An Empirical Investigation of the Impacts of Government Subsidies on Farmland Rental Rates

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1. Introduction

Land is an essential input for crop farming and ranching. Farmland is also the main asset in U.S. agriculture as it accounted for 68% of all farm assets between 1960 and 2001 on average (Lence and Mishra 2003). Farmland renting is a common practice in the U.S. About 58% of farmers own the land on which they are operating. The remaining 42% rent cropland in accordance to their production plans. In addition, about 18% of farm operators rent more than three quarters of the land they are farming and 7% of farmers can be considered pure tenants; i.e. they do not own any of the land on which they are operating (Goodwin, Mishra, and Ortalo-Magné, 2003).

Meanwhile, agricultural subsidies have become a major source of farm income in the U.S. over the years. For instance, over 70% of the net farm income of North Dakota agricultural producers came from government subsidies in 2001-2002 (Schmitz and Just, 2003). A wide variety of income support programs exist and are explicitly intended to raise and stabilize farm incomes. In the 2008 Farm Bill, financial support to agriculture over the life of the Bill was estimated to be about $300 billion (Goodwin, Mishra, and Ortalo-Magné, 2005).

Given the fact that a large share of U.S. farmland is rented out and operated by tenant farmers, it is important to evaluate how landlords and tenants share government subsidies. The greater is the share of the payments that go to landlords (through for example increases in the land rental rates), the less effective are subsidies to support

farmer incomes (Patton, *et al*. 2008). Some economists (see for example Sherrick and Barry, 2003) argue that the benefits of agricultural programs accrue entirely or almost entirely to operators who own all or part of the land and to non-operator landlords. Pure tenants gain little or nothing from programs since they have to pay higher rents as a result of the subsidies. In addition, pure tenants do not benefit from the capital gains generated by farmland value increase. Conversely, other researchers, (e.g., Kirwan, 2009), suggest that tenant operators, not landlords, are the main beneficiary of government subsidies. Goodwin, Mishra, and Ortalo-Magné (2003, 2005, 2009), and Patton *et al*. (2008) questioned this 'general statement'. They argue that the effects of program benefits vary substantially across the types of programs because production and risks associated with these payments are different. Generalizations with regard to the overall effects of program benefits on land values or rental rates can be misleading.

The purpose of this paper is to analyze the impacts of different farm programs on farmland rental rates under different leasing arrangements. We consider cash rental and sharecropping contracts. The literature has generally focused on cash leasing arrangements. However, other land tenure contracts are commonly used in the US. In 1999, about 59% of all leased farmland was under cash contracts. About 24% and 11% of leased land was, respectively, under share and hybrid contracts (USDA/NASS 2001). The latter type of contracts borrows elements of sharecropping and cash rental agreements such that the tenant pays part of the rent in cash and part as a share of crops or livestock. The benefit distribution of subsidies is complicated by the existence of different leasing arrangements. Legislation also involves restrictions and constraints on the distribution of program benefits. It requires in certain cases that payments be shared among producers

and landlords subject to the contract on a fair and equitable basis. For example, under a cash rental arrangement, direct decoupled payments are required to be distributed entirely to the farm operator. Naturally, landlords may indirectly capture a share of those payments by raising the cash rental rates. Under a share contract, government payments are designed to be distributed to both the tenant and landlord according to the proportion that they share the output. Hence, a landlord may capture program benefits directly as well through the monetary terms of the leasing contract. Changes in tenure arrangements (e.g., from a share contract to a hybrid contract) may also reflect a redistribution of benefits between tenants and landlords (Qiu, Goodwin, and Gervais 2009). Goodwin, Mishra, and Ortalo-Magné (2009) provide empirical evidence that government subsidies significantly impact farmland rental rates and the capitalization rates vary according to cash rental and sharecropping contracts.

The contribution of this study is twofold. As in most empirical exercises, the devil is in the details. We refine the construction of the dependent variable in studying the relationship between rental rates and government payments. We construct a "pure rental rate" which controls for the impact of a landlord's input cost sharing at the estimation stage. More than one third of landlords in the US share variable costs with their tenant operators. It has been a customary practice in the literature to construct a measure of the rental rate that does not account for this feature. We show that this may generate significant biases at the estimation stage. Second, the empirical procedure corrects for potential selection issues. Qiu, Goodwin, and Gervais (2009) have shown that government payments have a significant impact on the institutional arrangement chosen by the landlord and tenant operator. Because one observes a specific type of rent only if

this farmer first chooses this particular type of contract, there will likely to be some correlation between the error term in the regression explaining rental rates as function of government payments and the choice of a particular leasing arrangement. Different methods are investigated to detect and correct the selection bias.

