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# Credit market access and profitability in Tunisian agriculture

Jeremy D. Foltz*

*Department of Agricultural & Applied Economics, University of Wisconsin, 427 Lorch Street, Madison, WI 53706, USA*

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**Abstract**

This work develops an econometric model that links credit access with agricultural profitability and investment. Using data collected from rural Tunisia, this work provides direct estimates of credit access and its effects. Econometric estimates are run for agricultural investment and profitability as a function of credit access. The investigation of credit access and its effect suggests that the presence of credit market constraints does impinge significantly on farm profitability, but not on investments.


*JEL classification:* O12; C25; O53

*Keywords:* Credit markets; Switching regression; North Africa; Farm profits

## 1. Introduction

> How can we [Tunisians] pretend to have food self-sufficiency as our objective when we invest so little and accord so little credit [in agriculture] and we exclude the vast majority of the small and medium peasantry? (Sethom, 1992, p. 154)

Agricultural credit access has particular salience in the context of Tunisian rural development. Improving agricultural production and exports is a government policy objective. Recent structural adjustment loans to Tunisia from the World Bank (World Bank, 1996) have pushed the Tunisian government to reduce agricultural subsidies and price interventions, and let the private sector control marketing of agricultural products. Government investments in agriculture have been declining, with the private sector supposed to pick up the slack. Simultaneously, the government has started restructuring its banking sector to make it internationally viable through a program of privatisation and subsidy reduction. Evidence suggests that while the Tunisian structural adjustment program has been a success in most respects, it has not created an increase in private investment (Jayarajah et al., 1996). Private agricultural investment, which has lagged government expectations, requires adequate access to credit by farmers.

This study applies recent theory and methods on credit market disequilibrium to investigate the links between credit access and agricultural profitability and investment in Tunisia. While much of the literature on credit has been content to search for credit market imperfections, the work presented here seeks to push the analysis to another level. Having investigated the existence of credit market rationing, this work

* Tel.: +1-608-2626871; fax: +1-608-262-4376.
*E-mail address:* foltz@aae.wisc.edu (J.D. Foltz).

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asks the question: if credit markets work imperfectly, what effect does this have on agricultural investment and productivity? Using carefully collected data[1] from rural Tunisia, this work can directly estimate credit rationing and its effects. Direct estimates allow one to circumvent the problem of identifying empirically both the selection process of farm credit rationing and its effects on resource allocation. A second strength of the work presented here is that it uses disaggregated figures on agricultural production. This allows a focus on the relationship between credit rationing and investment as well as productivity.

## 2. Credit market literature

Recent theoretical and empirical work in economics has established that credit markets in developing countries work inefficiently due to a number of market imperfections.[2] The literature cites a number of market imperfections which lead some potential borrowers to be rationed out of the credit market. These imperfections include: (1) interest rate ceilings usually imposed by the government; (2) monopoly power in credit markets often exercised by informal lenders (Bell et al., 1997); (3) large transaction costs incurred by borrowers in applying for loans; and (4) moral hazard problems (Carter, 1988). In many cases a number of these imperfections combine to ration farmers out of the loan market.

While much of the literature (Conning, 1995; Kochar, 1997; Mushinski, 1999) concentrates on the determinants of access to formal loans with the idea of valuing the benefits to a future formal loan program, here we are primarily interested in how access to capital affects agricultural profits and investment. Some studies have measured the incidence of credit rationing and its effects (e.g. Jappelli, 1990; Feder et al., 1990; Barham et al., 1996). A few studies, notably work on loan programs by Zeller et al. (2001) in Bangladesh and by Diagne and Zeller (2001) in Malawi, have succeeded in both quantifying the degree of credit rationing (i.e. how much the borrower was rationed by the lender) and estimating its effects. Zeller et al. (2001) found that in Bangladesh credit access had a significant and strong effect on both income and food consumption. In contrast, Diagne and Zeller find that micro-finance institutions in Malawi had a negative impact on net farm income for participants. The results from Malawi suggest that interest rates may not be as important as loan program details, such as use restrictions, in determining the benefits of loan programs.

The literature (e.g. Carter, 1989; Feder et al., 1990) suggests that credit rationing can cause a misallocation of resources in farm production. This misallocation of inputs can then cause the credit rationed farmer to have lower profit levels than his unconstrained neighbour. The lower profit levels can come from a number of sources including lower investment levels and a misallocation of variable inputs. Although credit is often found to be a determining factor in profits and investments, multiple market failures outside of the credit market (e.g. labour or land markets or access to transportation) may overwhelm the effect of credit, as was found for example in Malawi by Diagne and Zeller (2001).

At the beginning of a production period, farm households need to allocate their available resources between current period consumption, purchase of variable inputs for production and investment. The household unconstrained in the capital market can separate consumption decisions from farm production decisions. Households can then choose production inputs optimally for the production process they face. In this case the levels of inputs in production and investment will not be affected by the level of credit they receive. The credit constrained household, however, will have to choose among the investments they make and the inputs they buy dependent upon the level of credit they receive. This will have a potentially detrimental impact on production in constrained households.

