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TESTING THE POSSIBILITY OF PRIVATE CROP INSURANCE AND REINSURANCE MARKETS

H. Holly Wang, Joseph L.Krogmeier and Bingfan Ke[1]

Risk theory tells us if an insurer can effectively pool a large number of individuals to reduce the total risk, he then can provide the insurance by charging a premium close to the actuarially fair rate. There is a common belief that only when the random loss is independent, the risk can be effectively pooled, therefore because crop yield is not independent among growers crop insurance market cannot survive without government subsidy. In this paper, a weaker condition, asymptotic nonpositive correlation (a.n.c), is presented as sufficient for effective risk pooling. US crop yield data are used to test the hypothesis and we cannot reject that US yields are a.n.c. As a result, private crop insurance/reinsurance markets are expected to be able to exist.

Introduction

In the wake of the historic 1996 farm legislation, mandating a decoupling of gradually diminishing government support from agricultural commodity price levels, the need exists for more thoroughly understanding the altered agricultural risk environment. Recent attention has focused on the potential role in farm income stabilization of federally-reinsured crop and revenue insurance programs. However, well established market yield risk instruments are not available currently. The CBOT yield futures are thinly traded, and several crop and revenue insurance programs are subsidized heavily by the government which is subject to change. Thus, the insurance itself is risky for long-run farm management.

Given the history of actuarial problems with federal crop insurance programs (Knight and Coble,1997; Skees, Black and Barnett,1997; Goodwin and Smith,1995; Wright,1993) and the renewed interest among researchers and policy makers in modern variants of crop insurance as an income stabilization tool (Hennessay, 1997; Wang et al, 1998), it is prudent to explore if private crop insurance and reinsurance markets can exist, what conditions are required, and whether these conditions are satisfied in the current situation. The objectives of this research are to study the theoretical conditions of risk pooling, to explore a weaker condition than the commonly believed independency on the random losses, and to empirically test these conditions with US crop yield data so as to discuss the potential for a privatized crop insurance market. Each of the objectives is pursued in one of the following three sections, and the last section is conclusion.

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Theoretical Foundations of Insurance

The institution of insurance has evolved in modern economies as one means of risk shifting in the face of uncertainty (Arrow, 1974). The primary function of insurance in this regard is risk pooling. Mehr, Cammack and Rose (1985, p.32) offer the following definition, "*[i]nsurance may be defined as a device for reducing risk by combining a sufficient number of exposure units to make their individual losses collectively predictable.*" In what follows, we will consider the statistical foundations of insurance first, drawing upon the illuminating framework provided by Cummins (1991).

Consider the following model of an insurance pool

$$S_N = \sum_{i=1}^{N} X_i , \tag{1}$$

where S_N is the total loss (claims) of the pool in a given period of time, X_i is the loss experienced by the i^{th} exposure unit, and N is the number of exposure units in the pool. In the context of agricultural insurance, an exposure unit may be a particular farm or parcel of land. In the model given by equation (1), each individual loss is conceptualized as a random variable and the total loss experienced by the pool is random as well. For crop insurance, the loss can be defined as a production shortfall from some prespecified level. If, for convenience, we assume the loss distributions of all exposure units are identical with mean μ and variance σ^2, the expected total loss of the pool is:

$$E(S_N) = N\mu \tag{2}$$

and the variance of the total loss of the pool is:

$$Var(S_N) = N\sigma^2 + 2\sigma^2 \sum_{j=2}^{N} \sum_{i=1}^{j-1} \rho_{ij} , \tag{3}$$

where ρ_{ij} is the correlation between the i^{th} and j^{th} exposure units.

To be of much use, additional information regarding the distribution of the X_i's must be included in the model. Elementary discussions of risk pooling often assume the X_i's are independently and identically distributed (i.i.d.). While this set of distributional assumptions allow application of the simplest version of a law of large numbers, it is unnecessarily strong and generally not realistic (Bühlmann, 1970). Identically distributed exposure units may generally be desirable, but this condition is by no means a necessary statistical condition for effective pooling. Cummins contends that a more compelling argument for homogeneity as a necessary condition for insurability involves information asymmetries between insurers and insureds. Our focus remains on the statistical foundations of insurance.

Similarly, independence is not a necessary condition for the application of more general forms of laws of large numbers and thus for potentially effective risk pooling. In this initial stage of our discussion, however, we shall begin with the simplest case, i.i.d. losses, which simplifies equation (3) since $\rho_{ij} = 0$ for all $i \neq j$. We will then present the more realistic case, especially in the context of agricultural insurance, in which the random losses tend to be positively correlated. A particular form of statistical dependency needs, then, be assumed.