2. Conceptual framework

This study uses the income approach proposed by Goodwin, Mishra, and Ortalo-Magné (2005) and Patton *et al*. (2008) to determine the impacts of payments on land rental rates. It is based on the premise that agricultural land rents reflect the profitability of the rented assets. Let r_t denote a farmland rental rate in period t. Farmland rental rates are expressed as a linear function of all variables that affect profitability, namely expected net agricultural market returns (denoted Mkt_{t+1}) and sum of expected future government subsidies (denoted $GovP_{t+1,j}$):

$$(1) \quad r_t = \alpha + \beta E_t\left[Mkt_{t+1}\right] + \sum_{j=1}^{J} \gamma_j E_t\left[GovP_{t+1,j}\right],$$

where $E_t[\cdot]$ denotes the expectation operator using a time t perspective. Benefits of government program j are measured on a per-acre basis. The parameter α is a constant which can be interpreted as the average effect of unobserved factors, such as potential non-agricultural earnings. The coefficients β and γ_j can be interpreted as the degree of capitalization in rental rates from market earnings and government subsidies, respectively.

The framework in (1) implies that the incidence of government subsidies vary by programs. We investigate three major government programs under the 2002 Farm Bill: *i*)

Loan Deficiency Payments (LDP); *ii*) direct payments (including direct decoupled payments and counter-cyclical payments); and *iii*) disaster payments. We also pool all remaining program payments into one category which includes conservation reserve program payments and subsidies from local governments. Eq. (1) can be rewritten as:

$$(2) \quad r_t = \alpha + \beta E_t[Mkt_{t+1}] + \gamma_1 E_t[LDP_{t+1}] + \gamma_2 E_t[DirP_{t+1}] + \gamma_3 E_t[DisP_{t+1}] + \gamma_4 E_t[Other_{t+1}]$$

Eq. (2) is the equation of interest. The next section deals with the construction of the database and discusses the empirical challenges at the estimation phase.

3. Data and Empirical Strategy

We use the 2002-2007 confidential farm-level Agricultural Resource Management Survey (ARMS) data and unpublished 1998-2007county level government program payment data from the United States Department of Agriculture (USDA). We also use data from the 1988-2007 Regional Economic Information Systems (REIS) and the county level farmland data from the 1997 and 2002 Census of Agriculture.

The ARMS is a national survey that provides observations of farm-level production practices, economic attributes, and operator households' characteristics. We utilize this dataset to calculate farm-specific rental rates, as well as obtain information about additional farm and operator characteristics that may impact leasing choices. The REIS contains annual estimates of personal income at the national, state, metropolitan area, and county levels. We obtained the county level net agricultural market returns (cash receipts from marketing minus the production costs) from REIS.

From ARMS data, we select farms (referred to as target farms hereafter) that rented at least part of their land. The selected ARMS data are merged with other county-

level variables using the five-digit Federal Information Processing Standards (FIPS) code. The ARMS questionnaire asks tenant operators to provide the information about farmland leasing arrangements. Farmland contracts are divided into three categories: cash rental (both fixed and flexible amounts), share (also include hybrid contracts), and rent for free. This study focuses on cash and share leases which account more than 95% of the rental arrangements. Tenant operators report acres rented from landlords under each leasing arrangement. They also report total cash rents paid to and the shares of production that go to their landlords. In our database, target farms are classified into one of three categories according to their leasing arrangements: “cash-only”, “share-only”, and both. For example, the target farms that rented land from others only using cash rental contracts are considered as cash-only farms. In our sample set, 68% farms are cash-only farms, 12% are share-only farms, and the remaining 20% include both. We exclude outliers (less than 2% of the available sample) from the analysis because they represent atypical situations (for example, farms reporting rents exceeding $1,000 per acre). After this selection procedure, a total of 60,981 observations are used for the empirical analysis. Table 1 presents the definitions of key variables and summary statistics.