Two key hypotheses spring from the literature:

*Profit-liquidity effect:* Access to credit allows farmers to optimise input usage for a given set of fixed assets in the short term. Credit rationed farmers will use inputs only up to their capital

[1] The data used in this paper come from a 1995 survey of irrigated farms in the Cap Bon region of Tunisia. The randomly chosen sample consisted of 142 farmers who were asked about farm production, household financial status, irrigation technology adoption and access to credit markets.

[2] For a sampling of the theoretical literature see Stiglitz and Weiss (1981), Carter (1988) and Milde and Riley (1988).

availability. In particular the amount of liquidity a rationed household has will influence the overall profit level.

*Investment demand effect:* Farmers with credit access problems will invest less in capital assets and their land. Credit rationed farmers will not be able to smooth their expenses over time implying that they will not make long-term investments, especially those which entail sunk costs.

## 3. The econometrics of rationing

A farm household will be credit rationed when it demands more loans than the combination of the formal and informal markets is willing to supply. When markets do not clear fully through price adjustments, farmer credit status will be a function of farm and farmer characteristics effecting both supply and demand of credit. Let the notional demand curve of an individual be represented by $L^{D}(R_{f}, K, \theta, u_{D})$ where $R_{f}$ is the formal sector interest rate $(1 + r_{f})$, $K$ represents farm capital, $\theta$ represents farmer ability, and $u_{D}$ is a variable representing unobserved latent qualities. Similarly let the net supply of credit for that individual from all lending sectors, formal and informal, be given by

$$L^{S}(R_{f}, K, \theta, u_{s})=L^{f}(R_{f}, K, \theta, u_{s}) + L^{i}(R_{f}, K, \theta, u_{s})$$

where $L^{f}(\cdot)$ represents the formal sector loan supply and $L^{i}(\cdot)$ represents informal sector supply. The criterion for whether households are credit rationed is whether households demand more credit than lenders will supply to them. Define a variable $G^{*}$ as the reduced form excess demand for credit:

$$G^{*} = L^{D}(R_{f}, K, \theta, u_{D}) - L^{S}(R_{f}, K, \theta, u_{S})$$

Since the econometrician cannot directly observe the amount of excess demand, one moves to a reduced form estimation by defining an index variable for the credit rationed. Let $G$ take on the values of zero and one as follows:

$$G = \begin{cases} 1, & \text{if } G^{*} > 0\,(\text{rationed}), \\ 0, & \text{otherwise}\,(\text{non} - \text{rationed}). \end{cases}$$

In order to understand the determinants of credit status we are interested in characteristics of farmers and farms which influence the probability that $G^{*} > 0$. Define $Z$ as a vector containing observable farm and farmer characteristics influencing either supply or demand ($K$ and $\theta$). If $G^{*}$ were observable we could write it as a function of $Z$ in the following manner: $G^{*} = \gamma' Z + \varepsilon$, where $\gamma$ is a parameter vector to be estimated and $\varepsilon$ is a random disturbance term. With that formulation we can write the probability that $G^{*} > 0$ in the following manner:

$$\text{Prob}(G^{*} > 0) = \text{Prob}(\gamma' Z + \varepsilon > 0)$$

where $\varepsilon$ is an error term assumed to be normally distributed with mean zero and variance equal to one. The error term $\varepsilon$ represents both of the unobservable latent qualities of farmers and lenders, $u_{s}$, $u_{D}$, as well as potential noise in the data.

This formulation leads to a standard probit model to estimate the probability that a household is credit rationed. Assuming $\varepsilon$ has a standard normal distribution [$\sim N(0, 1)$], the log likelihood function for a probit will be

$$\ln L = \sum_{G_i=0} \ln(1 - \Phi_i) + \sum_{G_i=1} \ln \Phi_i$$

where $\Phi$ is the standard normal distribution evaluated at $\gamma' Z$.

## 4. Data implementation

The data used in this work comes from a 1995 survey conducted by the author of randomly selected households engaged in irrigated farming in the Cap Bon region of northeastern Tunisia. Table 1 shows the types of credit sources and percent of farmers who borrowed by town. Enumerators interviewed a total

Table 1
Credit sources

| Zone | No debts (%) | Debt from private sources | Bank debt | Private and bank debt |
|---|---|---|---|---|
| Korba | 12 | 72 | 4 | 12 |
| Tazerka | 28 | 64 | 4 | 4 |
| Diar Hjaj | 28 | 69 | 3 | 0 |
| Lebna | 39 | 47 | 7 | 7 |
| Teffaloune | 10 | 80 | 0 | 10 |
| Survey | 24 | 64 | 4 | 7 |

Table 2
Financial and credit status of households

| Financial or credit status | Percent of sample households (%) |
|---|---|
| (1) Have access to US$ 2000 in their extended family | 46 |
| (2) Have access to US$ 10,000 in their extended family | 12 |
| (3) Have no access to US$ 2000 anywhere | 4 |
| (4) Have no access to US$ 10,000 anywhere | 55 |
| (5) Have asked for a loan in the last year | 26 |
| (6) Would take a US$ 2000 loan at 13% interest | 62 |
| (7) Would take a US$ 10,000 loan at 13% interest | 52 |
| (8) Defined as credit rationed (those who agreed to take both loans: #6 and #7) | 45 |

of 142 farmers in five different towns with different cropping and technology patterns.[3] Along with a full survey on agricultural production and assets, respondents answered an extensive series of questions on their financial and credit status. These questions were designed to elicit the credit liabilities of families and their access to their own capital. The data shows 76% of the farmers with some sort of loan, with the majority of these loans coming from outside the formal banking sector.

