

THE EFFECTS OF OPTIONAL UNITS ON
CROP INSURANCE INDEMNITY PAYMENTS

by
John Kuhling

A thesis submitted in partial fulfillment
Of the requirements for the degree

Of

Master of Science

In

Applied Economics

MONTANA STATE UNIVERSITY-BOZEMAN
Bozeman, Montana
January, 2002

APPROVAL
of a thesis submitted by
John Kuhling

This thesis has been read by each member of the thesis committee and has been found to be satisfactory regarding content, English usage, format, citations, bibliographic style, and consistency, and is ready for submission to the College of Graduate Studies.

Myles Watts

(Signature)

Date

Approved for the Department of Agricultural Economics and Economics

Myles Watts

(Signature)

Date

Approved for the College of Graduate Studies

Bruce Mcleod

(Signature)

Date

STATEMENT OF PERMISSION TO USE

In presenting this thesis in partial fulfillment of the requirements for a master's degree at Montana State University-Bozeman, I agree that the Library shall make it available to borrowers under rules of the Library.

If I have indicated my intention to copyright this thesis by including a copyright notice page, copying is allowable only for scholarly purposes, consistent with "fair use" as prescribed in the U.S. Copyright Law. Requests for extended quotation from or reproduction of this thesis in whole or in parts may be granted only by the copyright holder.

Signature _____

Date _____

ACKNOWLEDGEMENTS

I wish to thank all the members of my graduate committee for their time and guidance. Dr. Myles Watts read countless drafts of this paper, always responding with encouragement and helpful insights. Dr. Joe Atwood spent time explaining to me the critical points of the thesis. Dr. Randy Rucker and Dr. Dean Lueck were both generous with their time.

I would like to thank the Risk Management Agency of the USDA for allowing access to crop insurance data files and for their funding. I would like to thank Saleem Shaik for kindly allowing me to use SAS codes he had written. A special thanks to my parents, Richard and Stacy Kuhling, for their support.

TABLE OF CONTENTS

	Page
LIST OF TABLES	vii
LIST OF FIGURES	viii
ABSTRACT.....	ix
1. INTRODUCTION	1
2. DEVELOPMENT OF HYPOTHESES	6
Hypothesis One	7
Hypothesis Two	11
Procedures	13
3. HISTORY AND LITERATURE REVIEW	14
A Brief History of Crop Insurance in the U.S.	14
Mechanics of MPCl	18
Unit Structure	23
Literature Regarding Crop Insurance and Optional Units	26
4. THEORETICAL ARGUMENTS	28
Hypothesis One	28
Hypothesis Two	33
5. DATA AND METHODS	37
Data Sets	37
Tests of Hypothesis One	45
Monte Carlo Simulation	45
Regression Analysis	46
Tests of Hypothesis Two	52
6. RESULTS	56

TABLE OF CONTENTS—Continued

Hypothesis One	56
Regression Estimates	57
Monte Carlo Simulation Results	62
Hypothesis Two	62
7. CONCLUSION	74
APPENDICES	78
APPENDIX A	79
Inconsistency of OLS Coefficients When the Data Set is Censored	80
APPENDIX B	81
Standard Setup of the Tobit Model	82
APPENDIX C	83
Figures	83
REFERENCES CITED	90

LIST OF TABLES

Table	Page
1. Percent of Farmers Eligible for Optional Units Who insure Optional Units	10
2. Subsidy Amounts Based on the Yield Election	24
3. State Wide Average Yield, Acreage, and Yield Standard Deviation	37
4. Enterprise Unit Tobit Estimates, Model A	60
5. Enterprise Unit Tobit Estimates, Model A	61
6. Enterprise Unit Tobit Estimates, Model A	62
7. Optional Unit Tobit Estimates, Model 1, Model 2, and Model 3	65
8. Optional Unit Tobit Estimates, Model 1, Model 2, and Model 3	66
9. Optional Unit Tobit Estimates, Model 1, Model 2, and Model 3	67
10. Optional Unit Tobit Estimates, Model 1, Model 2, and Model 3	68
11. Optional Unit Tobit Estimates, Model 1, Model 2, and Model 3	69
12. Optional Unit Tobit Estimates, Model 1, Model 2, and Model 3	70
13. Likelihood Ratio Tests	74

LIST OF FIGURES

Figure	Page
1. The RMA's Current Premium Schedule	3
2. Participation in Multiple Peril Crop Insurance	19
3. Structure of Insurable Farmland	27
4. The Probability of Receiving an Indemnity Payment	53
5. Current and Proposed Premium Schedule	78
6. Texas Irrigated Cotton	84
7. Texas Dryland Cotton	85
8. Iowa Dryland Soybeans	86
9. Illinois Dryland Soybeans	87
10. Texas Irrigated Cotton Monte Carlo Results	88
11. Texas Dryland Cotton Monte Carlo Results	89

ABSTRACT

The federal government supports numerous agricultural programs including Multiple Peril Crop Insurance (MPCI). Currently, the Risk Management Agency of the United States Department of Agriculture is legally required to determine actuarially sound premiums for MPCI. A variety of producer and production characteristics affect expected MPCI indemnity payments and thereby actuarially sound premiums. Qualifying producers may choose to split their cropland into "optional units." For the purposes of calculating crop insurance indemnity payments, the Risk Management Agency considers each optional unit a separate insurance contract. This thesis investigates whether the number of optional units a farmer is eligible to insure affects expected per acre indemnity payments.

Using regression analysis and Monte Carlo simulations, this thesis statistically tests whether the number of optional units insured, separately or consolidated into a single parcel, affects expected indemnities per acre. The predictions are that the number of optional units insured separately does not affect expected indemnities per acre and that the number of optional units consolidated into a single parcel does affect expected per acre indemnities.

This thesis finds a per acre premium decrease for farmers who insure available optional units as a single parcel and that the decrease in the premium be positively related to the number of optional units consolidated. In addition, this thesis finds the RMA's practice of increasing per acre premiums for farmers who separately insure two or more optional units versus the premium per acre for farmers who insure a single optional unit.

CHAPTER 1

INTRODUCTION

The federal government, and therefore taxpayers, supports numerous agricultural programs including Multiple Peril Crop Insurance (MPCI). Currently, the Risk Management Agency (RMA) of the United States Department of Agriculture is legally required to determine actuarially sound premiums for MPCI.¹ A variety of producer and production characteristics effect expected indemnity payments and thereby actuarially sound premiums. Qualifying producers may choose to split their cropland into “optional units”. For the purposes of calculating crop insurance indemnity payments, the RMA considers each optional unit a separate insurance contract. Therefore, a low yield on one optional unit may generate an indemnity payment even if other optional units had high yields. This thesis investigates whether the number of optional units a farmer is eligible to insure affects expected per acre indemnity payments.

The passage of the 1938 Crop Insurance Act established Multiple Peril Crop Insurance. MPCI was designed to “... provide individual farmers with some degree of protection against financial hardship caused by adverse growing conditions that resulted in substantial crop losses” (Goodwin and Smith 1995). The nature of farming is risky.

¹ Actuarially sound premiums are premiums that are equivalent to the expected return (indemnity payment) from the insurance policy.

Farmers can do little to guard against, or foresee, natural events such as flooding and drought that may financially ruin a farmer. The purpose of MPCCI is to transfer some of the natural disaster risk from farmers to taxpayers (Goodwin and Smith 1995).

Optional units were created with the passage of the 1980 Federal Crop Insurance Act. Optional units are subdivisions of farmers' land that may be separately insured. If farmers meet certain requirements, they may divide their land into optional units.

"Specifically, for most commodities optional subdivision is allowed for land in different sections under rectangular survey, and for irrigated versus dryland production" (Knight and Coble 1999).² The opportunity for farmers to divide their cropland into optional units creates a variety of ways farmers may organize their cropland under an MPCCI policy.

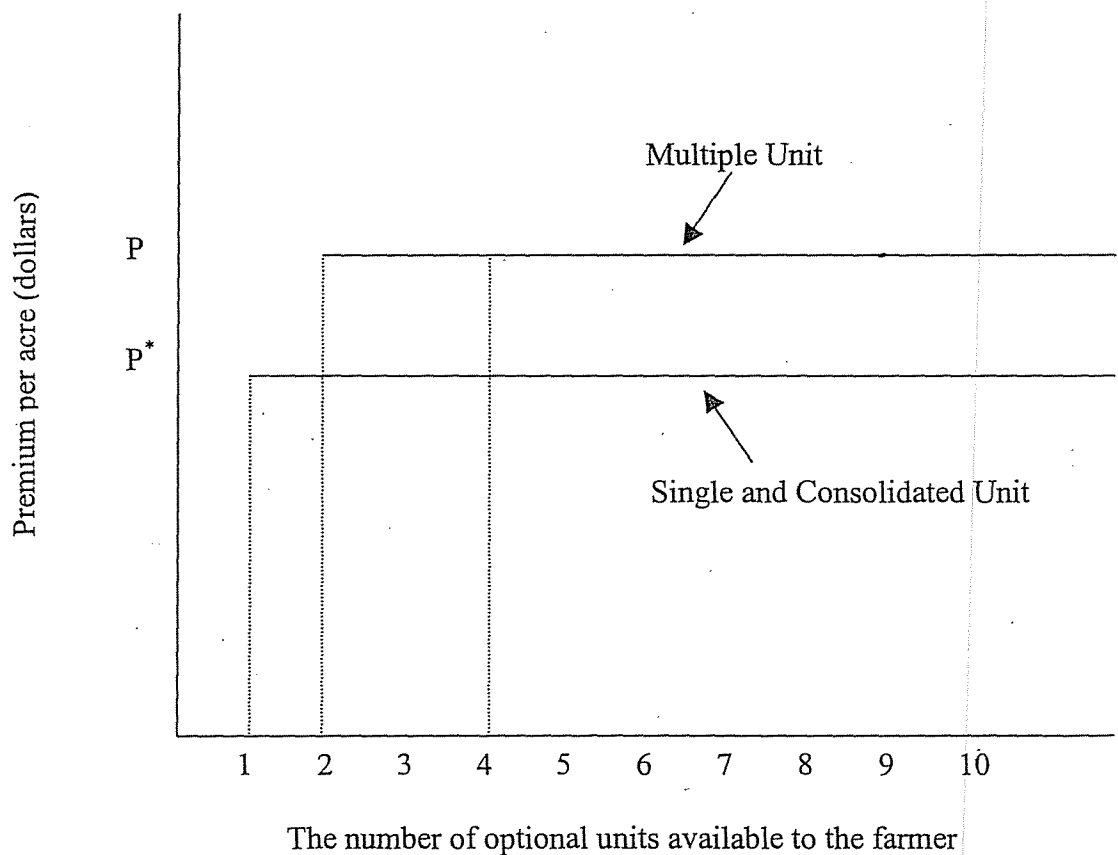
For the purposes of this thesis, there is three ways farmers can insure their cropland: 1) as a single unit, 2) as a multiple unit or 3) as a consolidated unit.³ Farmers eligible for one optional unit are said to insure as a single unit. There are two classifications for farmers eligible to insure two or more optional units. Those farmers who separately insure two or more optional units are said to insure as a multiple unit. Those farmers who are eligible for two or more optional units but choose not to insure them separately (i.e., they combine their optional units into one parcel) are said to insure as a consolidated unit. Farmers who insure as a single or consolidated unit insure their cropland as a single parcel.

² A section is one square mile or 640 acres and "Where legal descriptions are not based on rectangular survey, alternative criteria such as Farm Agency Service farm serial number and non-contiguity are used to define insurable units" (Knight and Coble 1999).

³ These classifications are used for pedagogical reasons in this thesis. RMA uses a more complex system of optional, basic, and enterprise units, which are based on ownership structure and cropland location.

Since 1987, farmers who insure as a single or consolidated unit receive a ten percent discount on their MPCCI premiums. They receive a discount because the RMA has determined that on average, farmers who do not separately insure two or more optional units collect ten percent less in indemnities.⁴ Figure 1 illustrates the RMA's premium schedule.

Figure 1. The RMA's Current Premium Schedule



⁴ Knight and Coble (1999) support the ten percent discount for soybean producers. Knight and Coble state, "for corn and wheat ... a discount of approximately half the current level is justified."

In Figure 1, P and P^* are not assigned dollar amounts because they are intended for illustrative purposes only, P^* is equal to ninety percent of P . The premium per acre to separately insure two or more optional units (i.e., insure as a multiple unit) is P . P^* is the premium per acre for farmers who insure as a single or consolidated unit. Farmers eligible for five optional units who separately insure their optional units pay P per acre. Note both premium schedules are flat. This implies that the number of available optional units insured as a multiple unit or as a consolidated unit do not affect the per acre premium. This is an important point because it means that only the method (as a consolidated or multiple unit) by which farmers eligible for two or more optional units insure their optional units affects expected per acre indemnities.

This thesis will statistically test whether the number of optional units insured, separately as a multiple unit or combined as a consolidated unit, effects expected indemnities per acre. The predictions are that the number of optional units insured separately as a multiple unit does not affect expected indemnities per acre and that the number of optional units insured as a consolidated unit does effect expected indemnities per acre. In other words, the RMA's practice of determining premiums based upon how optional units are insured may be correct for farmers who insure as a multiple unit but may not be correct for farmers who insure as a consolidated unit. Premiums adjusted for the number of optional units insured independently or combined into a consolidated unit have not been discussed in the crop insurance literature.

The data used in the empirical analysis are extracted from the RMA's 1999 yield history files. The data sets contain entries for each optional unit that was eligible to be

insured under MPCCI in 1999. The data sets contain the reported yields for each optional unit from 1979 through 1998. None of the optional units have a complete yield history. The lack of a complete yield history for the optional units led to the use of a bootstrapping process to simulate missing yields from 1987 through 1994.

The remainder of the thesis is broken into six chapters. Chapter 2 discusses the background information that led to development of this thesis and details the hypotheses to be tested. Chapter 3 discusses the history and mechanics of MPCCI as well as reviewing the literature on MPCCI. Chapter 4 develops the theoretical arguments that led to the hypotheses. Chapter 5 describes the methods used to test the hypotheses and the data used in the statistical procedures. Chapter 6 discusses the results and chapter 7 summarizes the findings and implications of the research.

CHAPTER 2

DEVELOPMENT OF HYPOTHESES

The goal of this thesis is to evaluate the RMA's current premium schedule, as illustrated in Figure 1, for both multiple and consolidated units. This chapter develops the logical basis for the concerns about the premium schedules and discusses the formation of testable hypotheses regarding RMA premium determination. The concern with the RMA's current premium schedule is that the number of optional units insured is ignored for farmers who insure two or more optional units. Accurate measures of expected indemnity payments are necessary for the development of actuarially sound premiums. If premiums are not actuarially sound, there may be unintentional wealth transfers from taxpayers to farmers or farmers may not participate in MPCl.

This chapter develops two testable hypotheses regarding the effect of the number of optional units insured on expected indemnities per acre. The first hypothesis is based on farmers insuring available optional units as a consolidated unit. The second hypothesis is based on farmers insuring available optional units as a multiple unit. The hypotheses are developed from theoretical considerations about the relationship between the number of optional units available and expected indemnities. The following discussion outlines these theoretical considerations and introduces the hypotheses.

Hypothesis One

Hypothesis one pertains to farmers who insure available optional units as a consolidated unit. Farmers eligible for two or more optional units have the choice to insure as a consolidated unit or as a multiple unit. The method farmers choose to insure their cropland may affect expected per acre indemnities. If the optional units combined to form the consolidated unit are “spatially diverse”, it may be that the expected per acre indemnities for the consolidated unit are less than the expected per acre indemnities if the optional units are insured as a multiple unit. Spatial diversity refers to the spatial difference within an insured parcel of land. Spatial diversity could be measured as the average distance between the optional units used to form the consolidated unit. Recall that when farmers insure as a multiple unit, each optional unit is considered a separate entity for calculating indemnity payments. Therefore, the relative location of the optional units may not affect expected per acre indemnity payments when insured as a multiple unit. Nevertheless, the relative location of the optional units may affect expected indemnities per acre if the optional units are insured as a consolidated unit.

A negative relationship between expected per acre indemnities and spatial diversity may exist for farmers who insure as a consolidated unit. The more spatially diverse a consolidated unit, the less the yields of the optional units may be correlated. The less the correlation between the yields of the optional units the lower the average per acre yield variance of the consolidated unit. The lower the average yield variance, the

lower the expected per acre indemnities.⁵ Therefore, if the optional units used to form consolidated unit A are more spatially diverse than the optional units used to form consolidated unit B, the expected indemnities per acre for consolidated unit A should be less than the expected indemnities per acre for consolidated unit B. One problem with developing premiums based on the spatial diversity of optional units is that the geographic location of the optional units must be known.

Unfortunately, the RMA does not have consistent legal descriptions or any other location identifiers to identify the geographic location of each optional unit. Without the geographic location, the spatial difference between optional units cannot be measured. In addition, if the information regarding the location of optional units were available, it is conceivable that it would be prohibitively costly to measure the spatial difference between optional units for every farm that is eligible for optional units across states and crop types.

Although the exact spatial difference between optional units cannot be measured, it may be the case that the number of available optional units can be used as a proxy for spatial diversity. Because most optional units must be located in different sections, the number of optional units may be a crude indicator of the spatial diversity of farms who insure as a consolidated unit. If the number of optional units are a proxy for spatial diversity, the number of optional units available to a farmer is critical information. Using the number of optional units as a proxy for spatial diversity would not be viable if the

⁵ Proof found in chapter four.

number of optional units available to farmers is unknown for a large portion of the farms in the data sets.

The data sets used in this thesis do not provide direct information whether farmers' are eligible for optional units. The data sets only indicate if farmers insure as a multiple unit. When farmers insure as a multiple unit, the number of optional units available is recorded in the data sets. Therefore, the number of farmers who separately insure two or more optional units is known, but it is not known how many farmers eligible for two or more optional units insure as a consolidated unit. If it is the case that a large number of the farms eligible for two or more optional units insured as a consolidated unit, the number of separately insured optional units will be a poorer proxy for spatial diversity.

