The Importance of Income Risk in Labor Allocation Decisions

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Abstract

Previous research has found that on-farm income variability helps determine off-farm labor supply. However, unobserved heterogeneity of farms or regions may have biased earlier results. In this study, we use an exogenous increase in Federal crop insurance subsidies as a natural experiment to identify the importance of risk in off-farm labor supply. The subsidy increases induced greater participation in crop insurance programs and thereby reduced farmers' financial risks. By merging county-level crop insurance participation data with farm-level Agricultural Census data from 1992 and 1997 we can compare the off-farm labor decisions of individual farms before and after the subsidy and thereby control for unobserved heterogeneity. Unlike previous studies, we find that on-farm risk does not affect the labor allocation decisions of farm households.

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1. Introduction

The standard model of off-farm labor supply maintains that risk-neutral agents allocate their endowment of labor to the farm and labor market so as to equate the value marginal product of labor on-farm with the wage rate off-farm (e.g., Kerachsky, 1977; Sumner, 1982). Empirical analyses of off-farm labor supply usually involve estimating a reduced form supply equation in which off-farm labor (total household or operator and spouse separately) is a function of wages, prices, and characteristics of the utility and production functions (Goodwin and Holt, 2002; Howard and Swidinsky, 2000; Lass and Gempesaw, 1992; Huffman and Lange, 1989; Sumner, 1982; Huffman, 1980).

If farmers are risk-averse, then an increase in the variation of returns to labor lowers the certainty equivalent wage. Mishra and Goodwin (1997) show that an increase in the variation of on-farm income causes farmers with constant absolute risk aversion to increase labor supplied off-farm. In general, however, the off-farm labor response to on-farm income risk is ambiguous: with increasing absolute risk aversion, an increase in the variation of on-farm income risk may reduce labor supplied off-farm (Fabella, 1989; Hartwick, 2000).

There have been several empirical investigations of the relationship between on-farm risk and off-farm labor supply, all of which have tested whether measures of on-farm income variability help explain off-farm labor supply. Mishra and Goodwin (1997) estimate simultaneous Tobit equations for farm operator and spouse's supply of off-farm labor. Based on a cross-sectional sample of Kansas farms, they find that the coefficient of variation of gross farm income is a significant determinant of the operator's (but not the spouse's) supply of off-farm labor. In a second paper, Mishra and Goodwin (1998) estimate an off-farm labor supply equation using county-level panel data from two states. They find that that the coefficient of variation of farm income has a significant and positive effect on off-farm labor supply in both states. Kanwar (2000) uses panel data on farms in India to estimate a two-stage labor supply model. The first stage explains the operator's decision whether to supply labor off-farm, and the second stage explains the quantity of labor to supply, conditional on market participation.

Kanwar finds that standard deviation of net returns is positive and significant in the first stage, but not in the second.

Two major differences distinguish our study from previous work. First, previous studies use estimates of income variation to measure risk. To some extent, however, income variation is endogenous: farmers can adjust their income variation by altering their crop mix, applications of fertilizer and pesticides, machinery investment, labor allocation, or location. If a farmer's environment becomes riskier (e.g., prices or yields become more stable or crop insurance becomes cheaper) then a farmer may alter his behavior to maintain the same variation in income or output. Because income variation is endogenous, estimates of the relationship between income variation and labor allocation may be biased. Our approach is to examine how farmers' labor allocation decision (to work on or off the farm) changes in response to an exogenous change in the costs of bearing risk (caused by a large exogenous increase in insurance premium subsidies).

Second, in the earlier studies using cross-sectional data or panel data with no individual fixed effects, unobserved heterogeneity of both farms and regions might have biased the results. For example, unobserved factors that could affect the off-farm labor supply decision include the type of crop grown, the share of land allocated to each particular cropping activity (rather than just total farm size), the level of irrigation, climate, and agricultural and labor market characteristics. Because variation in farm income may be correlated with these unobservable factors, income risk may be spuriously correlated with off-farm labor supply.