Measurement errors in dependent variables

Because the true per-acre rental rates are usually unobserved, the first step of the empirical analysis is to compute a proxy for this dependent variable. We argue that most analyses that relied on a constructed measure of the rental rate may suffer from measurement errors that may ultimately cause important biases at the estimation stage. Consider the Census of Agriculture data as an example. The census does not directly record the per-acre rental rates. It only records the total cash rent paid as well as total

acres rented. The latter variable includes all rented acres under cash, share, and other types of leasing arrangements. Previous studies (e.g., Roberts, Kirwan, and Hopkins, 2003; Kirwan, 2009) computed a per-acre cash rent by dividing total cash rent by total acres rented. Because the dependent variable includes rented acres under sharecropping (and other hybrid forms of leasing arrangements) and because acreage or contract choices are correlated with government subsidies (Goodwin and Mishra, 2006; and Qiu, Goodwin, and Gervais, 2009), there may be an important endogeneity bias at the estimation stage.

In case of the ARMS data, the 2002-2007 survey reports rented acres under cash contracts and share contracts separately. The latter category however includes hybrid contracts which are rent payments based on a fixed cash payment and along with some shared production. Given the presence of hybrid contracts, it may be misleading to compute the per-acre cash rent as the total cash amount (paid as rent) divided by total acres (rented under cash contracts). The per-acre cash rent would be overestimated while the per-acre share rent would be underestimated. Such a proxy for the rental rate may suffer from a measurement error.

We solve this problem by distinguishing hybrid contracts from share contracts, and then calculate the rental rates under different types of contracts separately. First consider cash-only farms. There are no reporting issues for these farms, i.e. the per-acre cash rent can be calculated as total cash rent divided by total acres under contracts. We then turn to farms in our database which classified as share-only (i.e. only use share contracts). Recall that in the ARMS dataset, the farms that report share leasing contracts include pure sharecropping and hybrid contracts. For these share-only farms, if reported

cash payments are zero, we can conclude that pure share contracts were used. The per-acre share rent can be calculated as the total value of the landlords' share of production divided by the acres under share contracts, and these observations will fall under the pure-share category. Finally, a share-only farm can report share contracts as well as positive cash payments; which would indicate a hybrid contract has been used. The number of farms that use only hybrid contracts in the sample is small (a little over 1% of the target farms); and thus are excluded from the estimates of rental rates on government subsidies.

Most studies use directly the computed rents as the dependent variable. This may also be problematic because it is quite common that landlords share part of the production inputs (other than land) with tenants, especially under share leasing arrangements (Allen and Lueck 2002). In our sample, more than 33% of the landlords reported sharing production costs with their tenants. In that case, the computed rents reflect the payments made to landlords in order to compensate for the costs of the non-land inputs. When these payments are correlated with one or more of the regressors (e.g., market returns), it introduces measurement problems which may result in biases at the estimation stage. To account for this, we subtract the landlords' share of variable costs from the calculated rental rates; thus providing a more reliable proxy of the dependent variable at the empirical model.

Measurement errors in independent variables

Theoretically, farmland rental rates are functions of expected cash flows from various sources. Certain payments (such as direct decoupled payments) are known with certainty prior to signing a contract. However, payments such as disaster and LDP payments are

not predetermined. They are triggered by market or production conditions. Measurement issues arise if actual reported payments are used to represent expectations as noted in Goodwin, Mishra, and Ortalo-Magné (2003). In addition, realized individual payment data may be correlated with unobserved factors, such as land productivity. This can cause endogeneity problem and result in bias as well. To control for the potential errors-in-variables and endogeneity problems, we follow the Goodwin, Mishra, and Ortalo-Magné (2005) approach and use a historical five-year county average of payments instead of the realized individual data reported in ARMS. As such, expected net market earnings are measured by a historical five-year county average value per acre. We also include a measure of 'farming risk' by computing a ten year average coefficient of variation of the farms' market returns.

Selection bias

One can observe a specific type of rent only if the farmer first chose this type of leasing arrangement. Sample selection bias occurs if unobservable characteristics (e.g., land productivity) that affect the rental rates are correlated with factors that affect the contract (e.g., farm level government subsidies and production risk).

There exists an extensive literature on the detection and correction of selection issues. Bourguignon, Fournier, and Gurgand (2007, BFG hereafter) provide an overview of the methods available to account for selection issues in the context of the multinomial logit model. BFG conduct a set of Monte Carlo experiments and find that in many cases the approach introduced by Dubin and MacFadden (1984) is the preferred method in comparison to the most commonly used procedure proposed by Lee (1983). BFG also develop an estimator that gets around the restriction of the Dubin and MacFadden

(1984)'s estimator and which is more robust. Our empirical analysis mainly focuses on the DMF approaches. However, the results from alternative methods will be reported as well.