Since an exact percent of those eligible for two or more optional units who choose to insure as a multiple unit is not available from the data sets, a method to estimate the percentage is developed. If a farmer insures more than 640 acres they are known to be eligible for two or more optional units. Table 1 provides an estimate of the percent of farmers known to be eligible for multiple units who insure two or more optional units separately. The table shows that most farmers known to be eligible to insure as a multiple unit choose to insure their optional units separately. Therefore, the number of optional units is used as a proxy for spatial diversity.

Table 1. Percent of Farmers Eligible for Optional Units Who Insure Optional Units

State / Crop	Number of farms	Number of farms with more than 640 acres	Number of farms with more than 640 acres that insured as a multiple unit	Percent of farms with acres > 640 that insured as a multiple unit
North Dakota / Barley	2639	31	26	83.87
Illinois / Corn	35754	665	497	74.74
Georgia / Cotton	1633	139	120	86.33
Texas / Cotton	17401	1305	1223	93.72
Iowa / Soybeans	45663	363	337	92.84
Illinois / Soybeans	31240	403	321	79.65
Minnesota / Soybeans	21627	440	417	94.77
Missouri / Soybeans	8315	196	164	83.67
Kansas / Wheat	35783	1575	1531	97.21

Source: RMA yield history files 1995-98

The preceding discussion explains why the number of optional units insured as a consolidated unit may affect expected per acre indemnity payments and shows that the data sets contain ample observations to test the hypothesis that:

There is a negative relationship between the expected per acre indemnity payment and the number of optional units insured as a consolidated unit.

The hypothesis is tested empirically using regression analysis and Monte Carlo simulations. In addition, chapter four provides theoretical arguments which support hypothesis one.

Hypothesis Two

The second hypothesis pertains to farmers who separately insure two or more optional units. Currently, the RMA charges a higher per acre premium to farmers who separately insure two or more optional units than to farmers who insure a single optional unit. However, a farmer who separately insures six optional units pays the same premium per acre as a farmer who separately insures two optional units. The question arises as to why there is an increase in premium per acre when farmers move from insuring one optional unit to separately insuring two optional units but not an additional increase in premiums if farmers insure more than two optional units? The following discussion outlines the arguments for why there should not be an increase in premiums per acre when farmers move from insuring a single optional unit to separately insuring two or more optional units.

When farmers insure as a multiple unit, optional units divide the insured cropland into parcels that are considered separately for calculating indemnity payments. As a multiple unit farm divides into more optional units, the optional unit per acre yield variance is expected to increase because each optional unit includes less acreage. This logic suggests that as the number of separately insured optional units increase, the expected per acre indemnity payments and thereby the per acre premiums should increase. If the yield were identical across the acreage each year (perfect correlation in yields across acres) then the number of acres included in an optional unit would not affect the expected per acre indemnity payments and per acre premiums.

On the other hand, suppose that multiple unit farms with larger acreage divide their farms in such a manner that each optional unit is the same size as each optional unit in a smaller farm. Of course, the smaller farms would have fewer optional units (and in the case of the smallest farmers a single unit). In such a situation, the number of optional units on a farm may not affect the expected per acre indemnity payments. Thus, if yields are homogeneous across acreage or optional units are homogeneous in acreage with single units, then theoretically there are no reasons why farmers who insure optional units should pay more in per acre premiums than farmers who insure a single unit.⁶ A plausible reason for higher indemnity payments for farmers who insure as a multiple unit than for farmers who insure a single unit is that some farmers who insure as a multiple unit generate fraudulent indemnity payments.

Chapter four further discusses whether expected indemnity payments should be higher for multiple versus single unit farms. Hypothesis two tests the theory that there should not be an increase in per acre premiums against the RMA's practice of increasing per acre premiums when farmers move from insuring a single optional unit to separately insuring two or more optional units. More formally hypothesis two is:

There is a positive relationship between the expected per acre indemnity payments and the number of separately insured optional units.

The hypothesis is stated so that rejecting the hypothesis supports the theoretical arguments presented in chapter four and accepting the hypothesis acknowledges the RMA's current practice of increasing premiums for farmer who insure as a multiple unit.

⁶ Knight and Coble (1999) have shown that optional units and single units are homogeneous in acreage. See also table 3 on page 32.

Procedures

Regression analysis is used to evaluate the hypotheses. In addition, a Monte Carlo simulation evaluates the proposed negative relationship between the number of optional units insured as a consolidated unit and expected per acre indemnities. Chapter four discusses the theoretical considerations that led to the development of the hypotheses.

The data used are extracted from the RMA's yield history files. The data set contains entries for most farmers that were eligible to participate in MPCCI in 1999 and includes all reported yields from 1979 to 1998.⁷ The data is broken down to the optional unit level. If the farmer insured as a single unit then there is one entry for that farmer and if the farmer insured as a multiple unit there are two or more entries for that farmer, one for each optional unit the farmer separately insured. As discussed previously, there is no indicator in the data set for farmers who were eligible for two or more optional units who insured as a consolidated unit. In addition, the RMA does not calculate losses as if farmers who insured as a multiple unit had insured as a consolidated unit. Therefore, consolidated unit losses are calculated using yield information from farms that insured as a multiple unit. Combining yield information for all optional units reported under a farm policy number formed consolidated units. For example, if a farmer separately insures four optional units then aggregating the yields and acreage of the four optional units, using a share-weighted average, produces a yield for the consolidated unit. From the consolidated unit yield indemnity payments are calculated.

⁷ Most farms did not have complete yield records for 1979 to 1998.

CHAPTER 3

HISTORY AND LITERATURE REVIEW

A Brief History of Crop Insurance in the U.S.

In 1938, Congress established the Federal Crop Insurance Corporation (FCIC) to administer MPCI. At the time, private insurance companies offered farmers coverage for hail and fire but did not offer insurance for risks such as flooding, insects or drought. Miranda and Glauber (1997), suggest that insurance for flooding, insects and drought were not offered by the private market because the risk from these perils is systematic.⁸ Other studies point to adverse selection and moral hazard as potential reasons for the lack of private markets for crop insurance (Chambers 1989; Skees and Reed, 1986; Goodwin and Smith, 1995). Because of the lack of private markets for crop insurance, the federal government offered farmers MPCI to shift "... the risk of growing crops from at least some farmers to taxpayers" (Goodwin and Smith 1995).

The participation rate in MPCI and the average annual loss ratio of the program are two statistics frequently cited to evaluate the performance of MPCI. Participation rates are measured as the percent of eligible acres that are covered by MPCI. The average

annual loss ratio is calculated as the total indemnities paid over a specific period of time divided by the total premiums collected over that time period. A loss ratio greater than one means that the amount of indemnities paid out to farmers is greater than the amount of premiums collected. From its inception through 1956, the annual average loss ratio remained above one and participation was low. From 1956 to 1973, the loss ratio improved to 0.86 but participation remained low (Atwood et al. 1999).

Initially, MPCl covered a limited number of crops. Some of the initial crops covered were wheat, cotton, and flax. By 1973, MPCl had expanded to cover most agricultural commodities for which countywide yield data were available yet participation remained low. During the 1970's, the government started disaster relief programs in which spending was about ten times as much as the spending on MPCl subsidies (Atwood et al. 1999).⁹ The Agricultural and Consumer Protection Act of 1973 and The Rice Production Act of 1975 established disaster programs that covered wheat, upland cotton, feed grains and other crops. Under these acts, farmers received payments if their yields fell below 66.67 percent of the farm's established yield (Goodwin and Smith 1995).¹⁰ In 1974, 580,000 applications for disaster relief were filed and 476,000 were approved, a success rate of 82 percent (Goodwin and Smith 1995). During the 1970's, disaster relief likely reduced the incentive for farmers to enroll in MPCl.

⁸ Systematic risks are risks where the probability of affecting one policyholder is highly correlated with its probability of affecting all policyholders.

⁹ Expenditures under the Disaster Relief Program were \$339 billion versus \$314 million for MPCl subsidies between 1973 and 1980 (Atwood et al. 1999).

¹⁰ The established yield is basically the farms average yield over the previous four to ten years. The established yield is also known as the APH yield and is discussed in detail in the following section 'Mechanics of MPCl'.

In response to low participation, the Federal Crop Insurance Act of 1980 mandated that crops covered and participation rates for MPCCI increase. "The 1980 Act set a target participation rate of fifty percent of all eligible production acres" (Atwood et al. 1999). The FCIC responded successfully to the goals and participation rates increased during the 1980's. One method used to bolster participation rates was allowing farmers to subdivide their land into optional units. With the expansion in participation and the introduction of optional units the loss ratios increased. Between 1984 and 1990, the annual average loss ratio was 1.32 versus 1.21 from 1981-1983 (Atwood et al. 1999).

In 1994, Congress passed the Crop Insurance Reform Act, which required innovations in federal insurance products (Atwood et al. 1999). According to Goodwin and Smith (1995), "The clear intent of the proposed legislation [was] to reduce federal budgetary outlays on agricultural disaster programs while guaranteeing farmers consistent and reliable access to government support payments when they experience very poor crops." Since 1994, the annual average loss ratio has decreased and participation rates have increased. The 1994 Act made major changes in the yield elections, premium subsidization, and ad hoc agricultural disaster bills.

Starting in 1996 the maximum yield election farmers could choose increased. The yield election is the percentage of the farmers' historical average yield insured. The yield election is comparable to a deductible on auto insurance. When farmers choose a yield election of 75 percent, they have a deductible of 25 percent of their average yield. If the farmer's yield does not fall below the product of their average yield and yield election, they do not receive an indemnity payment. For some crops, the maximum yield election

increased from 65 percent to 80 percent in regions that historically displayed little yield variation (Goodwin and Smith 1995).¹¹ Increasing the yield election may have increased participation because farmers who rarely had their yields fall below 65 percent of their average yield but faced risk of having yields fall below 80 percent of their average, had new incentive to purchase MPCI.

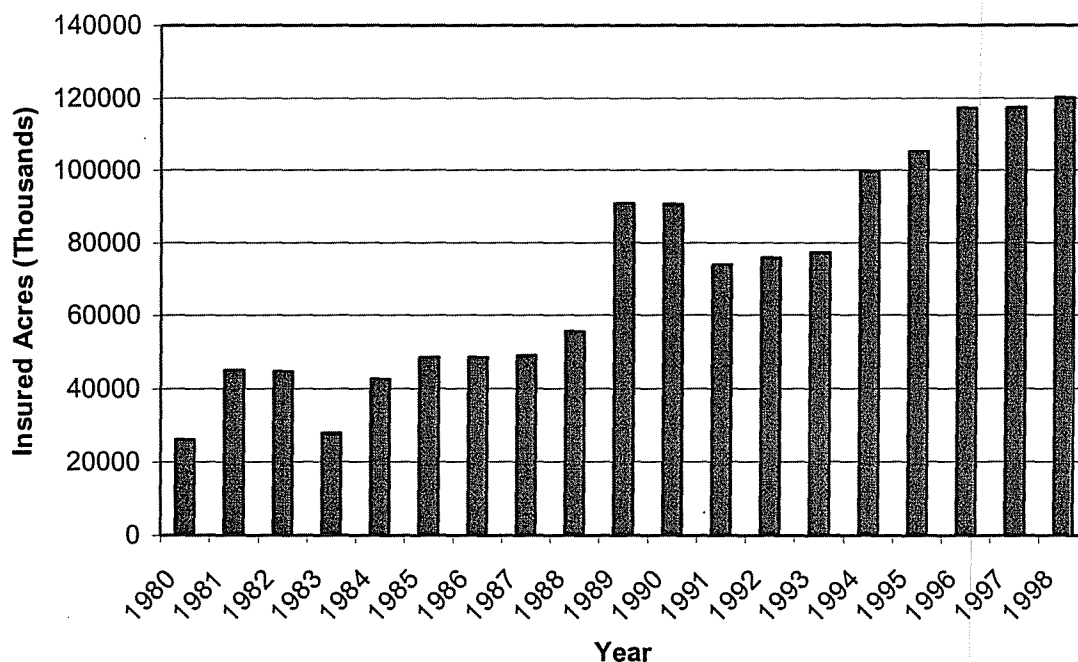
In addition, the 1994 Act increased the premium subsidy for farmers who choose yield elections above 65 percent. Hence, not only could farmers insure more of their average yield; they received an increased subsidy if they did so. The 1994 Act also made it more difficult for Congress to pass ad hoc agricultural disaster relief bills (i.e., temporary rather than permanent measures). As noted earlier, during the 1970's farmer's expectations of federal disaster relief may have depressed MPCI participation rates. Before 1994, there had never been more than one million insured acres in the United States. Since the enactment of the 1994 Act, participation has never fallen below one million acres (Atwood et al. 1999). Unlike during the 1980's, as participation increased after 1995, the loss ratio did not increase. Between 1994 and 1998, the annual average loss ratio was 0.79.¹² Figure 2 shows the number of acres covered under MPCI since 1980.¹³

¹¹ In 2000 the maximum yield election a farmer could choose increased to 85 percent (Agriculture Risk Protection Act of 2000).

¹² This value is misleading because federal subsidies were included in the calculation of the annual average loss ratio (Atwood et. all 1999). Including subsidies biases the statistic downwards. According to Goodwin and Smith (1995), the average premium subsidy was 25 percent over the time period. If this were the case, the annual average loss ratio for 1994-1998, less 25 percent of premiums, would be 1.05.

¹³ In 1998 the number of acres insured under MPCI approached 70 percent of acres planted (Schnapp et. all).

Figure 2: Participation in Multiple Peril Crop Insurance (Source: Atwood et al. 1999)



This brief outline of the recent history of MPCI indicates that the program has been plagued with problems from the outset. It is only in recent years that the problems have begun to be successfully addressed. The next section discusses the mechanics of MPCI.

Mechanics of MPCI

The mechanics of MPCI are considered from two points of view: the view of farmer and the view of the RMA. From the perspective of the farmer, the information provided and the type of MPCI policy chosen are important. The RMA is responsible for determining premiums.

Farmers choose among various crop insurance options and provide the RMA various data including their yield history. Yield history data are averaged to calculate an annual actual production history (APH) yield. The APH yield, as defined on the RMA's website (www.rma.usda.gov), "... is determined from producer production records for a minimum of four, up to ten consecutive crop years."¹⁴ After farmers establish their APH yield, they make decisions concerning the amount of coverage and the price level they would like to insure. Farmers choose a yield election to determine the percent of their APH yield that will be insured. The available yield election choices are 50, 55, 60, 65, 70, 75, 80, and 85 percent. When farmers insure 65 percent of their APH yield, they "trigger" an indemnity payment if the yield in the following year falls below 65 percent of the APH yield. The trigger level is the yield election farmers choose multiplied by their APH yield.

Next, farmers choose a price election. The price election ranges from 50 percent to 100 percent of the base rate (BASE RATE_t). The base rate is anywhere from 90 percent to 100 percent of the FCIC's forecast of expected market prices (Goodwin and Smith 1995). The price election (PRICE ELEC_t) is used to determine the amount of the indemnity payment if a farmer's yield falls below the trigger. In the following equation, a

¹⁴ Farmers may use a weighted transition yield to determine their average yield if a farmer does not have 4 years of yield data. "A transitional yield is equal to the farm's Agricultural Stabilization and Conservation Service (ASCS) program yield multiplied by a 't-factor.' The t-factor is the county average ASCS program yield divided by the National Agricultural Statistics Service county average yield. As the t-factor is typically less than the ASCS program yields and lower than expected yields in most cases, transitional yields are typically much smaller than county average yields (Goodwin and Smith 1995)."

farmer's indemnity payment per acre in year t (IND_t) is calculated, assuming the yield in year t is 40 percent of the APH yield ($APHYIELD_t$) and the yield election ($YIELD ELEC_t$) is 65 percent.

$$(3.1) \quad IND_t = (0.65 - 0.40) * APHYIELD_t * PRICE ELEC_t * BASE RATE_t$$

The indemnity payment per acre is equal to the product of the farmer's APH yield, price election, and the base rate, multiplied by the difference between the yield election and the percent of the APH yield that was realized in year t . With a yield election of 65 percent the farmer receives an indemnity payment per acre equal to 25 percent, $(0.65 - 0.40)$, of the product of his APH yield, price election, and the base rate.

The RMA determines premiums by focusing on two statistics, the farmers' liability and the loss cost ratio history of the county where the farm is located.¹⁵ The loss cost ratio history is used because it is a measure of the indemnity payments, as a proportion of liability, which are typically paid to farmers in a given year. The loss cost ratio measures the expected risk associated with a county. The loss cost ratio is the ratio of indemnities farmers in a county receive in a given year divided by the sum of the farmers' liabilities. The liability (LIABILITY) is the amount of the indemnity payment per acre a farmer receives if his entire crop is destroyed and so is the maximum indemnity payment a farmer could receive. The liability depends on three factors: the farmer's APH yield, the chosen yield election, and the chosen price election. Liability is calculated on a per acre basis as the product of the APH yield, yield election, base rate, and price election.

$$(3.2) \quad \text{LIABILITY} = \text{APHYIELD} * \text{YIELD ELEC} * \text{BASE RATE} * \text{PRICE ELEC}$$

The liability is then used in the calculation of the farmer's premium.

The premium per acre is calculated as the liability times the premium rate and is later adjusted for the federal subsidy. The premium rate used in the calculation of premiums is dependent upon the countywide historical loss cost ratio and an adjustment factor used to generate a prudent insurance reserve. Currently, the RMA uses data from 1975 to the present, or as far back as there are data from MPCl in the county to determine the premium rate. The premium rate for a county is based on the average of the annual loss cost ratios of farmers who purchase MPCl in the county. The premium rate is an average and therefore may be "... adversely affected by a few years with much higher than average LCRs. The RMA has developed a procedure that is intended to reduce the impact that a single year will have on the average loss cost [ratio] of each county (Josephson et. al. 2000)." This procedure is known as the excess loss adjustment. The excess loss adjustment caps the LCR for any single year at the 80th percentile LCR of all years (Josephson et. al. 2000). In addition, the RMA adds a "base rate load" to the premium rate. The base rate load is made up of four components: a disaster reserve factor, a state excess load, a prevented planting load, and a unit division load (Josephson et. al. 2000). The complex details of the processes used to form the excess loss adjustment and the base rate load are not discussed here but can be found in Josephson et. al. (2000). Next, premium rates are smoothed across county lines. This process reduces premium difference between neighboring farmers across county lines.