To illustrate, consider two observationally identical farms where each farm's operator earns the same expected on-farm income. Unobservable to the econometrician, one farm grows a crop with little revenue variability that requires a lot of labor to produce, while a second farm grows a crop with greater revenue variability that requires less labor to produce (both crops provide the same expected income so presumably the labor-intensive crop uses fewer non-labor inputs). If the farmers are risk neutral, then the farmer of the risky crop should supply more labor off-farm, simply because the crop is less labor-intensive. In this case, the observed correlation between on-farm income risk and off-farm labor supply is spurious. Hence, methods that do not control for all the factors correlated with labor supply (which is often impossible in a cross-sectional analysis, since many factors are unobservable) may falsely attribute off-farm labor supply to the riskiness of farm income.

One way to control for unobserved individual heterogeneity is to examine differences in labor supply by the same farmers across time. In this study, we use an exogenous policy change as a natural experiment to identify the importance of risk in off-farm labor supply. In 1994, the Federal government passed the Federal Crop Insurance Reform Act (FCIRA), which markedly increased subsidies for premiums paid in crop insurance – fully subsidizing low levels of insurance (Catastrophic Insurance) and partially subsidizing higher levels of insurance. Congress passed FCIRA in an attempt to provide a risk management tool for farmers while weaning them off their reliance on ad-hoc disaster assistance packages. These subsidies induced greater participation in crop insurance programs and thereby reduced some farmers' financial risks. By comparing changes over time in the off-farm labor supply of farms that faced different levels of insurance subsidies before and after the policy change, we can observe whether changes in farm income risk affect off-farm labor supply, holding factors common to the farm household constant.

Data for the study originate from several sources. Information on farm labor allocation and farm characteristics is from the farm-level Agricultural Census for 1992 and 1997. We combine this information with county-level data on crop insurance participation from the Risk Management Agency of the USDA, unemployment rates from the Bureau of Labor Statistics, and average wage rates from the Bureau of Economic Analysis. Unlike previous studies, we find that changes in the risk faced by farm households have no significant effect on the labor supply decisions of these households. The findings suggest that programs that reduce grower risk, such as crop or revenue insurance, may have few implications for farm household labor allocation decisions.

2. Methods

In this study, we examine how off-farm labor supply changed following a large 1995 increase in crop insurance subsidies. These subsidies caused a large increase in insurance adoption (coverage per acre) that reduced the amount of risk faced by many farmers. We explore two empirical relationships between risk and off-farm labor: pooled cross-sectional time differences and difference in differences.

For farm operator *i* in time *t*, the desired off farm labor supply L_{it}^{*} is assumed to be a function of factors X_{it} that influence the propensity to supply labor off farm, and factors W_{it} that influence the expected return to off-farm labor.[1] The propensity to supply off-farm labor may also depend on the operator's per acre crop insurance coverage level C_{it}:

$$(1) \qquad L_{it}^{*} = \alpha + \beta_X X_{it} + \beta_W W_{it} + \gamma C_{it} + \varepsilon_{it} .$$

We use per acre coverage rather than total coverage in (1) because total coverage is simultaneously determined with labor supply – total coverage (premium per acre times total acres) depends on farm size and farm size is closely related on-farm labor demand. Hence we would expect a negative relationship between total coverage and off-farm labor supply regardless of whether risk influences labor supply.

The "continuous" variable, desired off-farm labor supply, is double-censored: on the left at zero, and on the right at the maximum possible number of workdays per year, 225. Census respondents could report that the number of days they worked off farm were 0, 1-49, 50-99, 100-149, 150-199, or more than 199. We assigned the midpoint of each interval as the observed number of days, with 225 being the midpoint of the top category. Hence, we observe:

[1] The farm operator's off-farm labor supply may be a joint decision with his or her spouse. However, we have no information about the spouse of the operator, so it is not included in the analysis.

$$(2) \qquad L_{it} = \begin{cases} 0 \text{ if } L_{it}^* \le 0 \\ L_{it}^* \text{ if } 0 < L_{it}^* < 225 \\ 225 \text{ if } 225 \le L_{it}^* \end{cases}$$

We could estimate parameters associated with the pooled cross-sectional relationship given by (1) and (2) using maximum likelihood methods.

