rental markets will thus transfer land from bigger farms to smaller farms. Therefore, land markets may also serve important equity or social goals.

We use estimates of producers' agricultural ability (α_i) to test for Proposition 1. Including households' land endowment in the determinants regression allows us to test for Proposition 2.

The introduction of transaction costs however allow us to closely link the conceptual framework to our specific empirical estimation strategy which is discussed next.

Estimation Strategy

Determinants of rental market participation

We estimate the determinants of rental market participation by an ordered probit model, a same approach that was used by Deininger, Jin and Nagarajan (2008) and Jin and Deininger (2009). These previous papers have provided detailed discussions on why ordered probit is a more appropriate approach. Specifically, the three rental participation regimes follow:

$$\left.\begin{array}{lll} \text{I. Rent-in regime } (\mathrm{A}^* < \overline{\mathrm{A}}): & & MP(\overline{A}) + \varepsilon_i \geq r + T^{in} \\ \text{II. Autarky regime } (\mathrm{A}^* = \overline{\mathrm{A}}): & & r - T^{out} < MP(\overline{A}) + \varepsilon_i < r + T^{in} \\ \text{III. Rent-out regime} (\mathrm{A}^* < \overline{\mathrm{A}}): & & MP(\overline{A}) + \varepsilon_i \leq r - T^{out} \end{array}\right\} \qquad (1)$$

where $A*$ is the optimal operational land size, $\overline{A}$ is land endowment, r is the market rental rate, T^{in} and T^{out} are transaction costs associated with renting in and renting out, respectively. $MP(\overline{A})$ + ε_i is the marginal value product of cultivating an extra unit of land evaluated at the level of autarkic land endowment, and $\varepsilon_i \sim$ N(0,1) representing part of the marginal value of product that the farmer observes but is not observed by the econometrician. Let $MP(\overline{A}) = \boldsymbol{X\beta}$ where **X** is a vector of household and village characteristics that determine the marginal product of cultivating an extra unit of land, $\boldsymbol{\beta}$ is a vector of parameters to be estimated, and let y be an ordered indicator for rental participation status taking on the values {0, 1, 2} respectively for {rent-in, autarky and rent-out}.

One difference of our strategy compared to that in the relevant literature (Deininger, Jin, and Nagarajan 2008; Jin and Deininger 2009) is that rather than specifying T as including variables specifying village-level differences in land policies or regulations, we account for T using village-level dummies. While this approach does not allow us to identify the effects of T explicitly, this is not our objective, and the use of village dummies arguably accounts for village-level differences in transaction costs more comprehensively than specific policy terms, which may not capture other spatial differences in transaction costs. Modeling transaction costs in this manner is likely to improve the accuracy of our estimates of the variables of interest, to the extent that they may be correlated with unobserved village-level differences in transaction costs in land rental markets.[6]

Let the upper and lower bounds be $C^{in} = r + T^{in}$, and $C^{out} = r - T^{out}$ respectively. The maximum likelihood function corresponding to equation (5) can be written as:

$$l = \sum_{y_i=0} \log\{1 - \Phi[(C^{\text{in}} - X\beta]\} + \sum_{yi=1} \log[\Phi(C^{in} - X\beta) - \Phi(C^{out} - X\beta)] + \sum_{y_i=2} \log[\Phi(C^{out} - X\beta)] \qquad (2)$$

Equation (2) is a well behaved maximum likelihood function and can be easily estimated (Wooldridge 2002).

The two key variables in ***X*** are the household's farming ability and land endowment, the coefficients of which will allow us to test for Propositions (1) and (2) derived from the conceptual model. We follow previous studies (Jin and Deininger 2009) to estimate the farming ability variable (α_i) based on estimation of household fixed-effect model of crop production. Specifically, α_i is obtained as the estimated fixed effect parameter in the panel production function.[7] Other variables that are expected to affect household's rental participation decision include household member composition, households' agricultural assets, the education of the household head, a dummy for female-headed households, distance to markets, village population, weather shocks, and a dummy for topology and soil type. Finally, agro-ecological zone and village dummies are included.

Impact of rental market on household's income

Finally, we specify a dynamic panel model to quantify the impact of rental participation on household income and poverty status. The dynamic model has the following desired features: (1) it accounts for the fact that household income and poverty status are likely to be persistent over time, i.e., the level of current income is likely to be affected by income in prior years; (2) it is possible for the endogenous variables such as rental participation to be instrumented by the same variable in prior years; and (3) the potential estimation bias caused by many omitted time-invariant variables can be eliminated through the first-differencing process. Specifically, the dynamic income regression equation can be defined as:

$$Y_{ijt} = C + \eta Y_{ijt-1} + \delta R_{ijt} + \gamma X_{ijt} + \tau V_j + \varphi D_t + \psi V_j * D_t + \lambda_i + \varepsilon_{ijt} \qquad (3)$$

where Y_{ijt} is alternately the household's per capita total income, per capita agricultural income, or poverty status, with the subscript i, j, t respectively denoting household, village, and time. Y_{ijt-1} is lagged income (to capture the persistent effect of past income), R_{ijt} is a dummy for renting in land, X_{ijt} is a vector of household and village characteristics that are expected to affect household i's income or poverty status; V_j and D_t are respectively village and time dummies; $\boldsymbol{\lambda_i}$ is the time invariant and unobserved fixed factor.

Estimating equation (3) using OLS will yield inconsistent estimates due to the fact that Y_{ijt-1}, R_{ijt} and λ_i are included on the right hand side of the equation. We can eliminate λ_i by transforming equation (3) from level form to first-difference form as:

$$\Delta Y_{ijt} = \eta \Delta Y_{ijt-1} + \delta \Delta R_{ijt} + \gamma \Delta X_{ijt} + \varphi \Delta D_t + \psi \, \Delta D_t * Vj + \Delta \varepsilon_{ijt} \qquad (4)$$

We also add village and time interaction terms ($\Delta D_t * V_j$) to control for the impact of localized shocks that could affect household income over time. OLS estimation of equation (4) will still yield inconsistent estimates because both ΔY_{ijt-1} and ΔR_{ijt} are endogenous, i.e., $E[\Delta Y_{ijt-1}\Delta\varepsilon_{ijt}]\neq 0$ and $E[\Delta R_{ijt}\Delta\varepsilon_{ijt}]\neq 0$. To obtain consistent estimates, we estimate equation (4) through an instrumental variable (IV) approach. Anderson and Hsiao (Anderson T.W. and Hsiao 1981) propose using Y_{ijt-2} to instrument ΔY_{ijt-1} in equation (4) under the assumption that the error term is not serially correlated. Y_{ijt-2} is a valid IV because it is expected that $E[\Delta Y_{ijt-1}Y_{ijt-2}]\neq 0$ and $E[\Delta Y_{ijt}Y_{ijt-2}]=0$ under the assumption of no serial correlation. Other right-hand-side endogenous variables can be similarly instrumented, e.g. by using $R_{ij,t-2}$ to instrument ΔR_{ijt}. The exogenous right-hand-side variables in first differencing form, $\Delta \boldsymbol{X}_{ijt}$, can be their own instruments. Additional lagged dependent variables (e.g. $Y_{ij,t-3}, Y_{ij,t-4}, \ldots$) can also be used as instruments for ΔY_{ijt-1} (Arellano and Bond 1991). This comes at the cost that the more distant lagged dependent variables (or other endogenous variables) and the change in dependent variable (or change in other endogenous variables) may be only weakly correlated or not correlated at all. Prior studies following this approach have adopted various instrumenting strategies using lags of different lengths as instruments depending partially on the number of available panel waves.