Empirically we are interested in a measure of whether or not a household is credit rationed. Discerning this rationing is complicated by the fact that many households who do not take out loans may have zero demand for credit. Therefore, one must distinguish between those who have no credit because they have no demand and those who have no credit because they received an insufficient supply. Similarly, households with a positive supply of credit may not have received the full amount of credit they wanted. Thus, one must partition those who received credit into those who received sufficient credit and those with excess demand who did not.

One can parameterise the dependent variable $G$ using the data described in Table 2 on whether farmers would accept a new loan for agricultural production. This defines farmers as credit rationed if they would take 2000 and 10,000 (US$) loans offered to them (classification #8). In the survey procedure the loans were offered to farmers at the going formal sector interest rate for short-term loans. This represents an over-estimate of the actual costs of credit in the formal sector and was greater than the most often reported short-term loan rates in the informal sector. By requiring farmers to want both loans, this measure represents a conservative measure of credit rationing.[4]

The relevant variables in $Z$ will specify farm fixed capital ($K$) and farm quality measures ($\theta$). The fixed capital measures include owned land,[5] owned machinery, family labour per hectare, and household income. Measures of loan applicant quality include the number of titled hectares of land as a proxy for collateral, and the education level of the farm manager. In addition two variables are included to describe a farmer's history in the formal loan market: whether they had ever received a formal bank loan (36% of the sample) and whether they had ever defaulted on a bank loan (8% of the sample). A single regional dummy variable is used to distinguish the towns of Korba and Tazerka, which had more intensive cropping systems (higher demand) than the other three towns in the sample. Table 3 presents variable definitions and their averages broken down by credit status.

Because they are the outcome of an imperfect market equilibrium, no unambiguous predictions on the

[3] Of the 142 original farmers surveyed, 6 were dropped for missing data and unanswered questions. No significant patterns were found among those dropped from the data set.

[4] Other measures both stricter and looser were tried with results that corroborate the bifurcation of the sample used. Stricter measures, such as classifying as rationed only those who applied for loans in the last year (26% of sample), produced estimations in which only financial variables determined rationing. Looser measures, such as classifying as rationed those who would take any loan (74% of the sample), did not produce estimates that distinguished rationed from un-rationed farmers.

[5] Owned land is used as a measure of fixed assets, rather than operated land area which would be a measure of production potential. Potential production does enter into both borrower and lender decision making. However, operated land size may be endogenous to credit availability and land not owned by the farmers has no value as collateral to affect loan decisions. Bankers in Tunisia did claim that it was technically possible to use rented land as collateral on a loan if the rental contract was registered, with a duration of 3 years or more. When asked, few of the respondents knew of this possibility. Also, most rental contracts were informal and for a 1-year-period, suggesting that few farms if any were eligible for these types of loans.

Table 3
Variable means by credit status

| Variable | Full sample | Credit rationed | Non-rationed |
|---|---|---|---|
| Education level (1–5) | 2.28 | 2.42 | 2.16 |
| Family size | 6.9 | 6.9 | 6.9 |
| Years farming (farm manager) | 24 | 21 | 26 |
| Agricultural equipment owned (US$)[a] | 4,889 | 5,648 | 4,273 |
| Owned land (ha) | 2.7 | 1.2 | 4.0 |
| Household expenditure (US$/month) | 256 | 209 | 295 |
| Debts owed (US$)[b] | 4,941 | 4,064 | 5,654 |
| Loan default (1 if default, 0 otherwise) | 0.08 | 0.11 | 0.05 |
| Past bank loan (1 if had previous bank loan, 0 otherwise) | 0.36 | 0.33 | 0.39 |
| Titled land (hectares) | 2.21 | 0.64 | 3.5 |
| Profits (US$) | 11,702 | 13,624 | 10,139 |
| Farm investment: (US$ of technology, equipment, manure) | 2,055 | 2,110 | 2,010 |

[a] All monetary values are converted from Tunisian Dinars into US$ using the exchange rate in 1995 which was US$1 = 0.98 TD.
[b] Debts owed includes all debts taken on by farmers for the current agricultural season and any amounts still owed at the beginning of the season on long term loans from previous years.

signs of the reduced form estimation of excess credit demand can be made. The reduced form estimation will tell us most about which factors are more important to either supply or demand. A positive estimated coefficient, $\gamma$, signifies a characteristic which increases demand more than supply. Among the characteristics used, family labour and education level are expected to have a greater influence on demand than supply. Having title to land is expected to have a greater influence on supply than demand, because it increases collateral, creating a direct relationship to supply, while the increase in demand due to land titles moves indirectly through an investment demand equation. Having had previous bank loans and no bank loan defaults is likely to increase credit supply from the formal market and potentially also the informal market. The other variables (household expenditure, owned land and agricultural machinery) have indeterminate signs a priori depending on the strength of their influence on either supply or demand.