Identification based on time difference

One problem with estimating (1) and (2) is that in any one time period, per-acre coverage is endogenous – farms make their labor allocation decisions and insurance coverage decisions simultaneously – so that a relationship between the two may not be causal. However, we can exploit the large increase in the coverage per acre between 1992 and 1997 that was caused by an *exogenous* policy change. Because the change in coverage per acre over time was large, a sizeable portion of the variation in coverage per acre in a pooled cross-section will be due to the policy change. We could therefore estimate the following relationship:

$$(3) \qquad L_{it}^* = \alpha + \beta_X X_{it} + \beta_W W_{it} + \delta Y_t + \varepsilon_{it}$$

where L_{it}^* is censored as before in (2) and where Y_t is a dummy variable equal to one in 1997 and zero in 1992. If the crop insurance policy change was the only aggregate change affecting farm labor decisions that occurred between 1992 and 1997, then the parameter on the year dummy would provide an estimate of the average effect of the increase in the insurance subsidy on off-farm labor.

Identification based on difference in differences

It could be that unobserved aggregate factors caused off-farm labor supply to change between our sample years. To address this issue, we make use of a second source of identification: different growth rates in insurance coverage across crops and farms induced by FCIRA. Specifically, we estimate the average change in labor supply of individual farms between periods in response to *changes* in the exogenous variables *X* and *W* and to changes in their coverage levels. Subtracting the first time period (t = 0) from the second (t = 1), we have:

(4) $$\Delta L_i^* = \tilde{\alpha} + \tilde{\beta}_X \Delta X_i + \tilde{\beta}_W \Delta W_i + \tilde{\gamma}\Delta C_i + \varepsilon_{it}.$$

The censoring is now more complicated as the dependent variable is a difference of a double censored variable. The desired difference in labor supplied off-farm ΔL_{it}^* is censored depending on the censoring of the desired off farm labor in time zero L_{i0}^* and time one L_{i1}^*. That is, we observe:

(5) $$\Delta L_i = \begin{cases} \Delta L_i^* \text{ if } 0 < L_{i0}^* < 225 \text{ and } 0 < L_{i1}^* < 225 \\ \text{unobserved if } \left(L_{i0}^* = 0 \text{ and } L_{i1}^* = 0\right) \text{or} \left(L_{i0}^* = 225 \text{ and } L_{i1}^* = 225\right) \\ L_{i1} - L_{i0} \text{ otherwise} \end{cases}$$

Note that the desired difference in labor supplied off-farm is censored unless the farm operator's desired supply of labor is uncensored in *both* periods. If the operator supplies no labor in both periods, or supplies the maximum possible in both periods, then the desired change in the quantity supplied cannot be observed. If the desired supply of off-farm labor is censored in either period, then the desired change in supply is censored at the level of observed change in the supply $L_{i1} - L_{i0}$. The censoring will either be right censored (if $\Delta L_i > 0$) or left censored (if $\Delta L_i < 0$).[2]

[2] For example, if $L_{i0} = 0$ and $L_{i1} = 25$, then ΔL_i^* is right censored at 25. If $L_{i0} = 25$ and $L_{i1} = 0$, then ΔL_i^* is left censored at -25.

In addition to controlling for aggregate changes that affected the labor supply of farms, differencing also controls for omitted variable bias. To illustrate, suppose the error term in (1) contains time invariant factors that are correlated with the regressors: $\varepsilon_{it} = u_i + v_{it}$. These unobservable factors may include prices of inputs and outputs, specific features and locations of the land on which the farms are situated such as soil types and climate, and characteristics and preferences of the farm operators. If the omitted variables are correlated with the regressors then OLS estimates of (1) will be biased. For example, the labor intensity of the crops grown may be positively correlated with coverage levels, which results in an inverse correlation between coverage and off-farm labor, causing γ to be biased downward. After differencing, the error no longer contains u_i so there is no longer correlation between the regressors and the error term.[3]

Estimation of the difference equation (3) takes advantage of an identifiable, exogenous source of variation in coverage – the FCIRA caused insurance coverage to increase more for some crops and regions than for other crops and regions. In general, we find a negative correlation between the level of coverage prior to FCIRA and the growth in coverage following FCIRA. This relationship makes sense: farmers who were already insured did not have to change their behavior to obtain the newly increased insurance subsidies. But those who were not previously insured had to adopt insurance in order to obtain them.

