# MODELING CHANGES IN THE U.S. DEMAND FOR CROP INSURANCE DURING THE 1990S

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***Paper prepared for presentation at the American Agricultural Economics Association Annual Meeting, Montreal, Canada, July 27-30, 2003***



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## Introduction

The U.S. federal crop insurance program has been an important agricultural policy instrument since the 1930s. During the 1990s, the program underwent considerable changes. In particular, the introduction of new revenue insurance products in the late-1990s coupled with a substantial increase in government subsidies constituted rather profound changes in the operation of the programs. A large literature has attempted to estimate the extent to which the demand for insurance is sensitive to the price of coverage (i.e., the premium). Although most studies have implied rather inelastic responses, these studies are all based upon what was really a very different program than what currently exists. In particular, the introduction of catastrophic (CAT) insurance in 1995 and the subsequent introduction of a number of revenue insurance programs throughout the late 1990s have resulted in a very different insurance program from what was in existence prior to this period.

U.S. crop insurance programs traditionally based protection and indemnities on yields alone. A given percentage (which was selected from a range of available options) of an estimate of a producer's average yield could be insured. Indemnities would be paid at a predetermined price if realized yields fell beneath the guaranteed level. In the late 1990s, a number of new insurance products that provided revenue protection were introduced. These products include Crop Revenue Coverage (CRC), Revenue Assurance (RA), and Income Protection (IP). Though each of these programs has its own unique characteristics, the revenue plans all function to guarantee a given level of crop revenue and thus pay indemnities on the basis of low yields and/or low prices. The addition of alternative crop insurance products was a result of both a mandate of the 1994 Federal Crop Insurance Reform Act to develop a cost of production insurance pilot and the development of insurance products by private companies that would base indemnity payments on harvest time prices. Revenue insurance products were launched in 1996 through restricted pilot programs and became widely available in the following years.

The federal government has long provided generous subsidies to producers that paid a proportion of the overall insurance premium. In 1999 and 2000, the U.S. government significantly increased these crop insurance premium subsidies. Though the increases affected all products and coverage levels, some aspects of the subsidy increases were especially beneficial to revenue insurance products and coverage at higher guarantee levels.[1] Substantial premium discounts were also added to the already existing premium subsidies. With the 2000 Agricultural Risk Protection Act, premium subsidies expanded to cover, not only the yield portion of the premium of the revenue insurance, but the whole premium, which essentially provided an additional subsidy for revenue coverage.

In addition to increases in premium subsidies, crop insurance programs were significantly affected by the 1994 Crop Insurance Reform Act. Among other things, the Act contained a provision that required farmers to buy at least a minimum level of insurance (the CAT policies) in order to be eligible for any other farm program benefits. The intent was to eliminate reliance on ad hoc disaster relief and to instead encourage farmers to obtain their protection against widespread yield shortfalls from the crop insurance program. This requirement proved to be unpopular among farmers and thus was rescinded in the 1996 Farm Bill, although the provision of CAT (which is essentially provided free, excepting a small administrative charge) was continued.

Program changes during the nineties are likely to have had considerable effects on the demand for crop insurance. First, relevant increases in liability coverage have been realized. Second, there has also been a strong shift in coverage toward revenue insurance products (Goodwin 2001), indicating that many operators may find revenue insurance more attractive than yield insurance alone.

Comprehending and understanding the demand for crop insurance is important for forecasting the potential growth of the insurance market and to better understand and predict the actuarial performance of the crop insurance program. In particular, knowledge of how producers may respond to premium rate or subsidy changes is important and an understanding of the demand for crop insurance plays an essential role.

[1] Over its history, most insurance was purchased at the 65% yield election level. In recent years, however, "buy-up" coverage at levels of 75-85% of the guaranteed yield has become more popular, in part due to increased premium subsidies at those coverage levels.

Most of the empirical research that has investigated the demand for crop insurance has utilized aggregate (county or state level) data. This approach, while providing valuable inferences regarding the demand for insurance, may obscure many important characteristics of demand. In addition, as we have noted above, nearly all of the studies on the demand for insurance have addressed participation issues prior to the 1994 Reform Act and prior to the introduction of revenue coverage.[2] It is likely that the introduction of these new revenue products may have substantially shifted the overall demand for insurance. In addition, the overall farm policy environment underwent substantial changes with the passage of the 1996 FAIR Act, with producers being afforded much greater planting flexibility. In addition to a direct increase in the variability of planted acreage, some have argued that this planting flexibility may have increased the volatility of market prices, thus raising revenue risk (Ray et al. 1998). To the extent that such changes may have occurred, the demand for revenue products may have changed.

The objective of this analysis is to consider the demand for insurance during the 1990s—a period that experienced many changes in the policy environment. Farm-level data are utilized to study the evolution of the participation in the U.S. crop insurance program. Specifically, the analysis focuses on the probability that an individual farm purchases crop insurance products during each year of the period from 1993 to 2000, using a large sample of commercial Kansas farms.

## Econometric methods

Following previous analyses on the demand for crop insurance, we assume that insurance purchases will be influenced by a number of farm characteristics, as well as by parameters of the

[2] Barnett, Skees, and Hourigan (1990), Goodwin (1993) and Cannon and Barnett (1995) are examples of econometric analyses of crop insurance participation prior to the 1994 Reform Act that use aggregate data. Just and Calvin (1990), Calvin (1992), Smith and Baquet (1996), Smith and Goodwin (1996), Coble et al. (1996) and Just, Calvin and Quiggin (1999) analyze pre-1994 crop insurance participation through econometric studies based on farm-level data (see Knight and Coble (1997) for a thorough review of the literature on the demand for Multi Peril Crop Insurance). Goodwin (2001) carries out an econometric analysis of the demand for crop insurance after 1994 using aggregate data.

insurance program. As we explain in some detail below, we are only able to observe the discrete purchase decision for farms in our sample. If we define $y_i$ to be the (latent) desired level of liability coverage for farm i, we will observe a discrete indicator corresponding to the sign of $y_i$. In particular, we observe $y_i^* = 1$ if $y_i > 0$ and 0 otherwise. To model the likelihood that crop or revenue insurance will be purchased by an individual farm, we adopt a standard probit specification.

