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Getting Institutions 'Right' for Whom: Credit Constraints and the Impact of Property Rights on the Quantity and Composition of Investment

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Getting Institutions 'Right' for Whom?
Credit Constraints and the Impact of Property Rights on the Quantity and Composition of Investment

Abstract

The effects of property rights on investment are typically hypothesized to occur through a security-induced investment demand and a collateral-based credit supply. Using a two period model, this paper shows that for farms that are constrained in their access to liquidity, the investment demand effect will itself induce an increase in the endogenous shadow price of liquidity. Other things equal, this induced increase in the price of liquidity will discourage capital accumulation, and that the desired stock of expropriation-immune movable capital may decrease with tenure security. Empirical analysis of farm-level data from Paraguay corroborates this proposition and reveals that the underlying pattern of wealth-biased capital access creates a world in which property rights reform has differential effects across producer wealth classes and gets institutions "right" and agriculture moving for only for a wealthier subset of producers.

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Section 1 Introduction

The proposition that legally secure and complete individual property rights over land boost investment and growth has been examined in historical, theoretical and empirical literatures. This proposition suggests that property rights are a key to unlocking potential economic growth in low-income and transitional economies. In addition, if legally insecure property rights weigh most heavily on low-income households, then public policy designed to enhance the security of individual property rights over land would seem to be a 'win-win' policy that promotes both economic growth and income equality.

However, this optimistic 'win-win' scenario is correct only if property rights reform suffices to relax the constraints that limit small farm investment and productivity growth. If not, then the results of property rights reform could be disappointing if reform suffices to get institutions 'right' for medium and large-scale producers, but not for small.

The effects of property rights on investment are typically hypothesized to occur through a security-induced investment demand effect (households increase investment when they perceive a reduction in the likelihood that land in which they might sink, attached, long-lived investment will lost); and, a collateral-based credit supply effect (lenders become more willing to make loans when assured that land pledged as collateral is secure and free of competing claims).[1] While the former effect should be at least as favorable for small scale as for large-scale producers, the latter effect will tend to favor larger-scale producers if there are intrinsic wealth-

1 In addition, it has been hypothesized that the right to transfer land to other producers relaxes a terms of trade or investment regret effect, especially in areas where off-farm opportunities are expanding quickly.

biases in rural credit markets as a body of theoretical and empirical work suggests.[2] Separate identification of investment demand and credit supply effects is thus needed to explore the hypothesis that property rights reform has differential effects across producer wealth classes and gets institutions “right” and agriculture moving only for a wealthier subset of producers.

To date the empirical analysis of property rights reform has had to settle for reduced form methods that cannot distinguish investment demand from the credit supply effects. In an effort to more fully identify the impacts of property rights reform and its possibly socially differentiated effects, this paper builds on work that examines economic behavior under credit constraints (*e.g.*, Feder *et al.*, 1990) and develops panel data econometric methods to separately identify the demand and supply effects. Separate identification of these two effects promises to help make more intelligible the erratic support for the property rights proposition in the empirical literature. This literature tends to find that the impact of property rights on investment appear weakest in areas where both formal credit markets are weak and customary tenure institutions are strong, leaving open the question as to whether muted reduced form effects reflect the absence of investment demand or credit supply effects. (*e.g.*, Bruce and Migot-Adholla 1994; and, Feder and Akihiko, 1996).

In addition, the methods developed here permit exploration of the sensible but often overlooked proposition that if credit supply effects fail to relax binding credit constraints, then the investment demand effects of property rights may induce a portfolio effect in which the proportion of investment sunk in (at-risk immovable capital) increases, even as the overall

2 Carter (1988) develops theoretical rationale for wealth biases in rural capital market. Empirical work by Bell, Srinivasan and Udry (1997), Kochar (1997) and Mushinski (1999) find evidence that such biases are indeed at work.

portfolio size or mass of investment remains constant. Put differently, in the absence of liquidity-expanding credit supply effects, the cost of increased fixed investment induced by property rights reform is likely to be reduced investment in other (movable) assets. When such countervailing portfolio effects take place, the income effects of property will of course be muted compared to what would be expected based on the increase in fixed investment alone. These observations also suggest that the qualitative studies that have asked farmers what new investments they would undertake if they were to receive tenure security (*e.g.*, Prosterman and xxx, 1998) overstate the effects that tenure security would have because they fail to ask farmers what they would do *less* of if they increased fixed investment.

The remainder of this paper is organized as follows. Section develops a simple conceptual model to analyze the demand, supply and portfolio effects of property rights reform. Its main objective is to motivate the empirical strategy that is outlined later in that section. Section 3 introduces the data and presents the results of the econometric analysis. The panel structure of this data makes it possible to identify the impact of land tenure regimes stripped of the influence of latent farm and farmer characteristics (*e.g.*, land quality and farming skill), which may be correlated with tenure regimes. The chief findings of this analysis are that provision of land title would have strong credit effects, and weaker, but positive, investment demand effects. However, these effects are most pronounced for larger farms (above 20 hectares in size), as land title by itself does not appear sufficient to improve formal credit access and relax credit constraints for small farm units. Put differently, these results suggest that land tenure reform is sufficient to get institutions "right" for only a subset of producers. Section 4 concludes and outlines policy implications.

Section 2 Modeling the Demand, Supply and Portfolio Effects of Property Rights Reform

This section develops a model of the impact of tenure insecurity and capital constraints on the overall level as well as on the composition of agricultural capital. We assume that there are two types of capital. The first, K_a, is attached capital that is lost in the event that the land to which it is attached is forfeited in a land dispute. The second, K_m, is movable capital that is not lost in the event that land is forfeited or taken away. We assume that in every production period, the agricultural household takes its capital stocks as given and allocates variable inputs (labor, *etc.*) to maximize its income. Let $p^*(K_a,K_m)$ be the short run, restricted optimum value function that corresponds to the short run production problem. To keep matters simple, we assume that capital lasts two production periods and the decision to invest in capital takes place prior to the initial production period in accordance with the following problem:

$$\begin{aligned}
&\underset{\substack{K_a, K_m \\ B}}{Max} \quad p_1^*(K_a, K_m) + E[p_2^*(K_a, K_m)] - (K_a + K_m) - rB + r(W - K_a - K_m + B) \\
&\qquad\qquad subject\ to \\
&\qquad\qquad\qquad K_a + K_m \le W + B \\
&\qquad\qquad\qquad B \le S(h) \\
&\qquad\qquad\qquad E[p_2^*] = \hat{x}(h)p_2^*(0, K_m) + (1-\hat{x}(h))p_2^*(K_m, K_a)
\end{aligned}$$

where $\hat{x}$ is the probability that land (and its attached capital) will be lost in a dispute. This probability is a decreasing function of h, an indicator of tenure security. The term W is the household's initial wealth. Wealth that is not used to purchase capital goods earns a rate of return r. The term B is the household's borrowing from the financial system that take place also at interest rate *r*. On the assumption that the financial system is quantity, not price rationed, household borrowing is constrained to be no more than its non-negative ration *S(h)*, where *S¢* $\geq$*0.*

The inequality constraints for this problem define two important regimes. The first is the unconstrained regime in which the financing constraint does not bind such that $W + B > K_a + K_m$. The first order conditions for this problem are:

(1a) $$k_a \equiv \partial p_1^* / \partial K_a + (1-\hat{x}) \partial p_2^*(K_a, K_m) / \partial K_a = (1+r)$$

(1b) $$k_m \equiv \partial p_1^* / \partial K_m + \hat{x}\, \partial p_2^*(0, K_m) / \partial K_m + (1-\hat{x}) \partial p_2^*(K_a, K_m) / \partial K_m = (1+r)$$

where are k_a and k_m are the total marginal products of capital. Denote the solutions to this unconstrained problem as:

(2a) $$K_a^u = K_a^u(r, h)$$

(2b) $$K_m^u = K_m^u(r, h)$$

Note these unconstrained demand functions for agricultural capital depend only on prices and technologies and do NOT depend on household wealth endowments or rations of liquidity (cite Feder and Lau).