¹⁵ The loss cost ratio is different from the loss ratio used to evaluate the performance of

This is an overview of how the premium rates are calculated, the lengthy details of this process are not discussed. Nevertheless, the general structure has been outlined. The farmer's premium per acre (PREMIUM) is roughly calculated as the product of the farmer's liability and the premium rate (PRE RATE).

$$(3.3) \quad \text{PREMIUM} = \text{LIABILITY} * \text{PRE RATE}$$

For example, if a farmer's liability per acre is 100 dollars and the premium rate associated with the county in which the farm is located equals 0.25, the premium the farmer is charged, before subsidy, is equal to 25 dollars per acre.

The Agriculture Risk Protection Act of 2000 stipulates subsidy amounts that farmers receive toward their MPCl premiums. The subsidy amounts are dependent on the level of yield election farmers choose and represent the percent of the premium the government pays for farmers. Table 2 shows the subsidy amounts.

The amount of subsidy received, per acre, when insuring 65 percent of the APH yield (SUBSIDY⁶⁵) is equal to 59 percent of the per acre premium, from Table 2, which can be written as:

$$(3.4) \quad \text{SUBSIDY}^{65} = 0.59 * (0.65 * \text{APHYIELD} * \text{PRICE ELEC} * \text{BASE RATE} * \text{PRE RATE})$$

MPCI. The loss ratio is the ratio of indemnities over premiums.

Table 2: Subsidy Amounts Based on the Yield Election

Yield election	Subsidy as a percentage of liability
50%	67%
55%	64%
65%	59%
75%	55%
80%	48%
85%	38%

Source: Agriculture Risk Protection Act of 2000

Unit Structure

For pedagogical clarity, the unit structure classifications used here are single, multiple and consolidated units. RMA uses a somewhat different classification. There are four: whole farm unit, enterprise unit, basic unit and optional unit. These classifications are based on the ownership structure of the land and the relative geographic location of the land within a county.

The whole farm unit includes all insurable acreage in all insurable crops within a county that a farmer owns or rents (RMA CCIP 2001). A farmer can insure an enterprise unit for each insurable crop. For each insurable crop the enterprise unit must consist of all insurable acreage of the crop in the county and contain at least two or more basic or optional units, or one of each (RMA CCIP 2001). The basic unit is a subdivision based on ownership structure. Farmers can have one basic unit for each parcel of land they rent on

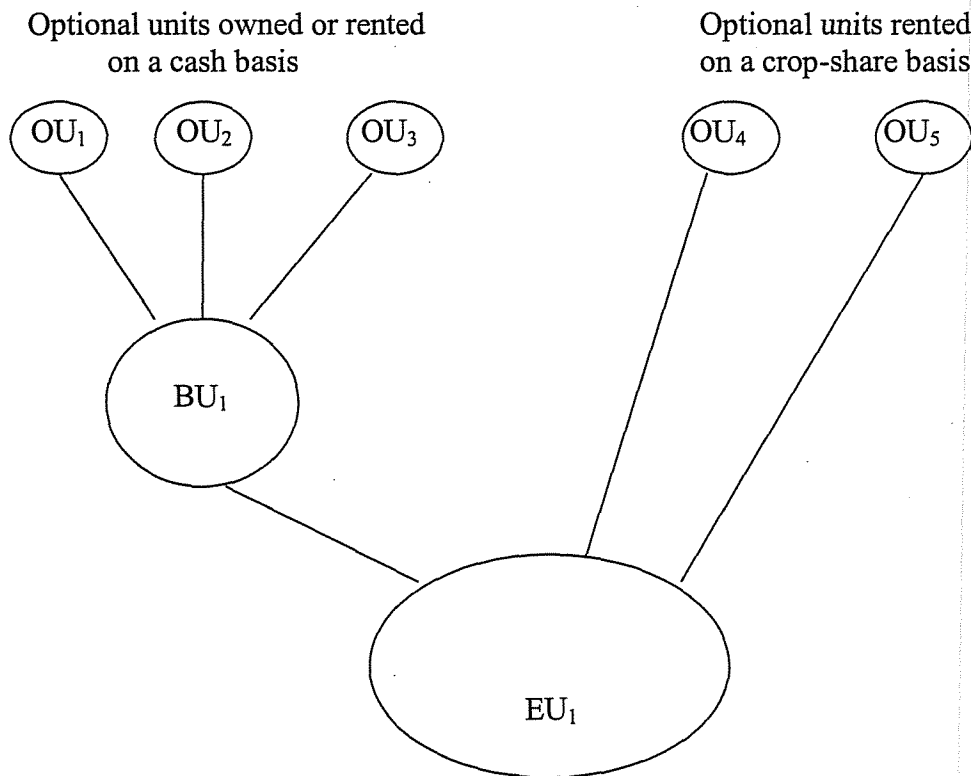
a crop-share basis and one for all parcels owned or rented on a cash basis. The optional unit subdivision is based upon the geographic location of the farmland. Recall that for a given crop, optional units must generally be located in different sections.

Figure 3 illustrates the possible ways to insure farmland for a single crop: optional units (OU), basic units (BU), and enterprise units (EU). The choice of how to organize the land is the farmers. In Figure 3, the optional units are the cropland to be insured. The basic unit and enterprise units represent ways in which the optional units may be grouped. Basic unit one is formed by combining the three optional units that are owned or rented on a cash basis into a single parcel. At this stage, the farmer insures three separate parcels. One parcel is basic unit one and the other two parcels are optional unit four and optional unit five. Under the classifications used in this thesis, the farmer would be said to insure as a multiple unit because he separately insures two or more optional units. Although basic unit one is not technically, according to the RMA, an optional unit, the data used in this thesis does not distinguish between a basic unit and an optional unit. Combining all the cropland in a county used to produce this particular crop, optional units one through five, into a single parcel forms enterprise unit one.

This thesis refers to enterprise units as consolidated units because an enterprise unit must consist of two or more basic or optional units. In addition, farmers who only insure a basic unit and are eligible for optional units are said to insure as a consolidated unit. Farmers currently receive a ten percent discount on their premiums if they do not

insure optional units.¹⁶ Therefore, if the farmer in Figure 3 insured basic unit one and optional units four and five separately, the farmer would receive a ten percent discount on the premiums for basic unit one but not for optional units four and five. If the farmer combined all the cropland into enterprise unit one, he would receive a ten percent discount on premiums for every acre of cropland.

Figure 3. Structure of Insurable Farmland



¹⁶ Pilot programs initiated in 2000 by the RMA provide discounts on MPCCI policies, in addition to the discount for not insuring optional units, for insuring an enterprise unit. The enterprise unit discount is based on the acreage included in the enterprise unit. The analysis of yield data in this thesis shows that acreage is not a reliable indicator of geographic diversity. Therefore, while RMA's new pilot program is addressing the situation of farmers being overcharged for consolidated units, it is not addressing the problem in a manner consistent with the findings of this thesis.

Literature Regarding Crop Insurance and Optional Units

The literature discussing crop insurance includes many studies on a wide variety of topics. These studies examine topics from adverse selection (Skees and Reed 1986; Just, Calvin and Quiggin 1999), moral hazard (Coble, Knight, Pope and Williams 1997; Horowitz and Lichtenberg 1993) and crop insurance demand (Smith and Baquet 1996) to alternative methods of administering crop insurance such as area yield (Miranda 1991) and revenue insurance (Skees 1998). There have also been investigations of the effects of optional units on indemnity payments (Atwood et al. 1999; Knight and Coble, 1999; Watts and Prince, 2000). The concept that premiums be adjusted for the number of optional units insured independently or combined into a consolidated unit has not been discussed in the literature. This thesis represents a new avenue of research in the crop insurance literature.

Knight and Coble (1999) discuss the ten percent discount on premiums for farmers who do not separately insure optional units. They conclude that the ten percent differential in rates is appropriate for soybeans and may over-estimate the difference in expected indemnities for wheat and corn. Their analysis uses loss history data, the same data that RMA uses to determine MPCl premium rates. Knight and Coble compare indemnity payments of producers who insured a single unit with producers who separately insured two or more optional units. A notable finding of Knight and Coble is that the average size (acres) of single units and optional units are not statistically different. In other words, the average size of a farm that insures a single unit is the same

as the average size of an optional unit. This thesis refers to those findings in the discussion of hypothesis two.

A paper by Atwood, Watts and Shaik (1999) explores reasons for reducing premiums for farmers who do not separately insure optional units. They conclude that the presence of fraud, caused by farmers misreporting yields, may increase loss cost ratios for farmers who separately insure two or more optional units. The method of fraud that Atwood et al. explore is called "yield switching". Yield switching occurs when farmers "switch" yields from one separately insured optional unit to another resulting in lower yields on one or more optional units that generate illegitimate indemnity payments.

A staff paper by Watts and Prince (2000) analyzes the magnitude of the difference in expected indemnities for farmers who separately insure two or more optional units versus those who insure as a consolidated unit. The conclusion of the paper posited an increase in expected per acre indemnities as farmers separately insure more optional units. Watts and Prince did not investigate the effect of consolidating optional units on expected per acre indemnities. As a result Watts and Prince did not correctly identify the source of much of the difference in expected indemnity payments between multiple and consolidated units. In addition, the variables used in the regression models have been redefined and the implications of the research have been further evaluated.

CHAPTER 4

THEORETICAL ARGUMENTS

This chapter discusses the theoretical arguments that led to the development of the hypotheses. Throughout the rest of this study, the loss cost ratio (LCR) is used instead of expected indemnity because LCR is a normalized value. As mentioned earlier the LCR is the ratio of indemnities over liability, indemnities are the dollar amount farmers receive from MPCCI when they qualify for a payment on their insurance policy, and the liability is the maximum indemnity farmers may receive. The hypotheses are restated as:

- 1) There is a negative relationship between the expected loss cost ratio and the number of optional units insured as a consolidated unit
- 2) There is a positive relationship between the expected loss cost ratio and the number of separately insured optional units.

Hypothesis One

The hypothesis is based on the following argument: The variance of a consolidated unit's average yield decreases as the number of optional units combined into a consolidated unit increases. As the variance of the average yield decreases, the

probability of receiving a yield below the trigger level also decreases. The trigger level is the quantity the current year's yield must fall below for farmers to receive an indemnity payment. Because, the probability of receiving an indemnity payments decreases as additional optional unit are combined, the expected loss cost ratio must also decrease (given that liability must increase with each additional optional unit).

The yield of a single optional unit in year t (Y_{it}) is defined as:

$$(4.1) \quad Y_{it} = \alpha_i + C_t + U_{it}$$

Where α_i is equal to the historical average yield of the unit, C_t represents a county disturbance term that identically affects all optional units in the county, and U_{it} represents a disturbance term that is unique to a particular unit. The random variables C_t and U_{it} have an expected value of zero across time which means the expected value of Y_{it} is equal to α_i , the historical average yield of the optional unit. Because C_t affects all optional units in a county equally and U_{it} affects each optional unit differently, C_t and U_{it} are independent. Because α_i is equal to the historical average yield of the optional unit, it is independent of the present disturbances C_t and U_{it} . Thus, the variables α_i , C_t , and U_{it} are independent by definition. Therefore, the variance of Y_{it} ($\sigma_{Y_{it}}^2$) equals the variance of the components.

$$(4.2) \quad \sigma_{Y_{it}}^2 = \sigma_{\alpha_i}^2 + \sigma_{C_t}^2 + \sigma_{U_{it}}^2$$

The consolidated unit yield (CUY) is the average yield of the optional units combined to form the consolidated unit, assuming the acreage of the optional units is homogeneous. Without this assumption, the CUY would be an acreage share-weighted average of the optional unit yields. N is the number of optional units combined.

$$(4.3) \quad CUY_t = (1/N) * \sum_{i=1}^N Y_{it} = (1/N) * \sum_{i=1}^N (\alpha_i + C_t + U_{it})$$

The mean of CUY_t across time, where T is the number of years considered, is expressed as:

$$(4.4) \quad E(CUY_t) = (1/T) * \sum_{t=1}^T CUY_t = (1/NT) * \sum_{t=1}^T \sum_{i=1}^N (\alpha_i + C_t + U_{it})$$

Given that the expected value of both C_t and U_{it} across time is equal to zero, the mean of CUY_t is:

$$(4.5) \quad E(CUY_t) = (1/NT) \sum_{t=1}^T \sum_{i=1}^N (\alpha_i) = (1/N) \sum_{i=1}^N \alpha_i$$

The expected value of the consolidated unit yield is equal to the average of the historical average yield (α_i) of the optional units that comprise the consolidated unit. To facilitate the calculation of the variance of the CUY_t , the value of $(CUY_t - E(CUY_t))$ is written as:

$$(4.6) \quad \begin{aligned} CUY_t - E(CUY_t) &= (1/N) \sum_{i=1}^N \alpha_i + C_t + (1/N) \sum_{i=1}^N U_{it} - (1/N) \sum_{i=1}^N \alpha_i \\ &= C_t + (1/N) \sum_{i=1}^N U_{it} \end{aligned}$$

Because the values of C_t and U_{it} are independent, the variance of the CUY_t may be written as:

$$(4.7) \quad \begin{aligned} \sigma_{CUY}^2 &= (1/N) \sum_{i=1}^N (CUY_t - E(CUY_t))^2 \\ &= (1/N) \sum_{i=1}^N (C_t + (1/N) \sum_{i=1}^N U_{it})^2 \end{aligned}$$

$$\begin{aligned}
&= (1/N) \sum_{i=1}^N (C_i)^2 + (1/N) \sum_{i=1}^N ((1/N)^2 \sum_{i=1}^N (U_{ii})^2) \\
&= \sigma_{C_i}^2 + (1/N) \sigma_{U_{ii}}^2
\end{aligned}$$

Therefore, the variance of the CUY_t is equal to the variance of the county disturbance term plus the average of the individual unit disturbance terms. As N increases, the variance of CUY_t decreases and approaches the variance of the county disturbance term ($\sigma_{C_i}^2$).

The LCR of the consolidated unit will decrease as N increases if the expected indemnities decrease and the liability increases or remains constant. Two assumptions facilitate showing that the LCR declines with an increase in N : 1) all optional units have an equal expected yield and 2) optional units are homogeneous in acreage. These assumptions are consistent with those in RMA's MPCIR rates. The assumption of homogenous acreage across units is also supported by the findings of Knight and Coble. Knight and Coble (1999) report "state-level means indicate that optional subdivision has a significant equalizing effect on insured unit size." Moreover, if farms with different average yields and acreage are grouped into different risk-pools, the results presented here hold for each of the different risk-pools.

Under most any plausible yield distribution, the expected indemnity payments decline as the variance of yields decline. Because consolidating optional units decreases yield variance, consolidation thereby also decreases expected indemnities. For example, if a farmer chooses a yield election of 65 percent and yields are normally distributed, the probability (P) of realizing a yield below 65 percent of the average yield is equal to:

$$(4.8) \quad P = \text{Prob}(0.5 - Z_{0.35\mu/\sigma})$$

where Z is the standard normal value evaluated at $0.35\mu/\sigma$. When the standard deviation of the yield (σ) decreases, the value of Z increases, which means the value of P is decreasing.¹⁷ Because the probability of receiving an indemnity payment (P) decreases as the number of optional units insured as a consolidated unit increases, the expected indemnity payment also decreases. Recall that the loss cost ratio is the ratio of indemnities over liability. To prove that the LCR is decreasing it is now only necessary to show that the consolidated unit liability is either constant or increasing as N increases.

Showing that the liability and the number of units insured are positively related is straightforward. The equation for calculating the liability is:

$$(4.9) \quad \text{LIABILITY} = (.65) * \mu_{\text{CUY}} * \text{ACREAGE} * \text{PRICE}$$

Where μ_{CUY} is the mean yield of the consolidated unit, ACREAGE is the acreage of the consolidated unit and PRICE is the dollar amount per unit of yield the farmer receives.

From the previous assumptions, ACREAGE can be written as $N*a$, where a is the acreage of each unit that is combined to form the consolidated unit. As previously defined: $\mu_{\text{CUY}} = (1/N) * \sum \mu_i$. Inserting these values for ACREAGE and μ_{CUY} into the equation for liability gives:

$$(4.10) \quad \begin{aligned} \text{LIABILITY} &= (.65) * (1/N) * \left(\sum_{i=1}^N \mu_i \right) * N * a * \text{PRICE} \\ &= (.65) * \left(\sum_{i=1}^N \mu_i \right) * a * \text{PRICE} \end{aligned}$$

This shows that the value of LIABILITY increases as N increases if the average yields of the optional units combined to form the consolidated unit (μ_i) are greater than zero. Thus as N increases, expected indemnities decrease and the liability increases, which means that the expected loss cost ratio of the consolidated unit decreases.

The preceding has shown that as farmers increase the number of optional units combined to form a consolidated unit, their expected LCR decreases. An implication is that the expected loss cost ratio of a consolidated unit is less than or equal to the expected loss cost ratio of a single unit because consolidated unit farmers must insure two or more optional units while a single unit consists of one optional unit.

Hypothesis Two

The impetus for hypothesis two is the RMA's practice of providing a ten percent discount to farmers who do not separately insure optional units. Recall earlier discussions that optional units whether they are insured as a single unit or as a multiple unit are of approximately the same size (acreage). If each optional unit is homogeneous in acreage, then the number of optional units a farmer separately insures should not affect legitimate expected per acre indemnity payments and therefore the expected LCR. The following discussion dismisses two arguments for offering single unit farmers a ten percent discount on their premiums versus the premium farmers who separately insure two or more optional units pay. For clarification, it is not being argued that the premiums for

¹⁷ It is assumed that the average yield (μ) is constant between units.

farmers consolidating units should not be discounted versus those farmers who insure as a multiple unit.