The parameter $\tilde{\gamma}$ is an unbiased estimate of the effect of the insurance policy change so long as factors *correlated with the change in insurance coverage* did not simultaneously alter crop insurance decisions. In fact, around the same time as FCIRA, another policy change occurred that might have been correlated with insurance coverage. The 1996 Federal Agricultural Improvement Reform Act (FAIR) dramatically altered the structure of agricultural income support payments. This Act, sometimes called the "freedom to farm bill," decoupled most payments from farmers' current planting decisions. Prior to FAIR, most government

[3] As an alternative to differencing, one can include fixed effects for each farm in equation 1. Due to the large number of farms in our sample and the non-linear statistical methods that we employ, this approach was computationally infeasible.

payments to farmers were tied to commodity prices, and farmers were required to limit current plantings to a share of historical plantings to qualify for payments. The FAIR Act lifted nearly all planting restrictions and decoupled payments from price levels. In effect, the Act scheduled lump-sum payments to land units based on pre-Act participation in government farm programs. If the Act caused changes in off-farm labor decisions in a way correlated with changes in insurance coverage, it could bias our estimates.

To control for the effects of the FAIR Act, we include each farm's level of 1997 government farm payments as an explanatory variable in the vector X_{it}.[4] The level of these payments was determined in advance according to parameters laid out in the FAIR Act. The larger these payments, the more a farm is engaged in pre-1996 farm programs, and the greater the effect of the policy change on income variability and insurance coverage, all else the same.

3. Data

Figure 1 shows total subsidies, total premiums, and total acres enrolled in the crop insurance program from 1990 to 1998. The figure was constructed using county-level data obtained from the Risk Management Agency (RMA) of the U.S. Department of Agriculture. The figure shows the marked increase in crop insurance subsidies beginning in 1995, the season following the FCIRA of 1994. We present separate plots for all crops and for the three largest individual crops (in acreage): corn, soybeans, and wheat. In 1997 these three crops made up 78.9% of the acreage insured, 55.5% of the subsidies, and 51.7% of the total premiums paid. As of 1997, these three crops also made up 53.8% of cultivated cropland in the U.S. (excluding hay).

For each of the ten crops that account for the most in total premiums, table 1 reports 1992 and 1997 levels of premiums, acres harvested, share of acres insured, premiums per acre harvested, premiums per insured acre, and subsidies per insured acre. These ten crops make up 85% of the premiums paid in 1997. The table illustrates the dramatic increase in premiums

[4] We do not include Conservation Reserve Program payments in government payments. In 1997, nearly all payments to farmers (net of conservation payments) were payments scheduled by the 1996 Act.

across most crops. Some crops, however, increased more than others. For barley, potatoes, and dry beans, premiums per acre harvested increased by about 1/3, whereas for wheat and sorghum the ratio increased by about ½, and for cotton, corn, soybeans the increase was almost 2/3. The most extreme cases were peanuts, which showed little increase (the crop was heavily insured before the policy change), and tobacco, for which no federal crop insurance was available in 1992.

The data obtained from RMA are county population values for crop insurance enrollment. These data, however, do not include information on production practices. We obtained data on individual farm operations from the micro files of the 1992 and 1997 Agricultural Censuses.[5] The Census micro files contain limited information about almost every farm operation in the U.S. and somewhat more detailed information, elicited in the "long form," for about one third of farm operations, aimed more heavily at large farm operations than smaller ones. We then merged all Census records from 1992 and 1997 by farm operation to obtain a panel data set. We restricted this data set to include all farms that received the long form in both 1992 and 1997 and received more than $100,000 in sales in both 1992 and 1997. Because large farms are more likely than small farms to receive the long form, these farms also are more likely to receive the long form in two consecutive censuses.