The Weak Law of Large Numbers (WLLN) states that if we assume all random losses in the insurance pool are independently and identically distributed with a finite mean, the realized average loss will be arbitrarily close to the mean loss with probability approaching 1 as the size of the pool approaches infinity.

(4)
$$\lim_{N \to \infty} Pr[|\bar{x}_N - \mu| < \varepsilon] = 1 \quad \forall \varepsilon > 0$$

where $\bar{X}_N = \frac{1}{N}\sum_{i=1}^{N} X_i$, and $EX_i = \mu < \infty$, $\forall$ i .[2]

It would then seem that if the true mean of the loss distribution could be accurately estimated from past loss experience; beginning each insurance period this amount (the net premium) could be collected from every insured thus ensuring the insurance pool would have sufficient funds available to pay any realized indemnities. This is essentially what risk pooling is all about, although as we shall see immediately below (equation (5)) it is not the whole story.

While the WLLN is an important theoretical result, it is of more practical interest to consider insurance pools with a finite number of members. Assuming i.i.d. random losses, implying $\bar{X}_N \sim (\mu, \sigma^2/N)$, we can use Chebyshev's Inequality to write the following bound

(5)
$$Pr[|\bar{x}_N - \mu < k\sigma/\sqrt{N}] \geq 1 - 1/k^2 \quad \forall k > 0 ,$$

i.e., a lower bound of $1-1/k^2$ exists on the probability that the realized average loss will fall in the interval
$(\mu - k\sigma/\sqrt{N},\ \mu + k\sigma/\sqrt{N})$. For some choice of k, which makes the lower bound on the probability given in (5) as close to 1 as we wish, the width of this "confidence" interval is determined by the standard deviation, $\sigma/\sqrt{N}$, of the average loss distribution, and k.

---

[2] A capital letter denotes a random variable and its lower case for a particular outcome.

In the insurance literature, the event of an aggregate loss occurring (a realization of S_N) which is so large as to deplete the insurance fund is captured by the concept of ***ruin***. It has been suggested in the literature that a possible objective criterion for the management of an insurance pool is to minimize the probability of ruin in a given time period or perhaps maximize returns subject to maintaining a specified probability of ruin (Bühlmann). The aggregated premium surplus above the expected value of the aggregate loss ($N\mu$ under homogeneity) required to maintain a particular probability of ruin is referred to as the ***buffer fund***.

Thus, to avoid ruin with probability $1-\frac{1}{k^2}$, the insurance fund must have a liquid buffer fund of the size $k\sigma\sqrt{N}$ for N individuals. Since the size of the buffer fund is proportional to the square root of N; as the size of the pool grows, the buffer fund amount allocated to each policy $k\sigma/\sqrt{N}$ (buffer load), decreases.

Premium rate setting is approached in various ways in the insurance literature, but the discussion generally begins with the concept of "pure" or "net" premium (Hogg and Klugman, 1984; Borch, 1974; Goodwin and Smith, 1995). The net premium is simply the expected indemnity per exposure unit. The gross premium, the amount paid by the insured per exposure unit in order to be eligible for coverage, is larger than the net premium by an amount referred to as the loading factor. We can examine the components of the loading factor by decomposing the gross premium in the following manner

$$P = P_N + A + L, \tag{6}$$

where P is the gross premium, P_N is the net premium, A is an administrative cost load, and L is the buffer load. Assuming identically distributed losses and that there is no deductible or cap on the maximum indemnity, then it is clear that the expected indemnity equals the expected loss, i.e., $P_N = \mu$.

If we assume economies of scale in the administration function, then the administrative cost per exposure unit, A, will decline as the size of the insurance pool grows. Furthermore under the conditions above (i.i.d. random losses), the buffer load per exposure unit, $L = k\sigma/\sqrt{N}$, will decline as the size of the pool increases. Thus for a sufficiently large insurance pool, the amount $A + L$ will not be large, i.e., the risk premium a risk averse insured must pay to obtain coverage will be small.

From our discussion heretofore, it is apparent that the critical ingredient for risk pooling is that the variance of the average loss diminishes as the size of the insurance pool increases, thus providing a statistical basis for predicting future losses. Furthermore under this condition, the buffer load also declines as the pool grows larger, thus ensuring (assuming economies of scale in administration) that the gross premium will exceed the actuarially fair net premium by a relatively small amount. At this point we want to know if weaker conditions of the random losses can ensure $var(\bar{X}_N) \to 0$ *as* $N \to \infty$ besides i.i.d. The section below examines one such condition, asymptotic nonpositive correlation, which has some intuitive appeal in the agricultural context.