In addition to this unconstrained, the problem above admits a constrained case in which the finance and borrowing constraints bind. In this case, $W + S = K_m + K_a$, permitting movable capital to be determined residually from the choice of attached capital and the liquidity constraint: $K_m = W + S - K_a$. In this case, the problem above reduces to a single choice variable and the first order conditions reduce to:

(3) $$k_a - k_m = 0 .$$

Subject to the available liquidity, investment is allocated to equalize the expected marginal returns to the two types of capital. Denote the solution values for this constrained problem as:

(4a) $$K_a^c(W, S, h)$$

(4b) $$K_m^c(W, S, h)$$

Note that these depend on wealth endowment, the credit ration, *S*, and the degree of tenure insecurity.

Figure 1 displays this equilibrium. The width of the horizontal axis is given by the available liquidity, $W + S$. K_a is measured from the left origin, and K_m is measured from the right. Any point along the axis exactly exhausts the available liquidity constraint. For a relatively low level of tenure security given by h_0, the optimum constrained capital portfolio is denoted by point *A* in the figure where expected rates of return are equalized between the two types of capital. An increase in tenure security (from h_0 to h_1) increases expected returns to attached capital such that k_a shifts as shown by the dashed line. While the increase in tenure security potentially has a secondary effect on k_m, we assume away this effect for purposes of the figure. Also for purposes of the figure, we assume that the increased tenure security has no liquidity (or credit supply) effect and that the width of the diagram does not change.[3] Under these assumptions, the new liquidity constrained equilibrium will shift to point *B* and be characterized by a higher level of attached capital but a lower level of movable capital. As discussed in the introduction, when credit supply effects do not relax a binding liquidity

3 Note that a credit supply effect that increased *S* would pull k_m to the right.

constraint, the increased investment induced by tenure security may come at the cost of reduced levels of other investment.

Figure 1 can also be used to illustrate the liquidity unconstrained equilibrium. Since the available liquidity no longer matters for this case, we will drop the interpretation of the width of the horizontal axis as indicating the total available liquidity.[4] Given a market price of capital of *r*, desired stocks of the two capitals will be as shown. As can be seen, capital stocks of these levels are not affordable for the household whose liquidity would be given by the sum $W+S$.

Finally, in preparation for the econometric work to follow, as Figure 1 shows, there exists an endogenous virtual or shadow price of liquidity, $\tilde{r}$, that just equates the constrained and unconstrained demands for agricultural capital:[5]

(5a) $$K_a^c(W, S, h) = K_a^u(\tilde{r}, h)$$

(5b) $$K_m^c(W, S, h) = K_m^u(\tilde{r}, h).$$

Also note that $\tilde{r}$ is the price that would just make the sum of the two unconstrained demands equal to the available (constrained) liquidity, $W+S$. In Figure 1, $\tilde{r}_0$ is the shadow price that equates demand and supply for liquidity in the household. More formally, note that we can implicitly define $\tilde{r}(W,S,h)$ by the following condition:

(6) $$K_a^u(\tilde{r}, h) + K_m^u(\tilde{r}, h) = W + S$$

and that This virtual shadow price interpretation thus lets us see more clearly the full effects of property rights reform on investment. When tenure security increases to h_1, the

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4 A better, but more space-intensive, graphical representation would be to pull the the right y-axis and the k_m curve so far to the right that the k_m falls below *r* prior to intersecting with the k_a curve.
5 This approach of defining a virtual or shadow price that equates internal supply and demand follows Singh, Squire and Strauss' analysis of the agricultural household in the absence of labor markets.

demand for capital shifts out as shown in Figure 1. However, without any compensating credit-supply-induced increase in liquidity, this shift in demand increases the endogenous, equilibrium shadow price of liquidity to $\tilde{r}_1$. The increase in attached capital is both lower than it would be if the shadow price of liquidity held constant at $\tilde{r}_0$. In addition, desired levels of movable capital actually decrease. In this case, tenure security shifts the portfolio mix of agricultural capital, but not the overall portfolio size. More generally, the total impact of tenure security of the desired liquidity constrained holding of capital stock of type *j* can be decomposed as:

(7)

$$\frac{dK_j^c}{dh} = \partial K_j^u \big/ \partial h + [\partial K_j^u \big/ \partial r][\partial \tilde{r} / \partial h] + [\partial K_j^u \big/ \partial r][\partial \tilde{r} / \partial S \, \partial S / \partial h].$$

Note that the first term on the right hand side of (7) is the direct, security-induced demand effect. The second term is the investment-depressing, endogenous shadow price of capital effect. The third term is the credit supply effect of land title. Note that for movable and other types of capital for which the direct, demand effect is small, the sign of (7) will depend the latter two terms. If credit supply effects are negligible for a credit constrained household ($\partial S / \partial h \approx 0$), then (7) would be expected to be negative, indicating that tenure security crowds out investment in movable capitals.

2.2 The Econometric Model and Estimation Strategy

This section derives a strategy for estimating the theoretical model of desired capital stocks using panel data. For household *i* in period *t*, consider the following linear approximation to the liquidity-unconstrained capital demand equations given in (2) above:[6]

(8) $$K_{it}^{u} = r_{it}y_{1} + X_{it}y_{2} + g_{i} + e_{it},$$

where y_1 and y_2 are vectors of structural parameters, and the row vector X_{it} denotes observable characteristics that affect the economic returns to capital (tenure security, farm size, and indicators of relative prices, technology and market access). The term g_i denotes latent, but time-invariant individual characteristics that influence desired capital stock; and, e_{it} is a random disturbance such that $E[e_{it}|r_{it},X_{it}]=0$, $E[e_{it}r_{it}]=0$ and $E[e_{it}X_{it}]=0$.

In most empirical situation (including that analyzed below), tenure status is not determined by an experimental or other exogenous process. Instead, individuals choose to invest in the legal and other procedures necessary to obtain secure tenure. The latent characteristics captured by g_i (*e.g.*, farming skill) may increase returns to investments in both tenure security and agricultural capital, suggesting that $E[g_i\,|r_{it},X_{it}]\neq 0$, $E[g_i r_{it}]\neq 0$ and $E[g_i X_{it}]\neq 0$. In this circumstance, it becomes vital to control for latent characteristics in order to avoid contaminating the estimates of the effect of tenure security with the effect of these latent characteristics.