Insurance coverage

We construct a measure of the insurance coverage based on total premiums paid for the insurance. The premium includes the farmer's contribution plus the government subsidy, and should equal the premium that would be charged by a private insurance provider. The estimated total "coverage" p_{ikt} for farm i located in county k at time t is:

[5] These data are confidential, so we were required to perform our statistical analysis on site at the data laboratory of the National Agricultural Statistics Association (NASS) of the U.S. Department of Agriculture, the agency that currently administers the Agricultural Census.

$$p_{ikt} = \sum_j \frac{P_{jkt}}{a_{jkt}} a_{ijkt}$$

where P_{jkt} is the total premiums paid for crop *j* in county *k* at time *t* from RMA, a_{ijkt} is farm *i*'s area planted in crop *j* in county *k* at time *t* from the Agricultural Census, and $a_{jkt} = \sum_i a_{ijkt}$ is the total area planted in crop *j* in county *k* at time *t*.

Although it may seem that farm level insurance data would be preferred over merging county-level insurance data with individual farms as described above, there are certain benefits to our approach. Idiosyncratic variability of individual farm coverage changes that is correlated with idiosyncratic variability of changes in crop shares and or input applications could bias our regression estimates. In using county-level coverage levels, proxied by the average premium per acre harvested, we limit our source of identification to between-county variation in growth rates, which should reduce biases of this kind. In essence, county-level coverage changes serve as an instrument for farm-level changes.

Table 2 reports summary statistics for the main variables used in the analysis for 1992 and 1997. There was a slight decline in the number of days worked off-farm. Some of this decline likely resulted from the fact that operators were five years older in the second period. Note that the average coverage per acre in the sample more than doubled from $2.97/acre in 1992 to $7.05/acre in 1997. The average change in total premiums paid was $5,987.7

4. Results

Table 3 presents the results of estimating (2) and (3) the double-censored cross-sectional analysis of the days of labor supplied off-farm with the year dummy variable. The age of the operator is very significant in explaining the number of days supplied off-farm. Younger operators supply more labor off-farm than do older operators. For example, farmers younger than 30 years supply 78.9 more days off-farm compared to farmers over 60 years (the omitted category "agecat5"). Wage is statistically significant in the regression and the estimated coefficient implies that an increase of $1000 in the annual wage results in an additional 2.7 days supplied off-farm. The unemployment rate is not significantly correlated with labor supply.

The year dummy variable is not significantly different from zero, which implies that the there was no significant average change in the quantity of labor supplied off-farm between 1992 and 1997 after controlling for local wages and unemployment rates. Moreover, the standard error of the year fixed effect is small. Despite the large increase in insurance coverage for these farms between 1992 and 1997, the 95% confidence interval for the change in off-farm labor between 1992 and 1997 is (-1.89, 0.69) days, all else the same.

Table 4 presents the estimates from three specifications of (4) and (5), the difference in differences regression. Out of a possible 22,334 difference observations, 18,552 of these farms had unobservable changes in their supply of off-farm labor for both periods (because they supplied no labor in both periods, or supplied all their labor in both periods), leaving us with only 5782 observations. Of these, only 1462 were not censored (the farm supplied between 0 and 225 days of labor in both years), and the remaining were either left or right censored.

Explanatory variables in the first regression (column 1) include changes in the wages and unemployment rates, the age categories in 1992, the change in insurance coverage per acre, and government payments per acre in 1997. As shown in the table, age is the only significant determinant of the change in labor supplied off-farm – being in one of the youngest three age categories is associated with an statistically significant increase in the supply of labor off-farm. The change in the wage, the change in the unemployment rates, the change in coverage per acre, and government payments per acre are not statistically significant.

The second difference in differences regression (column 2) includes controls for lagged cropland harvested, and lagged cropland harvested interacted with the change in coverage per acre – to allow the effect of a change in coverage per acre to vary with farm size. The results of this regression are very similar to those in column 1.

The final regression (column 3) adds state fixed effects. The test of the joint significance of these fixed effects reveals they improve the fit of the model. After controlling for effects that are common to the state in which the farms are located, the change in the unemployment rate now is a small but statistically significant determinant of off farm labor supply – an increase in the unemployment rate of 1% results in a decrease in labor supplied by 0.22 days.