The first argument for a positive relationship between the number of separately insured optional units and the expected LCR is that when a farmer separately insures an additional optional unit, the probability of incurring a loss increases. When farmers separately insure an additional optional unit the probability of incurring a loss increases because the amount of acreage insured increases.¹⁸ While it is true that the probability of incurring a loss increases, it is not true that the farmer's expected LCR should also increase. When farmers increase the number of units insured, not only does the probability of incurring a loss increase but the dollar amount of liability also increases. Recall that a farmer's liability is the amount of indemnity payment a farmer will receive if his entire crop is lost. When a farmer separately insures an additional optional unit the maximum indemnity payment he could receive must increase because the total number of acres insured has increased. Furthermore, the increase in liability is directly proportional to the increase in the probability of receiving an indemnity payment, if each acre in a county has an equal probability of incurring a loss. While it is true that farmers who separately insure more optional units have a larger expected indemnity because they have a higher probability of incurring a loss, theoretically it is not true that the LCR will increase with the number of separately insured optional units.

¹⁸ This argument is based on each acre in a county having an equal probability of incurring a loss. This assumption is made by the RMA when they determine premiums.

A second argument for a positive relationship between the number of separately insured optional units and the LCR is that the average yield variance is greater for separately insured optional units than for a single unit. Recall that as the average yield variance decreases the expected LCR also decreases. It is also true that as the average yield variance increases the expected LCR also increases. Table 3 shows that the average acreage, yield, and yield standard deviation of single units and separately insured optional units are nearly equivalent within states.

Table 3. State Wide Average Yield, Acreage, and Yield Standard Deviation

State / Crop	Single Unit Farms			Multiple Unit Farms		
	Yield	std dev	acreage	Yield	std dev	acreage per unit
North Dakota / Barley	52	17	108	53	17	115
Illinois / Corn	138	27	106	138	28	86
Georgia / Cotton	653	235	152	658	248	121
Texas / Cotton	269	160	126	255	158	134
Iowa / Soybeans	44	9	82	45	8	77
Illinois / Soybeans	44	7	91	44	7	82
Minnesota / Soybeans	39	9	93	40	10	94
Missouri / Soybeans	36	9	97	37	9	96
Kansas / Wheat	35	15	86	35	15	88

Source: RMA yield history file 1999

Table 3 shows that based on average standard deviation there should be no difference in expected LCR's between single units and multiple units. Nevertheless, if some farmers are involved in illegitimate indemnity enhancing activity such as "switching" production between optional units, then perhaps the expected LCR will be larger for multiple unit farmers than single unit farms.

This chapter has provided arguments supporting hypothesis one and criticized arguments commonly used to support hypothesis two. The predictions are that the empirical work will support hypothesis one and the reject hypothesis two.

CHAPTER 5

DATA AND METHODS

This chapter discusses the data and methods used to empirically test the previously mentioned hypotheses. The first section discusses the data used in the estimation processes. The second and third section present the methods used to test the hypotheses.

Data Sets

The panel data used in this thesis are extracted from the RMA's 1999 yield history file. The data sets contain information on each optional unit's yield history reported to the RMA in 1999. There is one observation per optional unit for farmers who insure as a multiple unit and one observation for each single unit farm. The state crop combinations included in the data sets are: Montana and North Dakota barley; Illinois and Indiana corn; Georgia, Florida, North Carolina, South Carolina, Tennessee, Virginia and Texas cotton; Alabama, Arkansas, Illinois, Iowa, Louisiana, Michigan, Minnesota, Missouri, North Carolina, South Carolina and Tennessee soybeans; Kansas, Montana and North Dakota wheat.

The data files are “cleaned”, organized, and the required variables are calculated. The first cleanup program removes inapplicable variables. The yield history files contain many variables that are not pertinent to this thesis; these variables are removed to reduce the size of the data files. The first cleanup program also removes observations whose coefficient of variation (standard deviation / mean) for yield is less than 0.025 and observations that reported yields greater than 10,000 lbs. per acre for cotton.¹⁹ Observations are removed by their coefficient of variation and yield because it is unlikely that these circumstances would exist and are therefore treated as entry errors.

A second cleanup program accomplishes four tasks. First, observations that did not report yields for 1995 through 1998 and farms that participated in the CAT program are removed.²⁰ Second, yields for missing entries in the yield history from 1987 to 1994 are simulated. Third, the indemnity payment and liability for each observation from 1995 to 1998 are calculated and then aggregated to form an observation for each multiple unit farm. Fourth, the observations from each farm that insured as a multiple unit are aggregated to form consolidated units.

Observations that did not have a complete yield history for 1995 to 1998 are removed because a loss cost ratio cannot be calculated for years with a missing yield. Missing yields before 1995 may be simulated but it would be inappropriate to simulate the yields from which indemnities and liability are calculated. Therefore, the loss cost

¹⁹ Appropriate boundaries are set for the various crops. The values 0.025 and 10,000 pounds for cotton were arbitrarily determined.

²⁰ The CAT program is a catastrophic insurance policy that insures “... 50 percent of the farm’s individual crop at 60 percent of its expected market price at a very low cost (Smith 1995).”

ratio is calculated for 1995 through 1998, not for years prior to 1995. Participants in the CAT program are removed because these farmers cannot separately insure optional units.

Next, the second cleanup program calculates the loss cost ratio from 1995 to 1998 for each observation in the data sets. The first step in calculating the LCR is to determine the trigger values for each observation. The trigger value is calculated from the observations yield history. The trigger for 1995 is determined by averaging the reported yields for the previous eight years. Most observations in the data sets, however, do not have complete yield histories for 1987 through 1994. Two options were considered for calculating the trigger value of each observation for 1995 through 1998. The first option is to average the reported yields regardless of how many years for which there is a reported yield. The second option is to simulate yields from 1987 to 1994 for those missing entries so that each observation has a complete yield history from 1987 through 1998. The latter option is used because on average the simulation methods should return a yield equal to the average of the reported yields. Hence, on average, the average of the simulated yields should be no different than the mean of the reported yields. However, by simulating the yields there is a chance that the average of the simulated yields will be greater than or less than the mean of the reported yields. This adds an element of realism into the data sets. It allows an observation's average yield to be affected by a 'good year' or a 'bad year'. A bootstrapping process simulates missing yields from 1987 through 1994.

The goal of the bootstrapping is to randomly generate yields for observations that are missing entries in their yield history from 1987 through 1994. Generating yields is

accomplished by randomly selecting, with replacement, error terms from the county pool and then adding these error terms to the observation's average yield (μ_{yi}) to get a simulated yield. The process is as follows:

- 1) μ_{yi} is calculated as the mean yield of each observation, using the reported yields from 1987 through 1998.
- 2) Next, the error terms from all reported yields for each observation are collected.²¹

$$(5.1) \quad E_{it} = y_{it} - \mu_{yi}$$

Where E_{it} is the error term and y_{it} is the yield, of observation i in year t

- 3) An error term (E_{it}^R) is randomly selected from the pool of county error terms and added to the mean yield (μ_{yi}) of each observation. The result is a simulated yield ($y_{sim_{it}}$) for observation i in year t . Where year t is a year with a missing yield observation. The pool of error terms consists of all error terms collected for a county.

$$(5.2) \quad y_{sim_{it}} = E_{it}^R + \mu_{yi}$$

This bootstrapping process simulates yields for missing entries in the yield history from 1987 through 1994. Once the data set contains yields from 1987 through 1998 for each observation, indemnity payments can be calculated.

²¹ There is minimum of four error terms collected from each unit because each observation must have a reported yield from 1995 to 1998.

For simplicity, losses are based on a yield election of 65 percent. The first step in simulating indemnity payments is calculating the trigger yield ($TRIGGER_{it}$) for observation i in year t . The trigger yield is 65 percent of the mean yield for the previous eight years.

$$(5.3) \quad TRIGGER_{it} = 0.65 * ((\sum_{l=1}^8 y_{i(t-l)}) / 8) \quad t = 1995-1998$$

The total indemnity for observation i in year t ($INDEMNITY_{it}$), measured in lbs. (for cotton), is calculated as the difference between the trigger value and the realized yield, multiplied by the acreage of the observation (a_i).²²

$$(5.4) \quad \begin{aligned} INDEMNITY_{it} &= (TRIGGER_{it} - y_{it}) * a_i && \text{if } TRIGGER_{it} > y_{it} \quad t = 1995-1998 \\ &= 0 && \text{if } TRIGGER_{it} \leq y_{it} \quad t = 1995-1998 \end{aligned}$$

If the current yield is greater than or equal to the trigger value then the indemnity payment equals zero. Liability ($LIABILITY$), the maximum indemnity observation i could receive in year t , is calculated as the trigger amount multiplied by the acreage of the observation.

$$(5.5) \quad LIABILITY_{it} = TRIGGER_{it} * a_i \quad t = 1995-1998$$

The data sets now contain the liability and indemnity payment for 1995 through 1998 for each observation. For farms that insured as a multiple unit, the liability and indemnities are aggregated across the optional units that comprise each multiple unit. The liability and indemnities are summed across the optional units using an acreage weighted procedure. The result is a multiple unit liability and indemnity.

²² Indemnities are calculated in terms of pounds of cotton and not dollars since the price does not affect the loss cost ratio.

In the following equation, LIABILITY and INDEMNITY are summed across N optional units to calculate the LCR for multiple unit j in year t (LCR MULTIPLE_{jt}). The acreage of optional unit i is a_i and the acreage of the multiple unit j is $A_j = \sum_{i=1}^N a_i$.

$$(5.6) \quad \text{LCR MULTIPLE}_{jt} = \frac{\sum_{i=1}^N (\text{INDEMNITY}_{it} * a_i)}{\sum_{i=1}^N (\text{LIABILITY}_{it} * a_i)} \quad t = 1995-1998$$

This process is different from the aggregation process used to form consolidated units, described below.

The RMA yield history files do not distinguish observation's insured as a single unit from those insured as a consolidated unit.²³ Therefore, to obtain information about consolidated unit farms, the optional units insured by a multiple unit farmer are aggregated as if the farmer insured as a consolidated unit. In other words, if a farm separately insured three optional units then the three optional units are aggregated to form a consolidated unit that consists of three available optional units.

Aggregating optional units insured as a multiple unit creates data for consolidated units. Indemnities and liability are calculated as if farmers who insured as a multiple unit combined their optional units into a single parcel and insured as a consolidated unit. The first step is to calculate the yields of the consolidated unit. An acreage weighted procedure is used to calculate the yields for consolidated unit j in year t (CONSOL

²³ The data sets distinguish between multiple units and single units because for multiple units there are more than one observation with the same farm identification number and for single units there is only one observation per farm identification number.

$YIELD_{jt}$). The yield of each of the N units (y_{it}) is multiplied by its fraction of the total acreage of the consolidated unit (a_i/A_j) and then the weighted yields are summed across the optional units. The acreage of the consolidated unit (A_j) equals the sum of the acreage of the optional units, $A_j = \sum_{i=1}^N a_i$.

$$(5.7) \quad CONSOL\ YIELD_{jt} = \sum_{i=1}^N y_{it} * (a_i / A_j) \quad t=1987-1998$$

The TRIGGER, INDEMNITY and LIABILITY are calculated the same as previously shown (equations 5.3, 5.4 and 5.5) except $CONSOL\ YIELD_{jt}$ is used instead of the optional unit yield y_{it} .

$$(5.8) \quad TRIGGER_{jt} = 0.65 * ((\sum_{l=1}^8 CONSOL\ YIELD_{j(t-l)}) / 8) \quad t = 1995-1998$$

The total indemnity for consolidated unit j in year t ($INDEMNITY_{jt}$), measured in lbs. (for cotton), is calculated as the difference between the trigger value and the realized yield, multiplied by the acreage of the consolidated unit (A_j).

$$(5.9) \quad INDEMNITY_{jt} = (TRIGGER_{jt} - CONSOL\ YIELD_{jt}) * A_j$$

$$\begin{aligned} & \text{if } TRIGGER_{jt} > CONSOL\ YIELD_{jt} \quad t = 1995-1998 \\ & = 0 \quad \text{if } TRIGGER_{jt} \leq CONSOL\ YIELD_{jt} \quad t = 1995-1998 \end{aligned}$$

If the current yield is greater than or equal to the trigger value then the indemnity payment equals zero. Liability (LIABILITY), the maximum indemnity consolidated unit j could receive in year t , is calculated as the trigger amount multiplied by the acreage of the consolidated unit.

$$(5.10) \quad LIABILITY_{jt} = TRIGGER_{jt} * A_j \quad t = 1995-1998$$

The data sets now contain the liability and indemnity payment for 1995 through 1998 for each consolidated unit. The LCR for consolidated unit j in year t (LCR CONSOLIDATED) is calculated as:

$$(5.11) \quad \text{LCR CONSOLIDATED}_{jt} = \frac{\text{INDEMNITY}_{jt}}{\text{LIABILITY}_{jt}} \quad t = 1995-1998$$

This process is different from the aggregation used to form multiple unit farms because the indemnities and liability are calculated after the yields of the optional units are aggregated.

The results are data sets that have the yields, acreage, indemnity payments and liability as if the farm insured as a multiple unit and as if the farm insured as a consolidated unit. For single unit farms in the data set the values for indemnity payments and liability will be identical between the two forms of insuring cropland because these farmers insured a single parcel of land. An observation represents a single farm as identified by the insurance policy number. Recall that the original data sets contained an observation for each optional unit. The ending data sets have aggregated the optional units to the farm level and therefore contain fewer observations than the original data sets.

The data sets for each crop are divided by state. States with less than 5,000 farmers who purchased MPCl are grouped together. The states are grouped based on geographic location and similar statewide average loss cost ratios. Some states for which data are available are not included in this study because of limited observations and lack of geographic neighbors with similar loss cost ratios.

Tests of Hypothesis One

This section discusses the methods used to test the hypothesis that there is a negative relationship between the number of optional units combined into a consolidated unit and the expected LCR. Two methods are used: Regression analysis and a Monte Carlo simulation.

Monte Carlo Simulation

A Monte Carlo procedure is used to create multiple data sets that contain farms with varying numbers of optional units. If the hypothesis holds, the LCR of the randomly determined farms will show an inverse relationship to the number of optional units combined to create the consolidated units.

The data used to perform the Monte Carlo simulation contain one observation for each separately insured parcel of cropland for which a yield history was reported in 1999. The observations have not been aggregated to form farm level multiple units or consolidated units. The liability and indemnity payment was calculated for each separately insured parcel of land or “unit”. The simulation begins by randomly drawing “units” from all observations within a county. The term “units” is used for simplicity.

The units are randomly drawn with replacement in sets of varying quantity. The size of the sets ranges from two to ten units per set. Drawing of the sets is repeated one thousand times for each county and set quantity. Next, the units are aggregated to form a

consolidated unit. If the set contained four units then the four units are aggregated and the LCR is calculated. The LCR is the sum of indemnities (INDEMNITY) from 1995 through 1998 divided by the sum of the liability (LIABILITY) over the same period.

$$(5.12) \quad \text{LCR} = \Sigma(\text{INDEMNITY}_t) / \Sigma(\text{LIABILITY}_t) \quad t = 1995, \dots, 1998$$

Next, the average LCR for each set quantity is calculated at the county level and then the LCRs are averaged across counties to calculate a state wide LCR for each set quantity.

Regression analysis

The dependent variable in the model is the loss cost ratio (LCR CONSOLIDATED) of each farm, calculated as if a farmer chooses to insure as a consolidated unit. The independent variables in the model are the number of optional units combined (N) to form the consolidated unit, the average yield (YIELD) of the consolidated unit, and the average acreage per optional unit (ACREAGE).

$$(5.13) \quad \text{LCR CONSOLIDATED} = f(\text{N}, \text{YIELD}, \text{ACREAGE})$$

The number of optional units combined to form the consolidated unit is included in the regression model because it is the focus of the hypothesis. The average yield is included because premiums, as determined by RMA, are differentiated by average yields. The average acreage per optional unit is included in the model because acreage may also indicate spatial diversity. Although it is not expected that acreage will affect the expected LCR, the RMA is involved in pilot programs that provide premium discounts to farmers with large acreages.

The number of units enters the regression equation as $1/N^{\gamma}$. Entering the variable into the regression in this form allows substantial flexibility in the relationship between the dependent variable and the number of available optional units. The variable representing the average yield of the consolidated unit (YIELD) is calculated as the average of the consolidated unit yields from 1979 through 1998. The consolidated unit yields are a share-weighted average of the reported yields of the optional units combined to form the consolidated unit.²⁴ The average acreage per unit (ACREAGE) is calculated as the average acreage of the consolidated units from 1996 through 1998, divided by the number of optional units contained within the consolidated unit. For example if there are four optional units combined to form a consolidated unit and the average total acreage of the optional units is 400 acres, the average acreage per optional unit is 100 acres ($400 \div 4$).

Equation 5.14 shows the LCR CONSOLIDATED as a function of the number of optional units (N), the average yield (YIELD), and the average acreage per unit (ACREAGE). Model A is:

$$(5.14) \quad \text{LCR CONSOLIDATED} = \alpha_1 + \alpha_2 * (1 - (1/N^{\gamma})) + \alpha_3 * \text{YIELD} + \alpha_4 * \text{ACREAGE}$$

Model A may be simplified. The process of transforming model A into a form that is simpler to estimate is straightforward. The model is:

$$(5.15) \quad \text{LCR CONSOLIDATED} = \beta_0 - \alpha_2 * (1/N^{\gamma}) + \alpha_3 * \text{YIELD} + \alpha_4 * \text{ACREAGE}$$

Where $\beta_0 = \alpha_1 + \alpha_2$. Equation 5.15 is estimated empirically. The transformation does not affect the estimated coefficients or the intercept term. In the results chapter, all references

²⁴ The simulated yields for 1987 through 1994 are not used in the calculation of the average yield.

to the intercept term are the value of $\alpha_1 = \beta_0 - \alpha_2$, not the value β_0 which is actually estimated.