A Generalization: Asymptotic Nonpositive Correlation

In what follows we will explicitly relax the assumption of independence. Given the reliance of crop yields on large scale weather patterns, the assumption of independently distributed losses is particularly tenuous in the present context. In place of the independence assumption we will assume asymptotic nonpositive correlation (a.n.c.). Mittelhammer (p. 266, 1996) defines an a.n.c. random sequence as follows:

> The sequence of random variables $\{X_i\}_{i=1}^N$, where $var(X_i) = \sigma_i^2 < \infty \ \forall i$, is said to be *asymptotically nonpositively correlated* if there exists a sequence of constants $\{a_t\}_{t=1}^\infty$ such that $a_t \in [0,1] \ \forall$ t, $\sum_{t=1}^{\infty} a_t < \infty$, and

$$cov(X_i, X_{i+t}) \le a_t \sigma_i \sigma_{i+t} \ \forall t > 0 \tag{7}$$

The two conditions, $a_t \in [0,1] \ \forall$ t and $\sum_{t=1}^{\infty} a_t < \infty$, imply that $a_n \to 0$ *as* $n \to \infty$. a_t represents the upper bound of the correlations between X_i and X_{i+t}. The definition implies that X_i and X_{i+t} cannot be positively correlated when $t \to \infty$. In the context of crop insurance, the natural way to order exposure units in a sequence is spatial so that t represents an ordinal measure of physical distance between exposure units.

The concept of a.n.c. seems quite sensible when applied to crop losses in an agricultural context. Under this assumption, for example, the correlation between the losses of two fields 500 miles apart will generally be less positive than the correlation between two adjacent fields. The definition places no lower bound on the amount of negative correlation between the losses of any two exposure units.

Assuming identically distributed losses for notational convenience, the variance of the average loss in the pool may be written as

$$Var(\bar{X}_N) = \frac{1}{N^2}(N\sigma^2 + 2\sum_{i=1}^{N-1}\sum_{t=1}^{N-i} cov(X_i, X_{i+t})) , \tag{8}$$

where the X_i are ordered spatially and t represents an ordinal measure of physical distance. If we assume that losses are a.n.c. then using the relationship given in equation (7) we can place the following bound on the variance

$$
(9)\qquad
\begin{aligned}
Var(\bar{X}_N) &\le \frac{1}{N^2}(N\sigma^2 + 2\sigma^2 \sum_{i=1}^{N-1}\sum_{t=1}^{N-i} a_t) \\
&= \frac{\sigma^2}{N} + \frac{2\sigma^2}{N^2}\sum_{t=1}^{N-1}(N-t)a_t \\
&\le \sigma^2\left(\frac{1}{N} + \frac{2(N-1)}{N^2}\sum_{t=1}^{N-1} a_t\right)
\end{aligned}
$$

If we take the limit as $N \to \infty$ of (9), recognizing that $\lim_{N\to\infty}\sum_{t=1}^{N} a_t$ is bounded by our definition of a.n.c., then we see that $var(\bar{X}_N) \to 0$ *as* $N \to \infty$ As mentioned above, that the variance of $\bar{X}$ collapses as

N goes to infinity is a sufficient condition for a WLLN result to obtain (see Mittelhammer, Theorem 5.22).

Using arguments similar to those used above, we see that under conditions of a.n.c. an insurer's relative risk (bounded from above by the square root of the bound given in (9) declines as the size of the insurance pool grows. Similarly, the total loss will increase as the size of the pool expands; but the buffer load needed to cover this loss will decrease with pool size.

Testing the Feasibility for US Crop Insurance

1. Testing the a.n.c assumption

Empirical analysis is conducted to test whether yields of several major grain crops throughout US show asymptotic nonpositive correlation, ie. $\lim_{t\to\infty} a_t = 0$, the pairwise correlations decrease as the two regions move away from each other.

County level crop yields obtained from NASS from 1972 to 1997 are used in this analysis. Three major crops, corn, soybeans, and wheat are studied. There are 2,591 counties for corn, 2,000 counties for soybean, and 2,660 counties for wheat in the US(Counties with less than 3 years of observations are dropped.).

These yields are first detrended by a log quadratic trend and then modeled as lognormal. Quadratic trends are used by Miranda and Glauber, and log linear trends are used by Wang et al. which takes care of the heteroskadestic problem. The trend parameters are estimated for each county, and the correlations are calculated based on the detrended error terms (see (A1) in appendix).

Calculating pairwise correlations for all of the several thousand counties are too cumbersome and not necessary. Since our interest is to test whether the correlations between two counties are decreasing while the distance between them are increasing, we select two to three reference counties for each crop and the correlation between the reference county and all other counties are calculated.