An additional complication is present in our analysis, however. It has been established elsewhere (Smith and Goodwin 1996) that crop insurance purchase decisions are likely to be made jointly with other production decisions. In particular, it has been shown that production methods (which we represent using fertilizer and agricultural chemical expenditures per acre) may be endogenous to the crop insurance purchase decision.[3] Smith and Goodwin found that farms that purchased insurance actually tended to use less fertilizer and agricultural chemicals, suggesting a form of moral hazard whereby producers with insurance have different production practices from those that do not insure. This suggests the need for a simultaneous equations version of the probit model, whereby fertilizer and chemical expenditures, a potentially relevant right hand side variable, is allowed to be endogenous to the insurance purchase decision.

The insurance demand indicator is discrete, taking a value of 1 if crop insurance is purchased and zero otherwise. Nelson and Olson (1978) have proposed a two-stage estimation procedure that is appropriate in such cases. Their estimator is utilized here to estimate the parameters of the model. As Amemiya (1979) and Lee, Maddala and Trost (1980) have noted, this two-stage method does not yield reliable estimates of the variances of the parameters since it ignores the fact that the right-hand side endogenous variable is represented using instrumental variables (i.e., standard variance estimates ignore the estimation of the reduced-form equations in the first stage). In light of this problem and following Efron (1979), Monte Carlo bootstrapping procedures are used here to derive consistent variance-covariance estimates for the parameters of the model. We utilize 1,000 pseudo-samples of the same size

[3] As Smith and Goodwin (1996) note, this simultaneity does not require the actions to be contemporaneous. It only requires simultaneous planning of both the decision to purchase insurance and the decision to apply chemicals.

as the actual sample, drawn with replacement, to provide a sample of parameter estimates from which we estimate the parameter covariance matrix. For each pseudo-sample of data, Nelson and Olson's two-step method is applied to estimate the parameters of the model. The covariance matrices are derived from the distribution of the replicated estimates generated in the bootstrap process.

After estimating the parameters of the model and their respective covariance matrices, the Wu-Hausman specification test is used to evaluate the exogeneity of the fertilizer and chemical expenditures variables in the insurance equation against the alternative of endogeneity. The test statistic is given by:

$$q = (\beta_0 - \beta_I)'\left[V(\beta_I) - V(\beta_0)\right]^{-1}(\beta_0 - \beta_I)$$

where $\beta_0$ is a vector of the parameter estimates derived from independent estimation of the first and second equation, $\beta_I$ is a vector of the parameter estimates derived from Nelson and Olson's two-step method, and V(.) is a covariance matrix of the parameters. The Wu-Hausman test is distributed as a $\chi^2$ with degrees of freedom equal to the number of coefficients evaluated.

**Data and Empirical Application**

The data used in this analysis are drawn from three sources. First, annual data for individual farms in Kansas observed between 1993 and 2000 are obtained from farm account records from the Kansas Farm Management Association. Though the analysis is based on individual farm data, county aggregates are needed to define several important variables that do not appear in the Kansas data set. These county aggregates are derived from the National Agricultural Statistics Service (NASS) data and Risk Management Agency (RMA) data.

Using these sources, the variables used to estimate the system of equations are constructed. Our selection of relevant variables follows a number of earlier analyses of the demand for crop insurance (see for example Just and Calvin (1990), Horowitz and Lichtenberg (1993), Goodwin (1993), Smith and

Baquet (1996), Smith and Goodwin (1996) and Goodwin (2001)).[4] The dependent variable is a discrete indicator of participation in crop insurance programs.[5] Specifically, this is a discrete dependent variable equal to 1 if the farm insures any of its crop production in year t and equal to zero otherwise.[6] Our explanatory variables include variables conceptually relevant to the crop insurance purchase decision. We consider the influence of the lagged expenses in crop insurance on the dependent variable. Lagged crop insurance expenses may be highly correlated with farmers' degree of risk aversion and thus with the likelihood that they purchase crop insurance over the period studied. We include a measure of input usage, defined as the annual expenses on fertilizer and herbicide per crop acre. The farmer's initial wealth is also considered. We define initial wealth as the lagged value of farms' net worth. Wealthier households may be less risk adverse relative to poorer ones, which may reduce their incentives to purchase crop insurance. Crop acres are defined as the total acres planted to wheat, corn, sorghum and soybeans, the predominant crops in Kansas. Total crop acres, a scale indicator, may be relevant to producers' yields and hence their insurance premium rates and insurance purchases. A farm's size might also influence the marketing effort carried out by insurance sellers working on commission, thus affecting the likelihood farms purchase crop insurance. The expected market net income per crop acre is included as a measure of the amount of farm revenue at risk. We define this variable as the average value of the difference between farms' income and farms' operating expenses over the preceding 4 years.

[4] Because the main interest of this paper is to analyze changes in the demand for crop insurance, details of the variables used to evaluate the input usage equation are not given here. However, this information is available from the authors upon request.

[5] In light of the recent changes to the crop insurance programs, Goodwin (2001) has proposed alternative methods to refine the way insurance participation is measured. Specifically, Goodwin proposes to use the ratio of actual liability to possible liability. It should be noted here that the lack of information on specific insurance purchases in the present analysis does not allow us to define insurance participation according to the previous proposal. In particular, liability data are not available from the Kansas Farm Management Association's Data Bank.

[6] It should be noted here that Kansas Farm Management Association's Data Bank only defines a single aggregate variable that represents total expenditures on crop insurance realized by each individual farm, without distinguishing between different crops, coverage levels and insurance plans. This substantially limits our modeling of insurance purchases in that we only observe the discrete purchase decision.