The relatively long time gap between two consecutive surveys (either 3 or 4 years) in our case is likely to make the correlation between the further lags for some of the endogenous variables and the endogenous variables themselves weak, which directly affects our choice of instrument variables. Specifically, we use Y_{ijt-2} and Y_{ijt-3} to instrument ΔY_{ijt-1}. And we use R_{ijt-1} rather than R_{ijt-2} to instrument ΔR_{ijt}. The choice of R_{ijt-1} rather than R_{ijt-2} is justified because land rental decisions are normally made at the beginning of the crop season, so R_{ijt} is more likely to be affected by Y_{ijt-1} (or ε_{ijt-1}) but not Y_{ijt} (or ε_{ijt}), which are not yet observed at the beginning of the crop season. In other words, we can reasonably assume $E[R_{ijt}\varepsilon_{ijt-1}]\neq 0$, but $E[R_{ijt}\varepsilon_{ijt}]=0$. In this case, R_{ijt-1} is qualified as a valid IV (Bond 2002). We use the Hansen's J statistics to check whether the instrument variables as a group are exogenous (Caselli *et al.* 1996). Other variables on the right hand side of equation (4) are either treated as pre-determined or exogenous.

Data and Descriptive Evidence

The data used for this paper come from four rounds of rural household surveys (1997, 2000, 2004 and 2007). The variables used in the analysis are defined in Table 1. These include socio-demographic household characteristics (such as the number of household members, adult equivalents, age and educational attainment of the household head, and whether the household incurred the death of an adult over the prior 3 years), indicators of income and asset wealth (such as landholding size, value of non-land assets per capita, total household per capita income, and the shares from crops and livestock, etc.), and variables describing the household participation in land rental markets.

Survey and sample design

The panel household survey was designed and implemented under the Tegemeo Agricultural Monitoring and Policy Analysis Project (TAMPA), implemented by Egerton University/Tegemeo Institute, with support from Michigan State University. The sampling frame for the panel was prepared in consultation with the Kenya National Bureau of Statistics (KNBS) in 1997. Twenty four (24) districts were purposively chosen to represent the broad range of agro-ecological zones (AEZs) and agricultural production systems in Kenya. Next, all non-urban divisions in the selected districts were assigned to one or more AEZs based on agronomic information from secondary data. Third, proportional to population across AEZs, divisions were selected from each AEZ. Fourth, within each division, villages and households in that order were randomly selected. A total of 1,578 households were selected in the 24 districts within eight agriculturally-oriented provinces of the country. Farms over 50 acres and two pastoral districts were excluded from the sample.

The initial survey was implemented in 1997. Subsequent panel surveys were conducted in 2000, 2004 and 2007. The initial number of observations in 1997 was 1,500. Our analysis focuses on the 1,142 panel households who were interviewed in all the 4 survey rounds and for which data were available on the complete set of variables used in the analysis. The average attrition rate between two consecutive rounds (about 3 years between each round) is about 5%, which is roughly comparable to attrition rates of most other household panel surveys in developing countries.[8]

Attrition bias is a potential problem in panel estimation. If sample attrition occurs randomly, then we do not need to worry about selection biases caused by attrition, although efficiency will be lost because of a reduced sample size. But if sample attrition occurs systematically, then attrition may create selection bias. We estimate reinterview models to assess the degree to which sample attrition is a problem and use the inverse probabilities of being reinterviewed as weights to control for attrition in the subsequent analyses. We follow Wooldridge's (2002, pp. 587-590) two-step estimation procedure to control for attrition. In the first step, probit models are used to estimate the probability that observation *i* remains in the next survey round and all subsequent survey rounds. Regressors include household and community characteristics and survey team dummies from the subsequent panel round (see appendix table 1). For t=2,…,T, let $\hat{\pi}_{it}$ be the fitted probability for household *i* to remain in year t. Then a set of probability weights $\hat{p}_{it}$ can be constructed as the product $\hat{p}_{it} = \hat{\pi}_{i2}\hat{\pi}_{i3}...\hat{\pi}_{it}$. In the second step, equations of main interest can be estimated using $\hat{p}_{it}$ as the weights for household i and year t.

Most of the household and community variables are statistically insignificant between households who were re-interviewed and those who dropped out in the sample for all the consecutive panel periods, suggesting that the attrition bias is likely small (appendix table 1). The variables that are significantly different include the interview team dummies, household land and non-land assets, and age composition of household members though they tend to vary from one panel period to another. The descriptive findings are further supported by the probit re-interview model (appendix table 1). Very few variables are statistically significant in the probit model for attrition in different periods. For example, for the 1997 and 2000 panel, compared to survey team 1 (base group), households that were interviewed by survey team 4 in the previous period are 5.7 percentage points less likely to remain in the next round of survey, households with more dependent members (those < 14 or >60) are more likely to be re-interviewed in the next round of survey. And for the 2000 and 2004 panel, households who were interviewed by survey team 3 in 2000 is more likely to remain in the panel (4.8 percentage points more likely compared to survey team 1). Households with more durable consumer goods and land in 2000 have a high tendency of remaining in the sample although the magnitude is very small. Finally, for the last available survey periods (2004-2007), land is the only variable that is significant with negligible magnitude of impact.

In light of the fact that very few variables are significantly different between panel households and households of attrition in both the descriptive and probit analyses, we expect the bias caused by attrition to be small. Consistent with our expectation, the results based on the two-step attrition correction estimation procedure are very close to those based on the standard estimates (i.e., ignoring potential attrition bias), and neither the signs nor significance levels of the main variables of interest change. Nevertheless, we report in this paper only the regression estimates corrected for possible attrition bias.

Household Characteristics and rental participation

Household characteristics and rental participation across regions are reported in table 1. We divide the total sample into 4 main zones according to their agro-ecological conditions and agricultural productivity potential – Eastern and Western Lowlands, Western Transitional and West Highlands, High Potential Maize zone and Central Highlands (see table 1) with the former two being of relatively lower potential and the latter two of relatively high potential.[9] On average, household size and adult equivalence have both declined considerably (by almost 1 member) during the past 10 years, with household size in Central Highlands Zone (4.2 adult equivalents) much smaller than the other zones (6.0 to 6.5). About 24% of households were headed by women, but this varied from 19% in the High Potential Maize Zone to 31% in the Eastern and Western Lowlands. Roughly 60% of household heads completed primary school, while a quarter of heads completed secondary school. Educational attainment was higher in the high potential zones than in the low potential zones. About 12% of households lost at least one adult member between 2004 and 2007; 5% of households suffered the death of their heads during the same period. Poor health and medical infrastructure, the prevalence of AIDs, and other diseases in rural Kenya are the main causes of adult mortality.