The shadow price of liquidity for households in constrained regime that is implicitly defined by (6) can be linearly approximated by:

(9) $$\tilde{r}_{it} = X_{it}^{r}a + g_{i}^{r} + e_{it}^{r},$$

where the explanatory variables have again been partitioned into observable (X_{it}^{r}) and latent time invariant factors (g_{i}^{r}) and a is the unknown parameter vector. It is assumed that ε_{it}^{r} is

6 The parameters and variables in (8) should include an additional subscript indicating type of capital (attached or movable). To reduce notational clutter, we suppress those subscripts, though the empirical work below does not restrict parameters across types of capital.

orthogonal to X^r_{it} and therefore $E\left[\varepsilon^r_{it} \middle| X^r_{it}\right] = 0$ for all X^r_{it}. However, because g^r_i may not be orthogonal to X^r_{it}, we may have that $E\left[\gamma^r_i \middle| X^r_{it}\right]$ depends on X^r_{it} and therefore does not always equal zero. Taking advantage of the relationship given by (5), we can substitute (9) into (8) and derive a reduced form expression for liquidity constrained demand for capital:

$$K^c_{it} = [X^r_{it}a + g^r_i + e^r_{it}]y_1 + X_{it}y_2 + g_i + e_{it}, \tag{10a}$$

or collecting terms:

$$K^c_{it} = Z_{it}b + \tilde{g}_i + \tilde{e}_{it} \tag{10b}$$

where the vector Z_{it} is the concatenation of X^r_{it} and X_{it}; b is the vector of reduced form parameters; and, $\tilde{g}_i = g_i + y_1 g^r_i$ and $\tilde{e}_{it} = e_{it} + y_1 e^r_{it}$. Note that the reduced form parameter that relates tenure security to desired investment corresponds to the total derivative defined by (7) above.

Together, equations (8) and (10b) define an endogenous switching regression for desired capital stock:

$$K_{it} = \begin{cases} K^c_{it} = Z_{it}b + \tilde{g}_i + \tilde{e}_{it}\text{, } \textit{if liquidity constrained} \\ K^u_{it} = r_{it}y_1 + X_{it}y_2 + g_i + e_{it}\text{, } \textit{otherwise} \end{cases} . \tag{11}$$

Three difficulties confront the estimation of the parameters in (11):

1. *Unobserved Regime Switching* as liquidity constraints and regime switching are not directly observable.[7]

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7 The fundamental problem is that in the possible presence of non-price loan rationing, observed loan transactions, or lack thereof, cannot be used to impute liquidity constraint status. While Barham *et al.* (1996) discuss questionnaire strategies that may yield credibly observable indicators of liquidity constraint status, the data available for this study contain only the more routinely available variables that do not permit the construction of such indicators.

2. *Heterogeneity Bias* as each regression regime contains a latent, individual-specific, time invariant variable that is likely to be correlated with property rights and other variables; and,

3. *Selection Bias* as the (unobserved) regime switching process is economically endogenous based on the demand and supply of liquidity.

We will develop and implement a two-stage estimation strategy to deal with these problems. In the first stage, we will use estimated liquidity demand and supply parameters that determine the constraint regime to replace the unobserved regime-switching variable with a consistent estimate of its expected value. In the second stage, we estimate the parameters in (11) via two alternative methods. The first method uses a conventional fixed effects differencing technique to control for heterogeneity and selection biases. The second, more conservative, method adopts the regression trimming and weighting methods suggested by Honore, et al., 2000 and Kiriazidou, 1997 to eliminate any residual bias associated with the first difference estimator.

The latent regime-switching variable, d_{it}, takes on the value one when demand for credit (D_{it}) exceeds the supply (S_{it}):

(12) $$d_{it} = \begin{cases} 1 \; if \;\; D_{it}(V_{it}^{D}, \upsilon_i^{D} \mid \delta^{D}) - S_{it}(V_{it}^{S}, \upsilon_i^{S} \mid \delta^{S}) = V_{it}\delta + \upsilon_i + \mu_{it} \geq 0 \\ 0, \; otherwise \end{cases},$$

where demand and supply are functions of both observable (V_{it}^{D} and V_{it}^{S}) and latent, time-invariant effects (u_i^{d} and u_i^{s}); and, d^D and d^S are the parameter vectors. For notational convenience, we use V_{it}, u_i and d as shorthand for the full set of demand and supply variables and parameters, while m_{it} is the residual white noise regression error. Note that:

$$E(d_{it}) = \Pr(d_{it} = 1) = \Pr(m_{it} \geq V_{it}d + u_i) \equiv r_{it},$$

where r_{it} is the liquidity constraint probability. The appendix below summarizes related work that derives $\hat{r}_{it}$, the consistent estimator of r_{it} that we will use in the analysis here.

Assuming for the moment that the d_{it} and regime switching are observed, the switching regression (11) can be written as:

(11′) $$K_{it} = d_{it}[Z_{it}\beta + \tilde{\gamma}_i + \tilde{\varepsilon}_{it}] + (1 - d_{it})[r_{it}\psi_1 + X_{it}\psi_2 + \gamma_i + \varepsilon_{it}] .$$

As mentioned above, a key part of the identification strategy here is to control for the latent individual characteristics that would otherwise contaminate the estimation of the impact of the property rights variables. Applying a conventional fixed effects or first difference transformation to (11′) yields:

(13) $$K_i^* = [Z_i^* b + \tilde{g}_i\, d_i^* + \tilde{e}_i^*] + [r_i^* y_{\,1} + X_i^* y_{\,2} - d_i^* g_i + e_i^*]$$

where the *'s indicate time-differenced variables such that $K_i^* = K_{it} - K_{it-1}$;

$Z_i^* = d_{it}Z_{it} - d_{it-1}Z_{it-1}$; $d_i^* = d_{it} - d_{it-1}$; $X_i^* = (1 - d_{it})X_{it} - (1 - d_{it-1})X_{it-1}$,

$r_i^* = (1 - d_{it})r_{it} - (1 - d_{it-1})r_{it-1}$, $\tilde{\varepsilon}_i^* = d_{it}\tilde{\varepsilon}_{it} - d_{it-1}\tilde{\varepsilon}_{it-1}$, $\varepsilon_i^* = (1 - d_{it})\varepsilon_{it} - (1 - d_{it-1})\varepsilon_{it-1}$, *etc.* Noting from (10b) that $\tilde{g}_i = g_i + y_{\,1} g_i^{\,r}$, the time-invariant, latent variable terms can be combined and simplified as $d_i^* \tilde{g}_i - d_i^* g_i = d_i^* y_{\,1} g_i^{\,r}$, and (13) can be therefore rewritten as:

(13′) $$K_i^* = Z_i^* b + r_i^* y_{\,1} + X_i^* y_{\,2} + \left(d_i^* y_{\,1} g_i^{\,r} + \tilde{e}_i^* + e_i^*\right).$$

As inspection of (13′) reveals that this first difference approach sweeps away the time invariant components that directly affect desired investment, but does not eliminate the latent individual effects that affect the shadow price of liquidity (γ_i^r) and indirectly affect investment demand (unless $d_i^* = 0$). Put differently, unless all observations stay in the same regression regime both periods, the first differencing method commonly used to control for latent characteristics in linear models will not completely eliminate the time invariant latent terms in (11′), exposing ordinary least squares estimation of (13′) to a residual heterogeneity bias.