The relationship between the number of units insured as a consolidated unit and expected indemnities is estimated as a Tobit model. The Tobit model is used because the dependent variable is truncated at zero. The dependent variable, LCR CONSOLIDATED, is censored because the LCR must be greater than or equal to zero. The data set is a censored data set because the values of the independent variables corresponding to a dependent variable whose value resides at the limit (i.e., the LCR CONSOLIDATED equals zero) are known. In other words, if the value of the dependent variable is equal to the limiting value, the values of the corresponding independent variable are known. Therefore, the data set is a censored data set. If the values of the independent variable were unknown for values of the dependent variable that reside at the limiting value, the data set would be a truncated data set. Because the dependent variable is censored, OLS would cause the coefficient estimates to be inconsistent.²⁵

James Tobin developed the Tobit model in 1958. Before Tobin's article, censored data sets were analyzed by including only those observations above the limiting value. This technique did not include potentially important information contained in those observations that reside at the limit. Appendix B discusses the standard Tobit model.

Johnston and Dinardo (1997) suggest a simple test to evaluate the appropriateness of the use of the Tobit model. The Tobit model may be inappropriate to use if the explanatory variables have different effects on the probability of the dependent variable

²⁵ The inconsistency is caused because the expected value of the least squares estimator of the coefficients (β_{ols}) does not equal the coefficients (β). Proof found in appendix A.

being above the limiting value and on the dependent variable conditional on the dependent variable being greater than the limiting value. They contend that the “likelihood estimates from the normalized Tobit coefficients should be the same as the Probit coefficients from the same data, treating nonzero values as 1 and zero values as 0 (Johnston 1997).” A comparison of the regression techniques found that the Tobit and Probit coefficients are similar. Therefore, it is concluded that the Tobit model is the appropriate statistical model.

The inclusion of the nonlinear γ parameter in model A presents minor difficulties in the estimation of the model. The model is estimated by iterating on values of γ until the log-likelihood of the regression is maximized. The estimation begins with a starting value of one for γ and searches over values for γ to maximize the log-likelihood. Randall J. Olsen (1978) has shown that the log-likelihood function of a Tobit model is strictly concave down which facilitates the search process and insures a global maximum.

After the value of γ that maximizes the log-likelihood function is found and a preliminary test provides assurance that Tobit regression is the appropriate method to estimate the model, it is important to note what information the Tobit results provide. Because of the censored data set, the regression coefficients estimated by Tobit regression cannot be directly interpreted as can the estimates from an OLS regression. The coefficients from a Tobit regression need to be transformed to calculate the unconditional expected value of the dependent variable. The transformation requires the calculation of an index value. The index value (I) is the conditional expected value of the dependent variable

$$(5.16) \quad I = \delta_1 + \delta_2(1-(1/N^\gamma)) + \delta_3 * \overline{YLD} + \delta_4 * \overline{ACR}$$

where $\delta_i = \alpha_i/\sigma$ and σ is the standard error of the estimate. Because the regression errors are assumed to be normally distributed, the unconditional expected value (LCR CONSOLIDATED^E) is calculated as:

$$(5.17) \quad \text{LCR CONSOLIDATED}^E = \sigma * [I * F(I) + f(I)]$$

Where $F(I)$ and $f(I)$ are the cumulative density function and probability density function of the standard normal distribution. After the transformation, LCR CONSOLIDATED^E may be plotted as a function of the number of units and the coefficients may be interpreted like those from an OLS regression.

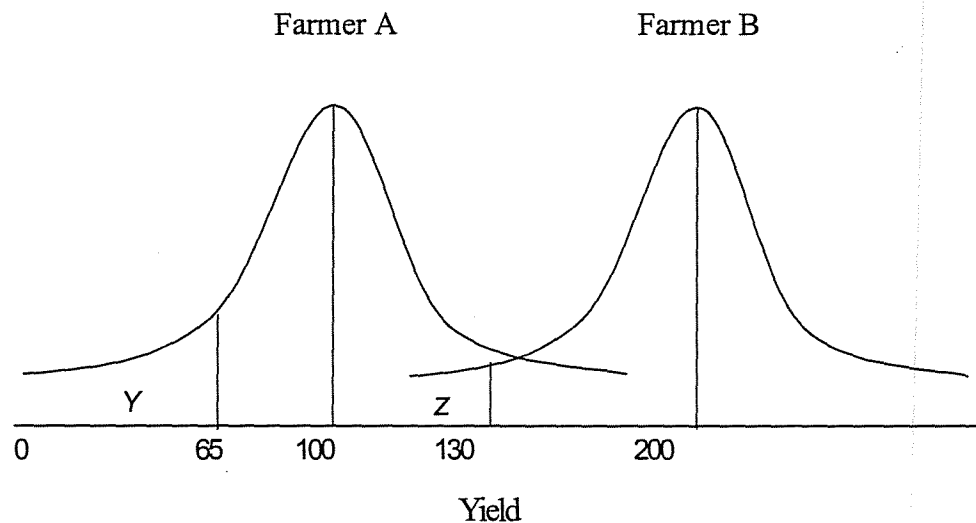
From chapter four, the prediction is that as an additional optional unit is combined into a consolidated unit the LCR CONSOLIDATED will decrease. Because N enters the equation as a nonlinear parameter, determining the expected sign of α_2 is not straightforward. The sign of α_2 that is consistent with the prediction is dependent on the sign of γ . Recall that the value of γ used in the estimation is predetermined. If γ is greater than zero, α_2 is expected to be less than zero. If γ is less than zero but greater than negative one, α_2 is expected to be greater than zero. If γ is less than negative one, the hypothesis is not supported.

The sign on α_3 is expected to be negative. If two farms have an identical average yield standard deviation and insure the same percentage of their yields, the farmer with

the greater average yield has a smaller probability of receiving an indemnity payment.²⁶

This is illustrated in Figure 4.

Figure 4: The Probability of Receiving an Indemnity Payment



In this example, both farmers choose a yield election of 65 percent and their average yield has a standard deviation of 35. If farmer A is to receive a payment his yield must fall below 65, a difference from the mean of one standard deviation, and this probability is represented by the area Y. If farmer B is to receive a payment, his yield must fall below 130, a difference from the mean of two standard deviations, and the area Z represents the probability of this occurring. Therefore, assuming identical standard deviations, the probability of an indemnity payment for farm A is greater than the

²⁶ Empirical work done at Montana State University has shown that the standard deviation is independent of the mean. Skees and Reed (1986) show empirically a constant standard deviation across farms with different expected yields for corn and soybeans grown in Illinois and Kentucky.

probability of receiving an indemnity payment for farm B. Hence, the expected negative sign on α_3 .

For α_4 , there is no expected sign. If yields are perfectly correlated across acres then acreage would not affect expected LCRs. If there is not perfect correlation of yields across acres then α_4 is expected to be negative. Further, the literature on crop insurance does not provide any guidance towards the determination of an expected sign. Therefore, no prediction is made for α_4 .

Tests of Hypothesis Two

This section presents the methods used to test the hypothesis that there is a positive relationship between the numbers of separately insured optional units and expected indemnities. The hypothesis is tested using three different regression models. The dependent variable in the models used to test hypothesis two (LCR MULTIPLE) is calculated as if farmers with available optional units insure those units separately rather than combining all available optional units into a single parcel as in LCR CONSOLIDATED. LCR MULTIPLE is greater than or equal to LCR CONSOLIDATED as shown in chapter four.

Model 1 is identical to model A except for the dependent variable. Model one is:

$$(5.18) \quad \text{LCR MULTIPLE} = \theta_1 + \theta_2(1 - (1/N^{\gamma})) + \theta_3 * \text{YIELD} + \theta_4 * \text{ACREAGE}$$

The independent variables N, YIELD, and ACREAGE are identical to those used in model A. The methods used to estimate model one are the same as those used to estimate model A.

The second model uses a dummy variable for the number of optional units insured. The dummy variable (DUM1) equals zero if a single unit is insured and is equal to one if two or more optional units are insured. Model two is:

$$(5.19) \quad \text{LCR MULTIPLE} = \lambda_1 + \lambda_2 * \text{DUM1} + \lambda_3 * \text{YIELD} + \lambda_4 * \text{ACREAGE}$$

In model two, YIELD and ACREAGE are calculated the same as before. Model 2 and model 3 are also estimated using tobit regression.

The third model uses two dummy variables to represent the number of units insured. The first dummy variable (DUM2) represents whether two units are insured and the second dummy variable (DUM3) represents whether three or more optional units are insured. Model three is:

$$(5.20) \quad \text{LCR MULTIPLE} = \Omega_1 + \Omega_2 * \text{DUM2} + \Omega_3 * \text{DUM3} + \Omega_4 * \text{YIELD} + \Omega_5 * \text{ACREAGE}$$

The intercept terms in models one, two, and three are equivalent. The intercept terms represent the insurance of a single unit. Another similarity between the models is that the dummy variable in model two, DUM1, is a linear combination of DUM2 and DUM3 in model three which is recognized in the statistical tests regarding the specification of model three.

Although the models used to test hypothesis two are very similar, the differences are important. The goal is to test whether or not a positive relationship exists between the number of optional units insured and expected indemnities. Models two and three will

provide a statistical test of the hypothesis since model two is a restricted form of model three. The null hypothesis will be whether the coefficients on DUM2 and DUM3 are equal. If the null holds then hypothesis two fails to be supported. In addition, model one allows a more flexible relationship to develop between the expected LCR and optional units than do the models that use dummy variables. Therefore, model one may provide a better estimator of the data than model two or model three.

The prediction from chapter 4 is that hypothesis two will not be supported empirically. Therefore, it is expected that θ_2 , λ_2 , Ω_2 , and Ω_3 will not be significant. Nevertheless, the findings of Knight and Coble (1999) support hypothesis two. In addition, the RMA's practice of providing a ten percent discount on premiums to farmers who do not insure optional units also supports hypothesis two. Therefore, it would not be surprising to find that θ_2 , λ_2 , Ω_2 , and Ω_3 are positive and significant.²⁷ Note that the findings of Knight and Coble (1999) and the RMA's practice of discounting premiums do not predict that the value of Ω_3 will be significantly different from the value of Ω_2 . Knight and Coble analyzed the difference in indemnity payments for farmers that insured a single parcel of land versus those farmers who separately insured two or more optional units separately (the specification of model 2). Knight and Coble did not analyze the difference in indemnity payments for farmers who separately insured two optional units versus those who separately insured three or more optional units (the specification of

²⁷ The sign of θ_2 that shows a positive relationship between the expected LCR and the number of separately insured optional units is dependent upon the value of γ used in the regression. If γ is greater than zero, θ_2 is expected to be greater than zero. If γ is less than zero but greater than negative one, θ_2 is expected to be less than zero. If γ is less than negative one, the hypothesis is not supported.

model 3). The RMA does not distinguish rates between how many optional units are separately insured, it provides a discount only if two or more optional units are not separately insured (the specification of model 2). Therefore, based on Knight and Coble and the RMA's practice of discounting premiums it may be the case that θ_2 , λ_2 , Ω_2 , and Ω_3 are positive and significant, if this is the case, it is predicted that Ω_3 will not be statistically different from Ω_2 .

CHAPTER 6

RESULTS

Currently the number of optional units a farmer insures, either as a multiple unit or as a consolidated unit, does not affect MPCCI per acre premiums. This implies that the RMA assumes the number of optional units either insured separately or combined into a consolidated unit does not effect expected per acre indemnity payments. Testing the two hypotheses allows this assumption to be evaluated because an implication of both hypotheses is that the number of optional units insured does affect expected per acre indemnities. It is important to note that the hypotheses are independent of each other. Each hypothesis can exist independently and can be confirmed independently. They specifically deal with different ways of organizing optional units under a MPCCI policy, as a multiple unit or as a consolidated unit.

Hypothesis One

A regression model and a Monte Carlo simulation are used to test hypothesis one. The prediction from chapter four is that the empirical results will support hypothesis one. The regression and Monte Carlo estimates show a decrease in the expected LCR as a

farmer combines an additional optional unit into a consolidated unit. Therefore, RMA's current premium schedule as illustrated in Figure 1 (pg. 3) may not reflect changes in the expected LCRs of farmers who combine an additional optional unit into a consolidated unit.

Regression Estimates

Model A tests the hypothesis that there is a negative relationship between the expected LCR and the number of optional units combined to form a consolidated unit. The regression estimates are shown in Tables 4 through 6. The estimates indicate that the signs of the coefficients for the intercept, the number of optional units ($1/N^\gamma$), and YIELD are as predicted for each state-crop combination and the t-statistics show significance at the 99 percent confidence level for the coefficients.²⁸ Therefore, the regression estimates show strong support for the acceptance of hypothesis one.

Recall that the sign of the coefficient for $1/N^\gamma$ that supports hypothesis one is dependent on the value of γ used in the regression. In the case of Texas dryland cotton, the value of γ used is 0.58 and for Iowa soybeans γ equals -0.44 . From chapter five, if γ is greater than zero then the predicted sign of the coefficient for $1/N^\gamma$ is less than zero and if γ is less than zero but greater than negative one the predicted sign is greater than zero. The sign of the coefficient for $1/N^\gamma$ for Texas dryland cotton is -0.365 and 0.59 for Iowa soybeans. Therefore, for Texas dryland cotton and Iowa soybeans, the greater the number

²⁸ The coefficient for the number of optional units in Georgia and Florida cotton is significant at the 90 percent confidence level.

Table 4: Consolidated unit Tobit estimates, **Model A**

Coefficients with respect to the independent variables (t-stats)								
Variable	Predicted Sign	Barley	Corn		Cotton			
		Montana and North Dakota - Dryland	Illinois - Dryland	Indiana- Dryland	Texas-Irrigated	Texas-Dryland	Georgia and Florida- Dryland	NC, SC, TN, and VA - Dryland
$1/N^{\gamma}$	- if $\gamma > 0$	-0.77	-2.13	-7.26	-0.497	-0.365	0.097	1.82
	+ if $-1 < \gamma < 0$	(-5.16)	(-18.54)	(-11.63)	(-14.34)	(-11.20)	(1.90)	(3.70)
YIELD	-	-0.037	-0.025	-0.026	-0.00149	-0.0043	-0.00277	-0.00405
		(-22.46)	(-66.51)	(-36.20)	(23.19)	(77.20)	(16.83)	(16.51)
ACREAGE	?	-0.57E-04	-0.28E-03	0.57E-04	-0.8E-03	0.47E-03	0.41E-03	0.59E-03
		(-0.38)	(-3.25)	(0.37)	(7.30)	(7.52)	(2.42)	(2.15)
Intercept	+	1.52	2.61	2.95	1.07	2.92	2.36	2.52
		(3.89)	(4.37)	(6.83)	(11.50)	(72.10)	(17.47)	(7.84)
γ		0.46	0.16	0.04	0.94	0.58	-0.50	-0.08
R-Squared		0.148	0.188	0.223	0.077	.320	.176	.225
Number of observations		3584	35754	9122	11278	17401	1749	1309

Table 5: Consolidated unit Tobit estimates, **Model A** cont.

Coefficients with respect to the independent variables (t-stats)							
Variable	Soybeans						
	Iowa	Illinois	Indiana and Michigan	Minnesota	Missouri	AL, AR and LA	NC, SC and TN
$1/N^y$	0.59 (15.69)	-3.21 (-12.57)	-0.21 (-4.63)	-1.47 (-13.57)	3.52 (9.47)	-0.41 (-4.08)	1.10 (3.92)
YIELD	-0.096 (-64.31)	-0.083 (-55.93)	-0.089 (-32.46)	-0.079 (-41.00)	-0.081 (-34.81)	-0.10 (-27.28)	-0.078 (-21.80)
ACREAGE	0.26E-03 (1.63)	0.29E-05 (0.023)	0.29E-04 (0.18)	-0.87E-04 (-0.52)	0.49E-03 (4.03)	-0.14E-03 (-2.03)	0.19E-03 (1.44)
Intercept	3.91 (46.09)	2.32 (3.67)	2.54 (24.52)	1.79 (2.59)	2.26 (10.89)	3.25 (20.24)	2.33 (8.20)
γ	-0.44	0.10	4.70	0.28	-0.08	0.70	-0.14
R-Squared	0.194	0.148	0.263	0.132	0.183	0.332	0.317
Number of observations	45663	3124	5080	21627	8315	1994	1599

Table 6: Consolidated unit Tobit estimates, **Model A** cont.

Variable	Coefficients with respect to the independent variables (t-stats)		
	Wheat		
	Kansas	Montana	North Dakota
1/N ^γ	-0.53 (-23.69)	0.47 (9.15)	1.69 (17.25)
YIELD	-0.066 (-79.30)	-0.043 (-25.39)	-0.047 (-45.70)
ACREAGE	0.15E-02 (19.47)	0.24E-03 (2.84)	0.14E-03 (2.22)
Intercept	2.60 (57.64)	1.01 (18.88)	1.32 (28.65)
γ	0.64	-0.38	-0.14
R-Squared	0.204	0.111	0.117
Number of observations	35783	7114	24466

of optional units a farmer insures as a consolidated unit the lower the expected LCR for the farm. This result holds for each of the state-crop combinations. In addition Figures 5 through 9 illustrate the inverse relationship between the expected LCR and the number of optional units combined to form a consolidated unit.

The predicted sign of the coefficient for YIELD is negative. The estimates show that for each state-crop combination the coefficient for YIELD is negative and significant at the 99 percent confidence level. This means that farmers with a higher average yield have a lower expected LCR. There was no prediction made for the coefficient for ACREAGE. It is unclear how acreage affects the expected LCR. The estimates show that the sign of the coefficient for ACREAGE is not consistent across state-crop combinations. Twelve of the seventeen state-crop combinations show a positive sign for the coefficient for ACREAGE. Of the twelve coefficients for ACREAGE that are positive, eight are significant at the 95 percent confidence level. Of the five coefficients for ACREAGE that are negative, two are significant at the 95 percent confidence level. Implications of the estimates of the coefficients for ACREAGE are unclear.

Hypothesis one does not make any assertions as to the magnitude of the effect of the number of optional units insured as a consolidated unit on the expected LCR. Therefore, the effects of marginal changes in the number of optional units on the expected LCR are not provided. There are only four years of LCR data used in the regression analysis. Using four years of data to make remarks about the magnitude of the effect of optional units on the expected LCR is not appropriate. Nevertheless, the support for a negative relationship is substantial even if only four years of data are used.