For all three models' specifications we find that changes in insurance coverage per acre were not statistically significant in explaining changes in labor supplied by individual operators. Furthermore, the standard errors for the estimated effects are very small. In all three difference regressions, the 95% confidence intervals for the change in days worked off farm induced by FCIRA, which increased total premiums in our sample by an average of almost $6,000 per farm, reflecting an increase in adopted insurance, lie in the range (-0.36, 0.33) days per year. Hence, the evidence strongly suggests that risk does not affect farmers' labor allocation decisions.

5. Conclusion

This paper used an exogenous increase in Federal crop insurance subsidies as a natural experiment to identify the importance of risk in off-farm labor supply. The crop insurance subsidy increase induced greater participation in crop insurance programs and consequently reduced some farmers' financial risks. By comparing changes over time in the off-farm labor supply of farms that faced different levels of insurance subsidies before and after the policy change, we observed whether changes in farm income risk affected off-farm labor supply, holding factors common to the farm household constant across time.

The study found that changes in crop insurance did not significantly influence the level of off-farm labor supply – either over time, as in the cross-section regression, or using the difference in differences regression that controls for time invariant heterogeneity. These results contradict earlier studies that found that risk, measured by historical variation in farm income, was negatively related to off-farm labor supply. It is possible that in earlier studies, using cross-sectional data or panel data with no individual fixed effects, unobserved heterogeneity of both farms and regions biased the results.

We can elaborate on these preliminary results in future work. One statistical concern is that changing prices or other unobserved factors besides FCIRA may have contributed to observed changes in crop insurance coverage. For example, coverage changes may have been caused to some degree by changes in the farm's crop mix, which depends on relative prices. To the extent that these factors affected all farms they do not bias our difference regressions. If, however, these factors affected some farms' crop-insurance adoption rates more than others, and the resulting pattern of adoption is correlated with the pattern of induced changes in off-farm labor, then our estimates could be biased. To eliminate possible biases we could include controls for prices and prices interacted with time-zero crop shares. We might also instrument the 1992-1997 change in the level of coverage with 1992 level of coverage. The 1992 coverage level provides a suitable instrument because the increase in subsidies caused farmers with little or no insurance in 1992 to increase coverage more than farmers with higher levels of coverage in 1992.