In addition to the problem of residual heterogeneity bias (which is caused by the presence of γ_i^r in the composite error term in 13′), OLS estimation of (13′) may be subject to selectivity bias caused by the correlation between the error terms in equations (12) and (13′). That is, if $E\left[(\tilde{g}_i+\tilde{e}_{it})(u_i+m_{it})\middle|X_{it},Z_{it},r_{it}\right]\neq 0$ and $E\left[(g_i+e_{it})(u_i+m_{it})\middle|X_{it},Z_{it},r_{it}\right]\neq 0$, we may have that:

(14a) $\tilde{\lambda}_{it}\equiv E[\tilde{\varepsilon}_{it}|d_{it},X_{it},Z_{it},r_{it}]\neq 0$;

(14b) $\lambda_{it}\equiv E[\varepsilon_{it}|d_{it},X_{it},Z_{it},r_{it}]\neq 0$; and,

(14c) $\omega_{it}\equiv E[\gamma_i|d_{it},X_{it},Z_{it},r_{it}]\neq 0$,

and therefore, the conditional expectation of the error term in (13') will not equal zero and, in addition, may depend on the conditioning variables $d_{it},X_{it},Z_{it},r_{it}$. That is, we may have that

(15)
$$\begin{gathered}E\left[\left(d_i^* y_1 g_i^r+\tilde{e}_i^*+e_i^*\right)\middle|r_i^*,X_i^*,Z_i^*,d_i^*\right]=\\ d_i^* y_1 w_i+d_{it}\tilde{l}_{it}-d_{it-1}\tilde{l}_{it-1}+(1-d_{it})l_{it}-(1-d_{it-1})l_{it-1}\neq 0\end{gathered}.$$

If condition (15) holds then estimation of (13’) via OLS will be biased and inconsistent.

However, if the correlation between the error terms in equations (12) and (13') comes from the time invariant error components $\tilde{g}_i, g_i$ and u_i—*i.e.*, if $E\left[\tilde{g}_i u_i|X_{it},Z_{it},r_{it}\right]\neq 0$ and $E\left[g_i u_i|X_{it},Z_{it},r_{it}\right]\neq 0$, but $E\left[m_{it}\tilde{e}_{it}|X_{it},Z_{it},r_{it}\right]=E\left[m_{it}e_{it}|X_{it},Z_{it},r_{it}\right]=0$—it is straightforward to show that $w_{it}\neq 0$, $\tilde{l}_{it}=l_{it}=0$. Thus, the OLS estimators of the parameters in (13’) will be biased due to the “omission” of the latent variable g_i^r (i.e., residual heterogeneity bias), but not due to the omission of the selection terms $\tilde{l}_{it}$ and l_{it}.

One approach to solve the problem of residual heterogeneity bias is to 'trim' the sample by eliminating all observations that change regime between time periods and for which $d_i^* \neq 0$ (see Honore, et al., 2000 and Kiriazidou, 1997). This approach not only solves the problem of residual heterogeneity bias (by sweeping the g_i^r's away from 13'), but it also relax the assumptions required for consistency under no selectivity bias. That is, under this approach what is required for consistency is that $\left(\tilde{l}_{it} - \tilde{l}_{it-1}\right) = \left(l_{it} - l_{it-1}\right) = 0$, which is a weaker condition than $\tilde{l}_{it} = l_{it} = 0$. Note that we can in general write the conditional expectations (14a-b) as a function of the linear index that predicts the shift between the liquidity constraint regimes:

$$\tilde{l} \equiv \tilde{L}(V_{it}d + u_i) \text{ and } l_{it} \equiv L(V_{it}d + u_i)$$

where the arguments of the functions are defined in (12) above. For those observations for which the linear index $V_{it}d + u_i$ does not change very much between the two periods, $\left(\tilde{l}_{it} - \tilde{l}_{it-1}\right)$ and $\left(l_{it} - l_{it-1}\right)$ in (15) will be small. These terms will also be small when there is little change in any monotonic transformation of the linear index.

For purposes here, a convenient transformation of this index is r_{it}, the liquidity constraint probability introduced earlier. Rewriting the selection terms above as $\tilde{l}_{it} \equiv \tilde{L}(r_{it})$ and $l_{it} \equiv \Lambda(r_{it})$, we have that $\left(\tilde{l}_{it} - \tilde{l}_{it-1}\right) = \left(l_{it} - l_{it-1}\right) \approx 0$ when $r_{it} \approx r_{it-1}$. Noting that the selection bias for this class of problem disappears as the linear index governing regime switching does not change over time, Kyriazidou (1998) suggests weighting the data in inverse proportion to the distance between the two selection indices for the two periods.

Combining these ideas we arrive at the two-stage estimation strategy to be used in this paper. We first estimate (12) and construct estimates of the liquidity rationing probability, $\hat{r}_{it}$.

Then we substitute these estimated $\hat{r}_{it}$'s for the d_{it}'s into (13'). In the analysis to follow, we will offer two alternative second stage estimators of the parameters in (13′):

1. *OLS First Difference Estimation*, in which (13′) is simply estimated using ordinary least squares. This estimator permits us to use all observations, but is subject to residual heterogeneity bias (as well as to selection bias if there are time-varying latent characteristics which enhance liquidity access and are correlated with property rights status and other explanatory variables).

2. *Trimmed Sample Estimation* in which we eliminate observations that are likely to have changed regression regimes over time; This is equivalent to a weighting scheme w_i given by:

   (16) $$w_i = \begin{cases} 1 \; if \; \hat{r}_{it} > 80\% \forall t, \; or \; \hat{r}_{it} > 20\% \forall t; \; AND, \left| \hat{r}_{it} - \hat{r}_{it-1} \right| < 20\% \\ 0, \; otherwise \end{cases}.$$

   Using these weights, only those observations that are very likely to have been in the same regime both periods are kept for the analysis.

We turn now to the implementation of this estimation strategy.

Section 3 How and For Whom Does Tenure Security Work in Paraguay?

In order to gauge how and for whom tenure security works, this section uses a farm-level panel data set collected in Paraguay. Paraguay remains one of Latin America's most highly agricultural economies, as well as one of its poorest. Similar to most of Latin America, the land distribution is highly dualistic, with numerous tiny farms co-existing with large production units that control most of the agricultural area. Carter and Galeano (1995) and Carter and Salgado (2000) give more detailed information on agricultural and land issues in Paraguay. After first presenting providing a descriptive overview of the available panel data, this section goes on to estimate the switching regression model of desired capital stock.

3.1 Property Rights, Capital and Credit in Paraguay

The panel data available for this study emerged from a stratified, multi-stage random sample of 300 producer households distributed across three distinctive regions of rural Paraguay: The traditional core “minifundia” zone of Paraguarí; the colonization zone of San Pedro; and, the department of Itapúa, located in the frontier region with Brazil and where there has been significant agro-export growth. These 300 producer households were interviewed in 1991 and 248 were successfully reinterviewed in 1994.[8] Both interviews collected full production and income information as well as a detailed accounting of the modes of land access, property rights and land transactions. Larger farms with more than twenty-five hectares of land reflects were purposefully oversampled relative to their weight in the population of farm units as these relatively few large farm units control a large portion of land resources. Adequate representation of this group was a prerequisite for the analysis of the hypothesis that credit access and the impact of property rights reform are wealth-differentiated.

Table 1 presents some basic indicators drawn from the available data. The columns of the table report mean values for farms in three tenure security or property rights categories. The rows divide observations based on formal credit status. It should of course not be assumed that farms without formal loans are necessarily credit rationed. Nonetheless, splitting the data in this way permits us to get a sense for the data and the primary hypotheses to be investigated.

8 Whenever possible, households that exited farming between 1991 and 1994 were replaced by successor units which were found in 1994 to be cultivating at least some portion of the land resources of the exiting household. Neither these successor units nor the sample causalities are included in the analysis that follows as it depends on multiple observations on each household.