Monte Carlo Simulation Results

Two of the resulting plots from the Monte Carlo simulation exercises are displayed in Figures 10 and 11. The plots exhibit a negative relationship between the expected LCR and the number of optional units combined to form a consolidated unit. Additional plots are not provided because the plots for each of the state-crop combinations show an inverse relationship between the expected LCR and the number of optional units combined. The Monte Carlo simulations support the acceptance of hypothesis one.

Hypothesis Two

Three regression models are used to test the hypothesis of a positive relationship between the expected LCR and the number of separately insured optional units. The prediction from chapter four is that hypothesis two will not be supported. Nevertheless, the regression estimates support hypothesis two. Tables 7 through 12 present the estimates from the regression analysis. Figures 5 through 8 show plots of the expected LCRs for each of the three models for some of the state-crop combinations. The regression estimates show an increase in the expected LCR when farmers move from insuring one optional unit to insuring two optional units. The estimates do not provide a clear indication whether the expected LCR increases when farmers move from separately insuring two optional units to separately insuring three or more optional units.

Table 7: Optional unit Tobit estimates

Coefficients with respect to the independent variables (t-stats)										
Variable	Predicted sign	Barley			Corn					
		Montana and North Dakota			Illinois			Indiana		
		Model 1	Model 2	Model 3	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
$1/N^{\gamma}$	+ if $\gamma > 0$ - if $-1 < \gamma < 0$	-0.19 (-2.35)	- -	- -	-17.11 (-28.41)	- -	- -	0.54 (9.68)	- -	- -
YIELD	-	-0.0362 (-22.70)	-0.0363 (-22.74)	-0.0362 (-22.74)	-0.0257 (-72.12)	-0.0253 (-71.48)	-0.0256 (-71.85)	-0.0279 (-40.29)	-0.0277 (-40.18)	-0.0278 (-40.25)
ACREAGE	?	-0.22E-03 (-1.52)	-0.15E-03 (-1.09)	-0.15E-03 (-1.09)	-0.43E-03 (-5.04)	-0.44E-03 (-5.19)	-0.43E-03 (-5.08)	-0.31E-04 (-0.21)	-0.29E-04 (-0.19)	-0.29E-04 (-0.19)
DUM1	+	- -	0.11 (2.30)	- -	- -	0.37 (25.34)	- -	- -	0.23 (9.06)	- -
DUM2	+	- -	- -	0.10 (18.13)	- -	- -	0.25 (59.18)	- -	- -	0.16 (36.35)
DUM3	+	- -	- -	0.11 (15.82)	- -	- -	0.48 (62.76)	- -	- -	0.28 (37.37)
Intercept	+	1.56 (11.56)	1.56 (19.49)	1.56 (19.49)	2.72 (23.82)	2.68 (56.86)	2.71 (57.23)	3.26 (35.82)	3.25 (36.63)	3.26 (36.71)
γ		-0.44	-	-	-0.02	-	-	0.52	-	-
R-Squared		0.139	0.139	0.139	0.159	0.162	0.160	0.203	0.205	0.203
Number of observations		3584	3584	3584	35754	35754	35754	9122	9122	9122

Table 8: Optional unit Tobit estimates cont.

Coefficients with respect to the independent variables (t-stats)									
Variable	Cotton								
	Texas - Irrigated			Texas - Dryland			Georgia and Florida - Dryland		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
$1/N^{\gamma}$	-1.64 (-8.45)	- -	- -	0.04 (2.18)	- -	- -	0.40 (6.29)	- -	- -
YIELD	-0.0016 (-25.72)	-0.0016 (-25.70)	-0.0016 (-25.72)	-0.0044 (-79.48)	-0.0044 (-79.47)	-0.0044 (-79.47)	-0.0030 (-18.28)	-0.0030 (-18.26)	-0.0030 (-18.26)
ACREAGE	-0.79E-03 (-7.69)	-0.79E-03 (-7.65)	-0.79E-03 (-7.61)	0.44E-03 (7.05)	0.44E-03 (7.05)	0.44E-03 (7.05)	0.35E-03 (2.11)	0.35E-03 (2.11)	0.35E-03 (2.09)
DUM1	- -	0.15 (7.54)	- -	- -	0.03 (2.12)	- -	- -	0.31 (6.11)	- -
DUM2	- -	- -	0.10 (28.38)	- -	- -	0.02 (101.63)	- -	- -	0.25 (21.15)
DUM3	- -	- -	0.19 (30.83)	- -	- -	0.04 (106.29)	- -	- -	0.35 (22.67)
Intercept	1.22 (2.07)	1.22 (29.02)	1.22 (29.02)	3.02 (109.65)	3.02 (120.79)	3.02 (120.80)	2.40 (22.88)	2.41 (21.15)	2.41 (21.15)
γ	-0.08	-	-	1.75	-	-	1.45	-	-
R-Squared	0.053	0.054	0.053	0.331	0.331	0.331	0.193	0.193	0.193
Number of observations	11278	11278	11278	17401	17401	17401	1749	1749	1749

Table 9: Optional unit Tobit estimates cont.

Coefficients with respect to the independent variables (t-stats)									
Variable	Cotton			Soybeans					
	NC, SC, TN, and VA - Dryland			Iowa			Illinois		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
$1/N^{\gamma}$	0.49 (5.76)	- -	- -	-1.13 (-24.54)	- -	- -	1.58 (22.59)	- -	- -
YIELD	-0.0044 (-18.51)	-0.0044 (-18.48)	-0.0044 (-18.51)	-0.0947 (-70.24)	-0.0938 (-69.90)	-0.0944 (-70.09)	-0.0843 (-61.46)	-0.0837 (-61.23)	-0.0842 (-61.38)
ACREAGE	0.65E-03 (2.51)	0.60E-03 (2.33)	0.65E-03 (2.52)	0.12E-03 (0.81)	0.12E-03 (0.84)	0.13E-03 (0.89)	-0.22E-03 (-1.86)	-0.20E-03 (-1.67)	-0.20E-03 (-1.67)
DUM1	- -	0.34 (5.47)	- -	- -	0.35 (21.40)	- -	- -	0.40 (20.80)	- -
DUM2	- -	- -	0.21 (16.94)	- -	- -	0.25 (50.30)	- -	- -	0.30 (43.72)
DUM3	- -	- -	0.39 (19.08)	- -	- -	0.45 (52.45)	- -	- -	0.50 (46.34)
Intercept	2.83 (19.21)	2.82 (16.89)	2.83 (16.89)	2.75 (23.35)	2.71 (48.94)	2.73 (49.21)	2.37 (44.10)	2.35 (41.34)	2.36 (41.55)
γ	1.00	-	-	-0.26	-	-	0.28	-	-
R-Squared	0.233	0.234	0.235	0.163	0.163	0.163	0.142	0.143	0.142
Number of observations	1309	1309	1309	45663	45663	45663	31240	31240	31240

Table 10: Optional unit Tobit estimates cont.

Coefficients with respect to the independent variables (t-stats)									
Variable	Soybeans								
	Indiana and Michigan			Minnesota			Missouri		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
$1/N^y$	0.63 (6.95)	- -	- -	-1.50 (19.08)	- -	- -	0.51 (9.58)	- -	- -
YIELD	-0.0894 (-34.04)	-0.0896 (-34.15)	-0.0894 (-34.03)	-0.0798 (-46.69)	-0.079 (-46.42)	-0.0797 (-46.61)	-0.0821 (-37.04)	-0.0818 (-36.97)	-0.0820 (-37.00)
ACREAGE	-0.14E-03 (-0.91)	-0.92E-04 (-0.61)	-0.13E-03 (-0.85)	-0.27E-03 (-1.83)	-0.19E-03 (-1.27)	-0.24E-03 (-1.58)	0.34E-03 (2.80)	0.34E-03 (2.83)	0.34E-03 (2.85)
DUM1	- -	0.26 (6.55)	- -	- -	0.34 (16.07)	- -	- -	0.27 (9.27)	- -
DUM2	- -	- -	0.19 (28.49)	- -	- -	0.22 (30.89)	- -	- -	0.23 (30.17)
DUM3	- -	- -	0.35 (29.35)	- -	- -	0.45 (33.64)	- -	- -	0.30 (30.14)
Intercept	3.06 (27.85)	3.06 (28.53)	3.06 (28.45)	1.90 (3.99)	1.87 (29.05)	1.90 (29.36)	2.34 (30.17)	2.33 (29.53)	2.34 (29.57)
γ	0.58	-	-	-0.20	-	-	0.76	-	-
R-Squared	0.259	0.261	0.259	0.109	0.111	0.110	0.164	0.165	0.165
Number of observations	5080	5080	5080	21627	21627	21627	8315	8315	8315

Table 11: Optional unit Tobit estimates cont.

Coefficients with respect to the independent variables (t-stats)						
Variable	Soybeans					
	AL, AR, and LA			NC, SC, and TN		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
1/N ^γ	0.31 (2.42)	- -	- -	0.14 (2.49)	- -	- -
YIELD	-0.101 (-27.97)	-0.101 (-27.95)	-0.1012 (-27.97)	-0.082 (-23.26)	-0.082 (-23.26)	-0.0820 (-23.27)
ACREAGE	-0.15E-03 (-2.27)	-0.16E-03 (-2.31)	-0.16E-03 (-2.30)	0.21E-03 (1.66)	0.21E-03 (1.66)	0.21E-03 (1.66)
DUM1	- -	0.11 (2.11)	- -	- -	0.14 (2.49)	- -
DUM2	- -	- -	0.03 (28.05)	- -	- -	0.16 (21.71)
DUM3	- -	- -	0.17 (29.14)	- -	- -	0.12 (22.76)
Intercept	3.36 (23.14)	3.36 (31.26)	3.37 (31.28)	2.56 (23.70)	2.56 (23.48)	2.56 (23.49)
γ	0.46	-	-	34.05	-	-
R-Squared	0.332	0.333	0.332	0.321	0.321	0.322
Number of observations	1994	1994	1994	1599	1599	1599

Table 12: Optional unit Tobit estimates cont.

Coefficients with respect to the independent variables (t-stats)									
Variable	Kansas			Wheat Montana			North Dakota		
	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3	Model 1	Model 2	Model 3
$1/N^y$	0.25 (6.56)	- -	- -	0.28 (5.50)	- -	- -	0.40 (12.83)	- -	- -
YIELD	-0.0690 (-83.66)	-0.0690 (-83.61)	-0.0690 (-83.64)	-0.0429 (-26.24)	-0.0430 (-26.27)	-0.0430 (-26.24)	-0.0453 (-46.30)	-0.0450 (-46.14)	-0.0453 (-46.29)
ACREAGE	0.13E-02 (17.80)	0.14E-02 (17.87)	0.13E-02 (17.82)	0.19E-03 (2.38)	0.20E-03 (2.44)	0.19E-03 (2.43)	0.10E-05 (0.16)	0.19E-04 (0.31)	0.12E-04 (0.20)
DUM1	- -	0.07 (5.80)	- -	- -	0.15 (5.31)	- -	- -	0.18 (12.09)	- -
DUM2	- -	- -	0.05 (86.32)	- -	- -	0.13 (20.45)	- -	- -	0.13 (43.09)
DUM3	- -	- -	0.09 (87.84)	- -	- -	0.16 (21.16)	- -	- -	0.23 (43.43)
Intercept	2.80 (64.70)	2.79 (90.42)	2.79 (90.45)	1.07 (20.37)	1.07 (20.87)	1.07 (20.86)	1.35 (40.83)	1.34 (43.64)	1.35 (43.80)
γ	0.28	-	-	0.76	-	-	0.64	-	-
R-Squared	0.196	0.197	0.196	0.092	0.092	0.092	0.093	0.093	0.093
Number of observations	35783	35783	35783	7114	7114	7114	24466	24466	24466

These results support the RMA's multiple unit premium schedule, as illustrated in Figure 1 (pg.3), is correct. Figure 1 shows an increase in per acre premiums when farmers move from insuring a single unit to separately insuring two optional units but no increase in per acre premiums when farmers move from separately insuring two optional units to separately insuring three or more optional units.

The intercept terms in model 1, model 2, and model 3 (Tables 7-12) should be similar within each state-crop combination. Reviewing the estimates shows this to be the case and shows that for each model and state-crop combination the intercept is positive and significant at the 99 percent confidence level. The predicted value of the coefficient for the number of optional units separately insured in model 1 is zero. The prediction is that the number of separately insured optional units ($1/N^\gamma$) would not affect the expected LCR. Nevertheless, the estimates show a coefficient for $1/N^\gamma$ that is significantly different from zero at the 98 percent confidence level.²⁹ In addition, the coefficients for $1/N^\gamma$ support the acceptance of hypothesis two. For example, the value of γ used in the regression analysis for Iowa soybeans is -0.26 . A positive relationship between the number of separately insured optional units and the expected LCR is estimated if the coefficient for $1/N^\gamma$ is less than zero. For Iowa soybeans the coefficient is -1.13 . Therefore, as a farmer separately insures an additional optional unit the expected LCR increases. This result holds for all of the state-crop combinations.

²⁹ The coefficient for the number of separately insured optional units for Texas dryland cotton is significant at the 95 percent confidence level.

The coefficient for YIELD in models 1 through 3 is significant at the 99 percent confidence level and of the predicted sign, negative. The predicted sign is the same as that for model A for the same reasons. The estimates show that YIELD has a similar coefficient across models 1 through 3 for each state-crop combination. If this were not the case, it would be an indication of problems with the regression analysis. The coefficient for ACREAGE has no predicted sign. The estimates show coefficient values that range from negative and significant to positive and significant with nine of the seventeen state-crop combinations not having a significant coefficient for ACREAGE.

The predicted value of the coefficient for the dummy variables is zero. Nevertheless, the t statistics show that the coefficients for DUM1, DUM2, and DUM3 are all significantly different from zero at the 95 percent confidence level. In addition, the sign of the coefficients for DUM1, DUM2, and DUM3 are all positive and the coefficient for DUM3 is greater in magnitude than the coefficient for DUM2. This provides support for the acceptance of hypothesis two.³⁰

Reviewing the regression results and plots does not provide guidance to which of the three models fits the data best. The relationship between the expected LCR and the number of separately insured optional units is flexible in model 1 but it is imbedded in model 2 and model 3. Model 2 is consistent with the RMA's practice of a flat rate premium per acre for farmers who separately insure two or more optional units and model 3 is consistent with the hypothesis that there is a positive relationship between expected LCR and the number of separately insured optional units. Because model 2 is a

³⁰ The coefficient for DUM3 is not greater than the coefficient for DUM2 for NC, SC and TN soybeans.

restricted form of model 3, a likelihood ratio (LR) test may be used to distinguish between the models.³¹ However, model 1 cannot be constrained to equal model 2 or model 3. Therefore, a LR test comparing model 1 to the other models is inappropriate.

The null hypothesis of the LR test is that the coefficients for DUM2 and DUM3 in model 3 are equal. In other words, there is not a positive relationship between the expected LCR and the number of optional units separately insured when two or more optional units are insured. The format of the LR test is:

$$(6.1) \quad LR = 2[LLF^{UR} - LLF^R] \sim \chi^2(q)$$

LLF is the log of the likelihood function for the unrestricted and restricted equations. The unrestricted equation is model 3 and the restricted equation is model 2. The LR test has a chi-square distribution with q degrees of freedom. Where q is equal to the number of parameter restrictions; in this situation, q is equal to one. Table 13 lists the results of the LR test between model 2 and model 3.

The results of the LR test indicate that the null hypothesis is accepted in nine of the seventeen state-crop groups. In the eight state-crop combinations where the null is rejected the difference in the magnitude of the coefficients for DUM3 and DUM2 is small, less than 0.20 and in most cases less than 0.05. Because when the null hypothesis is rejected the magnitude of the difference between the coefficients of DUM3 and DUM2 is small, it is concluded that model 2 is preferable to model 3.

The regression estimates for model 1 indicate that there is a positive relationship between the expected LCR and the number of separately insured optional units, as

³¹ The dummy variable in model 2 is a linear combination of DUM2 and DUM3 from model three. Hence, model 3 can be restricted to equal model 2.

Table 13. Likelihood ratio tests of the null hypothesis that the coefficients on DUM2 and DUM3 from model #3 are equivalent

State-crop	LLF-#2	LLF-#3	Chi-square stat	Critical value (95%) =5.024
Barley - MT and ND	-7703.9844	-7703.9817	0.0054	Reject
Corn- Illinois	-44807.125	-44756.08	102.09	Accept
Corn- Indiana	-16879.68	-16875.263	8.834	Accept
Cotton- GA and FL	-5503.5094	-5502.2655	2.4878	Reject
Cotton- NC, SC, TN and VA	-3318.1845	-3316.3285	3.712	Reject
Cotton- TX Dry	-71117.247	-71116.839	0.816	Reject
Cotton- TX Irig	-32913.427	-32909.659	7.536	Accept
Soy- AL, AR and LA	-6101.6095	-6100.2087	2.8016	Reject
Soy- Illinois	-22806.824	-22783.262	47.124	Accept
Soy- Iowa	-31377.199	-31343.152	68.094	Accept
Soy- Michigan	-7446.3215	-7443.4688	5.7054	Accept
Soy- Minnesota	-19125.715	-19093.099	65.232	Accept
Soy- Missouri	-13495.818	-13494.509	2.618	Reject
Soy- NC, SC and TN	-4362.71	-4362.5396	0.3408	Reject
Wheat-Kansas	-112714.32	-112711.19	6.26	Accept
Wheat-Montana	-16122.197	-16122.005	0.384	Reject
Wheat-ND	-59619.524	-59610.567	17.914	Accept

illustrated in Figures 4 through 8. It is not clear whether model 1 or model 2 provides a better fit of the data.

Comparing model one to model two is not straightforward. There are no statistical tests to indicate which model is preferable. However, the fact that the R-squares of model 1 and model 2 are nearly identical across state-crop combinations lends support to the conclusion that no distinction can be made between the models.³²

One reason why the three models cannot be statistically differentiated may be contained within the makeup of the data set. In each state-crop combination, over 80 percent of the farms contain four or fewer optional units. Since there are not many observations for five or more optional units, the residual sum of squares is dominated by the observations with four or fewer optional units. Hence, the similar R-squared values for model 1 and model 2. It is suspected that if there were more farms with five or more optional units that the R-squared value and corresponding t-stats would be significantly stronger for model 1 and model 3 than for model 2. However, this claim cannot be supported here. It is a potential topic for further research.