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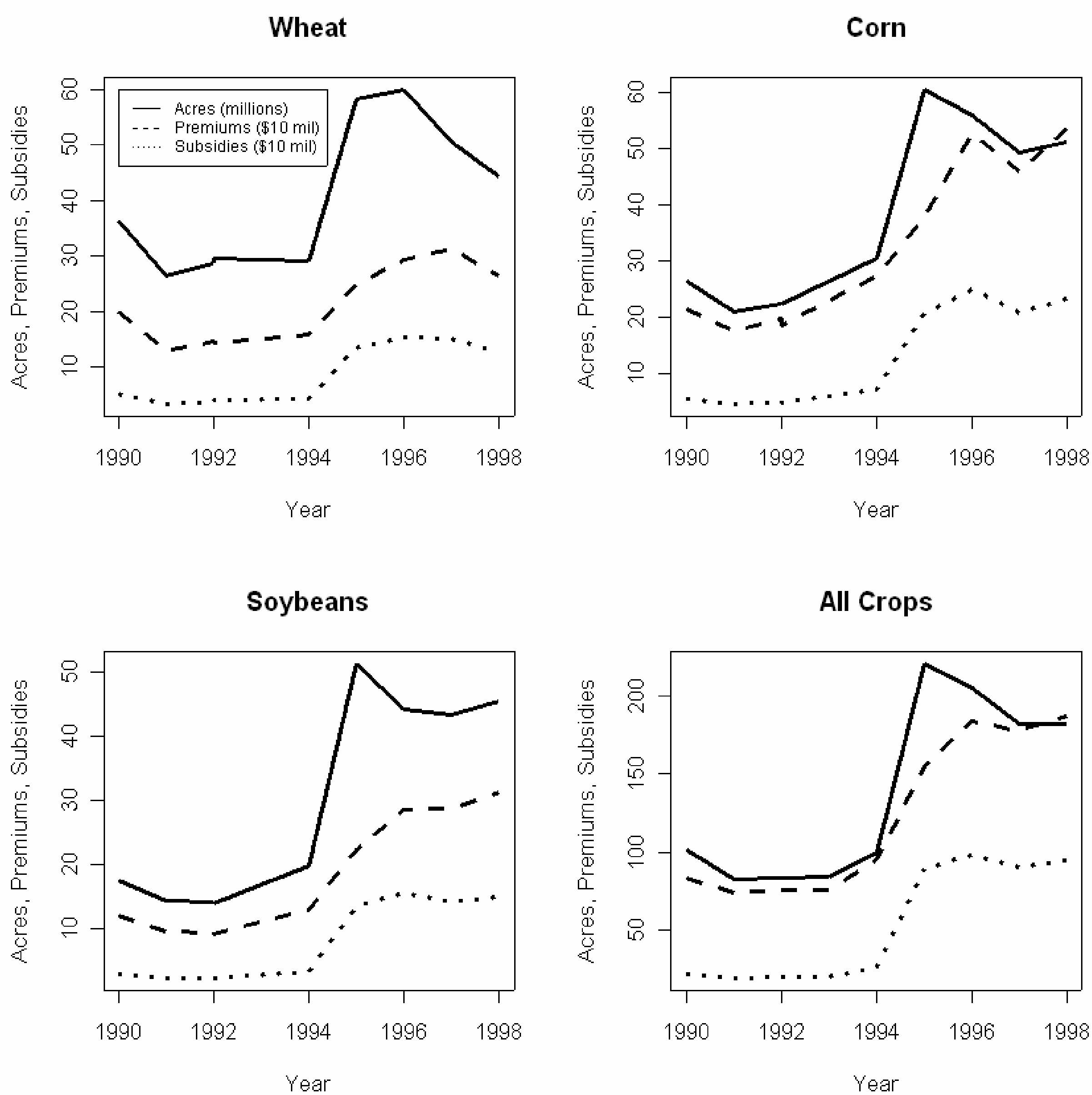
Figure 1. Insurance coverage of all crops and largest individual crops in years preceding and following the FCIRA of 1994



Source: Risk Management Agency, at http://www.rma.usda.gov/data/

Table 1. Insurance coverage before and after FCIRA of 1994

Crop	Total premiums ($1,000)		Total Acres Harvested (1,000)		Share of Acres Insured		Average Premium per Acre Harvested ($/acre)		Average Subsidy per Acre Insured ($/acre)		Average Premium per Acre Insured ($/acre)	
	1992	1997	1992	1997	1992	1997	1992	1997	1992	1997	1992	1997
Wheat	146,118	313,933	59,003	60,953	0.497	0.833	2.53	5.16	1.36	2.98	5.09	6.2
Cotton	90,657	252,676	11,742	13,787	0.371	0.835	7.86	18.36	6.22	12.84	21.21	21.98
Corn	196,412	460,662	68,905	70,371	0.327	0.702	2.87	6.55	2.23	4.18	8.78	9.34
Dry Beans	13,326	25,136	1,159	1,530	0.628	0.848	11.57	16.47	5.15	9.56	18.43	19.42
Sorghum	24,974	44,788	10,336	8,351	0.351	0.755	2.45	5.38	1.96	3.59	6.98	7.13
Peanuts	39,840	36,153	1,354	1,292	0.78	0.914	29.54	28.01	8.77	13.67	37.86	30.63
Soybeans	93,715	288,374	54,672	66,135	0.262	0.659	1.74	4.37	1.69	3.29	6.62	6.63
Potatoes	12,497	28,857	905	1,107	0.326	0.626	15.91	26.52	11.68	23.55	48.73	42.35
Barley	17,486	23,708	6,463	5,893	0.474	0.763	2.78	4.06	1.55	2.61	5.86	5.32
Tobacco	0	31,768	783	806	0	0.826	0	68.66	0	31.17	0	83.15