General description of the selection model

Consider a simple two-equation censored regression model. Let the subscript j define a categorical variable that describes the choice of a decision-maker among M alternatives based on utilities y_j^*. The variable of interest y_1 is observed if and only if alternative 1 is chosen:

$$(3)\quad \begin{aligned} y_1 &= \mathbf{x}\boldsymbol{\theta}_1 + u_1 , \\ y_j^* &= \mathbf{z}\boldsymbol{\lambda}_j + \eta_j , \qquad j = 1.....M , \end{aligned}$$

where vectors x and z are vectors of exogenous variables and u_1 is a disturbance that verifies $E(u_1 \mid x, z) = 0$ and $V(u_1 \mid x, z) = \sigma^2$. If one observes alternative 1 being chosen, it means that $\varepsilon_1 = y_1^* - \max_{j \neq 1} (y_j^*) > 0$. Assume η_j is independent and identically Gumbel distribution and define $\Gamma \equiv \{z\lambda_1, \ ..., z\lambda_M\}$. After some algebraic manipulations, BFG show that a consistent estimate of the vector $\boldsymbol{\theta}_1$ can be obtained based on the regression:

$$(4)\quad y_1 = \mathbf{x}_1\boldsymbol{\theta}_1 + \mu(\Gamma) + w_1 = \mathbf{x}_1\boldsymbol{\theta}_1 + \tilde{\mu}(P_1, ..., P_M) + w_1$$

where $\mu(\Gamma) = \tilde{\mu}(P_1, P_2, ..., P_M) = E[u_1 \mid \varepsilon_1 > 0, \Gamma]$ is the conditional mean of the disturbance in the equation of interest, P_j is the probability that the alternative j is chosen, and w_1 is a residual which is mean-independent of the regressors.[1] Bias correction models discussed in BFG paper differ in the restrictions imposed on $\tilde{\mu}(\cdot)$, or equivalently, $\mu(\Gamma)$. BFG summarize that restrictions can be expressed in terms of the

joint distribution of the residuals in eq. (4) $(u_1, \eta_1, ..., \eta_M)$ given the linearity assumption. Two types of restrictions exist on the covariance matrix of the error terms and linearity assumption.

Model specification

The empirical model includes a set of selection bias correction obtained from a multinomial logit model. The equation of interest is a regression of rental rates on expected market earnings and expected government payments. A target farm chooses a leasing contract among three alternatives: cash-only contracts, pure-share contracts, or both cash and share contracts. The decision is made conditional on a set of independent variables which includes the variables from the equation of interest and other democratic variables such as farm type and age of the operator.

The selection equation implicitly imposes the Independence of Irrelevant Alternatives (IIA) assumption which states that the probability of choosing among two alternatives is not affected by the presence of additional alternatives. The Chi-Square test statistic proposed by Hausman and McFadden (1984) are used to test the IIA. We are not able to reject the null hypothesis that the IIA assumption is valid at a high level of significance. In addition, the ARMS survey is conducted annually from a stratified random sample of farms. It is possible that individual farms have been sampled in multiple years. Correlation among observations from the same farm may exist. Therefore, clustered robust standard errors are used and based on the farm's unique id number in the database. The second step applies weighted least squares in the rental rates regression to account for heteroskedasticity presented in the model due to selection issues. We also use

a bootstrap with 50 replications to estimate the parameter standard errors in the rental rates regression.

4. Results

The empirical analysis includes two parts. First, we investigate the incidence of aggregate government subsidies on farmland rental rate - under both cash and share arrangements. This provides a general idea of the extent of capitalization into farmland rental rates of an extra dollar of government payments. Second, we disaggregate government payments into three distinct programs and evaluate the impacts of each program payments on rental rates. In each part, we report the results from four different models: *i*) a generalized linear regression (GLR); *ii*) Lee (1983) selection procedure; *iii*) Dubin-MacFadden original selection procedure (which we identify as DMF1); and *iv*) Dubin-MacFadden flexible selection procedure (labeled DMF2).

Aggregate program payments

Table 2 presents the estimates the impacts of aggregated subsidies on cash rental rates obtained from a generalized linear regression (model 1) and selection regression models (models 2-4).[2] The results from the generalized linear regression show that each additional farm subsidy tends to raise cash rents by $0.73 per acre. Landlords capture 73% of the total benefits and leave only 27% to tenant producers. Meanwhile, the results indicate that an extra $1 obtained from market returns increases the rental rates by $0.12. This suggests that landlords acquire higher benefits from government subsides than from market earnings. The estimated coefficient of the variable CV is -163.19. Given the mean value and standard deviation of 0.17 and 0.07, respectively, reported in Table 1, an

increase in CV of one standard deviation will decrease cash rental rates by $11.42 on average. The negative effect of CV on cash rental rates suggest that risk is an important determinant of rents.