Given that RMA does not increase per acre premiums when a farmer moves from separately insuring two optional units to separately insuring three or more optional units, the conclusion is that model 2 is the appropriate model.

³² The R-square values reported in Tables 7-12 are not exact. Rounding error caused a loss of accuracy in the numbers. Nevertheless, it is believed that the bias is not large and the R-squares are not used for hypothesis testing.

CHAPTER 7

CONCLUSION

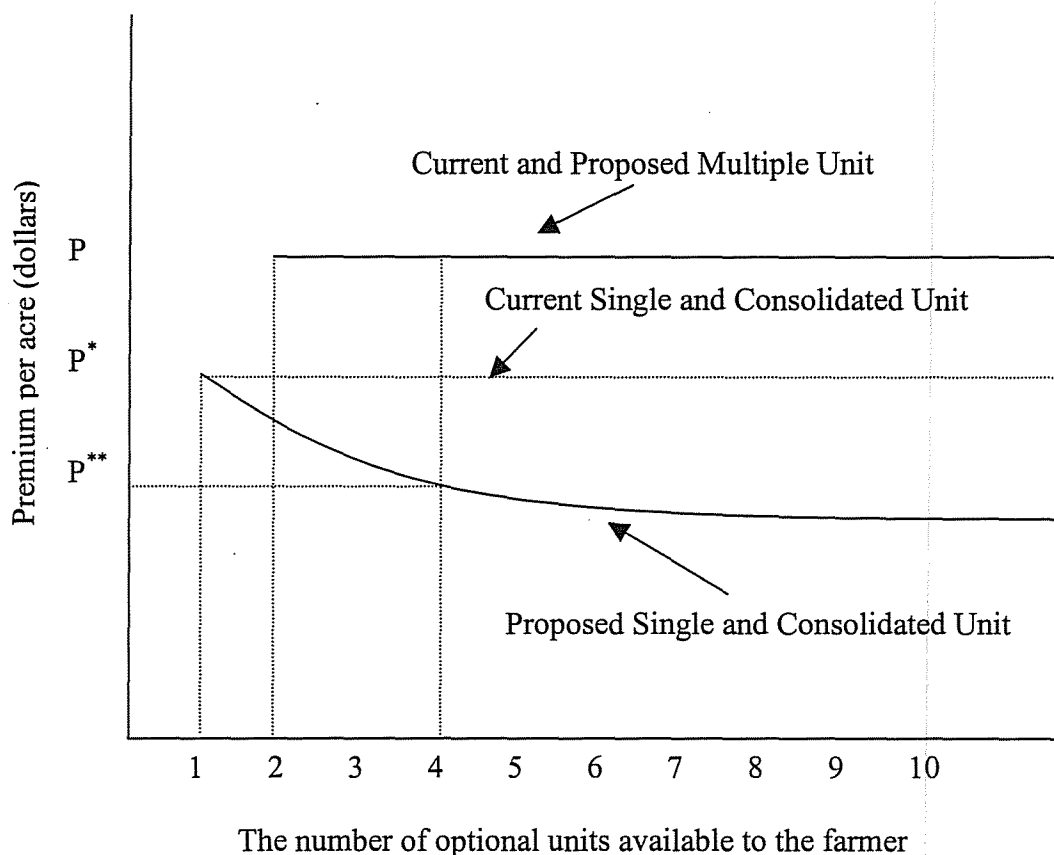
This thesis estimates the effects of optional units on expected per acre indemnity payments for farmers who insure as a consolidated or as a multiple unit. The RMA assumes, in its determination of MPCl premiums, that the number of optional units insured as a multiple or consolidated unit does not affect expected per acre indemnity payments. The empirical analysis shows that the assumption may be incorrect when farmers insure as a consolidated unit.

The starting point for hypothesis one is that the number of optional units a farmer is eligible to insure may act as a proxy for the spatial diversity of the farm. Recall that spatial diversity refers to the spatial difference within an insured parcel of land and that most optional units are located in different sections. The empirical analysis shows a negative relationship between the expected LCR and the number of optional units combined to form a consolidated unit, and therefore supports using the number of available optional units as a proxy for the spatial diversity of a farm.

The implication of using the number of available optional units as a proxy for spatial diversity is illustrated in Figure 5. The “proposed” schedule is based on the thesis analysis as shown in the plots for model A (Figures 6 through 9). The “current” schedules

represent premium schedules currently in use. Dollar amounts for P , P^* , and P^{**} are not included because they are meant for illustrative purposes only. The proposed and current schedules for multiple units are the same. The proposed schedule for farmers who insure as a consolidated unit is considerably different from the current schedule. The RMA's current schedule would charge a premium per acre of P^* for farmer who consolidated four optional units. The proposed premium schedule would charge a farmer P^{**} for insuring four optional units as a consolidated unit, a decrease in the premium per acre. The decline in the premium per acre is caused by the greater spatial diversity of the farm with four available optional units versus a farm that insures one optional unit.

Figure 5. Current and Proposed Premium Schedule



This thesis supports a per acre premium decrease for farmers who insure as a consolidated unit and that the decrease in the premium be positively related to the number of optional units combined to form the consolidated unit. Recall from chapter 2 that most farmers eligible for two or more optional units choose to insure the optional units as a multiple unit. Figure 5 shows that if the “proposed” premium schedule reflects actual expected LCRs then under the “current” premium schedule farmers would be charged a premium for insuring as a consolidated unit that is greater than their expected indemnity payment. If this is the case then farmers have an incentive to insure as a multiple unit. This incentive may help explain why most farmers eligible to insure two or more optional units overwhelmingly choose to insure as a multiple unit.

The theoretical arguments of chapter 4 predicted that hypothesis two would not be supported by the empirical analysis. Nevertheless, the empirical work supports hypothesis two. This confirms the RMA’s practice of increasing per acre premiums when farmers move from insuring as a single or consolidated unit to insuring as a multiple unit. Theoretically there is no justification for the RMA’s practice, the most plausible reason for the empirical support of hypothesis two is the presence of fraud among farmers who insure as a multiple unit. This conclusion raises the question: Why do farmers who insure as a multiple unit generate more illicit indemnity payments than farmers who insure as a single unit? It may be that farmers who insure as a multiple unit have a higher expected gain from illicit activities or that they have a lower probability of being caught.

This thesis does not suggest “how much” premiums per acre be decreased if the number of optional units is used as a proxy for spatial diversity for farmers who insure as

a consolidated unit. Recall that only four years of LCR data is used in the empirical analysis, which is not a long enough time series to accurately determine MPCI premiums. The RMA currently uses twenty years of data. A topic for further research is to gather a data set that consists of a longer time series to calculate premiums for farmers who insure as a consolidated unit using the number of available optional units as a proxy for spatial diversity.

APPENDICES

APPENDIX A

INCONSISTENCY OF OLS COEFFICIENT ESTIMATES
WHEN THE DATA SET IS CENSORED

When OLS is used on a censored data set, the coefficient estimates are inconsistent. The following is from the econometrics text of Johnston and Dinardo (Johnston 1997). The standard setup of a censored data set is:

$$y_i = y^* \text{ if } y^* > 0 \text{ and } y_i = 0 \text{ if } y^* \leq 0$$

The latent variable formulation is:

$$y^* = X\beta + e$$

Assuming that x and y^* are distributed jointly normal³³ the probability limit of the OLS coefficients, β_{ols} , is:

$$\begin{aligned} \text{Plim } \beta_{ols} &= \Sigma_{xy^*} / \Sigma_{xx} * \text{prob}(y^* > 0) + 0 * (\text{prob}(y^* \leq 0)) \\ &= \beta * \text{prob}(y^* > 0) \neq \beta \end{aligned}$$

Therefore, the OLS estimates of β , given a censored data set, are inconsistent.

³³ Although this assumption does not apply to the models used in this thesis, this example provides a general idea of the problems associated with using OLS on censored data sets.

APPENDIX B

STANDARD SETUP OF THE TOBIT MODEL

The standard set up of the Tobit model is:

$$Y_i = Y_i^* \text{ if } Y_i^* > 0 \\ = 0 \text{ if } Y_i^* \leq 0$$

The log likelihood function is split into two parts; one is the OLS maximum likelihood estimator (MLE) and the second considers the censored observations.

$$1) L = -(n/2) \text{Log}(2\Pi) - (n/2)\text{Log}(\sigma^2) - 1/(2\sigma^2)*\Sigma(Y_i - \alpha - \beta X_i)^2$$

$$2) L = \Sigma \text{Log}(F((- \alpha - \beta X_i)/\sigma))$$

F is the cumulative distribution function of a standard normal variable. A linear combination of these log likelihood functions gives the MLE for the Tobit model:

$$L = \Sigma \{ Z_i [-(1/2) \text{Log}(2\Pi) - (1/2)\text{Log}(\sigma^2) - 1/(2\sigma^2)*(Y_i - \alpha - \beta X_i)^2] + \\ (1 - Z_i) \text{Log} F((- \alpha - \beta X_i)/\sigma) \}$$

$$Z_i = 1 \text{ if } Y_i > 0$$

$$Z_i = 0 \text{ if } Y_i = 0$$

Amemiya (1973) has shown that this MLE for the Tobit model is both efficient and consistent.