The results from a probit estimate of the probability that a household is credit rationed are presented in Table 4. The model predicts 70% of the farmers correctly and distributes farmers between rationed and non-rationed in approximately the right proportions, suggesting reasonable explanatory power.

The estimated coefficients show many of the predicted signs, with the probability of being credit rationed decreasing with household expenditure and land title. Higher household income levels, proxied by expenditure, would seem to increase credit supply more than credit demand. This fits with our intuition that a wealthier household would be more likely to receive credit, yet also less likely to need it. Land title, as predicted, increases credit supply more than demand. This suggests that the benefits of land title

Table 4
Coefficients of probit model: the probability of being credit rationed

| Variable | Estimate coefficient (standard error) |
|---|---|
| Owned land (ha) | 0.094 (0.087) |
| Agricultural equipment owned (US$) | 0.00002 (0.000013) |
| Family members per hectare of land (number/ha) | −0.036 (0.039) |
| Household expenditure (US$/month) | −0.0034*** (0.001) |
| Education level (1–5) | 0.139 (0.107) |
| Loan default (0–1) | 0.495 (0.447) |
| Past bank loan (0–1) | −0.017 (0.256) |
| Titled land (ha) | −0.166* (0.098) |
| Korba/Tazerka (0–1) | 0.741** (0.295) |
| Constant | 0.106 (0.362) |
| Log likelihood | 80.79 |
| Percent correctly predicted ($N$ = 136) | 70% |
| Likelihood ratio test $\chi^2$, (9 degrees of freedom) | 25.5** |

The dependent variable is a 0–1 indicator of being credit constrained, where 1 = constrained. Standard errors are in parentheses with the estimates.
* Significance at a 10% level.
** Significance at a 5% level.
*** Significance at a 1% level.

in increasing credit supply may be stronger than the degree to which it increases the desire to invest. This proposition is tested below in the investment demand equations. Agricultural machinery, though not significant, points toward farmers with more equipment as more likely to be credit rationed. It may be that more mechanised farmers continue to need to purchase more equipment, leading to greater demand. As most agricultural equipment cannot be used as collateral, it does not increase supply. The credit history variables have estimated parameters of the predicted signs, but are not significantly different than zero. The significant regional dummy variable suggests a higher demand for credit in Korba and Tazerka, one that is not mitigated by the better formal banking infrastructure in those towns. Since the highest degree of credit rationing occurs in the regions with the lowest transaction costs (at least in travel time) this suggests that transaction costs for farmers may not be the major cause of credit rationing.

## 5. The effects of credit rationing

Based on the hypothesis set forth above, a switching regression model framework is used to test the relationship between credit access and farmer profits and investment choices. In this case, the credit status, rationed or non-rationed, determines the switch between two different regimes describing the dependent variable. By positing two types of farmers, the rationed and the non-rationed, one can estimate the difference between the parameters of the equations as a measure of the costs of credit rationing.

Let an individual farmer's expected net farm revenues for the non-rationed and the rationed be denoted as $y^{\mathrm{n}}$ and $y^{\mathrm{c}}$. In general, the expected farm profits will be

$$E(y_i^{\mathrm{n}}|G_i = 0) = \beta'^{\mathrm{n}}_y x_i + \eta'^{\mathrm{n}}_y P + \delta^{\mathrm{n}}_y L_i^{\mathrm{S}} + E(v_i^{y\mathrm{n}}|G_i = 0),$$

$$E(y_i^{\mathrm{c}}|G_i = 1) = \beta'^{\mathrm{c}}_y x_i + \eta'^{\mathrm{c}}_y P + \delta^{\mathrm{c}}_y L_i^{\mathrm{S}} + E(v_i^{y\mathrm{c}}|G_i = 1),$$

where $G_i$ is the credit rationing indicator variable, $x$ represents observable farm and farmer characteristics including fixed assets, $P$ represents prices and $L^{\mathrm{S}}$ is the loan amount supplied to that individual. The random variable $v$ represents latent qualities unobservable to the econometrician. We expect the common coefficients in these two equations to differ between the rationed and non-rationed: i.e. $\beta^{\mathrm{n}} \neq \beta^{\mathrm{c}}$, $\eta^{\mathrm{n}} \neq \eta^{\mathrm{c}}$. For the credit rationed, we also expect that net farm revenues will increase with the amount of credit they received, $L^{\mathrm{S}}$, implying $\delta^{\mathrm{c}} > 0$, while for the unconstrained money would be borrowed up to the point at which its shadow value equalled zero, $\delta^{\mathrm{n}} = 0$.

Investment decisions ($I_i$) follow in a similar fashion with the addition of variables describing farmers' time horizons. The investment demand equation for an individual will be

$$E(I_i^{\mathrm{n}}|G_i = 0) = \beta'^{\mathrm{n}}_I x_i + \eta'^{\mathrm{n}}_I P + \delta^{\mathrm{n}}_I L_i^{\mathrm{S}} + E(v_i^{I\mathrm{n}}|G_i = 0),$$

$$E(I_i^{\mathrm{c}}|G_i = 1) = \beta'^{\mathrm{c}}_I x_i + \eta'^{\mathrm{c}}_I P + \delta^{\mathrm{c}}_I L_i^{\mathrm{S}} + E(v_i^{I\mathrm{c}}|G_i = 1).$$

In general we expect $\beta$ to have non-zero coefficients whether or not markets work efficiently. Credit rationed farmers are expected to have lower investment levels, after controlling for farm and farmer characteristics, because of their inability to borrow against future returns. However, we expect a liquidity effect, implying that $\delta > 0$ for the rationed farmers, and no liquidity effect for the non-rationed farmers.