We have argued that a generalized linear regression framework that does not account for the role of government payments and other related factors on the types of leasing arrangements may result in an estimation bias because of selection issues. The results from models that account for selection issues are also reported in Table 2 (model 2-4). All three selection models reveal a statistically significant correlation coefficient between the error terms of the cash rental rate equation and multinomial logit model. This suggests that selection issues in the context of government payments and leasing arrangements are very relevant. As argued in the previous section, there exists evidence in the literature that the DMF models (DMF1 and DMF2) are superior than the most commonly used method of Lee (1983). In what follows, the discussion of results largely focuses on the two DMF approaches (models 3-4).

The DMF results suggest that an additional dollar of government subsidies tends to increase cash rental rates by $0.37-$0.38. The incidence of government payments into cash rents is much smaller than that under the GLR model, and can be explained the selection issues not captured by the generalized linear regression model. The results also show that an extra dollar from market returns raises the rental rates by $0.17-$0.18. In words, landlords capture 17%-18% of market returns under cash arrangements. This number is higher than those obtained by the GLR model. The impacts of CV on cash rents are similar under GLR and DMF models. The results from selection models confirm that government subsidies have a significant positive effect on cash rental rates.

However, the capitalization rates are not as large as the ones predicted by a simple linear framework when the selection issues are accounted for.

Table 3 presents the estimates of the share rental rates on aggregated government payments. The share rents in the empirical analysis include the value of the share of production that went to landlords (according to lease terms) and government payments directly distributed to landlords. In contrast to cash rents, the results from the generalized linear regression and selection models are similar. The results from the DMF models suggest that each additional $1 of aggregate government subsidy raises share rental rates by $0.86-$0.88. In the U.S., 50-50 share arrangements are common practices (Huffman and Just 2004). Allen and Lueck (2002) suggest that 50-50, 60-40 (the tenant obtain 60% of the production and leave 40% to the landlord), and 67-33 are the three most common sharecropping rules in their sample of Nebraska and South Dakota leasing arrangements. Our results suggest that landlords are able to capture benefits that are substantially higher than the common parameters of the leasing arrangements. Hence, it offer support to the argument that landlords are capturing program benefits both from direct program payments and from increased rental rates.[3] In other words, legal restrictions on benefit distribution between contracting parties are ineffective and benefits can be redistributed by other means such as adjusting/renegotiating the rental rates.

The results also indicate that landlords extract 22%-23% of the market returns under share leases. The proportion is larger than that under cash contracts (17%-18%). The difference probably reflects a risk premium that landlords share with tenant farmers. Non-land related input costs shared by landlords have already been accounted for in the construction of the dataset. When considering the impact of risk on rental rates, the

results suggest the rents will decrease $9.18-$9.73 per acre following a one standard deviation increase in CV. The negative effects of CV are found to be smaller than those under cash contracts ($12.56-$14.52).

In summary, the results confirm that aggregate government subsidies have substantial effects on rental rates. These effects vary across leasing arrangements. The incidence of program payments on share rental rates is larger than those on cash rental rates. Given that more than 40% of leased farmland is under share or hybrid contracts, it must be noted that results obtained from cash rental arrangements to proxy the entire distribution of benefits between landlords and tenants is likely to be misleading.

Disaggregate program payments

Table 4 presents the estimation results of the cash rent equation when payments are disaggregated according to the different programs. The results from the generalized linear regression model (model 9) show that all three program payments have significant and large impacts on cash rental rates. In particular, an increase of one dollar in LDP and direct payments increase the cash rental rates by $1.10 and $0.84, respectively. In contrast, the disaster payments are found to have a large and negative effect on cash rents (in a near two to one factor). Yet, results from the linear regression framework may suffer from selection biases. The results from the DMF models (models 11-12) are indeed different. LDP are estimated to have a negative impact on cash rents. The statistically significant negative coefficient is counter-intuitive.