The differences in households' land endowment, assets, and income across zones are even more pronounced. On average, per capita land endowment is 0.81 acre ranging from 0.64 acre in Western Transitional and Western Highlands to 1.0 acre in Eastern and Western Lowlands. The survey data reveal remarkable differences in income and wealth across regions. Mean per capita household incomes (productive assets) in Central Highlands are more than double (triple) that in Western Transitional and Western Highlands.

About 20% of households rented in land and 12% of households rented out land in 2007. We notice that the share of households who rented out land is much smaller than the share of those who rented in land. This phenomenon is commonly found in other studies (Deininger and Jin 2008). One of the main explanations is that some households who rented out land are absentee landholders living outside the sample area. Moreover, an individual landlord may lease land to multiple renters. Rental market activity also varies considerably across regions with the Western Transitional, Western Highlands, and High Potential Maize Zones being much more active than the other zones both in terms of participation and size of land area (relative to land endowment) being transferred.

Households' decision to rent land varies inversely with farm size. Among the 20 percent of farm households owning the least amount of land (mean of 1.02 acres), 30.6 percent of them rented land. By contrast, 13.5 percent of households in the top landholding size quintile (owning a mean of 11.42 acres) rented land. The proportion of households renting-out land in 2007 ranged from zero among the bottom landsize quintile to 25.4 percent for households in the largest landsize quintile. The share of cultivated land also varied inversely with farm size, accounting for 53 percent of the land cropped by farm households renting land in the bottom landsize quintile.

Evidence on determinants of rental participation

To understand the extent to which the survey data support our hypotheses, table 2 reports household characteristics, land and productive assets and income composition for four rental participation groups: those who rented out land; those who remained autarkic; those who rented in land; and those who rented in only during 2007 but not in any of the previous periods. The descriptive evidence provides some bi-variate support for our hypothesis about the functioning of the land rental markets. First of all, rental markets increase access to land for households with relatively little owned land. The average per capita land endowment for households who rented in land is only 0.56 acres, which is only half of the landholding size for those who rented out land (1.07 acres). Household renting in land also tend to have bigger household size or adult equivalence than those renting out land. Second, there is some indication that land markets also tend to transfer land from less efficient producers to more efficient producers as evidenced by the fact that the derived farming ability coefficient is higher for the rent-in group than the rent-out group (0.01 versus -0.04) although the difference is not statistically significant. Moreover, land rental markets tend to transfer land from female-headed households to male-headed households, which is probably related to the fact that female headed households have higher land-labor ratios than male-headed households (1.49 vs. 1.25 acres per person).

It is interesting to note that the proportion of household heads with primary or secondary education is higher for those who rented in land than for those renting out land. This finding is somewhat in contrast to empirical findings elsewhere, which often find that those with higher education are more likely to lease their land while they devote their labor to higher return off-farm jobs (Deininger and Jin 2005). On the other hand, it may suggest that there are either limited off-farm opportunities or the off-farm jobs are mostly low skill jobs. The proportion of households incurring the loss of an adult member is not higher among those renting out land compared to autarkic households, suggesting that adult mortality does not in general trigger labor shortages given small average farm sizes.[10]

It has been long debated whether land markets would exacerbate the concentration of land among the rich and bigger landholders. We have already showed that the households with the smallest farms are more likely to rent in land. Our data also indicate that on average, per capita income and the value of productive assets of those renting in land are slightly higher but not statistically significantly different from other households. To control for the fact that some households who rented in land in 2007 may have already benefited from renting in land during the previous periods, we report the characteristics in column 4 of table 2 for those who rented in land in 2007 but did not rent in land during any of the previous periods. By comparing the initial per capita income and per capita value of assets between the new group (roughly 43% of those renting in land in 2007) and the other groups, we found those who only rented in land during 2007 appear to have the lowest level in both per capita income and per capita assets. However, the difference is not statistically different from the other groups, suggesting that land rental markets in rural Kenya do not appear to concentrate land to among the rich and bigger landholders. Gini coefficients of landholdings across the full sample drop from 0.55 to 0.53 after including rented land in the computation. When computed on the basis of per capita landholdings, the Gini coefficient declines from 0.60 to 0.57 after including rented land. Finally, neither the level of nor the change in the share of households below national poverty line is significantly different across rental participation status though there is some evidence that the relatively higher share of households who rented in land the first time in 2007 improve their poverty status than autarkic households (-0.08 vs. -0.046). While these descriptive results provide some bivariate support for our hypotheses, they do not control for endogeneity or the effects of other factors affecting incomes and land rental decisions. In the next section, we are interested in examining whether these findings are also borne out in the econometric analysis.

Econometric Results

This section reports estimation results for three sets of regressions, namely the Cobb-Douglas production function that is used to derive the farming ability variable, the ordered probit model on determinants of rental market participation, and the dynamic income model to quantify the impact of renting land on household income and poverty status.

The results are largely consistent with our expectations. The coefficients of the production function for the main inputs are all statistically significant at 10% or more, have the expected signs, and are of plausible magnitude. Results from the ordered probit model are consistent with our hypothesis that land rental markets enhance productivity and transfer land from bigger farms to smaller farms. These results are also consistent with those based on simple probit or fixed-effect panel linear probability models. Lastly, the results from the income regression further support our expectation that participation in land rental markets by small and poor land holders is associated with significant income gains. On the other hand, the results from poverty regression indicate that the gain in income is not sufficient to reduce poverty. Each of these findings is discussed in detail below.

Production function

Following the discussion in section 2.3.2, the panel fixed effect estimation (column 3) is used to recover the farming ability coefficient. For purposes of comparison, the results from pooled OLS estimation with village dummies (column 1) and those from random effects estimation (column 2) are also reported in Table 3. The crop production function R^2 ranges from 0.67 for the pooled OLS estimation to 0.78 for the household fixed effect model. The results from the random effect model are extremely close to the pooled OLS regression, as expected. The coefficients for all the main factors of production have the expected sign and are statistically significant at the 10% level or higher. Land is by far the most important contributor to crop production; doubling the size of total crop area leads to a 50% to 58% increase in total output. Compared to land, returns to adult equivalent labor endowment is only moderate. A doubling of adult equivalent labor endowment leads to an increase in output of only 5.3% to 9.0%. A doubling of expenditure on seed leads to a 12% to 17% increase in total crop production. The productivity of a female-headed household is 9%-10% lower than a male-headed household based on pooled OLS and random-effect models, other factors constant. By contrast, the household fixed effect model indicates no significant gender difference, most likely because the headship status of a given household does not vary much over time. Finally, it is interesting to note that rainfall has an important and highly significant effect on production in all models. A change from the 25th to 75th percentile of annual rainfall over the 1990-2008 period (roughly a 36% increase) is associated with a 13.5% to 15.3% increase in total production, indicating the sensitivity of rainfed agriculture to weather variations.