We face particular challenges in defining the relevant premium rate for an individual farm. We are unable to measure this directly, though we are able to accurately measure the factors that determine this premium rate. In particular, the premium rate paid by an individual farm depends on the county-wide risk facing producers of the crop in each particular county and the average (APH) yield for the individual farm (which affects the individual farm's premium rate in an inverse fashion).[7] It is important to note that all farmers with an identical average yield for a particular crop and practice combination in a county will pay the same premium rate. However, as a farmer's average yields fall, the corresponding premium rate rises. To have a measure of the cost of insurance, the county average (farmer-paid) premium rate over all crops and insurance plans is used.[8] This variable is expressed as the ratio of the total premium minus government subsidies to the total liability, using data of the Risk Management Agency.[9]

An interaction term comprised of the county average premium rate and the mean normalized yields for each farm is included to account for the fact that this county-wide premium rate varies for an individual farm according to the farm's average yield.[10] Because we are working with multiple crops, any or all of which may be insured, we represent this effect by taking an average of the normalized yield across all crops on a farm, where the normalized yield is the farm's yield divided by the county-average yield. Such a metric is comparable across crops and farms and thus serves to adjust the overall premium rate for differences in an individual farm's production performance. County average yields are taken from the National Agricultural Statistics Service (NASS) data.

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[7] Premium rates in the federal program are set at the county level, and then are adjusted for an individual based on the farm's average yield. Higher average yields result in lower premium rates.

[8] Recall that Kansas Farm Management Association Data Bank only provides a single measure for crop insurance purchases, without distinguishing between crops and insurance plans.

[9] The Kansas data set only provides information about expenditures for crop insurance and not on total liability. Hence, it does not allow one to compute premium rates.

[10] We acknowledge that a problem inherent to working with aggregate data is that the farm's normalized yield may not fully correspond to the county average premium rate in situations where there are differences between the crop mix on the farm and the county.

A second interaction term comprised of the previous variable (the first interaction term) and the mean coefficient of variation of yields over the preceding 10 years expressed as a percentage is also considered. We expect that the coefficient of variation will be highly correlated with yield risk and thus will be a good indicator of the expected loss ratio (expected returns to insurance). This allows us to analyze adverse selection in crop insurance. In particular, it allows us to discern whether farmers with more variable (risky) yields tend to have less elastic responses to premium rate changes (as theory would suggest). A similar specification was considered in Goodwin (1993), Smith and Goodwin (1996), and Goodwin (2001). The mean normalized yields are also included as an explanatory variable. In addition, to approximate the producer's risk level, the mean coefficient of variation of yields over the preceding 10 years is used.[11]

Some additional elaboration on the specification of the model may be helpful. With the exception of the discrete purchase decision, other parameters of the crop insurance program that are relevant to demand cannot be directly observed. In particular, premium rates that individual farmers must pay to purchase coverage cannot be observed. We represent this effect by considering a series of interaction terms representing the factors that underlie premium rates and aggregate premium rates at the county level. In particular, rates are set at the county level on the basis of previous loss experience, which is captured using county-level average premium rates. In addition, rates vary inversely with farm's average yield, which is represented using the farm's mean normalized yields. Individual farms may have considerably different yield risk, which is represented using the coefficient of variation of the farm's historical yields. Interactions of these various factors are included in the empirical models to capture the effects of premium rates on participation, which are influenced by the overall rate, the farmer's average yield, and the amount of variation experienced by an individual farmer.

The risk environment of the farm is not only represented by the variability of a farm's yields. Other variables are also included. A measure of the diversification of total sales between crops and

[11] Recall here that, for insurance purposes, individual coverage levels and premium rates are computed based on the insured expected yield. The expected yield is usually calculated as the average yield over the preceding 4-10 years.

livestock is incorporated. This is represented using the share of total farm sales derived from livestock sales (livestock ratio). As Featherstone and Goodwin (1993) have recognized, relative crop yields may be lower on farms that diversify across crops and livestock, as a result of limited gains from specialization. Likewise, this may reflect an individual farm's comparative advantage—farms with riskier or lower yields may be more likely to specialize in livestock. Diversification may also reflect a lower degree of income risk and thus may affect the demand for crop insurance. The ratio of irrigated to non-irrigated crop acres is also taken into account. The crop production practice is included in the analysis as it is likely to affect the riskiness of an individual farm as well as the insurance premium faced by the farmer (premium rates vary by practice). We also include the ratio of the farm's debts to assets. This variable is included because creditors may impose an obligation to purchase crop insurance and thus more highly leveraged producers may face stronger pressures to insure. In addition, the leverage ratio is a relevant indicator of the overall financial risk a farm faces.

Summary statistics for the variables of interest are presented in table 1. The evolution of participation in crop insurance programs during the period of study seems to reflect two main changes. First, an important increase in crop insurance demand in 1994 and 1995 is apparent. This period corresponds to the implementation of the 1994 Federal Crop Insurance Reform Act and its provisions for mandatory participation in crop insurance programs.[12] Second, an increase in participation toward the end of the 1990s is also reflected in the statistics. This increase is consistent with the attractiveness of the new revenue insurance products and the increase in government subsidies that occurred near the end of the 1990s. Lagged crop insurance expenses register important increases until 1998 and a slight decrease by the end of the decade. The increase in crop insurance expenses may be suggesting that producers have tended to shift to higher coverage levels and to revenue insurance products as they have become

[12] As noted above, with the 1996 farm bill, producers who agreed to forego any disaster assistance were allowed to opt out of CAT insurance. As a result, CAT coverage fell considerably across most of the U.S. in 1996. However, as an anonymous referee notes, this was not the case in Kansas because, at the time the bill was passed (April 1996),

available.[13] The slight decrease in crop insurance expenditures by the end of the decade may have some connection with the substantial increase in crop insurance premium subsidies in 1999 and 2000. An increase in input usage from 1993 to 1997 also occurred. This may reflect changes in production practices or more intensive use of land. However, input usage fell back by the end of the 1990s, which is coincidental with the increase in crop insurance demand and is consistent with Smith and Goodwin finding that farms that purchase insurance tend to use less agricultural chemicals. Farms' average wealth also increased during the period of study. Figures in Table 1 indicate an increase in average crop acres planted to wheat, corn, sorghum and soybeans. Average net income per crop acre over the preceding four years (expected net income) fluctuated considerably during the period analyzed.