Direct payments (which include direct decoupled payments and counter-cyclical payments (trigged by low market price, but based on historic base acreages) are independent of current production activities. Qiu, Goodwin, and Gervais (2009) argue

that landlords are able to capture the entire benefits of payments if the wealth effect is negligible. Our results support this argument. The DMF models show that an increase of one dollar in direct payments yields a benefit of $0.85-$1.10 to the landlords, *ceteris paribus*. Lence and Mishra (2003), and Goodwin, Mishra, and Ortalo-Magne (2009) report the similar findings in their studies. Disaster payments do not have statistically significant impacts on cash rental rates. Disaggregating payments according to their type has no impact on the relationship between market returns and cash rents. Landlords claim $0.17-$0.19 in benefits for each additional $1 return from the market under cash leasing arrangements.

Table 5 repeats the analysis for share rental rates. The DMF estimates imply that share rents increase by $0.69-$0.77 for each dollar increase in of direct payments. Though effects are a bit lower than those for cash rental rates, this result still indicates that landlords capture most of the government program benefits. The results show that an additional dollar in LDP raises share rental rates by $0.95-$1.18. These estimates are different than the results under cash contracts. As in the case of cash contracts, disaggregating payments has no bearing on the relationship between market returns and rental rates. Landlords claim $0.23-$0.25 in benefits for each additional $1 return from the market under share leasing arrangements. The impacts of disaster payments are found to be statistically insignificant.

5. Conclusion

This paper provides an empirical investigation to the effects of government subsidies on rental rates. We use data from a variety of sources to analyze the degree of capitalization of government payments using selection bias correction models. We report the results of

four different estimation procedures. We find that government payments have statistically significant impacts on farmland rental rates. These impacts vary across leasing arrangements. More specifically, we find that landlords capture 37%-38% of aggregate subsidies under cash leases and 86%-88% under share contracts. Disaggregate farm programs are also found to have different impacts on rental rates according to the type of programs. In addition, the results also indicate that risk-sharing is an important determinant of farmland rents.

Given the increased reliance on contracting in agriculture and the complex mix of leasing arrangements that is emerging in US agriculture, this study should appeal to policy makers who try to understand the impacts of government programs under different institutional organizations. We illustrate the potential biases that may arise when restricting the subsidy incidence to only cash contracts. Introducing share contracts (as well as other types of leases) into the analysis is especially important in order to understand the impact of program payments on rental rates. Most existing empirical research that analyzes the distribution of program benefit between landlords and tenants focuses on the cash rental contracts (e.g., Lence and Mishra 2003). Future studies may find it helpful to consider different types of contracts, especially hybrid contracts.

References

Allen, D.W., and D. Lueck. 2002. *The Nature of the Farm: Contract Risk, and Organization in Agriculture*. Cambridge, Massachusetts: The MIT Press.

Bourguignon, F., M. Fournier, and M. Gurgand. 2007. "Selection Bias Corrections Based on the Multinomial Logit Model: Monte Carlo Comparisons." *Journal of Economic Surveys* 21(1): 174-205.

Gardner, B.L. 2003. "U.S. Commodity Politics of Farmland Values." In C. Moss, and A. Schmitz, eds. *Government Policy and Farmland Markets—The Maintenance of Farmer Wealth.* Ames. IA: Iowa State Press, pp. 81-96.

Goodwin, B.K., A.K. Mishra, and F. Ortalo-Magne. 2003. "What's Wrong with Our Models of Agricultural Land Values?" *American Journal of Agricultural Economics* 85:744-752.

Goodwin, B.K, A.K. Mishra, and F. Ortalo-Magne. 2003. "Explaining Regional Difference in the Capitalization of Policy Benefits into Agricultural Land Values." In C. Moss, and A. Schmitz, eds. *Government Policy and Farmland Markets—The Maintenance of Farmer Wealth.* Ames. IA: Iowa State Press, pp. 3-14.

Goodwin, B.K., A.K. Mishra, and F. Ortalo-Magne. 2009. "The Buck Stops Where? The Distribution of Agricultural Subsidies." Working paper, Department of Agricultural and Resource Economics, North Carolina State University. The authors have an early version in 2005.

Hausman J. and D. McFadden. 1984. "Specification Tests for the Multinomial Logit Model." *Econometrica* 52:1219–1240.

Huffman, W.E., and R.E. Just. 2004. "Implications of Agency Theory for Optimal Land Tenure Contracts." *Economic Development and Cultural Change Working paper* 54(3):617-642.

Kirwan, E.B. 2009. "The Incidence of U.S. Agricultural Subsidies on Farmland Rental Rates." *Journal of Political Economy* 117: 138-64.

Lence, S.H., and A.K. Mishra. 2003. "The Impacts of Different Farm Programs on Cash Rents." *American Journal of Agricultural Economics* 85:753-761.