Rental market participation

As indicated earlier, the ordered probit model is appropriate for jointly estimating a farmer's decision to participate in one of the three rental regimes. To compare the sensitivity of our findings to alternative estimation procedures, we also report the results from probit and linear probability models based on the panel data.[11] The main results, however, are extremely consistent across different modeling strategies. The estimated results from the ordered probit model are reported in table 4. To help interpret coefficients from ordered probit model, table 5 reports the change in probability with respect to the change of a few key variables that are shown to be important.

The results from ordered probit model strongly support our two hypotheses. First, the positive and statistically significant coefficient on farming ability suggests that relatively productive farmers are more likely to rent in land and are less likely to rent out land. The probability that a households at the top 10% productivity ranking will rent in land is 3.9 percentage points higher than a household of average productivity. On the other hand, a household of average farming productivity is 2.5% more likely to rent out land or 1.4% more likely to remain in autarky than a household at the top 10% productivity ranking. The second main finding is that the positive coefficient on adult equivalents and negative coefficient on acres owned are both statistically significant at 1% across all the three model specifications. The findings are consistent with expectations, indicating that land rental markets transfer land from land-rich and labor-poor households to labor-rich and land-poor households. According to the marginal probabilities (rows 2 and 3 of table 5), doubling the size of the farm would lead to a 10.7% percentage point decline in the likelihood of renting land. By contrast, doubling the amount of adult-equivalent labor in the household is associated with a 9.6 percentage point increase in the probability of renting land. A doubling of landholding size is also associated with a 6.9 and 3.8 percentage point rise in the probability of renting out land or remaining autarkic.

Our results also show that after controlling for other effects, land rental markets transfer land to households with more agricultural assets, to those whose heads are male, and to those whose heads are in their early 50s, which may be at the most productive stage of the typical household life cycle. Doubling productive assets would increase an average household's tendency to rent in land by 5.5 percentage points. Female-headed households are 5.4 percentage points less likely to rent in land than male-headed households. While statistically insignificant, the negative coefficient on the dummy variable for household heads who completed primary school across all the specifications tends to suggest that households with more education are more likely to rent out land and take on non-farm job, which is consistent with other findings in the literature. Perhaps surprisingly, participation in land rental markets appears unaffected by the loss of the household head during the prior 4 years. Given the small farm sizes and relatively high labor/land ratios typifying most of the sample areas of Kenya, labor does not appear to be the main constraining input on crop production (Mather et al. 2005).

To partially check the validity of our assumption of variable transaction cost and no credit constraints in our conceptual model is, we include squared terms for landholding size and assets, and the lagged household income term in one of the equation (column 3, table 4). The insignificant coefficient on lagged household income suggests that the decision to rent land is not strongly related to household income prior to renting. The positive coefficient on the squared term and negative coefficient on the linear term of land endowment is inconsistent with the fixed transaction cost argument as the small farmers are more likely to rent in land (with a turning point of 5 acres). This is further supported by the insignificance of the squared term of asset value. The results on all other variables nonetheless are highly consistent with the other specifications.

Finally, results from a separate panel probit, a panel linear probability model and a tobit model are reported in Appendix Table 1.[12] The results from the panel probit and the panel fixed-effect linear probability model are highly consistent with those from the ordered probit model. Because farmer-specific ability is constant over time, we are not able to include this variable in the fixed-effect model, but the results on the remaining variables, especially on labor and land endowment, female head and past income are again similar to those of the pooled probit model. They basically reconfirm all the main findings from the ordered probit model that rental markets transfer land from farmers with relatively low agricultural ability and agricultural assets, less labor, and more land to farmers with greater ability, more labor, and less land. The results from the tobit model (which are only cross sectional because area rented in/out is not available for early rounds of the survey) are consistent with those obtained in the ordered probit model reported in table 4.

Impact of rental markets on household income

Table 6 reports the results of the dynamic income model for per capita total income (columns 1-3) and per capita agricultural income (columns 4-6). We estimate equation (4) for per capita total income and per capita agricultural income in three ways. First, we present the OLS results (columns 1 and 4), then the Instrumental Variable Generalized Method of Moments (IVGMM) results, treating as endogenous only the lagged dependent variable (per capita total or per capita agricultural income) and the dummy for renting land (columns 2 and 5). Finally, using IVGMM again, we treat the other variables as predetermined and instrumented by their own values at *t-1* and *t-2* (columns 3 and 6).

Before discussing results, it is worth noting that the Hanson-J over-identification test does not reject the null hypothesis of all instrumental variables being uncorrelated with the error term. The tests for no serial correlation suggests that original error-terms (in levels) are serially uncorrelated in most of the equations, indicating that the moment conditions are valid (Loayza et al. 2000).

As discussed earlier, the coefficients from an OLS model would be biased. If our model specification is correct, we would expect the coefficient of the lagged dependent variable from our IVGMM regressions to be smaller than estimates from a biased OLS model (Loayza et al. 2000). Our results confirm this expectation as the OLS coefficient on the lagged total income of 0.18 (0.19 for crop income) is much greater than the corresponding figures (0.07 and 0.09) based on IVGMM regressions. The positive and significant coefficient of the lagged dependent variables in all models indicates that there is strong persistence in total and crop income among rural Kenyan households, as is the case in much of sub-Saharan Africa. Second, the coefficients on household endowments and demographic characteristics in the IVGMM estimations are also consistent with our expectations. Doubling total productive assets would lead to an 8 to 9 percentage point increase in per capita income and a 7 to 9 percentage point increase in crop incomes, respectively. Similarly, a household moving from land autarky to owning 1.73 acre of land (the average amount of land being rented by those who rented in) is associated with a 6 to 11% increase in total income per capita and a 9 to 15% increase in crop income per capita. It is interesting to note that while the household head's education is not significantly associated with crop income (which seems to be consistent with the earlier results in the ordered probit land rental participation models showing an insignificant coefficient on education), education is positively and significantly associated with total income, suggesting higher returns to education in non-agricultural jobs. Finally, the highly significant and positive coefficient of our main variable of interest, the household decision to rent land, suggests that access to additional land through rental markets would significantly improve rural households' welfare. Renting in land would raise household's level of per capita total income and crop income by 14-16% and 31-33% respectively.

Based on the estimated coefficients, a simple simulation is conducted to determine how the decision to lease land affects net household income. We first identified all households renting land in 2007 and stratified them into five landholding size quintiles as presented in Table 7. Column 1 reports mean land endowment for different landholding quintiles. Column 2 presents the mean gross crop revenue generated from rented fields based on household-specific quantities of land rented in 2007. Column 3 shows the net crop revenue from renting land after deducting the costs of fertilizer, seed, land preparation, hired labor, family labor valued at hired wage rates, and the land rental rate. Comparison of the rental payment in relation to net crop revenue indicates that the renting households receive the lion's share of the net revenue produced on rented fields. Column 4 shows that tenant households generated 2.19 shillings of net crop revenue on rented land for every 1 shilling paid to the owner of the land. There is an inverse relationship between farm size of the tenant and the net crop revenue generated by tenants per shilling of rental payment. The 20% of smallholders with the smallest farms were able to produce 2.77 shillings in net revenue from rented land per shilling paid to the landlord, while the largest 20% of farms produced only 1.62 shillings in net revenue per shilling paid for rented land.