County average premium rates register an important decline in 1995 that likely reflects the introduction of the compulsory CAT coverage. As participation in CAT falls, average county premium rates increase. Premium rates register another decline by the end of the decade (table 1). This is consistent with the previously mentioned increase in premium subsidies that occurred near the end of the 1990s. Individual mean yields seem to be similar to county yields in that the mean normalized yields are quite close to one. This suggests that the farms in our sample, at least in terms of production performance, are representative of the county as a whole. The coefficient of variation of yields fluctuates considerably, with values suggesting that the standard deviation of these yields is always above one-fourth of their mean.

An important fluctuation of the livestock ratio is suggested in table 1 (its value ranges from a low 0.19 in 1998 to a high 0.36 in 1994). Summary statistics also show a decrease of the ratio of irrigated over nonirrigated crop acres. This decrease may be consistent with changes in the distribution of crop acres, especially after 1996, which were consistent with the planting flexibility provisions of the 1996 FAIR Act

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Kansas farmers had already signed up for CAT for winter wheat and other crops planted in fall. As a result, the decline in CAT in Kansas occurred in 1997.

and the expansion of new short-season crop varieties, especially for corn, that made the production of certain crops more attractive, thus expanding their acreage. The level of debt held by the farms studied slightly decreased during the period of study.

**Empirical results**

Over the period analyzed, average crop insurance expenses per farm increased considerably from around $1,250 in 1993 to $2,850 in 2000.[14] Such a change likely reflects increased participation in crop insurance programs, as well as participation at higher coverage levels.[15] Results derived from this research help to better understand the evolution of these figures. For each year in the sample, a simultaneous equation probit model of the demand for insurance is estimated. A simultaneous equation probit model is also estimated for the entire panel of data.[16] Parameter estimates and summary statistics for the crop insurance participation equations are presented in table 2 for yearly models and in table 3 for the pooled data model.[17] Estimations show that farms that use more chemical inputs are less likely to purchase crop insurance (estimated coefficients representing input usage are negative and significant in most cases). Hence, results suggest that the application of chemical inputs reduces the expected return from crop insurance and thus reduces the probability that a farm insures its crops. This conclusion is consistent with the findings of previous research (see for example Smith and Goodwin (1996) and

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[13] In fact, the distribution of crop insurance expenditures suggests such a shift, with both the mean and quartiles increasing substantially over time (though these statistics are not presented here, they are available from the authors upon request).

[14] These values are expressed in constant 1993 dollars.

[15] Though an increase in crop insurance expenses could also reflect a higher cost of insurance, premium rates actually paid by producers declined over the period analyzed. In 1993 while the average base premium was 0.08, the premium actually paid equaled 0.057. In the year 2000, the base premium was 0.088 and the actual premium 0.042. Hence, although the base premium rates increased over the period analyzed, the increase in subsidies allowed a reduction in the actual premium rates paid by producers.

[16] We incorporate annual dummy variables in this model, in order to capture fixed annual effects.

[17] Results for the chemicals usage equation are available from the authors upon request.

Goodwin (2001)). This likely represents a moral hazard effect, whereby farmers may substitute chemical inputs for insurance and thus modify their tillage practices.

Contrary to what was expected and has been found in other analyses, the results of the Wu-Hausman test mainly suggest that input usage is an exogenous variable in the annual crop insurance equations (table 2). Hence, the test indicates that the parameter estimates may be less efficient than the ones derived from independent estimation of the equations of the model. This may also reflect a lower degree of statistical efficiency in the two-stage estimates, since we do not jointly estimate an equation for input usage.[18] However, when the model is estimated using the pooled set of data, endogeneity is implied by the Wu-Hausman test (table 3).

Results show that those farms that spent more money in insuring their crops over the last year are more likely to insure their production in the present year (tables 2 and 3). Parameters representing lagged crop insurance expenses are positive and a majority are statistically significant. As expected, results suggest that wealthier farms are less likely to insure. This may be indicating that farms' risk preferences are of the sort where risk aversion decreases as farms' wealth holdings increase. Parameter estimates indicate that farms with more crop acres are more likely to purchase crop insurance. Coefficients representing total planted acres are all positive though only a minority are statistically significant. Other analyses have also found a positive influence of farm size on crop insurance participation (see for example Barnett, Skees, and Hourigan (1990), Goodwin (1993) and Cannon and Barnett (1995)). This result may reflect the greater marketing effort concentrated by insurance marketers on larger farms (Goodwin 2001). Likewise, larger farms may have other scale advantages that make crop insurance more attractive. Coefficients representing the expected net farm income per acre indicate that net income has a negative effect on crop insurance purchases. A majority of these coefficients are statistically significant. A higher degree of income may make farms more likely to self-insure. This result is consistent with what would be expected if richer farms were less risk averse than smaller ones.

[18] Such joint estimation in systems with mixed continuous and discrete dependent variables is complex and thus was not undertaken here. This does serve as an important avenue for future research.