### 5.1. *Exogenous switching model*[6]

First assume that the unobserved effects on profit between the rationed and non-rationed farmers are independent of the unobserved effects on the credit rationing equation:

$$E(v_i^{\mathrm{c}}|G_i = 0) = E(v_i^{\mathrm{n}}|G_i = 0) = 0.$$

This implies that $v^{\mathrm{c}}$ and $v^{\mathrm{n}}$ are uncorrelated with each other or with $\varepsilon$, the error term from the probit estimate of the probability that $G = 1$. Here, the sample separation between rationed and non-rationed households is assumed to be exogenous to their behaviour.

[7] To simplify notation only the case for the farm profit function is shown. The investment demand equations follow the same formulas presented below.

Dividing households into credit non-rationed (n) and credit rationed (c), we have the following exogenous switching structure:

$$y_i^{\rm n} = \beta'^{\rm n} x_i + \delta^{\rm n} L_i^{\rm s} + \nu_i^{\rm n}\; y_i^{\rm c} = \beta'^{\rm c} x_i + \delta^{\rm c} L_i^{\rm s} + \nu_i^{\rm c}$$

where $x_i$ is a matrix of exogenously determined household variables; $\beta^{\rm n}$, $\beta^{\rm c}$, $\delta^{\rm n}$ and $\delta^{\rm c}$ are the parameter vectors to be estimated, and while $v^{\rm n}$ and $v^{\rm c}$ are random disturbance terms which are not correlated with one another.

### 5.2. Endogenous switching model

For lenders and borrowers, loan demand and supply are governed by farm assets and the farmer's latent productivity attributes, $\theta$. To the extent that these latent productivity attributes are unobservable to the econometrician, they will be among the elements of the disturbance term $v$. For example, one would expect that greater farmer skills would decrease the probability of being credit rationed as well as increase the realised farm profits. If one cannot control for farmer skill with observable characteristic, e.g. education or farming experience, the disturbance terms $v^{\rm n}$ and $v^{\rm c}$ will be correlated with $\varepsilon$ from the credit rationing equation. In this case the exogenous sample separation estimation will confound the output effects of farmer skills with their credit access enhancement effects. The resultant estimated coefficients will be inconsistent.

Relaxing the constrained version, the sample separation can be estimated endogenously in the estimation of the dependent variable. In other words, the fact that a household might be credit rationed is allowed to be potentially correlated with the investments or profits of that household. Following the discussion in Maddala (1983) for the endogenous switching model, one can assume two regimes with an endogenous switching equation. For any observation $i$, the relevant structure is

$$y_i^{\rm n} = \beta'^{\rm n} x_i + \delta^{\rm n} L_i^{\rm S} + \nu_i^{\rm n} \quad \text{iff}\; \gamma' Z + \varepsilon_i \le 0,$$

$$y_i^{\rm c} = \beta'^{\rm c} x_i + \delta^{\rm c} L_i^{\rm S} + \nu_i^{\rm c} \quad \text{iff}\; \gamma' Z + \varepsilon_i > 0,$$

where the switching Equation is the standard probit estimation of whether a household is credit rationed from the previous section. In practice one observes only one value of $Y$ dependent upon which regime that particular individual is in, rationed or non-rationed. The parameters of the probit equation can only be estimated up to a proportionality constant, so we assume that the variance of the random disturbance terms will be one: $\text{Var}(\varepsilon_i) = 1$. We further assume that the random disturbance terms $v^{\rm n}$, $v^{\rm c}$, $\varepsilon_i$ have a trivariate normal distribution, with mean vector zero, and the following covariance matrix:

$$\sum = \begin{bmatrix} \sigma_{\rm n}^2 & \sigma_{\rm nc} & \sigma_{{\rm n}\varepsilon} \\ & \sigma_{\rm n}^2 & \sigma_{{\rm c}\varepsilon} \\ & & 1 \end{bmatrix}$$

For the credit rationed, the implied econometric model will be as follows:

$$\begin{aligned} E(y_i^{\rm c}|G_i = 1) &= \beta'^{\rm c} x_i + \delta^{\rm c} L_i^{\rm S} \\ &\quad + E(\nu_i^{\rm c}|\varepsilon_i > -\gamma' Z_i)\, E(y_i^{\rm c}|G_i = 1) \\ &= \beta'^{\rm c} x_i + \delta^{\rm c} L_i^{\rm s} + \sigma_{{\rm c}\varepsilon}\sigma_{\rm c}\lambda(\alpha), \end{aligned}$$

where $\lambda(\alpha)$ is the inverse Mills ratio from the probit equation describing credit rationing. For the non-rationed, similar equations apply:

$$\begin{aligned} E(y_i^{\rm n}|G_i = 0) &= \beta'^{\rm n} x_i + \delta^{\rm n} L_i^{\rm S} \\ &\quad + E(\nu_i^{\rm n}|\varepsilon_i \le -\gamma' Z_i) E(y_i^{\rm n}|G_i = 0) \\ &= \beta'^{\rm n} x_i + \delta^{\rm n} L_i^{\rm S} + \sigma_{{\rm n}\varepsilon}\sigma_{\rm n}\lambda(\alpha). \end{aligned}$$

The second stage estimation for both rationed and non-rationed incorporates the corresponding Mills ratios into a corrected linear regression for each of the two regimes.[7] The unconstrained case follows directly from the constrained case, and second stage estimates for $\beta^{\rm c}$ and $\beta^{\rm n}$ will be consistent and asymptotically normal. The resultant variance–covariance matrix, however, needs correction for heteroskedasticity which is done using a procedure outlined in the appendix to this paper.