Columns 5 and 6 express the percentage change in net crop income and total household income resulting from the decision to rent land. Across the full sample, renting the mean amount of land contributed an average of 25.1 percent to renting households' crop income after deducting all production costs, and contributed 6.6 percent to total household income. However, among households with the smallest farms, i.e., the bottom landholding size quintile, their decision to rent land raised their net crop income by 41.6 percent and raised their total household income by 11.4 percent. This substantial increase in net crop income for the smallest farms is because the amount of land they own (0.70 acres on average) is a major limitation on their farm income. Land-constrained households' ability to rent land can double or triple the amount of land they cultivate and hence greatly improve their net crop income and ability to feed themselves. This is especially the case since the smallest farmers appear to generate roughly 2.7 times the net revenue per unit of land rented than the rental payment for that land. The impact on total household income, while positive, is less dramatic.

Conclusions

There is considerable controversy but a paucity of empirical evidence in Africa concerning the impact of land rental markets on the distribution of income within rural communities and on agricultural productivity. This paper examines the characteristics of smallholder farm households renting in and renting out land in Kenya, and the impacts of participation in land rental markets on farmers' income. Analysis is based on panel data on 1,142 households in 22 districts covering four waves over a 10-year period. To our knowledge, this is the first study that quantifies the impact of land rental market participation on farm incomes and poverty status using a relatively long panel. The paper also contributes to an understanding of the processes by which land rental markets affect resource reallocations and agricultural productivity within smallholder farm sectors.

The analysis highlights four main findings: First, rental markets contribute to agricultural productivity within the smallholder farming sector by transferring land from less efficient to more efficient households. In this way, land rental markets in Kenya appear to support national agricultural productivity objectives.

Secondly, we find little evidence to support the widespread concern that land markets may lead to land consolidation among the relatively rich and large landholders. In fact, land rental markets appear to promote the transfer of land from larger to smaller farms. Renting in land is inversely proportional to farm size, while renting out land is directly proportional to farm size. The Gini coefficient of landholding size per capita declines from 0.60 to 0.57 after accounting for the reallocation of land in the rental market.

Third, renters are able to generate roughly twice as much net crop revenue as the amount of rent payment to the owner of the land. The ratio of net revenue to the tenant : rental payment is inversely proportional to the farm size of the renter, being over 2.7:1 among smallholders with the smallest farms (prior to renting) and declining to 1.6:1 among the 20% of smallholders with the largest farms.

The fourth and potentially most important finding of the study is that participation in land rental markets is associated with a statistically significant contribution to farmers' crop and overall incomes. After the rental payment and other production costs are accounted for, leasing in land was found to increase households' crop and total income by an average of 25.1 and 6.5 percent respectively, compared to not renting. The percentage improvement in net crop and net total household incomes are highest for households with the smallest farms (41.6 and 11.4 percent) compared to 15.9 and 3.8 percent for households in the highest landholding size quintile. These findings would seem to endorse and encourage the Government of Kenya's current efforts to promote the development of land rental markets. Future research to identify the constraints on the functioning of land rental markets and understand how to better leverage the potential for land rental markets for the rural poor could contribute meaningfully to the design of future poverty reduction strategies.

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Table 1. Household characteristics by region

	1997	2007	2007			
	National	**National**	**Eastern & Western Lowlands**	**Western Transitional & West Highlands**	**High Potential Maize Zone**	**Central Highlands**
Demographic characteristics				Means		
Number of household members	6.66	5.86	6.04	6.27	6.47	4.22
Adult equivalents per hh	6.07	5.02	5.11	5.38	5.58	3.64
Household head's age	50.76	58.69	58.61	57.56	57.90	61.29
Female headed households (%)	0.13	0.24	0.31	0.23	0.19	0.20
Head with primary education (0/1)	0.54	0.60	0.56	0.59	0.61	0.64
Head with secondary education (0/1)	0.20	0.26	0.22	0.26	0.31	0.26
Death of adult since prior survey (0/1)	n.a.	0.12	0.15	0.09	0.12	0.09
Death of household head since prior survey (0/1)	0.00	0.05	0.07	0.04	0.05	0.03
Land, Assets and Income						
Acres owned	3.90	4.06	4.67	3.44	4.98	2.52
Acres owned per capita.	0.65	0.87	1.02	0.64	0.98	0.74
Acres cultivated	4.82	4.43	4.85	3.97	5.57	2.70
Value of assets per capita (Ksh)	31,432	53,228	50,703	18,445	61,327	85,875
Income per capita (Ksh)	42,723	40,065	30,495	25,253	47,546	61,189
Share of income from crops	0.41	0.44	0.38	0.51	0.38	0.55
Share of income from livestock	0.20	0.16	0.09	0.15	0.26	0.16
Share of crop production sold	0.42	0.45	0.27	0.48	0.52	0.60
Value of crop production per acre (Ksh)	35,787	22,387	10,881	19,382	20,670	45,811
Rental Participation						
% of households renting in	0.18	0.20	0.11	0.26	0.24	0.19
Acres rented in (by tenants)	n.a.	1.73	1.68	1.64	2.33	0.81
% of households renting out	n.a.	0.12	0.09	0.13	0.18	0.06
Acres rented out (by landlords)	n.a.	2.60	1.78	1.55	4.07	0.90
Number of observations	1229	1229	367	278	345	239

Authors' computation based on the data.
Note: Sampled districts within each zone are as follows. Eastern and Western Lowlands (Machakos, Mwingi, Makueni, Kitui, Siaya, Kisumu); Western Transitional and Western Highlands (Bungoma, Kisii, Vihiga; Kakamega); High Potential Maize Zone (Trans Zoia, Uasin Gishu, Bomet, Nakuru, Lugari, high-elevation Narok); Central Highlands (Nyeri, Muranga, Meru).

Table 2: Household Characteristics by Rental Participation Status, 2007

Variable	Rent-out	Autarkic	Rent-in	Rent-in, 1st time in 2007
Farming ability (estimated from model results in Table 3)	-0.04	0.00	0.01	
Household size	5.98	5.73	6.35***	6.31**
Members younger than 14	1.75	1.72	1.95*	1.89
Members between 14 and 60	4.42	4.61	4.93**	4.85*
Members older than 60	0.61**	0.76	0.50***	0.63*
Adult equivalence	5.12	4.89	5.50***	5.46**
Head’s age	56.61***	59.95	55.28***	58.41
Female head (%)	0.23	0.26	0.13***	0.18*
Head with primary educ.	0.64	0.57	0.68**	0.60
Head with secondary educ.	0.26	0.24	0.34***	0.33*
Loss of adult labor between 2004 and 2007	0.12	0.12	0.09*	0.08
Loss of head between 2004 and 2007	0.03	0.06	0.03*	0.03
Total area of land owned (Acre)	5.26***	4.10	3.17***	3.52*
Area owned p.c. (Acre)	1.07*	0.92	0.56***	0.62***
Value of assets p.c. (Ksh)	43,399	55,319	48,523	39,990
Income p.c. (Ksh)	36,445	40,480	40,818	41,372
Income p.c. in previous period (Ksh)	44,680	45,395	49,991	42,764
Change in income p.c. from prior survey (Ksh)	-8,235	-4,915	-9,173	-1,392
Value of total production (Ksh)	22,128	24,501	28,699***	26,272
No of Observation	146	842	240	102

Authors’ computation based on the data.
*, ** and *** denote significantly different from the mean of the autarkic group at 10%, 5% and 1%, respectively.