APPENDIX C

FIGURES

Figure 6. Texas Irrigated Cotton

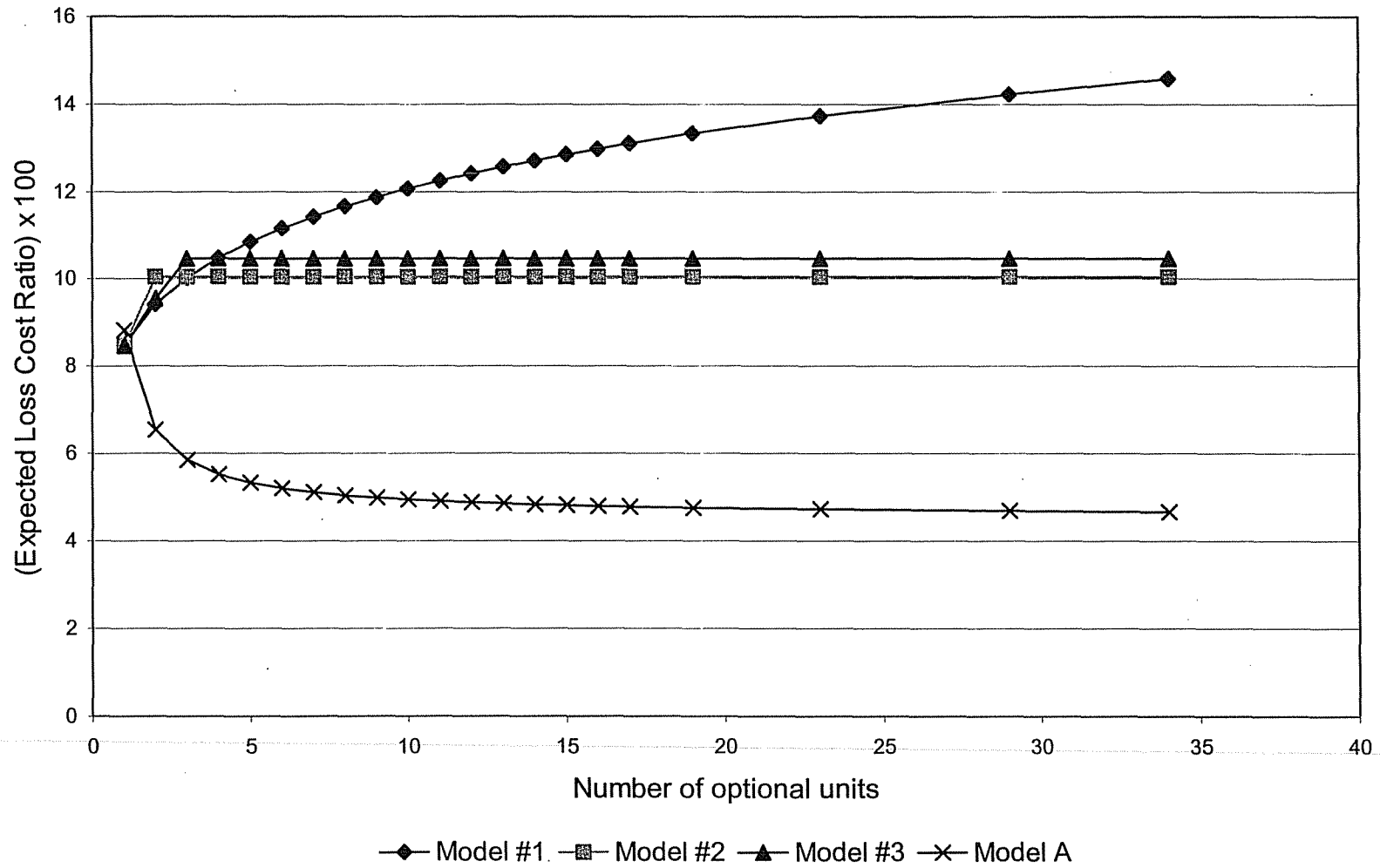


Figure 7. Texas Dryland Cotton

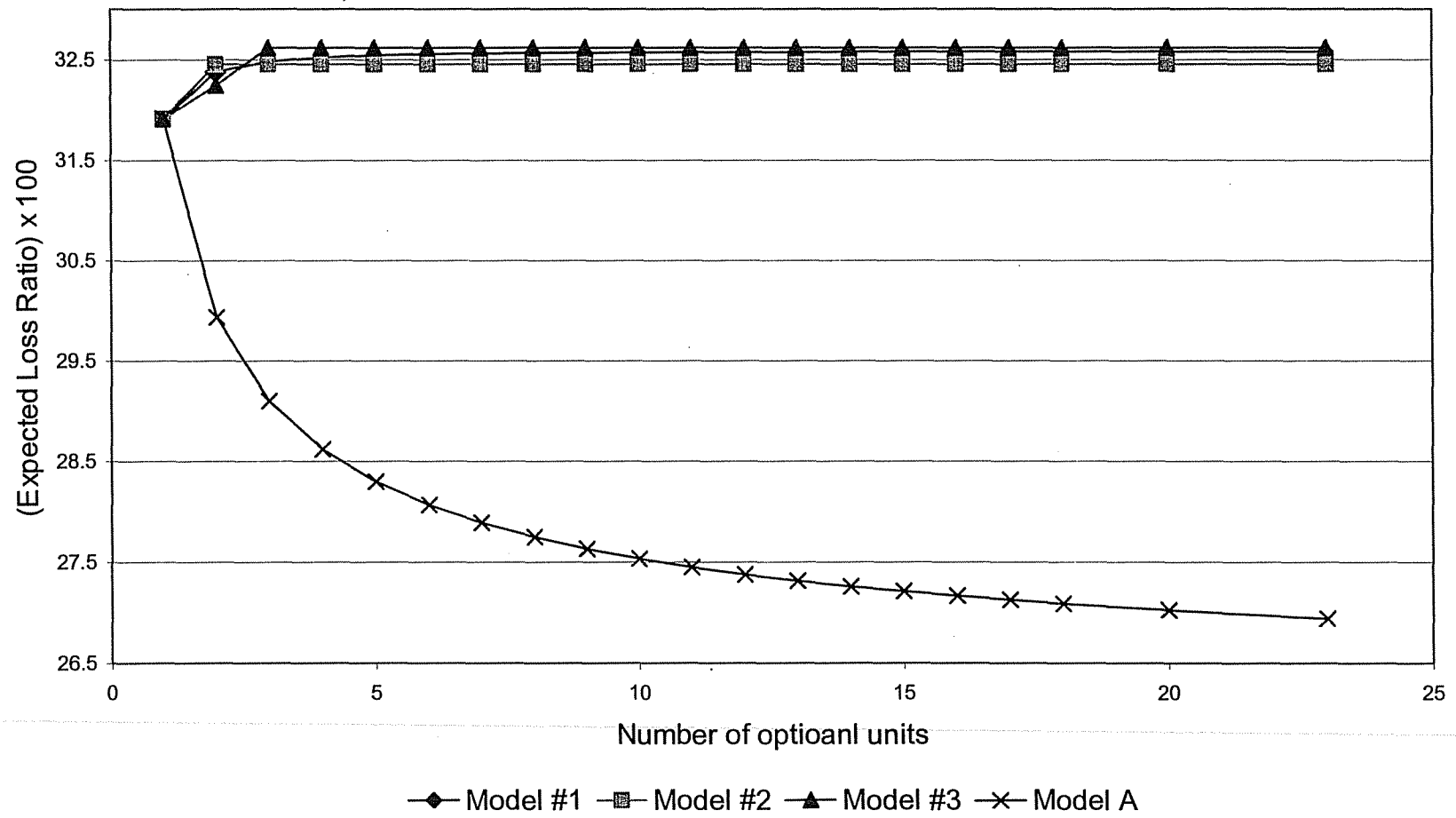


Figure 8. Iowa Dryland Soybeans

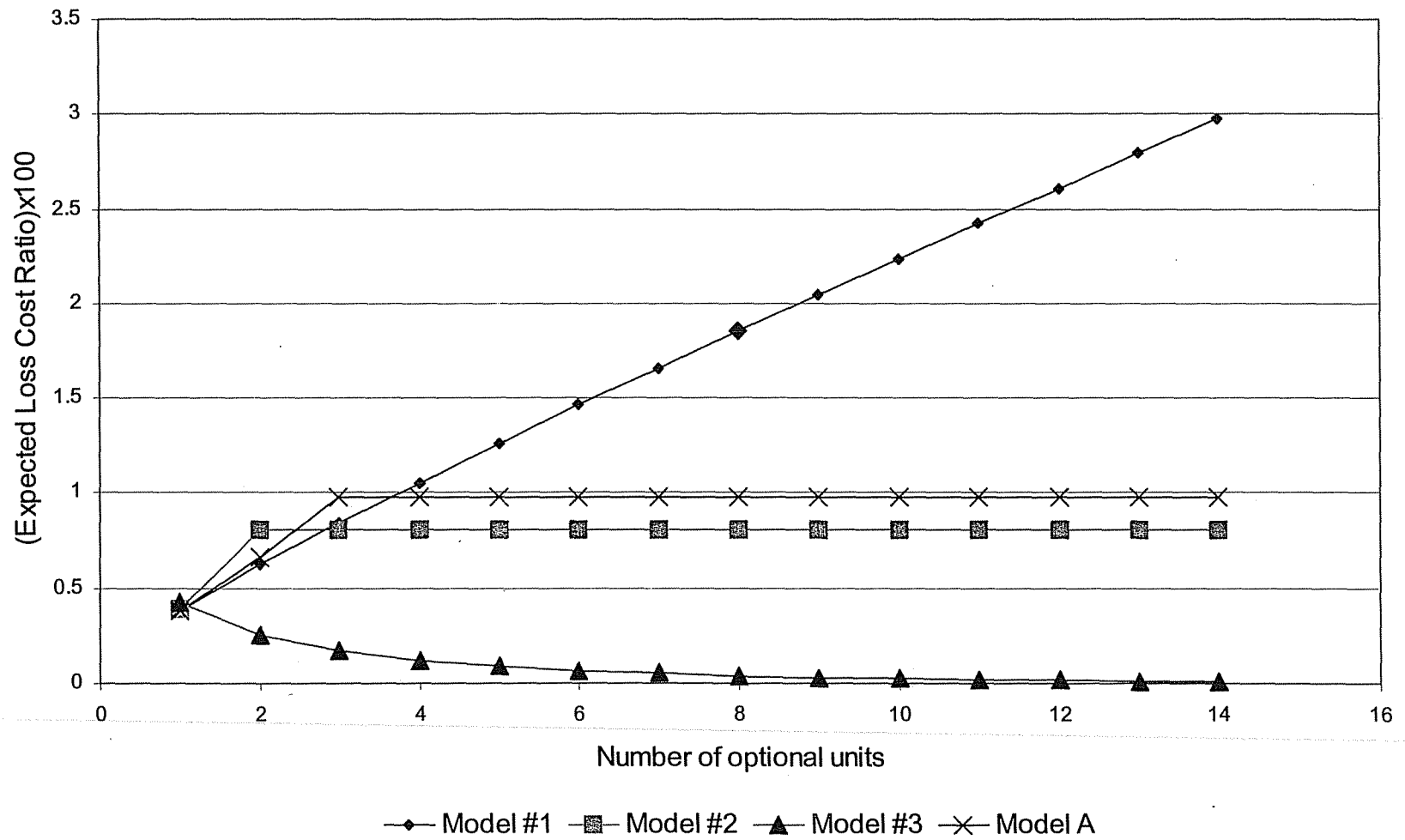


Figure 9. Illinois Dryland Soybeans

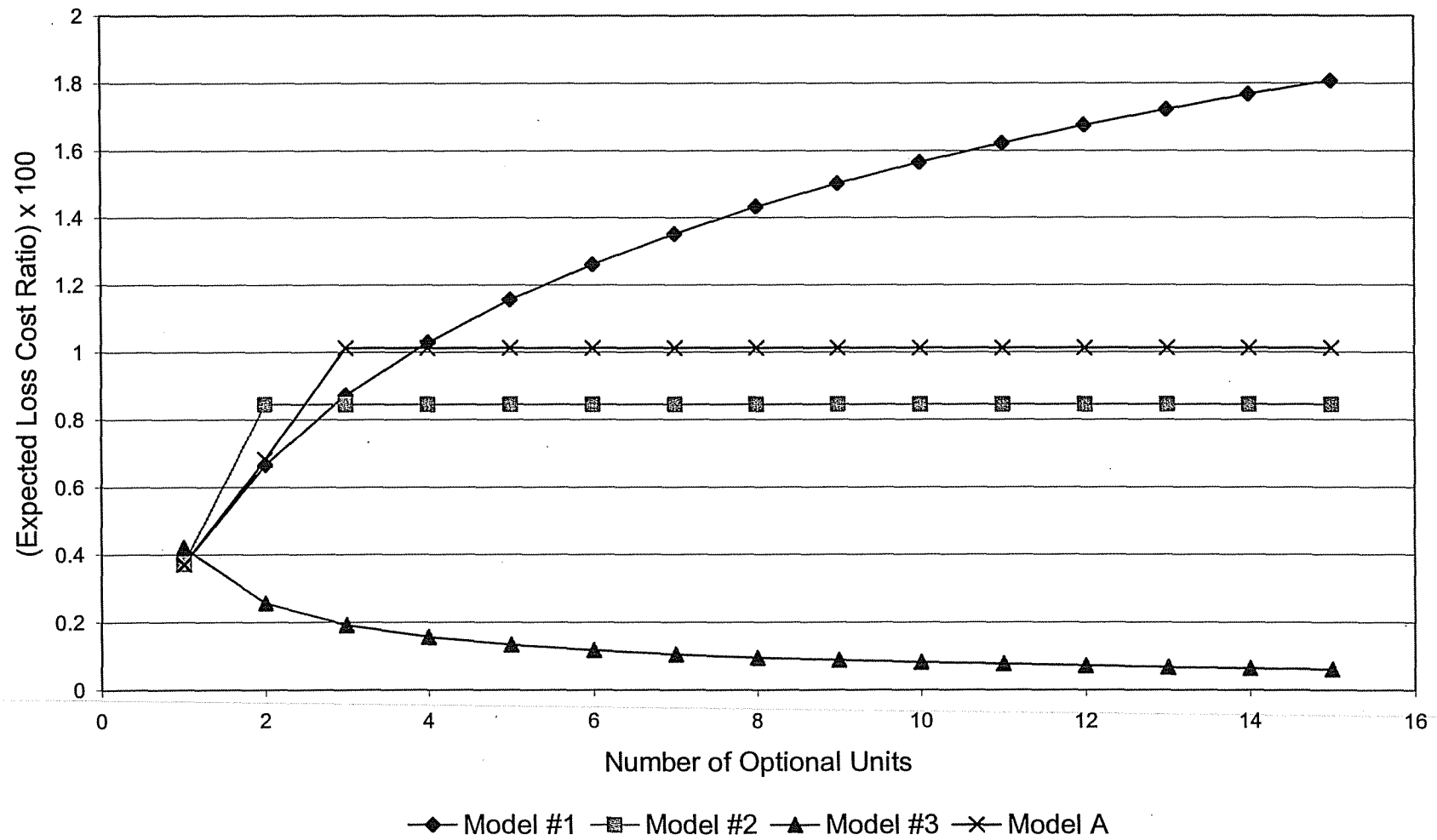


Figure 10. Monte Carlo Results - Texas Irrigated Cotton

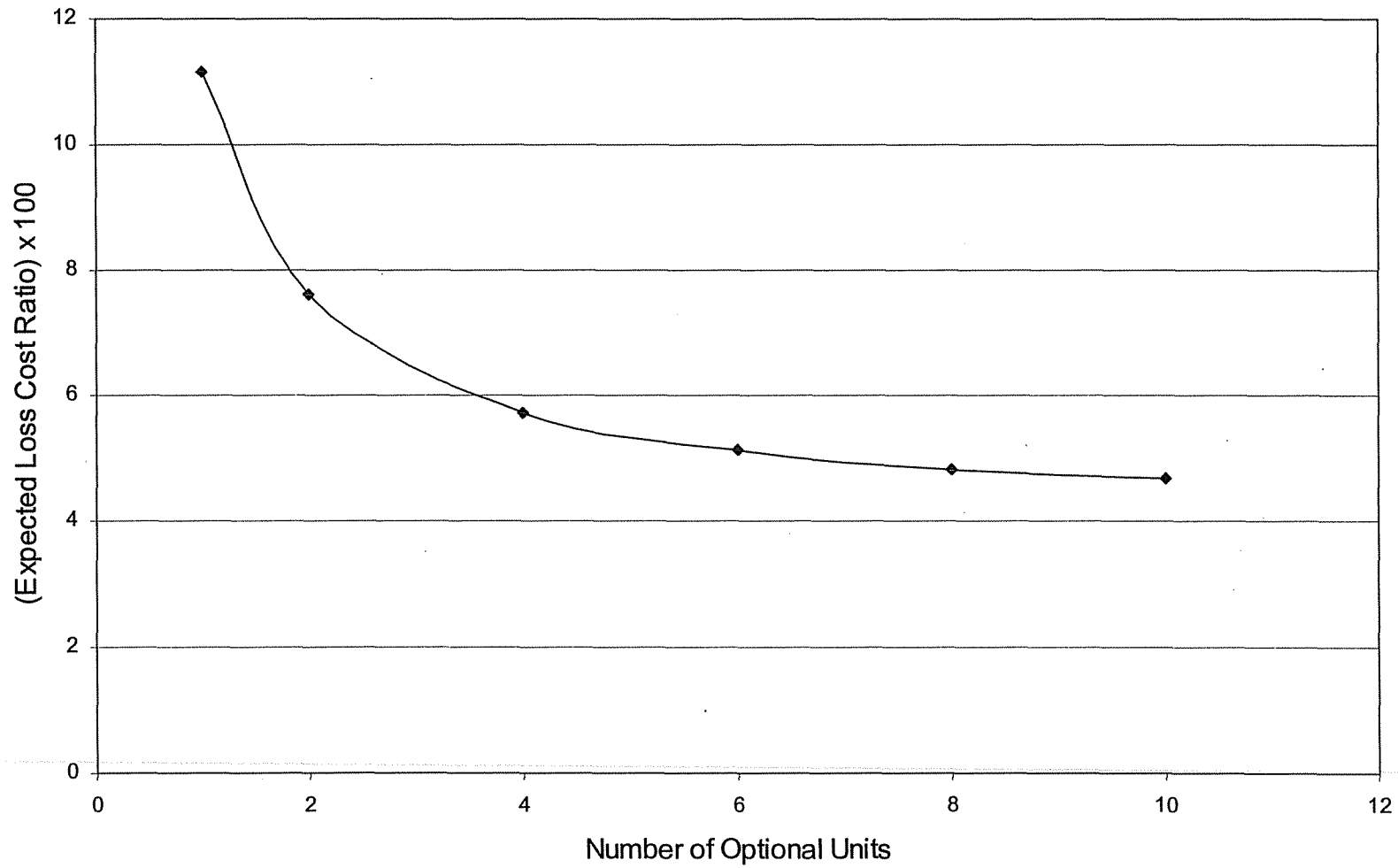
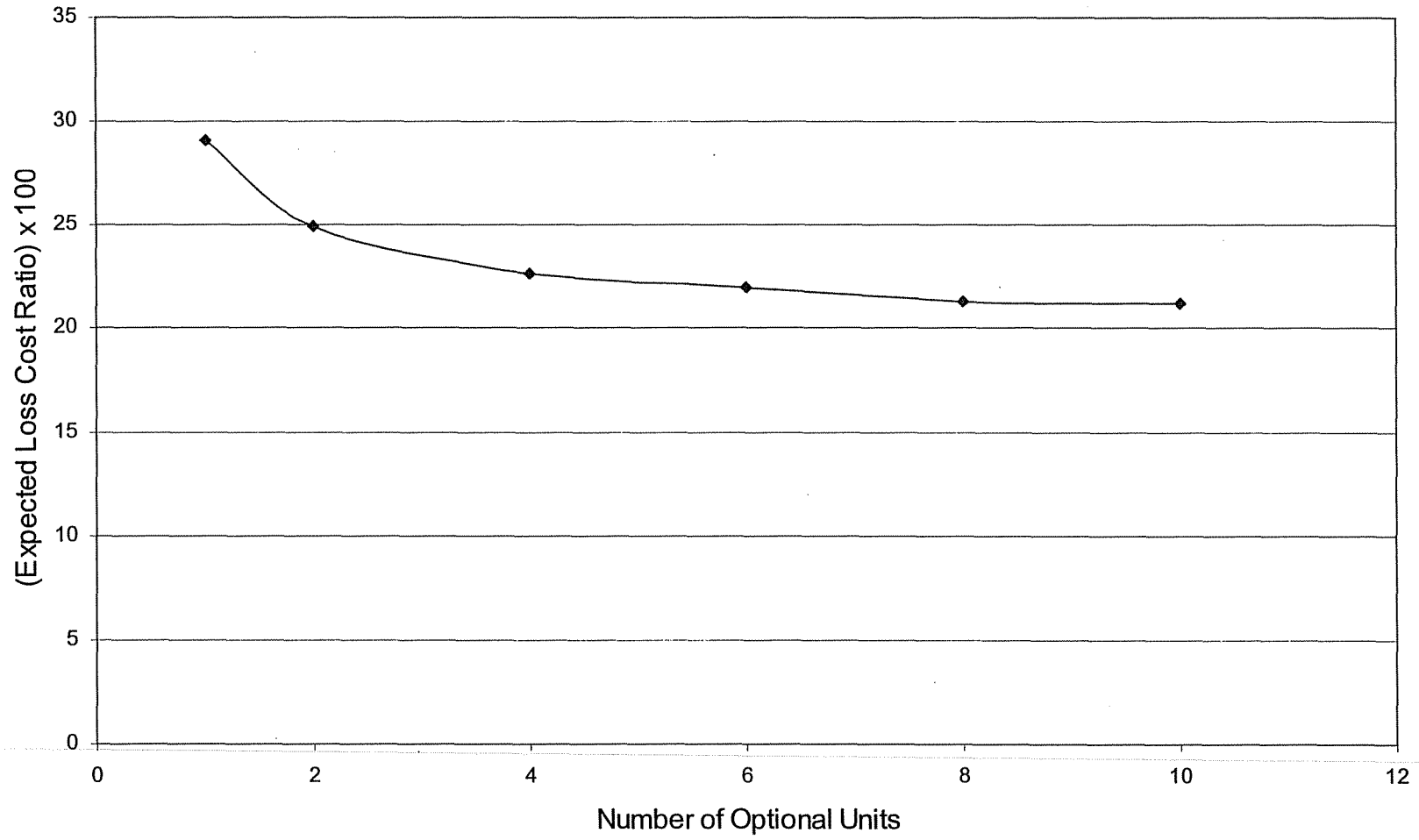


Figure 11. Monte Carlo Results - Texas Dryland Cotton



REFERENCES CITED

Agriculture Risk Protection Act of 2000. Public Law 106-224 June 20, 2000.

Amemiya, Takeshi. "Regression Analysis When the Dependent Variable is Truncated Normal." *Econometrica*, 41 (1973), 999-1012.

Atwood, Joseph, James Driscoll, Vincent Smith, and Myles Watts. "Federal Crop Insurance Programs: History, Practice, and Issues." Unpublished Manuscript, Montana State University, College of Agricultural Economics and Economics, May, 1999.

Atwood, Joseph, Myles Watts, and Saleem Shaik. "An Examination of Potential Fraud in Crop Insurance Due to Yield Switching between Optional Units in US Cotton." Unpublished Manuscript. Montana State University, Department of Agricultural Economics and Economics, 1999.

Chambers, Robert G. "Insurability and Moral Hazard in Agricultural Insurance Markets." *American Journal of Agricultural Economics*. 71 (August 1989) 604-616.

Coble, Keith, Thomas Knight, R. D. Pope, and J. R. Williams. "An Expected-Indemnity Approach to the Measurement of Moral Hazard in Crop Insurance." *American Journal of Agricultural Economics* 79 (February 1997) 216-226.

Common Crop Insurance Policy 2001. Risk Management Agency. Available at www.rma.usda.gov.

Goodwin, Barry K. and Smith, Vincent H., The Economics of Crop Insurance and Disaster Aid. Washington, D.C.: AEI Press, 1995.

Horowitz, John K. and Erik Lichtenburg. "Insurance, Moral Hazard, and Chemical Use in Agriculture." *American Journal of Agricultural Economics* 75 (November 1993) 926-935.

Johnston, Jack and John Dinardo. Econometric Methods. New York: McGraw Hill, 1997.

Josephson, Gary R., Richard B. Lord, and Charles W. Mitchell. "Actuarial Documentation of Multiple Peril Crop Insurance Ratemaking Procedures." Millman and Robertson, Inc. Brookfield, Wisconsin. Prepared for: USDA-Risk Management Agency. August 2000.

- Just, Richard E., Linda Calvin, and John Quiggin. "Adverse Selection in Crop Insurance: Actuarial and Asymmetric Information Incentives." *American Journal of Agricultural Economics*. 81 (November 1999) 834-849.
- Knight, Thomas O. and Keith H. Coble. "Actuarial Effects of Unit Structure in the U.S. Actual Production History Crop Insurance Program." *Journal of Agricultural and Applied Economics*. 31 (1999) 519-535.
- Miranda, Mario J. "Area-Yield Crop Insurance Reconsidered." *American Journal of Agricultural Economics* (May 1991) 233-242.
- Miranda, Mario J. and Joseph W. Glauber. "Systematic Risk, Reinsurance, and the Failure of Crop Insurance Markets." *American Journal of Agricultural Economics*. 79 (February 1997) 206-215.
- Olsen, Randall J. "Note on the Uniqueness of the Maximum Likelihood Estimator for the Tobit Model." *Econometrica* 46:5 (1978) 1211-1215.
- Schnapp, Frank, James L. Driscoll, Thomas P. Zacharias, and Gary R. Josephson. "Ratemaking Considerations for Multiple Peril Crop Insurance." Fedreal Crop Insurance Corporation. No date provided.
- Skees, Jerry R. and Michael R. Reed. "Rate Making for Farm-Level Crop Insurance: Implications for Adverse Selection" *American Journal of Agricultural Economics*. 68 (1986) 653-659.
- Skees, Jerry R. et al. "The Potential for Revenue Insurance Policies in the South." *Journal of Agricultural and Applied Economics*. 30 (July 1998) 47-61.
- Smith, Vincent and Alan E. Baquet. "The Demand for Multiple Peril Crop Insurance: Evidence from Montana Wheat Farms." *American Journal of Agricultural Economics* 78 (February 1996) 189-201.
- Watts, Myles and Collin Prince. "Optional Units." Unpublished Manuscript, Montana State University, College of Agricultural Economics and Economics, January 2000.