Parameter estimates representing county average premium rates suggest, with the exception of 1994, that higher premium rates correspond to a lower demand for crop insurance (tables 2 and 3). Although premium rates seem to exert a statistically significant effect on the demand for crop insurance when panel data are used (table 3), it should be acknowledged that the parameters in the annual models are, with a sole exception, not statistically significant. However, it should also be noted that these coefficients do not represent the overall effect of premium rates on the crop insurance participation decision (interaction terms of the premium rate with yields and their coefficient of variation are also relevant to the effect of premium rates). In order to calculate the overall effect, the elasticity of crop insurance purchases with respect to the premium rate is computed as the product of the net marginal effect of the premium rate on participation in crop insurance programs and the ratio of the county average premium rates to participation in crop insurance programs.[19] The standard error of the elasticity is derived using the replicated elasticity estimates from the bootstrapped samples. Results are presented in table 4.

Results suggest that price elasticity evaluated at the data means is negative and statistically significant for all years with the exception of the 1996 equation, which was negative but not statistically significant. As it has been noted above, the 1995-96 period corresponds to the implementation the 1994 Federal Crop Insurance Reform Act, which included provisions in these years for mandatory participation. Thus, it is not surprising that, in light of mandatory participation requirements, participation was not sensitive to premium rates. Results also indicate a reduction in the price elasticity toward the end of the decade, a finding that may reflect the considerable increase in government subsidies by the end of the 1990s.[20] In 2000, the elasticity of demand had fallen to -0.13. This result may also reflect the changing nature of crop insurance demand that came about as revenue insurance products were

[19] Variables are measured at their mean values.

[20] It should be noted here that, as in most empirical studies that assess market conditions, many competing hypotheses may underlie the revealed behavior implied by our estimates. In particular, apart from policy changes, low yields in one year may also increase insurance purchases in the subsequent year.

introduced.[21] In particular, the attractiveness of such products may have made producers less sensitive to premium changes.[22] Such an effect is expected and represents one of the key findings of our analysis—producers have become less responsive to premium rate changes as new products have been introduced. Such a finding is confirmed by an examination of unpublished summary of business data from the Risk Management Agency. The total amount of premium subsidy increased from $942 million to nearly $1.7 billion between 2000 and 2001. Yet, total net insured acres fell from almost 205 million acres to under 203 million acres. Liability remained relatively constant over this period, rising from $34.2 to $34.7 billion.[23] The overall implication is that the increase in premium subsidies that occurred between 2000 and 2001 did not evoke an especially large increase in participation, at least as measured through total liability. Through the 1990s, as revenue insurance and expanded opportunities for higher coverage levels increased, total liability rose substantially (up from $25.2 billion in 1997).[24]

The values reported for the price elasticity of the demand for crop insurance are largely consistent with results presented in other analyses. The pre-1994 Reform Act elasticity, -0.57 in 1993, is comparable to the values presented in other studies that use pre-1994 farm-level data. Coble et al. (1996) obtain a crop insurance participation elasticity of -0.65 for a group of Kansas wheat farms. Smith and Baquet (1996) derive an average price elasticity of liability per planted acre of -0.6 for a sample of Montana wheat farms. Goodwin and Kastens (1993) obtain a price elasticity of liability per planted acre equal to -0.5 for a group of Kansas wheat farms. For the period after 1996, once the mandatory participation to crop

[21] It should be noted here that, as an anonymous referee points out, the lower elasticity derived for the end of the 1990s may also reflect the fact that we are analyzing the participation in crop insurance programs strictly as a yes/no proposition. However, if coverage level data was available and used in the analysis, the response to increased subsidies would likely be more elastic.

[22] We should recall here that our database does not allow to distinguish between revenue and yield insurance programs. Hence, results should be interpreted with care.

[23] We should note here that crop price declines for crops relevant to Kansas may have also affected liability by holding it down.

[24] These figures were taken from the Risk Management Agency's Summary of Business Statistics, available from the RMA website (www.rma.usda.gov).

insurance programs was lifted, elasticity values reported are clearly below the pre 1994 Reform Act period. This decrease is also suggested when comparing insurance demand elasticity values computed in Goodwin (1993) and Goodwin (2001) using aggregate data.

Annual coefficients representing the adjustment of premium rates to individual yields are mainly positive (tables 2 and 3). Positive coefficients are contrary to what would be expected if low yield farms are truly subject to higher levels of yield variation. If this were the case, these farms would be expected to be less responsive to premium rate increases. However, it should be noted that the coefficients for this interaction term are rarely significant, which may suggest that premium rates adequately incorporate relative risk.

The estimated coefficients for the interaction of premium rates, yields and their coefficient of variation are mainly positive, indicating that higher expected returns to insurance reduce the demand price elasticity (table 2). The coefficient in the panel data model is also positive (table 3). However, it should be acknowledged that this interaction term is not statistically significant in any equation. This may suggest that adverse selection is not a strong problem among farms in this sample, though it may be difficult to adequately capture this result using our proxy measures of premium rates. Contrary to this result, several analyses have provided empirical evidence of adverse selection in the demand for crop insurance. Barnett, Skees, and Hourigan (1990), Goodwin (1993 and 2001) and Cannon and Barnett (1995) have shown, using aggregate data, that the expected rate of return to insurance can increase participation. Evidence of adverse selection using farm-level data has also been given (see for example Just and Calvin (1990), Coble et al. (1996), Smith and Goodwin (1996) and Just, Calvin and Quiggin (1999)).

Coefficients estimated for individual yields are, with a sole exception, not statistically significant (tables 2 and 3). A majority of the parameters are negative, including the coefficient in the pooled data model, which indicates that farms with higher yields are less likely to insure their production. This is consistent with the potential adverse selection problem suggested by the second interaction term. The results show that, contrary to the findings in other analyses (see for example Calvin (1992) and Goodwin

(2001)), farms with more variable yields are not necessarily more likely to insure. The variable representing yield variability is mainly positive, but is not statistically significant in any equation. It is important to also note, however, that such variability may exert an effect through its interaction with premiums.