A farm is designated as 'Titled' if at least some of its land is held with legally registered, mortgagable property rights. A farm is assigned to the 'Formal' category if it did not qualify as titled but owns some land emanating from colonization projects in the 1960s and the 1970s that assigned colonist legally secure, but inalienable and unmarketable property rights. Once an individual pays off his or her colonization debts, land held under formal tenure is titled and becomes fully marketable (and mortgagable). In the analysis to follow, land in the formal category will be treated as having equal tenure security as titled land but as having potentially different collateral value. However, over the study period Paraguay had no general program of land titling, and those with titled land had to go to some trouble and expense to obtain and maintain their titles. As mentioned in the introduction, this economic endogeneity of title is one of the challenges that this study must confront.

The 'Other' tenure category in Table 1 reflects the fact that rural Paraguay has been typified by a wide variety of tenure regimes, including significant informal squatting that developed given the country's historically long period of extreme land abundance. Land that is accessed under these more precarious regimes amounts to a significant share of land for farms in the smallest size strata. Less than 60% of the small farms sampled in 1991 either had, or were in the process of obtaining legal title to their land. This not atypical pattern in which legally unclear or insecure land access predominates in the small farm strata underlies the perspective that property rights reform policies should differentially advantage the less well-off even as it promotes aggregate growth.

Observations in the northwest corner of Table 1 have legal title to their land and were observed to have formal production loans. These farms have both high levels of attached capital[9] per-hectare compared to other farms, and much higher loan levels. The key question to be investigated is whether or not these relatively high levels of credit and attached capital are the result of these farm's legal status *per se*. For example, note also that these farms are by far and away the largest in the sample, raising the question as to whether their favorable capital and credit indicators are a result of their tenure status, or their size in markets that scale sensitive. It could of course also be that this descriptive association between title status, credit and attached capital is a spurious reflection of the fact that all three variables are caused by a fourth factor (*e.g.*, entrepreneurial zeal). In this case, the provision of title to randomly selected (not necessarily zealous) farmers would do nothing for credit access and agricultural productivity. Panel data methods, which can be used to control for latent entrepreneurial zeal and other characteristics that may differ between individuals in the different tenure, are a key part of the econometric identification strategy developed in section 3 below.

The panel data methods used here rely on the changes in household characteristics and behaviors over time to identify the impact of property rights on investment and credit access. While the sacrifice of between household variation to control for time-invariant, household specific effects is likely to be statistically expensive, the data do exhibit intra-household variation over time in the key property rights variables. As can be gleaned from Table 1, the number of farms in the title category rises from 112 to 132 from 1991 to 1994. In addition, some number of titled farms increased the amount of their farm area held under legally secure status.

9 Attached capital includes buildings, fences, land improvements and irrigation infrastructure.

3.2 Econometric Results

Table 2 presents first difference estimates for the switching regression functions for attached and movable capital given by equation (13). Following the regression strategy outlined above, results are presented for the full sample as well as for a trimmed sample that excludes observations that likely to have changed credit constraint status between 1991 and 1994. As just discussed, there are two types of legally secure land in rural Paraguay. The two types should offer similar security to the owner and hence have identical impacts on investment demand. Their credit supply effects should, however, be different. For households in the liquidity-unconstrained regimes, the two types of land should have identical effects and identify the investment demand effect of tenure security. For households in the liquidity-constrained regimes, the effects of these two types of land should be different (see expression (7) above). Initially, all liquidity-constrained regime equations were estimated without imposing any restrictions on the coefficients of titled and formal land. In three of the cases, however, the point estimates were very similar and it was impossible to reject the hypothesis that the coefficients of titled and formal land were the same.[10] Table 2 reports only the restricted estimates (shown in bold) for these cases.

The overall character of the results is quite striking using both the full and trimmed samples. Land tenure security appears to increase attached capital for both liquidity-constrained and liquidity-unconstrained farms. For unconstrained households, the shift of one hectare of land from the insecure to the secure category (holding total farm size fixed) is estimated to

10 Note that these coefficients would be expected to be similar for households for whom title has low marginal credit leverage.

increase attached capital by $134 to $187. For constrained households, the point estimates of this same effect range from $73 to $254.[11] As presented in section 3, the equation for unconstrained households is a structural, capital demand equation and the coefficient on secure land identifies the pure investment demand effect of tenure security. The equation for the constrained regime is a reduced form and the estimated tenure coefficients identify the both demand and marginal liquidity effects (see expression 7).

In contrast to these uniformly positive tenure security effects for attached capital, we see that for liquidity-constrained households, tenure security has a significant negative effect on movable capital (-$110 to –$184), whereas the effect for liquidity-unconstrained households is small and statistically insignificant.

As might be expected, the results using the trimmed regression soften up as nearly half of the observations are tossed out. Interestingly, these more conservative estimates show a strong and significant coefficient of secure land on attached capital for the unconstrained regime and a significantly negative impact of secure land on movable investment in the constrained regime. In the trimmed sample, the coefficient on secure land for the constrained regime remains positive, but has shrunk and lost statistical significance. Nonetheless, the basic portfolio effect story remains in the trimmed sample results.

As predicted by the theoretical model, the positive effects of tenure security on investment are dampened by an unfavorable liquidity constraint effect for credit-constrained

11 With the exception of the attached capital coefficient on secure land for unconstrained households, these primary coefficients are all statistically significant using the full sample. The coefficient on secure land for unconstrained equation for attached capital is only marginally significant with a p-value of 13%.

households that lead to reduced stocks of movable capital. But for which type of households do these dampening effects occur? To gain insight on this problem, the estimated coefficients from the full sample were used to calculate the following expected tenure security investment effects for farms of different sizes:

(15) $$E[D_j(x_{it})] = \left\lfloor r_{h=1} E(K^c_{jit}) + (1 - r_{h=1}) E(K^u_{jit}) \right\rfloor + \left\lfloor r_{h=0} E(K^c_{jit}) + (1 - r_{h=0}) E(K^u_{jit}) \right\rfloor$$

where $D_j(x_{it})$ is the expected change in the stock of capital of type j as a farm with characteristics given by x_{it} would experience if it moved from having all of its land under insecure tenure ($h=0$) to having all its land titled ($h=1$). Notation indicating the conditioning of the credit rationing probabilities and the capital stock functions on x_{it} has been suppressed.

Figure 2 graphs (15) over a range of farm sizes holding the other variables at their median values. The solid line shows the estimated positive impact of tenure security on attached capital for all farms. The dashed line shows the estimated impact of tenure security on total capital (defined as the sum of attached plus movable). As can be seen, despite the consistently positive effect of tenure security on attached capital, total capital does not increase until a farm size of approximately 15 hectares. For farms below that size, tenure security has a portfolio effect (increasing the ratio of attached to total capital), but movable capital decreases all most in proportion to the increase in attached capital over this range.

A large part of the reason for this result can be seen in the dotted line in Figure 2 that shows the expected credit supply effect of tenure security as a function of farm size. More formally, the credit markets estimates taken from Carter and Olinto (see the appendix below for details) are used to estimate the following expression:

(16) $$D_r(x_{it}) = r_{h=1}(x_{it}) - r_{h=1}(x_{it}),$$

where $D_r(x_{it})$ is the change in the credit rationing probability for a farm with characteristics x_{it}. Careful examination of Figure 2 shows that the change in the estimated impact of land title on the credit rationing probability is zero for farms of less than about 3 hectares in size. For these farms, the estimated credit rationing probability is approximately 100%. Beyond 3 hectares, the acquisition of land title begins to reduce that rationing probability, but it is not until a titled farm has in excess of 15 hectares that its rationing probability ($r_{h=1}$) falls below 50%. As can be seen in Figure 2, it is at this farm size that title begins to facilitate an increase in the total stock of capital and not just a shift in its composition. In terms of expression (15), it is at this farm size that titled farms are likely to be in the liquidity-unconstrained regime.