Table 3: Estimation of Cobb-Douglas Production Function (dependent variable is log of crop output in Kenyan Shillings)

	Village fixed effect	Random effect	Household fixed effect
Log of area cultivated	0.584***	0.570***	0.499***
	(18.92)	(17.33)	(12.16)
Log of adult equivalence	0.053*	0.064**	0.090***
	(1.97)	(2.24)	(2.52)
Log of agricultural assets	0.074***	0.075***	0.046***
	(7.07)	(6.98)	(3.58)
Log of total cost (except for seed)	0.030***	0.029***	0.014
	(4.02)	(3.60)	(1.60)
Log of total seed expenditure	0.170***	0.165***	0.119
	(8.17)	(7.70)	(5.42)***
Log of head's age	-0.475	-0.389	0.346
	(0.33)	(0.28)	(0.21)
Log of head's age squared	0.066	0.056	-0.051
	(0.37)	(0.32)	(0.23)
Head with primary education	0.021	0.020	-0.002
	(0.66)	(0.63)	(0.05)
Female headed household (=1)	-0.085***	-0.072**	-0.051
	(2.69)	(2.50)	(0.84)
Death of adult since prior survey (=1)	0.114	0.078	0.065
	(1.61)	(1.17)	(0.88)
Log of anual rainfall	0.419***	0.503***	0.516***
	(3.82)	(3.96)	(4.14)
Lagged log of annual rainfall	0.152**	0.158***	0.159***
	(2.59)	(2.66)	(2.83)
Distance to extension service	0.001	-0.000	0.000
	(0.34)	(0.12)	(0.07)
Distance to Motor Road	0.011	0.010	0.005
	(1.24)	(1.20)	(0.53)
Population Intensity	-0.075*	-0.115*	
	(1.67)	(1.88)	
Observations	4617	4617	4617
R-squared	0.67		0.32

* significant at 10%, ** significant at 5%; significant at 1%.
Standard errors adjusted for clustering effect at village level.
Village dummies were also included in the random effect model.
Inverse probability weighting is used to account for potential attrition bias (see methods section for description).

Table 4: Determinants of Participation in Land Rental Market, pooled data (Ordered Probit Model)

	(1)	(2)	(3)
Farming ability (estimated from model results in Table 3)	0.254**		0.194*
	(2.55)		(1.95)
Log of household income (lagged)			0.056
			(1.15)
Log of adult equivalents	0.405***	0.406***	0.338***
	(4.43)	(4.46)	(4.00)
Log of acres owned	-0.455***	-0.375***	-0.559***
	(7.04)	(5.18)	(6.36)
(Log of acres owned)^2			0.053*
			(1.74)
Log of rental rate	0.096	0.108	0.016
	(0.73)	(0.84)	(0.14)
Log of value of agricultural assets	0.233***	0.219***	0.475
	(5.06)	(4.75)	(1.49)
(Log of value of agri. assets)^2			-0.014
			(0.90)
Log of head's age	13.291***	13.410***	13.110***
	(3.93)	(4.00)	(4.12)
Log of head's age squared	-1.692***	-1.707***	-1.668***
	(3.95)	(4.01)	(4.13)
Head completed primary school (=1)	-0.134	-0.142	-0.099
	(1.30)	(1.38)	(1.00)
Female headed household (=1)	-0.244**	-0.237**	-0.221**
	(2.44)	(2.32)	(2.51)
Distance to extension services (kms)	0.009	0.006	0.012
	(0.86)	(0.56)	(1.35)
Death of household head since prior survey (=1)	0.090	0.086	0.125
	(0.41)	(0.40)	(0.63)
Distance to main road (kms)	0.017	0.017	0.040
	(0.42)	(0.43)	(1.19)
Population density (persons per km^2)	-0.032	-0.100	-0.138
	(0.06)	(0.19)	(1.26)
Log of rainfall during main growing season (mm)	-2.769	-2.518	1.143
	(1.35)	(1.25)	(1.12)
Log of rainfall during last main growing season (mm)	0.252	0.098	0.624
	(0.27)	(0.11)	(1.28)
Observations	1143	1144	1142

Robust z statistics in parentheses.
* significant at 10%; ** significant at 5%; *** significant at 1%.
All the standard errors are adjusted for clustering effect at village level.
Agro-ecological zone and village dummies are included in all the regressions.
Other controls include village soil types and topology features.
Inverse probability weighting is used to account for potential attrition bias (see methods section for description).

Table 5. Change in probability of falling into one of the three rental regimes with respect to changes in each of the key variables.

Variable	Change type	Rent-out	Autarkic	Rent-in
Adult equivalents	Elasticity	-0.062	-0.034	0.096
Acres owned	Elasticity	0.069	0.038	-0.107
Value of Productive assets	Elasticity	-0.035	-0.020	0.055
Farming ability	Min→Max	-0.125	-0.074	0.199
Female headed household	0 → 1	0.041	0.013	-0.054

The computation of probability changes are based on the estimated parameters in column 1 of Table 4.

Table 6: Effect of Renting Land on Household Per Capita Total Income and Crop Income (Dynamic Panel Regression)