[7] The inverse Mills ratio is defined for the constrained and unconstrained cases as follows:

$$\lambda(\gamma' Z_i) = \frac{\phi(\gamma' Z_i)}{\Phi(\gamma' Z_i)}, \quad \text{constrained;}$$

$$-\frac{\phi(\gamma' Z_i)}{1 - \Phi(\gamma' Z_i)}, \quad \text{unconstrained,}$$

where $\Phi$ and $\phi$ are, respectively, the cumulative and probability density functions of the normal distribution.

### 5.3. Estimations of the effect of credit rationing

Table 3 presents the independent and dependent variables used in the estimations. Those farmers unconstrained in the credit market on average have more titled hectares, marginally better credit histories, larger farms, higher expenditure levels, less agricultural equipment, lower debt levels and lower overall profitability. None of these differences is significant except for the expenditure levels which are different at the 5% confidence level. In general the variable means do not suggest that credit rationing influences profits or even the amount of credit received. One should remember, however, that these are unconditional means and that the estimations will provide a better view of whether credit rationing matters.

### 5.4. Profit equation

The farm profit estimates use net revenue functions, often called pseudo-profit functions (Carter, 1989), in order to account for possible imperfections in capital, land and labour markets. Net revenues differ from economic profits in that they do not account for depreciation costs and payments to fixed factors, including land, family labour and management. In an area of imperfectly operating markets, farmers will not be able to rent their fixed factors out at a 'market' price. The going market price for a fixed factor might well overstate the real opportunity costs of using that factor in production. Therefore, valuing profits using market prices for inputs might bias the results because deviations of profits will be due to differences in farmer endowments and access to markets, particularly the market for capital. Due to the imperfection of markets, farmers will make profit decisions based on the shadow values of their fixed assets, making quantity a reasonable proxy variable for the true price of those inputs.

As described above, the appropriate dependent variable describes net revenues of the farm, while the independent variables will describe farm and farmer characteristics which influence profits. Regressors in the two equations represent farm fixed assets (owned land and capital equipment), farmer characteristics (education and wealth), whether a farmer has title to the land he owns, and the amount of credit the farmer received. Since there was relatively little variation in prices among farms, price was dropped from the equations.[8] The dependent variable, pseudo-profits, is hypothesised to be increasing in farm fixed assets, farmer education, farmer wealth, and whether the farm has land title. The credit rationed are expected to have increasing profits in the liquidity variable with no liquidity effect seen for the non-rationed.

Table 5 shows estimated coefficients for a profit function estimated first assuming exogenous sample separation (first two columns) and then with endogenous sample separation (last two columns). Standard errors have been corrected for heteroskedasticity. The models produce fairly high levels of fit ($R^2$ equal to 0.49 and 0.74) for cross-section data with small datasets. A reasonable number of the coefficients are significant at common levels and have the expected signs. As predicted, several coefficients differ in magnitude between the credit rationed and non-rationed equations in both the exogenous and endogenous separation models.

The estimate of correlation between the error term in our switching equation and the profit equation is, as expected, negative. A farmer rationed when our model predicts he should not be, will have lower profits than we would otherwise predict. The endogenous separation model's estimate of $\lambda$, the inverse Mills ratio, does not turn out significantly different from zero suggesting that this formulation has not perfectly described all the variance in regression switching. The endogenous separation formulation changes the magnitude of a number of the coefficients in both equations, though it does not change any of their signs. The large but insignificant values of the endogeneity effect do not support our proposition of endogenous sample separation.

The results of our estimates for a pseudo-profit function provide support for the profit-liquidity hypothesis. The estimated coefficient on total debts owed for credit rationed farmers is significantly large in magnitude and statistically significantly different from zero. In addition the relatively smaller and non-significant

[8] Most of the differences in input and output prices were based on geographic differences, with two towns, Korba and Tazerka, having consistently higher labour costs and marginally lower input prices. These differences are picked up in the regional dummy variable.