Section 4 Summary and Policy Implications

While the literature on land tenure security and investment has often discussed the potential investment demand and credit supply effects of land titles or other provisions to increase tenure insecurity, this paper has put forward a simple model to show that for households that are constrained in their access to liquidity, the investment demand effect will itself induce an increase in the endogenous shadow price of liquidity. Other things equal, this induced increase in the price of liquidity will discourage capital accumulation. For movable and other types of capital that are relatively immune from expropriation in the event of land loss (and hence not directly influenced by investment demand effects), the net effect of an increase in tenure security could well be a decrease in the desired stocks of these capital goods. From the perspective of this model, exactly who will benefit from tenure security will depend centrally on the interactions between investment demand and credit supply effects.

While the underlying theoretical model is relatively simple, consistently estimating the endogenous switching regression that it implies confronts a number of difficulties. Taking advantage of the available panel data from Paraguay, this paper has used panel data procedures to control for time invariant characteristics that are likely to be correlated with both tenure status and with the endogenous selection into credit constraint regimes. In addition, credit constraint status is not directly observable. Instead it is estimated using the econometric strategy outlined in a companion paper that relies on both unobserved sample separation methods and ancillary sample information. Finally, an analogue to the trimmed regression procedures developed by Honore *et al.* (2000) and Kiriazidou (1997) are used in an effort to control for any biases that the fixed effect procedure may not eliminate when farms switch constraint regimes over time.

Emerging from this estimation procedure are three key results:

1. Tenure security has a strong effect on the demand for attached capital;

2. The credit supply effects of tenure security are non-existent for the smallest farms and only become large for farms in excess of 15 hectares.

3. As might be suspected based on points 1 and 2, tenure security induces a shift in the portfolio composition of capital for smaller farms (toward more attached capital), but only for larger farms is it estimated to enable an unambiguous increase in total investment.

For a country like Paraguay, where agricultural land is often held under a variety of legally tenuous arrangements and the small farm sector which remains home to a majority of the country's population, these results hold several nuanced implications. First, it is clear that provision of tenure security does not get institutions right for all farmers. In particular, a generalized policy of land titling would be expected to disproportionately benefit larger scale producers who experience both investment demand and credit supply effects and whose mass and composition of capital would be positively affected by such a policy. In this context it is important to note that in eastern Paraguay (the area exclusive of the large, arid and thinly populated Chaco region), average farm size is no more than 5 hectares. The finding that a farm must be of some 15 hectares to fully benefit from tenure reform implies that most households—especially low-income rural households—would at best experience only a muted set of benefits from tenure reform. Indeed, to the extent that the differential advantage created by tenure reform encourages land accumulation by larger farmers, the secondary or spillover effects could be negative for these poorer households.

Finally, it should be stressed that evidence that land titling would have socially skewed effects is not an argument against land titling. Instead it suggests that policy—if aimed at achieving broadly based agrarian growth—needs to be carefully sequenced, with prior, or at least simultaneous, attention given to credit market reform. Discussion of the options and institutions available to address financial market bias is beyond the scope of this paper. However, the analysis here strongly suggests that attention needs to be given to these issues as an intrinsic part of land tenure reform lest a set of policies be designed that get institutions right only for a small, and already privileged subset of producers.

Figure 1
Portfolio Effect of Property Rights Reform for Credit-Constrained Producer