	Per capita total income			Per capita crop income		
	OLS (1)	IVGMM (2)	IVGMM (3)	OLS (4)	IVGMM (5)	IVGMM (6)
Log of per capita total income	0.176***	0.065**	0.064**			
(lagged)	(7.97)	(2.03)	(2.05)			
Log of per capita agri. income				0.187***	0.093***	0.068**
(lagged)				(7.89)	(2.88)	(2.02)
Rented-in dummy (=1)	0.132***	0.159**	0.178**	0.248***	0.317***	0.333***
	(3.69)	(2.03)	(2.14)	(4.93)	(3.93)	(4.15)
Log of total agricultural assets	0.193***	0.080***	0.081***	0.122***	0.064***	0.073*
	(14.35)	(3.98)	(2.76)	(7.15)	(2.87)	(1.93)
Log of adult equivalents	-0.497***	-0.561***	-0.703***	-0.424***	-0.587***	-0.594***
	(14.29)	(10.04)	(5.80)	(9.38)	(10.39)	(3.92)
Land endowment p.c.	0.245***	0.205***	0.352***	0.389***	0.319***	0.494***
	(8.28)	(5.60)	(5.89)	(8.84)	(10.26)	(7.60)
Log of livestock value	0.058***	0.057***	0.056***	0.035***	0.028***	0.022***
	(12.82)	(10.18)	(8.53)	(8.76)	(5.49)	(2.84)
Female headed household (=1)	-0.174***	-0.073	-0.301	-0.099**	-0.150	-0.780
	(4.17)	(0.65)	(0.61)	(2.06)	(1.27)	(1.33)
Death of household head since	0.037	0.068	0.246	0.054	0.024	0.577
prior survey (=1)	(0.58)	(0.59)	(0.56)	(0.61)	(0.16)	(1.14)
Log of head's age	3.147**	2.583	5.240*	1.802	9.006***	12.836***
	(2.62)	(1.08)	(1.82)	(1.00)	(2.83)	(3.21)
Log of head's age squared	-0.418***	-0.291	-0.641*	-0.209	-1.141***	-1.654***
	(2.73)	(0.94)	(1.71)	(0.91)	(2.81)	(3.17)
Head completed primary school	0.140***	0.161***	0.137**	0.076*	0.015	-0.030
(=1)	(3.46)	(2.73)	(2.12)	(1.89)	(0.24)	(0.43)
Log of rainfall	0.031	0.122	0.182	0.935	1.351	1.600
	(0.04)	(0.12)	(0.17)	(1.50)	(1.08)	(1.23)
Log of rainfall (lagged)	0.584	0.358	0.718	-0.495	-0.009	0.236
	(0.81)	(0.46)	(0.86)	(0.76)	(0.01)	(0.18)
Hanson-J over-identification test (p-value)	-	0.14	0.21	-	0.83	0.73
Test for autocorrelation(p-value)[a]	-	0.13	0.13	-	0.53	0.53
R-squared	0.62			0.64		
Observations	3607	2381	2381	3513	2290	2290

Absolute value of t statistics in parentheses.

* significant at 10%; ** significant at 5%; *** significant at 1%.

All standard errors adjusted for clustering effect at village level.

Village dummies were added to control for local fixed factors that affect the change in income, time dummies were included to control for time trend of income growth, and interactions of village and time dummies are included to control for any policy change over time that may affect the income.

Columns (1) and (4): Pooled OLS regression with village dummies and year dummies included.

Columns (2) and (5): First differencing equation, ΔY_{ijt-1} and ΔR_{ijt} are treated as endogenous variable with the former being instrumented by Y_{ijt-2} and the latter being instrumented by R_{ijt-1} and R_{ijt-2}.

Columns (3) and (6): First differencing equation, ΔY_{ijt-1} and ΔR_{ijt} are again treated as endogenous and instrumented the same way as col. (2) and (5). Δ(log of value of agricultural assets)$_t$, Δ(land endowment)$_t$, Δ(log of livestock value)$_t$, Δ(female head)$_t$ are all treated as predetermined and they are instrumented by the level of themselves lagged by one period and lagged by two periods (i.e. x_{t-1} and x_{t-2}).

IPW approach was used to account for potential attrition bias in all the regressions.

[c] The autocorrelation of the original error term in the level equation is tested using Wooldridge's xtserial command in the Stata program. An alternative and more commonly used option is the test for the second order autocorrelation of the error term in the first differenced GMMIV estimation equation. This is not possible in a four time period panel.

Table 7: Simulation of impacts of renting land on tenant households' net crop income and total income by farm size category, based on 2007 data.

	(1)	(2)	(3)	(4)	(5)	(6)
		Revenue from renting in land			Impact of renting in land	
Landholding size group	Acres owned, tenant households	Gross crop revenue from rented fields	Net crop revenue from rented fields (col. 1 - all costs)	Ratio of net crop revenue to tenant : rental payment to landlord	Gain in net crop income compared to not renting	Gain in total household income compared to not renting
All tenant households	3.29	25,320	12470	2.19 : 1	25.13%	6.58%
Bottom 20%	0.70	21,612	12,311	2.77 : 1	41.58%	11.43%
2nd quintile	1.53	22,407	12,455	2.98 : 1	33.25%	10.41%
3rd quintile	2.25	21,554	10,931	2.29 : 1	29.37%	6.21%
4th quintile	3.38	25,498	11,754	1.76 : 1	23.24%	7.45%
Top 20%	8.60	35,491	14,835	1.62 : 1	15.90%	3.82%

Note: The simulation is based on the quantity of land rented and rental payments made in 2007 by each renting household. Gross crop revenues are derived from the coefficients on rent-in dummy in column (2) of table 6 for total income and column (5) of table 6 for crop income.

Appendix Table 1: Probit estimation of marginal probability of remaining in the next round of panel survey (Dependent variable: dummy for remaining in the panel)

	Panel period		
	1997-2000	2000-2004	2004-2007
adult equivalents in the initial period	-0.003	0.002	-0.014
	(1.32)	(0.77)	(1.13)
Rented in land in the initial period	-0.010	0.011	0.003
	(0.87)	(0.70)	(1.16)
Members < 14 in the initial period	0.016***	-0.002	0.009
	(4.42)	(0.65)	(0.23)
Members > 60 in the initial period	0.018**	0.008	0.003
	(2.32)	(0.75)	(0.49)
Female headed household in the initial period	-0.024*	-0.020	0.006
	(1.69)	(0.95)	(0.57)
Head with primary education	-0.016	-0.001	-0.002
	(1.61)	(0.05)	(0.17)
Head with secondary educ.	0.017	-0.014	-0.007
	(1.55)	(0.80)	(0.62)
Secondary educ. as the highest educ. of HH	0.002	0.025	0.012
members	(0.17)	(0.79)	(0.58)
Productive assets in the initial period	0.004***	0.002	-0.001
	(2.85)	(0.95)	(0.97)
Value of durable consumer goods in the prior period	-0.001	0.006**	0.001
	(0.78)	(2.44)	(0.33)
Share of income from livestock production	0.001	-0.001	0.012
	(0.05)	(0.77)	(1.38)
Share of income from off-farm activities	-0.009	-0.007	-0.010
	(0.63)	(0.50)	(0.73)
Land endowment in the intial period	-0.001	0.005**	0.003**
	(1.61)	(2.03)	(2.02)
Distance to tarmac road (kms)	0.001	-0.001	-0.001
	(1.28)	(0.88)	(1.55)
Distance from extension advice (kms)	-0.001	0.002*	0.000
	(0.72)	(1.85)	(0.27)
Survey team 2	-0.056	0.034	0.020
	(1.49)	(1.57)	(0.83)
Survey team 3	-0.032	0.048***	-0.050
	(1.44)	(3.29)	(1.47)
Survey team 4	-0.057**	0.026	0.006
	(2.49)	(1.19)	(0.41)
Predicted probability of remaining in the panel	0.95	0.93	0.96
Observations	1523	1439	1333