Coefficients estimated for the livestock ratio variable are always negative (tables 2 and 3), indicating that farms with more livestock production are less likely to insure. This result is consistent with the results in Barnett, Skees, and Hourigan (1990), Calvin (1992), Cannon and Barnett (1995) and Goodwin (2001). This may reflect the greater diversification associated with such farms, which may lower income risk (business risk). Livestock ratio coefficients, however, are not always statistically significant. Coefficients representing irrigation are, with a couple of exceptions, not statistically significant. Parameter signs are mainly positive, contrary to what would be expected if the irrigation practice actually reduced production riskiness.[25] This could be, however, entirely consistent with the fact that premium rates tend to be lower for irrigated production, thus confounding precise identification of the premium response in counties with a large degree of irrigated production. It is also possible that such farms have much more at stake in terms of the value of production to be lost due to a weather event (outside of precipitation shortfalls) or a failure of irrigation sources. A majority of coefficients estimated for the debt to assets ratio are positive, indicating that those farms with more debt are more likely to purchase crop insurance. However, parameters are rarely statistically significant. Previous analyses have found evidence that more indebted farms are more likely to participate in crop insurance programs (see for example Smith and Baquet (1996) and Smith and Goodwin (1996)). This likely reflects the pressures of lenders on highly leveraged borrowers to purchase insurance.

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[25] Barnett, Skees, and Hourigan (1990) find the percentage of county acreage irrigated to have a positive and significant effect on crop insurance participation, Cannon and Barnett (1995) report a negative effect, while Goodwin (1993) finds this variable to be statistically insignificant.

## Concluding remarks

Crop insurance purchase decisions during the 1990s were analyzed using individual farm data. The central question underlying our analysis pertains to whether the elasticity of demand for crop insurance has changed in recent years in response to substantial changes in the insurance programs. The results largely confirm the findings of existing research on the demand for crop insurance (see for example Goodwin (1993 and 2001), Smith and Goodwin (1996), Smith and Baquet (1996)). In particular, the results indicate that the insurance purchase decision responds in an inelastic fashion to changes in insurance premium rates.

Perhaps of greater interest is our finding that demand elasticities have undergone substantial changes over time as the programs have changed. The results suggest that the demand elasticity has experienced two main changes in the period analyzed. First, during 1995 and 1996, it registered an important drop in absolute value to the extent that, in 1996, it was not statistically different from zero. This can be explained by the implementation of mandatory participation in the crop insurance program during these years. Second, the results also indicate a reduction in the price elasticity by the end of the decade, which is coincidental with considerable increases in government premium subsidies toward the end of the 1990s. In addition, this may reflect the introduction of new revenue insurance products, which may have shifted the demand for crop insurance and made insurance more attractive and thus less price responsive.

The reduced demand elasticity suggests that further increases in crop insurance participation may be hard to achieve through premium subsidies or premium discounts. The design of more attractive products may play a more relevant role in increasing future demand for crop insurance. The 2000 ARPA Act contained a number of incentives for the introduction of new insurance products and thus one would anticipate that demand conditions may continue to change as new products become available.

Apart from premium rates, the variables that are more important in explaining crop insurance purchases are chemical input use (negative effect), expected farm net income per crop acre (negative), farms' wealth holdings (negative), the lagged value of crop insurance expenses (positive) and the relevance of the farm's livestock production (negative effect).

Results confirm that insurance participation may be negatively related to production input usage, a result consistent with moral hazard. Higher expected net incomes reduce the demand for crop insurance, which is in accordance with richer farms being less risk adverse than poorer ones. Farms' wealth holdings exert a negative influence on program participation, which is consistent with farms' preferences being of the sort where risk aversion decreases as farms' wealth increases. The lagged value of crop insurance expenses, a variable likely to be positively correlated with farms' risk aversion, positively influences the demand for crop insurance. Consistent with previous findings, the results show that livestock farms are less likely to insure. The significance of livestock sales and the expected net income in explaining the demand for crop insurance seems to diminish in the second half of the 1990s decade. This result is in accordance with the registered increase in demand motivated by changes in the U.S. crop insurance program.

The demand elasticity for insurance is a critical parameter for assessing the effects of policy changes on the actuarial performance and overall operation of the federal insurance program. An important implication of our analysis is that the elasticity estimates that policy analysts commonly work with are likely to be too high (in absolute value). These estimates are generally based upon what can only be described as a very different crop insurance program. Recent changes in the operation of the program, including the introduction of revenue insurance, seem to have had a substantial impact on the elasticity of demand. Policy evaluations should give consideration to the implications of smaller demand elasticities, which may better characterize the responsiveness of farmers to premium changes in the current policy environment.

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Table 1. Variable Summary statistics

| | Mean (Standard Deviation) | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | **1993 n=1,592[a]** | **1994 n=1,577** | **1995 n=1,564** | **1996 n=1,502** | **1997 n=1,554** | **1998 n=1,542** | **1999 n=1,456** | **2000 n=1,437** |
| Participation to crop insurance programs (=1 if crops are insured and = 0 otherwise) | 0.50126 (0.50016) | 0.63061 (0.48279) | 0.83269 (0.37337) | 0.80519 (0.39619) | 0.77920 (0.41492) | 0.76307 (0.42533) | 0.78807 (0.40882) | 0.82467 (0.38038) |
| Input usage[b] ($ per acre) | 21.88555 (12.20978) | 23.18898 (12.55149) | 23.90639 (13.08045) | 23.95523 (12.73714) | 26.65258 (12.86799) | 25.39059 (12.89546) | 23.24816 (12.08067) | 23.82365 (11.78041) |
| Lagged crop insurance expenses ($1,000)[b] | 0.01312 (0.25021) | 1.22367 (2.44343) | 1.91458 (3.32078) | 1.93036 (3.03292) | 2.50884 (3.62721) | 3.02963 (4.27609) | 2.86903 (4.03666) | 2.79550 (3.86556) |
| Initial wealth[b] ($10,000) | 41.18690 (40.20516) | 41.23131 (39.00387) | 40.02082 (38.70426) | 39.90609 (37.60988) | 39.58966 (38.44209) | 43.45182 (41.69335) | 43.30156 (41.19539) | 44.47440 (42.74391) |
| Total wheat, corn, sorghum and soybeans acres | 765.75258 (598.12558) | 812.69330 (645.33809) | 803.81149 (632.00688) | 891.83484 (712.99687) | 875.32683 (671.62006) | 924.97456 (756.93849) | 907.70329 (727.32385) | 978.33812 (778.05571) |
| Expected farm net income per crop acre[b] ($ per acre) | 49.42957 (73.06011) | 48.96083 (67.41160) | 41.70424 (53.37717) | 39.86863 (59.32141) | 40.17366 (57.07150) | 45.75548 (52.57076) | 40.63625 (115.03217) | 45.04496 (83.11371) |
| Premium rate (premium-subsidies) / liability | 0.05753 (0.01259) | 0.06105 (0.01441) | 0.03419 (0.00979) | 0.04000 (0.00975) | 0.04983 (0.00979) | 0.04587 (0.00828) | 0.03944 (0.00846) | 0.04244 (0.00859) |
| Lagged mean normalized yields (individual yields) / (country yields) | 0.97334 (0.25776) | 1.00280 (0.20855) | 1.01961 (0.25285) | 1.01127 (0.27213) | 1.01590 (0.19349) | 0.98712 (0.22703) | 0.99406 (0.23722) | 0.99643 (0.24788) |