Source: Risk Management Agency at http://www.rma.usda.gov/data/

Table 2. Variable Definitions and Summary Statistics for Sample

Variable Name	Definition	1992 Mean	1992 Std. Dev.	1997 Mean	1997 Std. Dev.
daysoff	Number of days worked off-farm by operator. (midpoint of age-bracket on questionnaire)	20.3444162	55.788414	19.902630	55.349636
wage	Average county annual wage per job (Bureau of Economic Analysis, December 2002)	13181.89	2350.42	13538.97	2375.41
agecat1	Age less than or equal to 30 years	0.1267095	0.3326532	0.0480635	0.2139042
agecat2	Age greater than 30 and less than 40 years	0.3147995	0.4644443	0.2477835	0.4317332
agecat3	Age greater than 40 and less than 50 years	0.2888833	0.4532517	0.3224452	0.4674207
agecat4	Age greater than 50 and less than 60 years	0.1928641	0.3945543	0.2540292	0.4353219
agecat5	Age greater than 60 years	0.0767436	0.2661890	0.1276787	0.3337377
uer	County unemployment rate (Bureau of Labor Statistics, Local Area Unemployment Statistics)	6.6776501	2.9560608	4.9189324	2.6036093
totalcov_acre	Estimated premium paid for insurance including government subsidies per acre – See text for details	2.9676425	4.1036094	7.0505998	6.3337094
gov97	Total government payments in 1997, excluding Conservation Reserve Program payments	18223.75	17114.41	18223.75	17114.41
cropland_harv	Cropland harvested (acres)	1179.25	805.24659	1346.14	946.95282

Source: All variables from the Census of Agriculture, 1992 and 1997, unless specified. There were 48668 observations.

Table 3. Censored Regression – Pooled Cross Section with Year Dummy

Variable Name	Estimate	Std. Err.
Intercept	895.71944	1319.4
Year-1997 dummy	-0.59886	0.66115
wage	**0.0026990**	0.0006461
agecat1	**78.97771**	7.46617
agecat2	**72.41579**	6.29476
agecat3	**69.74501**	6.22826
agecat4	**65.06487**	6.17452
uer	-0.73444	0.55953

Bold indicates significance at the 1% level.

Dependent Variable	daysoff
Number of Observations	48668
Noncensored Values	7207
Right Censored Values	2322
Left Censored Values	39139
Log Likelihood	-66963.47

Table 4. Censored Regression – Difference in Differences under Three Alternative Specifications

	(1)		(2)		(3)	
Variable Name	**Estimate**	**Std. Err.**	**Estimate**	**Std. Err.**	**Estimate**	**Std. Err.**
Intercept	**-1.94204**	0.43418	**-1.75297**	0.50542	-6.58697	5.76600
d_wage	0.00008627	0.0001388	0.00008426	0.0001388	0.00004356	0.0001412
agecat1	**1.92464**	0.57897	**1.89946**	0.58018	**1.99694**	0.57907
agecat2	**1.75883**	0.43200	**1.73644**	0.43343	**1.68116**	0.43397
agecat3	**1.63820**	0.42530	**1.62237**	0.42587	**1.57135**	0.42617
agecat4	0.79130	0.44781	0.77669	0.44830	0.75356	0.44713
d_uer	-0.06677	0.06855	-0.06466	0.06871	**-0.22338**	0.08297
d_totalcov_acre	0.01339	0.02520	-0.0050769	0.03769	-0.0033677	0.04306
gov97_acre	-0.0053469	0.0075197	-0.0060564	0.0075980	-0.0098019	0.0078172
l_cropland_harv	-	-	-0.0001503	0.0001930	0.00003585	0.0002015
l_land_cov	-	-	0.00002116	0.00003462	-9.6098 E-6	0.00003674
state fixed effects	no		no		yes	
log likelihood	-8472.69		-8472.37		-8425.53	

Bold indicates significance at the 1% level. yes/no indicates whether model included state fixed effects.

Dependent Variable	d_daysoff
Number of Observations	5782
Noncensored Values	1462
Right Censored Values	2014
Left Censored Values	2306