Table 5
Pseudo-profit function coefficients for rationed and non-rationed households

| Variable | Exogenous separation | | Endogenous separation | |
|---|---|---|---|---|
| | Credit rationed | Non-rationed | Credit rationed | Non-rationed |
| Education (1–5) | −1076 (2088) | 1739 (1510) | −1929 (2247) | 1663 (1491) |
| Years farming (farm manager) | −231 (175) | −121 (104) | −184 (160) | −123 (98.2) |
| Family size | −850 (655) | −758* (421) | −698 (609) | −752* (393) |
| Agricultural equipment owned (US$) | 0.36** (0.14) | 0.51*** (0.19) | 0.19 (0.23) | 0.489** (0.23) |
| Owned land (hectares) | 2458** (936) | 477*** (78.5) | 2569** (1020) | 471*** (83) |
| Household expenditure (US$/month) | 44.2** (18.1) | 13.4 (8.19) | 72.7** (29.7) | 15.2 (14.1) |
| Debts owed (US$) | 0.772** (0.324) | 0.096 (0.07) | 0.71** (0.31) | 0.096 (0.065) |
| Title (0–1) | 3577 (3626) | 829 (2670) | 6109 (3765) | 999 (2714) |
| Korba/Tazerka (0–1) | 3796 (4109) | 3854 (2539) | −1758 (6138) | 3407 (3723) |
| Constant | 5651 (9273) | 4462 (6323) | 13112 (11289) | 3505 (8517) |
| Lambda | | | −15405 (11677) | −1177 (7573) |
| $R$-square [$n$(cons) = 61, $n$(un) = 75] | 0.49 | 0.74 | NA | NA |

The dependent variable is defined as total farm profits. Standard errors in parentheses.
* Significance at a 10% level.
** Significance at a 5% level.
*** Significance at a 1% level.

estimate on total debts for non-rationed farmers provides some support for the bifurcation of the sample used. From the endogenous switching regression the elasticity of profits with respect to credit is 0.20 for the rationed and 0.04 for the non-rationed.

As might be expected under imperfectly operating land markets, the amount of owned land is a significant determinant of overall profits. The significantly larger coefficient on land for the credit rationed suggests that credit rationed farmers have a higher shadow price for land. These differences are statistically significant at a 10% level.[9] The increased profit (shadow value) a rationed farmer would receive from renting another hectare of land is greater than the average land rental cost, while that for the non-rationed is less than rental costs. This implies that liquidity may constrain the ability to rent or buy land, leading to the divergence of shadow prices. The estimated shadow price (US$ 2569 for the rationed and US$ 471 for the non-rationed) of an additional hectare of owned land brackets the rental values commonly seen in Cap Bon which ranged from US$ 500 to 3000.

Household expenditure level as a proxy for overall household permanent income also has a positive effect on profits. The effect appears stronger for the credit rationed as one might expect if liquidity effects were present, although the differences between the coefficients for the rationed and non-rationed are not statistically significant even at a 10% level.[10] Agricultural equipment increases farm profits in both the rationed and non-rationed samples. Neither the rationed nor the non-rationed gain any profit gain from having better access to family labour, with an estimated negative effect for the non-rationed. Having title to land has a positive, but not statistically significant, effect on overall profits for both the rationed and the non-rationed.

### 5.5. Investment demand

In order to test the effects of credit constraints on farm investments, the dependent variable is defined as purchases of capital equipment, new technologies and manure. Capital equipment includes investments in tractors and machinery, irrigation pumps and green houses. The new technology is drip irrigation, a

[9] The lower bound of the 90% confidence interval on the credit rationed coefficient is 890, while the upper bound for the non-rationed is 609.

[10] The lower bound of the 90% confidence interval on the credit rationed coefficient is 23, while the upper bound for the non-rationed is 38.

Table 6
Investments

| Variable | Exogenous separation | | Endogenous separation | |
|---|---|---|---|---|
| | Credit rationed | Non-rationed | Credit rationed | Non-rationed |
| Education (1–5) | −534* (311) | 936*** (332) | −593** (302) | 588 (641) |
| Years farming (farm manager) | −49.1 (59.1) | 161** (67.4) | −49.3 (53.5) | 147 (108) |
| (Years farming)$^2$ | −0.21 (1.06) | −2.09* (1.07) | −0.14 (0.96) | −2.02 (1.73) |
| Family size | 68.9 (99.8) | −257*** (92.6) | 82.1 (92.1) | −221 (152) |
| Agricultural equipment owned (US$) | 0.059** (0.023) | 0.012 (0.041) | 0.046 (0.029) | −0.073 (0.086) |
| Owned land (ha) | 86.6 (145) | 14.5 (17.1) | 94.5 (137.7) | −9.52 (39.9) |
| Household expenditure (US$/month) | 4.20 (2.89) | 8.1*** (1.75) | 6.16 (3.87) | 14.7** (5.85) |
| Debts owed (US$) | 0.015 (0.048) | 0.063*** (0.015) | 0.011 (0.044) | 0.059** (0.029) |
| Title (0–1) | −677 (539) | −1543*** (575) | −489 (552) | −792 (1047) |
| Korba/Tazerka (0–1) | 2453 (620) | 769 (553) | 2059** (806) | −924 (1574) |
| Constant | 2148 (1428) | −2443 (1507) | 2703 (1555) | −6053 (3354) |
| Lambda | | | −1109 (1559) | −4610 (3082) |
| $R^2$ [$n$(cons) = 61, $n$(un) = 75] | 0.41 | 0.60 | NA | NA |

The dependent variable is defined as new farm investments. Standard errors are in parentheses with the estimates.
* Significance at a 10% level.
** Significance at a 5% level.
*** Significance at a 1% level.

resource saving and production increasing technology. Finally, manure purchases represent a measure of long-term investments in the land quality of a farm.