[a] n represents the number of observations

[b] values have been deflated using the Consumer Price Index deflated to constant 1993 dollars.

Table 1. Variable Summary statistics (continuation)

| | Mean (Standard Deviation) | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | **1993 n=1,592**[a] | **1994 n=1,577** | **1995 n=1,564** | **1996 n=1,502** | **1997 n=1,554** | **1998 n=1,542** | **1999 n=1,456** | **2000 n=1,437** |
| Mean coefficient of variation yields ( %) | 0.35243 (0.31817) | 0.30438 (0.29706) | 0.27243 (0.30145) | 0.25518 (0.31554) | 0.26667 (0.28005) | 0.31103 (0.67125) | 0.35127 (0.34114) | 0.32493 (0.32666) |
| Livestock ratio (livestock sales) / (total sales) | 0.29946 (3.32490) | 0.35804 (3.54198) | 0.26225 (1.49740) | 0.26934 (1.77503) | 0.21290 (1.62611) | 0.18862 (0.77725) | 0.23082 (1.10332) | 0.20606 (0.86338) |
| Ratio of irrigated over nonirrigated crop acres (irrigated acres) / (nonirrigated aces) | 0.29940 (0.26856) | 0.29587 (0.25702) | 0.30408 (0.34010) | 0.29044 (0.27966) | 0.28051 (0.25964) | 0.27977 (0.26038) | 0.28281 (0.25321) | 0.27534 (0.25260) |
| Debt to assets ratio (debt / assets) | 0.35243 (0.31817) | 0.30438 (0.29706) | 0.27243 (0.30145) | 0.25518 (0.31554) | 0.26667 (0.28005) | 0.31103 (0.67125) | 0.35127 (0.34114) | 0.32493 (0.32666) |

[a] n represents the number of observations

Table 2. Empirical models of the crop insurance purchase decision

| | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 |
|---|---|---|---|---|---|---|---|---|
| Constant | 1.73000<br>(0.76893)*[a] | -0.07820<br>(0.93588) | 2.46070<br>(0.78766)* | 1.24770<br>(0.75221) | 1.88350<br>(1.35074) | 2.45260<br>(1.40153) | 4.49350<br>(1.18612)* | 1.44490<br>(1.01884) |
| Lagged crop insurance expenses | 0.02770<br>(1.18283) | 0.32560<br>(0.06523)* | 0.07060<br>(0.05986) | 0.12800<br>(0.08533) | 0.32280<br>(0.04701)* | 0.29720<br>(0.05482)* | 0.27060<br>(0.04993)* | 0.22880<br>(0.16935) |
| Initial Wealth | -0.00270<br>(0.00122)* | -0.00232<br>(0.00123) | -0.00036<br>(0.00129) | -0.00272<br>(0.00148) | -0.00414<br>(0.00163)* | -0.00382<br>(0.00170)* | -0.00064<br>(0.00148) | -0.00323<br>(0.00160)* |
| Total acres planted to wheat, corn, sorghum and soybeans | 0.00005<br>(0.00007) | 0.00007<br>(0.00009) | 0.00003<br>(0.00011) | 0.00019<br>(0.00009)* | 0.00024<br>(0.00010)* | 0.00000<br>(0.00011) | 0.00001<br>(0.00009) | 0.00003<br>(0.00009) |
| Expected farm net income per crop acre | -0.00184<br>(0.00065)* | -0.00137<br>(0.00060)* | -0.00183<br>(0.00085)* | -0.00243<br>(0.00067)* | -0.00195<br>(0.00079)* | -0.00097<br>(0.00095) | -0.00158<br>(0.00091) | -0.00297<br>(0.00101)* |
| County average premium rate | -19.80560<br>(11.59850) | 13.71230<br>(13.51930) | -32.36650<br>(18.17340) | -21.00200<br>(16.11850) | -25.96830<br>(22.66270) | -41.01880<br>(24.70070) | -82.18220<br>(25.28110)* | -0.03300<br>(23.60620) |
| County average premium rate * mean normalized yields | 3.89710<br>(11.81050) | -24.60850<br>(13.99330) | 31.32710<br>(19.41250) | 20.61710<br>(17.46450) | 2.81740<br>(22.63000) | 12.10180<br>(25.02050) | 57.57390<br>(26.95110)* | -28.99140<br>(25.18110) |
| County average premium rate * mean normalized yields * CV of yields | 0.05820<br>(0.14166) | 0.09070<br>(0.14937) | -0.27980<br>(0.26593) | -0.14600<br>(0.21739) | 0.22890<br>(0.25480) | 0.31160<br>(0.25606) | 0.01250<br>(0.32816) | 0.54860<br>(0.29487) |
| Mean normalized yields | -0.13410<br>(0.67439) | 1.20150<br>(0.84538) | -0.90620<br>(0.67108) | -0.58810<br>(0.65883) | 0.16280<br>(1.19217) | -0.38350<br>(1.21182) | -2.48790<br>(1.04514)* | 0.75060<br>(1.05994) |

[a] Numbers in parenthesis are standard errors. The asterisks indicate statistical significance at the $\alpha=0.05$ level.