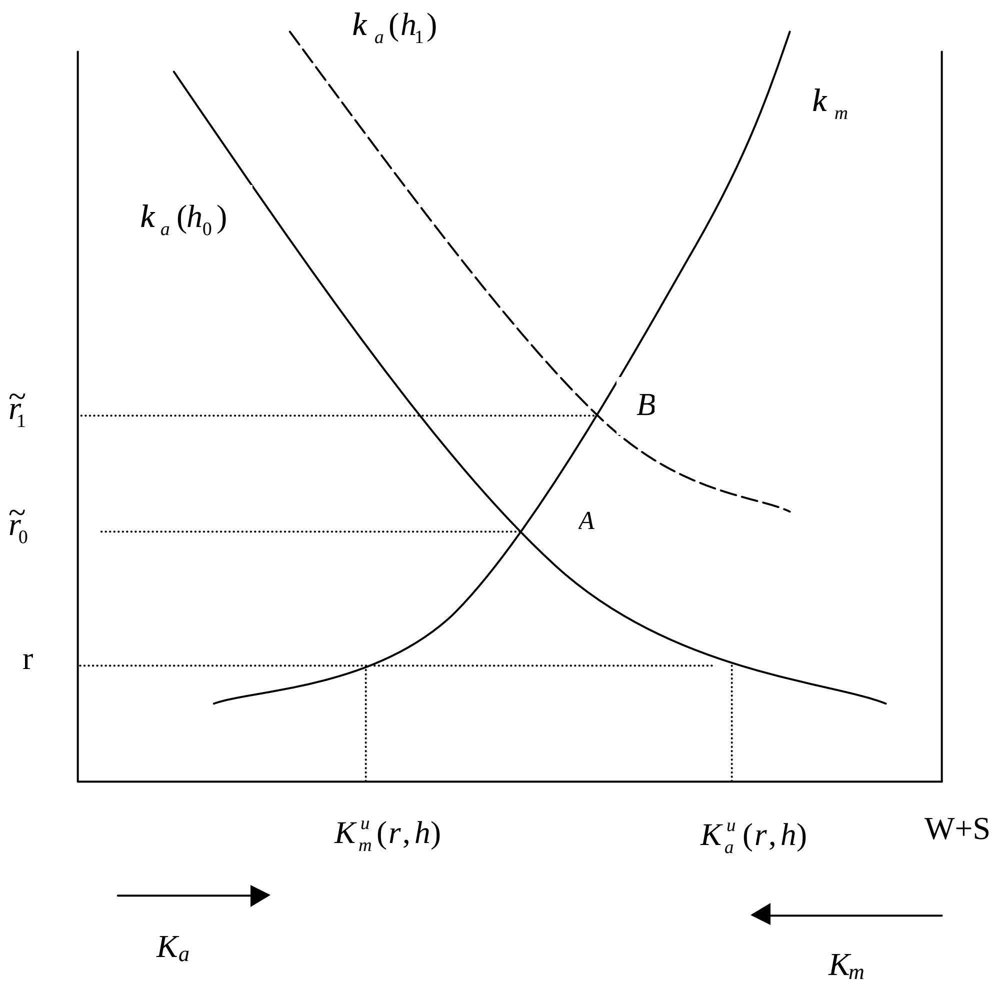

Figure 2:
Heterogenous Impact of Property Rights on Investment

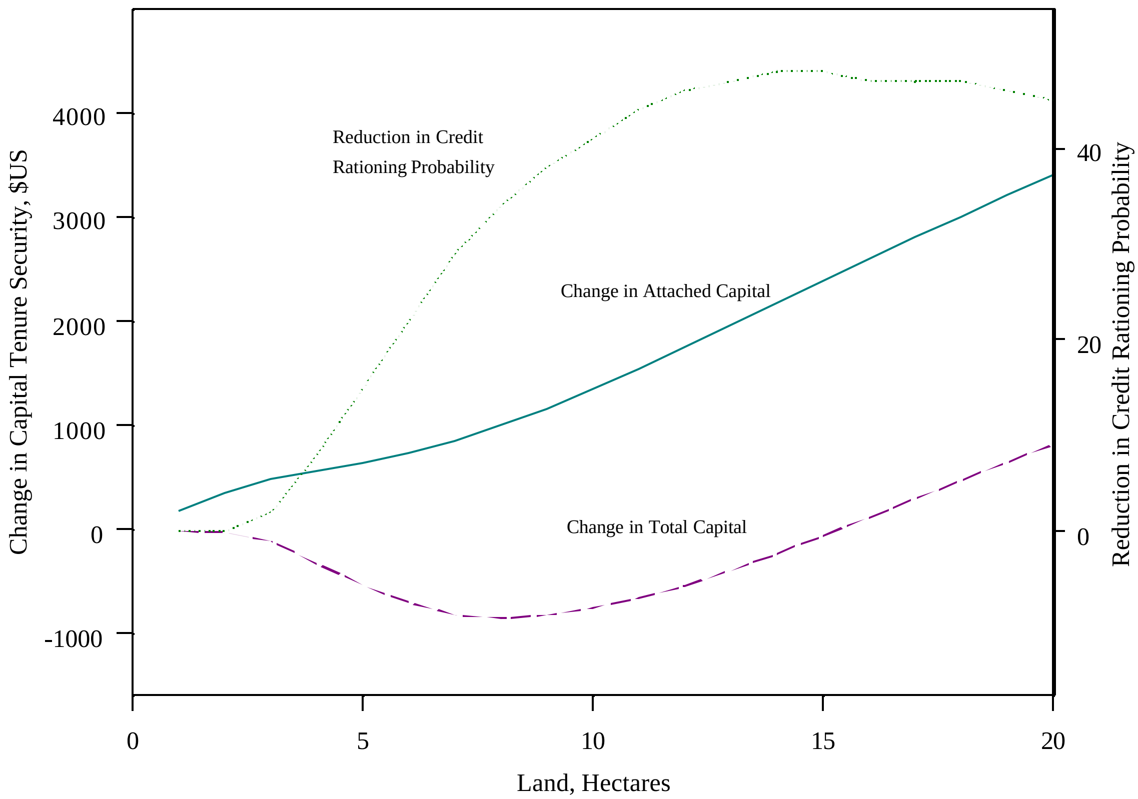

Table 1: Property Rights, Credit and Capital Stock
(current $US/hectare unless otherwise noted)

	Titled		**Formal**		**Other**	
	1991	*1994*	*1991*	*1994*	*1991*	*1994*
Has Formal Loan						
Attached Capital	133.2	103.0	25.3	23.2	108.2	77.9
Movable Capital	204.2	150.7	140.9	115.2	111.6	214.1
Formal Credit	58.9	50.2	32.2	28.9	31.5	38.6
Informal Credit	3.0	3.1	0.0	1.1	0.0	29.3
Farma Size (hectares)	141.6	111.3	41.1	43.6	47.9	42.2
% Titled or Formal	94%	97%	94%	85%	0%	0%
Number of Observations	34	48	5	6	4	5
No Formal Loan						
Attached Capital	82.6	73.9	103.2	92.7	53.9	51.7
Movable Capital	199.3	133.5	169.5	136.1	111.4	99.3
Informal Credit	11.4	8.3	7.5	5.5	10.2	10.8
Farma Size (hectares)	30.0	26.8	11.1	9.3	11.0	10.9
% Titled or Formal	93%	95%	90%	93%	0%	0%
Number of Observations	79	86	85	65	41	38

[*]For the purpose of this table, a household is considered titled if it owns at least some titled land, and formal if owns at least some formal document but no title.
[*] Other includes landless households (tenants) and squatters.

Table 2: Switching Regression Results

	FULL SAMPLE REGRESSION				TRIMMED REGRESSION			
	ATTACHED CAPITAL		**MOVABLE CAPITAL**		**ATTACHED CAPITAL**		**MOVABLE CAPITAL**	
	Coefficients	***Std Errors***	***Coefficients***	***Std Errors***	***Coefficients***	***Std Errors***	***Coefficients***	***Std Errors***
Liquidity Unconstrained Regime								
Legally Secure Farm Area (Ha)	134	87	-44	113	187*	91	-55	157
Farm Size (Ha)	-177*	85	165*	75	-317*	95	201**	107
Farm Size Squared	0.5*	0.2	-0.4*	0.2	0.78*	0.17	-0.43	0.28
Minifundia Region Dummy	-351	315	769	1139	-243	559	1159	1836
Frontier Region Dummy	-8022*	3001	-848	2305	-9617*	3743	-531	2844
1994 dummy	15	251	-1181	749	-66	315	-1490*	867
Liquidity Constrained Regime								
Formal Tenured Area (Ha)	254*	91	**-184***	**87**	**73**	**68**	**-110***	**57**
Titled Area (Ha)	173*	88						
Farm Size (Ha)	-337*	82	100	69	12	64	22	80
Farm Size Squared	0.76*	0.21	0.03	0.19	4.8	3.2	-2.0	2.4
Minifundia Region Dummy	-372	217	-283	330	268	187	-339	257
Frontier Region Dummy	770	503	-21	369	420	242	-207	282
1994 dummy	343	227	181	311	-269	199	237	239
Intercept Shift	173	797	2082*	1053	1952	4792	-6410	7716
R-Squared	0.44		0.14		0.62		0.12	
Number of observations	248		248		141		141	

*Different from zero at 10% level
**Different from zero at 5% level

Appendix: Credit Market Model

Full details on estimation of the disequilibrium credit market model summarized here are given Carter and Olinto 1998. The complete reduced form model for the formal credit market is:

(A-1) $$S_{it} = V_{it}^{s}\boldsymbol{d}^{s} + \boldsymbol{u}_{i}^{s} + m_{it}^{s}$$

(A-2) $$D_{it} = V_{it}^{d}\boldsymbol{d}^{d} + \boldsymbol{u}_{i}^{d} + m_{it}^{d}$$

(A-3) $$L_{it} = Min[Max(0, D_{it}), Max(0, S_{it})]$$

where D_{it} and S_{it} are household i's latent demand for and supply (or ration) of formal loans in period t. The parameter vectors $\boldsymbol{d}^{d}$ and $\boldsymbol{d}^{s}$ capture the marginal effects of household characteristics X_{it} (including tenure security) the demand for and access to formal loans. The terms $\boldsymbol{u}^{d}{}_{i}$ and $\boldsymbol{u}^{s}{}_{i}$ are mean zero household-specific effects that might be correlated with the observed explanatory variables. Time-variant, but cross-section invariant, variables (including the interest rate r_t) are subsumed into a time trend dummy variable that is included in the vectors of explanatory variables. The stochastic disturbances, m_{it}^{s} and m_{it}^{d}, account for household specific, time-varying omitted variables. We assume that they are bivariate normally distributed with zero means and variance covariance matrix given by:

$$(m_{it}^{s}, m_{it}^{d})' \sim N(0, \Sigma_{e}) \text{, such that } \Sigma_{\varepsilon} = \begin{bmatrix} \sigma_{d}^{2} & 0 \\ 0 & \sigma_{s}^{2} \end{bmatrix}.$$

Equation (A-3) is the observability condition that indicates whether the observed loan amount (perhaps zero) identifies a point on the supply or ration equation (A-1), or the demand equation (A-2). Together these three equations define an unobserved switching regression model with each of the regimes characterized by Tobit-like censorship.