While all farmers will invest based on farm needs, captured in this equation by farm and farmer characteristics, based on Feder et al. (1990) we expect that the levels of investment for the capital rationed will also be a function of the capital available to them. As would be typical of investment demand functions, investment should be increasing and concave in farmer age, increasing in education, and decreasing in fixed assets (capital equipment and family size). Both the rationed and non-rationed are expected to invest more on land for which they have title.

The estimates of an investment demand function, shown in Table 6, produce fairly high levels of fit, $R^2$ equal to 0.40 and 0.69. A reasonable number of the coefficients are significant at common levels, but many have unexpected signs. The liquidity variable coefficients are significantly positive for the non-rationed, and insignificant, though positive for the rationed farmers. Also surprising is the negative coefficient on land title in all the equations, although it is only significantly different from zero in one of the estimates. Since some of the untitled land is under long-term rental or sharecropping contracts, it may be that a title is not a necessary condition for having enough tenure security to make investments.

The estimations suggest a difference in the determinants of investment behaviour of the rationed and the non-rationed. Education has opposite effects, with educated farmers who are rationed investing less, while the non-rationed operate as one would expect with more educated farmers investing more. The non-rationed also seem to follow a life cycle behaviour in their investments which are estimated as concave in age. In other words, investments increase with age, but decreasingly so as a person gets older.

Of primary interest, the investment demand equations show no liquidity effect for the rationed farmers, while the non-rationed farmers have investments increasing in both total debts and household expenditure. Rationed farmers may, because of a lack of capital, be choosing to spend the capital available to them on necessary variable inputs rather than longer term investments. It is possible that uncertain credit access reduces rationed farmers' ability to make long-term plans, so they invest less. The fact that for rationed farmers credit significantly affects profits but not investments suggests that the farmers are using the money available to them in productive ways, but not for long-term investments.

## 6. Discussion

This work set out to test whether Tunisian farmers were rationed in credit markets and whether that rationing affects production and investment. The findings suggest that according to a number of different measures of being credit rationed, a significant number of the surveyed farmers did not receive as large loans as they demanded. This credit rationing was shown to have direct effects on farm profits. The estimate of the profit function showed a liquidity effect of credit access on profitability, as well as significant differences in shadow values of land between the rationed and non-rationed. In contrast, the investment equations did not show a liquidity effect, but did show differences in investment behaviour between the rationed and non-rationed. The evidence suggests that the effects of credit rationing on profits do not operate through lower investment levels, but instead through lower production outcomes from sub-optimal allocations of other factors of production such as land, labour and variable inputs.

The liquidity effect of credit rationing on profits has a number of implications for agricultural policy in Tunisia. Better access to the credit market will improve the profitability of a great number of farmers, though not necessarily the poorest. In addition if credit access were improved, it might activate the rural land markets by allowing farmers to rent in or buy the optimal amount of land. In fact many of those who claimed they would borrow when offered a loan said they would do so to buy or rent land. Given the potentially high cost of credit access programs and the relatively low elasticities of profit with respect to debt (0.20) for the credit rationed, however, it is unclear whether such benefits would justify the cost of improving credit access.

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## Appendix A. Correction for heteroskedasticity in an endogenous switching model

As Maddala (1983, pp. 227–228) shows, the estimated variances for the endogenous switching model will be underestimated if we use a simple heteroskedasticity corrected variances from OLS. This happens because we would be ignoring the effects of $\lambda(\alpha)$ on the variance of $\beta$. Instead we need to correct for the endogeneity of $\lambda(\alpha)$ in the variance covariance matrix. Let $W_1$ be the vector of $\lambda(\alpha)$ for the credit rationed. Define $C_1 = (X_1 - W_1)$ and the estimates of the parameters as

$$\begin{bmatrix} \beta_1 \\ \sigma_{1\varepsilon} \end{bmatrix} = (C'_1 C_1)^{-1} C'_1 Y_1,$$

where $\sigma_1\varepsilon$ is the estimated coefficient on the inverse mills ratio and $Y_1$ is the dependent variable. Let $N_1$ be the number of observations in the rationed sample out of a full sample size $N$ and define the matrix $D_1$ as an $N_1 \times N_1$ matrix whose $i$th diagonal element is $W_{1i}$ $(W_{1i} + \gamma = Z_i)$ where $Z_i$ is the $i$th element of the matrix $Z$. Also define $\delta$ as an $N \times N$ matrix whose $i$th element is equal to $\Phi_i/[\Phi_i(1 - \Phi_i)]$. Now define the estimated variance in the following way:

$$\begin{aligned} \operatorname{Var} \begin{bmatrix} \beta_1 \\ \sigma_{1\varepsilon} \end{bmatrix} &= \sigma_1^2 (C'_1 C_1)^{-1} C'_1 Y_1 - \sigma_{1\varepsilon}^2 (C'_1 C_1)^{-1} C'_1 \\ &\quad \times [D_1 - D_1 Z_1 (Z' \Delta Z)^{-1} Z_1 D_1](C'_1 C_1)^{-1}, \end{aligned}$$

where $Z_1$ is the partition of the matrix $Z$ for those who are credit rationed. For the non-rationed one can change the subscript to get the result.

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