Robust z statistics in parentheses
* significant at 10%; ** significant at 5%; *** significant at 1%
Provincial and Agro-ecological zone dummies are included in all the regressions
Clustering effect at village level are controlled for in all the regressions

Appendix Table 2: Determinants of Renting in Land (panel evidence)

	Probit Model[a]		Linear Probability Model (FE Mode)		Tobit Model[b]	
					Area rented in	Area rented out
Farming ability (estimated from model results in Table 3)	0.052***	0.063***			0.774**	-0.210
	(3.52)	(3.53)			(2.51)	(0.23)
Log of adult equivalents	0.062***	0.078***	0.055***	0.051**	1.261***	-1.357*
	(4.39)	(4.79)	(3.08)	(2.29)	(3.36)	(1.80)
Log of acres owned	-0.054***	-0.083***	-0.018*	-0.038***	-1.084***	0.282**
	(6.77)	(8.32)	(1.86)	(3.27)	(5.82)	(2.01)
Log of income in previous period		0.009		-0.006		
		(1.01)		(0.68)		
Log of rental rate	0.035	0.025			0.364	1.053
	(1.55)	(0.98)			(0.91)	(0.95)
Log of agricultural assets	0.015***	0.024***	0.004	0.009	0.622***	-1.165**
	(2.76)	(3.85)	(0.61)	(0.94)	(3.92)	(2.53)
Log of head's age	1.288**	2.257***	-0.465	0.849	48.739***	-60.533
	(2.21)	(2.98)	(0.58)	(0.76)	(3.83)	(1.51)
Log of head's age squared	-0.185**	-0.304***	0.067	-0.100	-6.360***	7.469
	(2.47)	(3.15)	(0.65)	(0.69)	(3.90)	(1.46)
Head completed primary school (=1)	-0.012	-0.023	0.014	0.026	-0.408	0.963
	(0.86)	(1.31)	(0.69)	(0.91)	(1.15)	(0.83)
Female headed household (=1)	-0.069***	-0.079***	-0.051*	-0.073**	-1.421***	0.051
	(4.22)	(4.28)	(1.66)	(1.99)	(3.96)	(0.05)
Distance to extension service	-0.001	-0.001	-0.001	-0.001	0.034	-0.047
	(0.51)	(0.36)	(0.67)	(0.57)	(0.99)	(0.53)
Death of head during the prior survey (=1)	-0.020	-0.002	-0.003	0.018	0.181	-1.088
	(0.53)	(0.04)	(0.09)	(0.44)	(0.23)	(0.47)
Distance to main road	0.007	0.011*	0.006	0.011*	0.017	-0.513*
	(1.47)	(1.83)	(1.33)	(1.79)	(0.14)	(1.83)
Population density	-0.030	-0.054**			-0.513*	-0.147
	(1.02)	(2.07)			(1.76)	(0.14)
Log of rainfall	-0.000	0.031	-0.008	0.010	0.098	-13.461
	(0.00)	(0.73)	(0.28)	(0.31)	(0.04)	(1.47)
Log of rainfall last year	0.017	0.026	0.018	0.024	4.394**	4.293
	(0.57)	(0.76)	(0.80)	(0.90)	(2.25)	(0.85)
Observations	4077	2997	4206	3171	1143	1158

Robust z statistics in parentheses.
* significant at 10%; ** significant at 5%; *** significant at 1%.
[a] The coefficients in the probit models are marginal probabilities.
[b]Area rented in and area rented out are only available for the last panel period.

[1] This is consistent with Yamano et al. (2009) who found that almost all the land rental transactions in Kenya are on fixed-rental basis.
[2] Alternatively, farmers could also choose to use sharecropping arrangement to reduce the credit constraint problem.
[3] While certain kinds of transaction costs associated with land rental transactions are fixed (Bell and Sussangkarn 1988; Skoufias 1995; Pender and Fafchamps 2006) others vary with the amount of land rented, resulting in aggregated transaction costs that increase with plot size albeit with several discontinuous jumps. The main way in which fixed transaction costs would change our hypotheses is that there would be no rental transactions for very small plots of land because the gains from such transactions would be less than the fixed transaction costs involved.
[4] The transaction costs include the costs of obtaining information on market conditions, negotiating and enforcing payments, and the presence of tenure insecurity or regulations that restrict transferability.
[5] We can also adopt the standard household model where the household is maximizing utility subject to full income and other resource constraints. Here we follow (Carter and Yao 2002) by assuming there is no substitutability between leisure and labor. In that case, maximizing total consumption, which is a function of total income, is equivalent to maximizing total income.
[6] For example, households in villages close to markets are expected to have a smaller transaction costs. Transaction costs in land rental markets may also vary according to village-level population densities and land endowments.
[7] The estimated farming ability based on household panel regression is likely to be the lower bound for the following reasons: The first source of downward bias in α is the lack of information on land quality in the production function. If, as is commonly assumed, land quality on rented plots is lower than on owned plots (Benin et al. 2006), this would impart a downward bias on the estimated ability of tenants while leaving the estimated ability of landlords unaffected, implying that the estimated ability would be a lower bound for tenants. This problem is likely to be smaller in fixed effect model than an OLS model as land quality is unlikely to change much from one year to another for majority of farmers who do not participate in the rental markets. Second, estimates of the fixed effects, even though unbiased, are not consistent – (e.g., see Wooldridge 2002, pp. 272-274). We also estimated the rental determinants models without including farming ability, the results for all other variables are highly consistent. And the potential bias of the ability also does not affect the results of the main part of our analysis which is the impact of renting in land on households' incomes and poverty status.
[8] For example, the average attrition rate for a widely studied LSMS surveys in Vietnam is approximately 11% between initial survey in 1992-1993 and the follow-up survey in 1997-98 (Do and Iyer 2008). For the well known KwaZulu-Natal Income Dynamics Study (KIDS) in South Africa, only 84 percent of the 1,393 households in the original 1993 sample were successfully re-interviewed in 1998 (Maluccio 2000).
[9] Eastern and Western Lowlands include the districts of Kitui, Mwingi, Makueni, Machakos, Siaya, and Kisumu. Western Transitional and West Highlands include Bungoma, Vihiga, Kisii, and Kakamega districts. The High Potential Maize Zone includes Trans-Nzoia, Eldoret, Nakuru, Bomet, Lugari, and Narok districts, while the Central Highlands is composed of Meru, Muranga, and Nyeri districts.
[10] Other evidence from Kenya (Yamano and Jayne 2004; Mather et al. 2005) indicates that most households tend to attract new adult members after incurring the death of an adult member, partially offsetting the loss in labor. Moreover, rural households suffering from recent adult mortality did not have lower labor-land ratios than unafflicted households.
[11] One limitation with the ordered probit model is that only 2007 data are used because the data on rent-out were not collected in early periods. To check whether the results from 2007 are consistent with other periods, we also estimate a panel probit and a panel linear probability model on farmers' decision in leasing in land. We use the fixed effect estimator to estimate the panel linear probability model to account for unobserved heterogeneity.
[12] We are not able to do the same for those who rented out land because data for rent-out are not available for the early rounds of survey.