Table 2. Empirical models of the crop insurance purchase decision (continued)

| | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 |
|---|---|---|---|---|---|---|---|---|
| CV of yields | 0.00234<br>(0.00780)[a] | 0.00314<br>(0.00931) | 0.01180<br>(0.01017) | 0.01460<br>(0.00905) | -0.00229<br>(0.01253) | -0.00508<br>(0.01198) | 0.00210<br>(0.01360) | -0.01200<br>(0.01344) |
| Total sales derived from livestock sales | -0.32040<br>(0.12171)* | -0.48620<br>(0.13092)* | -0.39190<br>(0.14786)* | -0.21300<br>(0.16089) | -0.21900<br>(0.16611) | -0.02460<br>(0.25949) | -0.30280<br>(0.14377)* | -0.06510<br>(0.14788) |
| Irrigated over nonirrigated crop acres | 0.07800<br>(0.03853)* | -0.00248<br>(0.04417) | 0.02830<br>(0.03755) | 0.07700<br>(0.12168) | 0.00899<br>(0.05817) | 0.33020<br>(0.13818)* | -0.03610<br>(0.08249) | 0.06460<br>(0.10895) |
| Debt to assets ratio | 0.42100<br>80.16244)* | 0.16180<br>(0.17332) | 0.00555<br>(0.15870) | 0.01020<br>(0.19180) | -0.01510<br>(0.24480) | -0.09030<br>(0.19728) | 0.34600<br>(0.23331) | -0.04830<br>(0.24609) |
| Input usage | -0.03410<br>(0.00638)* | -0.01960<br>(0.00625)* | -0.02170<br>(0.00697)* | -0.00641<br>(0.00690) | -0.02800<br>(0.00801)* | -0.02500<br>(0.00934)* | -0.02740<br>(0.01049)* | -0.01970<br>(0.01476) |
| Likelihood Ratio | 208.0562* | 381.5690* | 103.1080* | 146.3532* | 371.5017* | 407.5835* | 317.54710 | 260.74180 |
| Wu-Hausman test[b] | 0.07510<br>(0.78400) | 2.12700<br>(0.14470) | 1.52290<br>(0.21720) | 0.28230<br>(0.59520) | 3.14440<br>(0.07620) | 2.71210<br>(0.09960) | 0.34400<br>(0.55750) | 4.89610<br>(0.02690) |

[a]Numbers in parenthesis are standard errors. The asterisks indicate statistical significance at the $\alpha$=0.05 level.

[b]P-values of Wu-Hausman test of exogeneity of chemicals usage in the crop insurance demand equation are in brackets.

Table 3. Empirical models of the crop insurance purchase decision: pooled data

| | **1993-2000** |
|---|---|
| Constant | 1.85140 (0.28834)*[a] |
| Lagged crop insurance expenses | 0.20580 (0.03451)* |
| Initial Wealth | -0.00278 (0.00051)* |
| Total acres planted to wheat, corn, sorghum and soybeans | 0.00011 (0.00004)* |
| Expected farm net income per crop acre | -0.00189 (0.00026)* |
| County average premium rate | -17.91500 (4.58724)* |
| County average premium rate * mean normalized yields | 1.21570 (4.64422) |
| County average premium rate * mean normalized Yields * CV of yields | 0.11860 (0.06612) |
| Mean normalized yields | -0.08420 (0.22847) |
| CV of yields | 0.00086 (0.00340) |
| Total sales derived from livestock sales | -0.15810 (0.10867) |
| Irrigated over nonirrigated crop acres | 0.02360 (0.01649) |
| Debt to assets ratio | 0.08640 (0.07030) |
| Input usage | -0.02100 (0.00268)* |

[a]Numbers in parenthesis are standard errors. The asterisks indicate statistical significance at the $\alpha$=0.05 level.

Table 3. Empirical models of the crop insurance purchase decision: pooled data (continued)

| | **1993-2000** |
|---|---|
| 1993 | -0.47000 (0.07129)*[a] |
| 1994 | -0.23660 (0.06425)* |
| 1995 | 0.00747 (0.06517) |
| 1996 | -0.06110 (0.06117) |
| 1997 | -0.03680 (0.05842) |
| 1998 | -0.20440 (0.05723)* |
| 1999 | -0.22530 (0.05808)* |
| Log Likelihood | 2604.81480 |
| Wu-Hausman test[b] | 7.96270 (0.00480) |

[a]Numbers in parenthesis are standard errors. The asterisks indicate statistical significance at the $\alpha$=0.05 level.

[b]P-values of Wu-Hausman test of exogeneity of chemicals usage in the crop insurance demand equation are in brackets.

Table 3. Price elasticity of the demand for crop insurance

| | 1993 | 1994 | 1995 | 1996 | 1997 | 1998 | 1999 | 2000 | Panel |
|---|---|---|---|---|---|---|---|---|---|
| Elasticity | -0.57545 (0.11428)*[a] | -0.23453 (0.08231)* | -0.09362 (0.04557)* | -0.06502 (0.05613) | -0.22143 (0.06913)* | -0.26446 (0.08778)* | -0.28610 (0.06597)* | -0.12707 (0.06143)* | -0.21369 (0.02324)* |

[a]Numbers in parenthesis are standard errors. The asterisks indicate statistical significance at the $\alpha$=0.05 level.