Identification of the parameters in (A1)-(A3) is based upon the following assumptions:

1. The error terms m_{it}^{s} and m_{it}^{d} are orthogonal to the observed explanatory variables;

2. $\boldsymbol{u}_{i}^{d}$ and $\boldsymbol{u}_{i}^{s}$ are assumed to be related to the may be related to the other explanatory variables via the following linear projection which follows Mundlak (1978):

$$\boldsymbol{u}_{i}^{d} = (\textstyle\sum_{t} V_{it}^{d})\boldsymbol{a}^{d} + \boldsymbol{n}_{i}^{d}$$
$$\boldsymbol{u}_{i}^{s} = (\textstyle\sum_{t} V_{it}^{s})\boldsymbol{a}^{s} + \boldsymbol{n}_{i}^{s}$$

where $(n_i^d, n_i^s)' \sim N(0, S_n)$, $\Sigma_\nu = \begin{bmatrix} \eta_d^2 & \rho\eta_d\eta_s \\ \rho\eta_d\eta_s & \eta_s^2 \end{bmatrix}$, and r is the correlation coefficient between n^d_i and n^s_i.

3. *Partial Rationing Regime Information*. Information on informal loan transactions is used to provide partial information on rationing regime. Households that do not borrow in the formal market but do borrow from informal sources (mainly trader lenders) are assumed to be rationed in the formal market. Secondly, households that borrow both from formal and informal sources are assumed to be quantity-rationed in the formal credit market. Use of this information converts the problem from one with completely unobserved, endogenous regime-switching to one with partially observed, endogenous regime-switching.

Note that assumption 2 allows for correlation between the errors in equations (A-1) and (A-2) via the time invariant unobserved effects. That is, we assume that only unobserved characteristics that are time-invariant (e.g., entrepreneurial drive, talent, etc.) affect both demand and supply for formal credit simultaneously. Thus, it is assumed that time-variant effects that affect credit demand do not affect supply and vice-versa.

While the resulting likelihood function is analytically unmanageable, simulated maximum likelihood methods are used to estimate the parameters in (A-1) and (A-2) (see Gourieroux, C. and A. Monfort, 1993). For the simulation, the likelihood function is written conditional on the (n_i^d, n_i^s) pair. Maximization of mean conditional likelihood takes place over randomly generated error pairs of these error terms.

Estimation Results

Table A.1 displays the results for the credit market model estimated using the simulated maximum likelihood method. The table presents the results only for the full, unrestricted model. The hypotheses that the household effects are unrelated to the explanatory variables ($a^s, a^d = 0$) were rejected. The tenure security variable is defined as the ratio of titled to total land controlled by each household *i* at period *t* (*t* =1991, 1994). In order to allow for the possibility that title affects credit supply differently for different sized farms, the tenure security variable is interacted with the logarithm of total farm size (and the square of the logarithm of total farm size in the supply equation)[12]. The estimates of (16) that are displayed in Figure 2 are derived from these unrestricted estimates.

12 The dependent variable is total short-term formal credit received in 100 US dollars. Family labor is in adult equivalent, and Land is farm size in hectares.

Table A.1
Simulated Maximum Likelihood Estimates of Liquidity Constraints

EXPLANATORY VARIABLES	Unrestricted Model	
	Demand	*Supply*
Constant	-1.57 (6.23)	-117 (88.9)
Time Effect	2.731 (1.81)	11.89 (8.57)
Colonization Region Dummy	-3.77* (2.10)	33.9** (14.0)
Minifundia Region Dummy	-14.4** (2.66)	40.8** (16.9)
% Titled Land	35.2** (17.2)	-36.71 (106)
% Titled X ln(Land)	-15.8** (6.40)	47.0 (58.6)
Ln(Land)	3.73** (1.55)	-7.18 (51.6)
Family Labor Stock	0.469 (1.03)	
Ln(Land)^2		7.57 (5.59)
Titled X ln(Land)^2		-4.21 (5.49)
Household Specific Effect Projection:		
Ln(Land)	-0.31 (1.46)	-10.1 (33.7)
Family Labor Stock	-1.50 (1.19)	
% Titled X ln(Land)	19.9** (6.73)	-25.8 (43.0)
% Titled Land	-49.2** (18.6)	47.3 (109)
Variance Terms:		
σ	40.2** (13.0)	14.7** (1.78)
U_1	0.37 (1.04)	-1.10 (4.50)
U2		2.72 (4.39)
Log-Likelihood	-773.6	
Observations	496	

Figures in parentheses are estimated standard errors.

References

Barham, Bradford, Steve Boucher and Michael Carter (1996). "Credit Constraints, Credit Unions and Small-Scale Producers in Guatemala," *World Development* 24(5):792-805.

Bell, C., T.N. Srinivasan and C. Udry (1997). "Rationing, Spillover and Interlinking in Credit Markets: The Case of Rural Punjab," *Oxford Economic Papers* 49: 557-585.

Bruce, J. and S. Migot-Adholla, editors (1994). *Searching for Land Tenure Security In Africa.* (Kendall Hunt).

Carter, Michael "Equilibrium Credit Rationing of Small Farm Agriculture," *Journal of Development Economics*, February 1988.

Carter, M.R. and L. Galeano (1995). *Campesino, Tierra y Mercado* (Asuncion: CPES).

Carter, Michael and Ramón Salgado (2000) "Land Market Liberalization and the Agrarian Question in Latin America," in A. de Janvry, J.-P. Platteau and E. Sadoulet (eds) *Land Access, Rural Poverty and Public Action* (Oxford University Press).

Feder, G. and N. Akihiko 1996. "The Benefits of Land Titling and Registration: Economic and Social Perspectives," *Land Policy Studies*.

Feder, Gershon, T. Oncham, Y. Chalamwond and C. Hongladaron (1987). "Land Ownership Security and Farm Investment in Thailand," *American Journal of Agricultural Economics* 69:311-320.

Feder, Gershon, T. Oncham, Y. Chalamwond and C. Hongladaron (1989). "Land Ownership Security and Farm Investment: Reply," *American Journal of Agricultural Economics* 71(1):215-217.

Gourieroux, C. and A. Monfort (1993). "Simulation-Based Inference: A Survey with Special Reference to Panel Data Models," *Journal of Econometrics*, 59:5-33.

Honore, Bo E. and Ekaterini Kyriazidou (2000) . "Estimation of Tobit-Type Models with Individual Specific Effects." *Econometric Reviews* (forthcoming).

Kochar, A. (1998). "An Empirical Investigation of Rationing Constraints in Rural Markets in India," *Journal of Development Economics* 53(2): 339-372.

Kyriazidou, Ekaterini (1997). "Estimation of a panel data sample selection model." *Econometrica*, 65(6):1335-1364

Mundlak (1978). "On the Pooling of Time Series and Cross-Section Data," *Econometrica* 46:69-85.

Mushinski, D. (1999). "An Analysis of Loan Offer Functions of Banks and Credit Unions," *The Journal of Development Studies* 36(2):88112.

Prosterman, Roy (1998). "Efficiency and equity implications of different land tenure options: Evidence and the way ahead," paper presented at the Land Tenure, Land Market, and Productivity in Rural China conference, Beijing, May 15-16.