Patton, M., P. Kostov, S. McErlean, J. Moss. 2008. "Assessing the influence of direct payments on the rental value of agricultural land." *Food Policy* 33(5): 397-405.

Roberts, M.J., B. Kirwan, and J. Hopkins. 2003. "The Incidence of Government Program Payments on Agricultural Land Rents: The Challenges of Identification." *American Journal of Agricultural Economics* 85:762-769.

Schmitz, A. and R.E. Just. 2003. "The Economics and Politics of Farmland Values." In C. Moss, and A. Schmitz, eds. *Government Policy and Farmland Markets—The Maintenance of Farmer Wealth.* Ames. IA: Iowa State Press, pp. 53-80.

Sherrick B.J. and P.J. Barry. 2003. "Farmland Markets: Historical Perspectives and Contemporary Issues." In C. Moss, and A. Schmitz, eds. *Government Policy and Farmland Markets—The Maintenance of Farmer Wealth.* Ames. IA: Iowa State Press, pp. 27-49.

USDA/NASS (U.S. Department of Agriculture, National Agriculture Statistics Service). 2001. *Agricultural Economics and Land Ownership Survey (AELOS 1999).* Internet Website: http://www.nass.usda.gov/census/census97/aelos/aelos.htm

Table 1. Summary Statistics (N=60,981)

Variable	Definition	Frequency	Percentage
Cash-only	Farms use only cash leases	41,583	68.19
Share-only	Farms use only share leases	7,225	11.85
Both	Farms use both cash and share leases	12,173	19.96
Variable	**Definition**	**Mean**	**Std. Dev.**
Cash rent	Cash rental rate (exclude the LL's non-land cost share)	72.44	91.73
Share rent	Share rental rate (exclude the LL's non-land cost share)	92.03	100.31
	Historical county average payments		
Total Payments	Total government subsidies received by tenants and landlords ($/acre)	31.25	5.56
LDP	Loan deficiency payments ($/acre)	7.39	7.98
Direct	Direct payments (including countercyclical payments) ($/acre)	16.71	14.98
Disaster	Disaster payments ($/acre)	2.77	3.28
Other	All other government payments ($/acre)	4.41	5.56
Market returns	Net returns from markets ($/acre)	20.34	134.73
CV	10-year county average coefficient of variation of cash receipts from market ($/acre)	0.17	0.07

Table 2. Effects of Aggregate Subsidies on Cash Rental Rates Models (N=60,981)

Cash rent	Model 1. GLR		Model 2. Lee		Model 3. DMF(1)		Model 4. DMF(2)	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
Total Payments	0.73*	0.02	0.50*	0.02	0.37*	0.09	0.38*	0.03
Market Returns	0.12*	0.01	0.14*	6.58E-3	0.18*	0.01	0.17*	0.01
CV	-163.19*	6.76	-223.04*	8.31	-207.40*	21.48	-179.47*	10.47
Constant	74.72*	1.39	61.04*	1.52	36.36*	8.21	38.41*	2.15
σ^2	--	--	7031.54*	176.53	46303.89*	5407.42	30250.12*	2612.68
ρ_1	--	--	-0.70*	0.03	--	--	-0.17*	0.04
ρ_2	--	--	--	--	0.78*	0.20	0.42*	0.09
ρ_3	--	--	--	--	-1.15*	0.19	-1.52*	0.07

1. σ^2 equals $Var(u)$, where u is the disturbance of interest function; ρ_i (i=1, 2, 3) is the correlation coefficient between u and ($v_j - v_1$), where v is the disturbance of selection function.
2. The asterisks (*) indicate that a coefficient is significantly different from zero at 0.05 or smaller level.

Table 3. Effects of Aggregate Subsidies on Share Rental Rates Models (N=60,981)

Share rent	Model 5. GLR		Model 6. Lee		Model 7. DMF(1)		Model 8. DMF(2)	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
Total Payments	0.90*	0.06	0.92*	0.05	0.88*	0.06	0.86*	0.05
Market Returns	0.21*	0.03	0.20*	0.03	0.22*	0.04	0.23*	0.03
CV	-169.99*	18.17	-139.13*	23.29	-138.96*	25.82	-131.10*	20.02
Constant	92.14*	4.19	66.67*	9.18	48.08*	10.69	8.02	24.82
σ^2	--	--	9482.87*	784.66	12273.24*	1678.26	13485.30*	2431.05
ρ_1	--	--	-0.14*	0.06	0.39*	0.06	-0.73	0.16
ρ_2	--	--	--	--	--	--	0.32*	0.06
ρ_3	--	--	--	--	-0.62*	0.08	-1.02*	0.17

1. σ^2 equals $Var(u)$, where u is the disturbance of interest function; ρ_i (i=1, 2, 3) is the correlation coefficient between u and ($v_j - v_1$), where v is the disturbance of selection function.
2. The asterisks (*) indicate that a coefficient is significantly different from zero at 0.05 or smaller level

Table 4. Effects of Disaggregate Subsidies on Cash Rental Rates Models (N=60,981)

Cash rent	Model 9. GLR		Model 10. Lee		Model 11. DMF(1)		Model 12. DMF(2)	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
LDP	1.10*	0.09	1.07*	0.13	-1.03*	0.41	-0.35*	0.16
Direct	0.84*	0.05	0.23*	0.08	1.10*	0.21	0.85*	0.09
Disaster	-1.97*	0.18	-0.73*	0.23	-0.15	0.49	7.32E-5	0.27
Other	0.47*	0.08	1.46*	0.09	1.90*	0.24	1.98*	0.13
Market Returns	0.13*	0.01	0.14*	0.01	0.19*	0.01	0.17*	0.01
CV	-156.58*	6.87	-211.92*	11.12	-212.57*	35.11	-171.48*	12.16
Constant	78.12*	1.38	57.38*	1.69	31.24*	6.56	27.55*	2.73
σ^2	--	--	7011.32*	155.30	40783.74*	4735.78	27936.20*	2142.93
ρ_1	--	--	-0.75*	0.04	--	--	-0.23	0.05
ρ_2	--	--	--	--	0.97*	0.15	0.35*	0.10
ρ_3	--	--	--	--	-1.32*	0.17	-1.64*	0.06

1. σ^2 equals $Var(u)$, where u is the disturbance of interest function; ρ_i (i=1, 2, 3) is the correlation coefficient between u and ($v_j - v_1$), where v is the disturbance of selection function.
2. The asterisks (*) indicate that a coefficient is significantly different from zero at 0.05 or smaller level

Table 5. Effects of Disaggregate Subsidies on Share Rental Rates Models (N=60,981)

Share rent	Model 13. GLR		Model 14. Lee		Model 15. DMF(1)		Model 16. DMF(2)	
Variable	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**	**Estimate**	**Robust Std. Err.**
LDP	2.11*	0.21	1.94*	0.25	1.18*	0.32	0.95*	0.30
Direct	0.36*	0.11	0.52*	0.13	0.69*	0.14	0.77*	0.14
Disaster	-0.05	0.48	-0.60	0.6	0.22	0.62	0.70	0.54
Other	-0.59*	0.38	0.40	0.48	0.04	0.35	0.24	0.42
Market Returns	0.23*	0.03	0.22*	0.03	0.23*	0.03	0.25*	0.03
CV	-136.47*	19.47	-102.34*	19.44	-111.15*	24.63	-105.83*	20.85
Constant	94.86*	4.51	66.38*	12.23	49.03*	13.05	23.93*	24.91
σ^2	--	--	9482.87*	784.66	12424*	1818.69	14520.89*	2359.36
ρ_1	--	--	-0.14*	0.06	0.41*	0.07	-0.20	0.17
ρ_2	--	--	--	--	--	--	0.30*	0.05
ρ_3	--	--	--	--	-0.64*	0.09	-1.18*	0.20

1. σ^2 equals $Var(u)$, where u is the disturbance of interest function; ρ_i (i=1, 2, 3) is the correlation coefficient between u and ($v_j - v_1$), where v is the disturbance of selection function.
2. The asterisks (*) indicate that a coefficient is significantly different from zero at 0.05 or smaller level

Endnotes

[1] More detailed information about the calculation can be found in Bourguignon, Fournier, and Gurgand (2007) paper. An overview of sample selection models can also be found in Shadish, Cook, and Campbell (2001); Stolzenberg and Relles (1997); and Winship and Morgan (1999).

[2] For linear regression models, clustered standard errors have been used. For the selection bias correction models, we use the bootstrap standard errors. When reporting the results, we present the second-step regression equation only. Results from the multinomial logit estimated in the first-step are available upon request.

[3] Under share contracts, legislation implies that payments to producers and landlords are made according to the contract terms on a fair and equitable basis. However, landlords can always capture extra benefits by raising the share rates by charging extra payments.