The reference counties for corn are Adair, IA, and Redwood, MN, for soybeans are Adair IA and Marshall, IA, and for wheat are Sumner, KS, and Chouteau, MT. These are major production counties whose output levels are at the top in recent years in the top producing states. An additional county, Whitman, WA is selected for wheat because it is the top wheat producing county for soft wheat while the other two counties produce primarily hard wheat. An additional county, Lancaster, PA, is also selected, because both the other two corn counties are in the heart of corn belt and there are not many observations beyond 1200 miles away from them.

The centroid latitude and longitude of each county are obtained from GIS System. These spherical references are then transformed into plain coordinates using ArcInfo, so that the distance equation, $d_{ij} = \sqrt{(x_i - x_j)^2 + (y_i - y_j)^2}$ where x_i and y_i are the coordinates for county i, can be applied.

The correlations between the reference county and any other counties are then ranked by the corresponding distance between them from the nearest to the farthest. These relationships are plotted in figure 1 through 8 for each crop/reference county. Each dot in a particular graph represents the correlation between the reference county and a county that is *d* miles away, *d* is on the x-axis. The pattern is clear for all the eight cases that as the counties are farther away from each other, the crop yield correlation tends to get smaller.

Because the correlations lie between [-1,1] are not strictly decreasing with distance, we are interested in finding the central tendency and the converging speed. The converging speed affects the actuarial soundness of insurance and the necessity of reinsurance. A flexible central tendency function is estimated with MLE assuming normal errors:

(10)
$$\rho = \frac{1 - b_1}{\text{[illegible]}} + b_1 + u$$

where ρ is the vector of correlations, *d* is the corresponding distance, *u* is error vector assuming iid standard normal, and *b*s are parameters. This functional form is flexible in the sense of allowing the trend to be either increasing or decreasing, either concave or convex, and converging point to be either positive, zero or negative. The converging point is b_1, if b_2 and b_3 are positive, and the correlation is one when distance is zero. The estimators are reported in table 1, and the fitted curves are also plotted in figure 1 through 8 correspondingly.

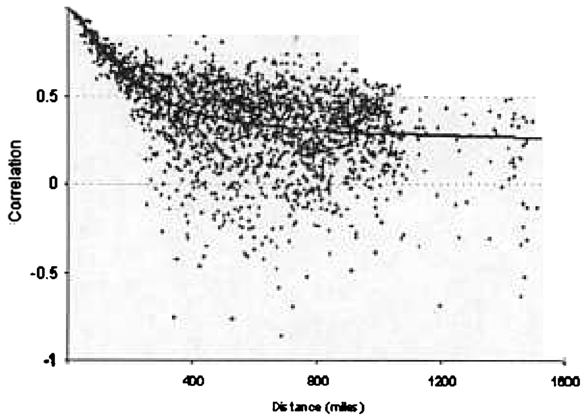

Figure   Corn Yield Correlation, Adair, IA

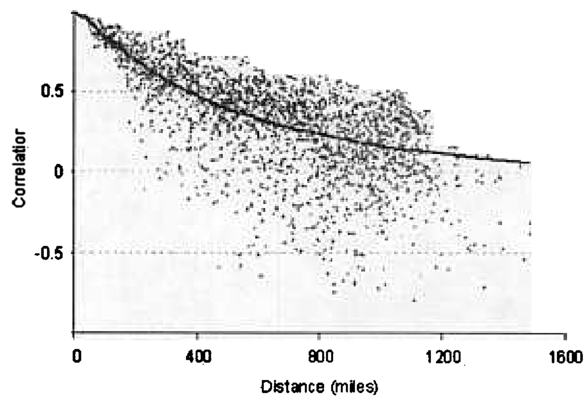

Figure 2 Corn Yield Correlation, Redwood, MN

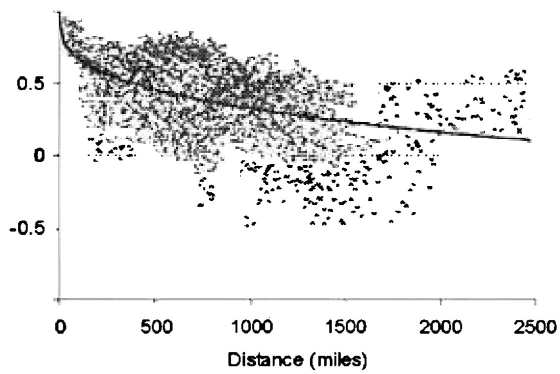

Figure 3 Corn Yield Correlation, Lancaster, PA

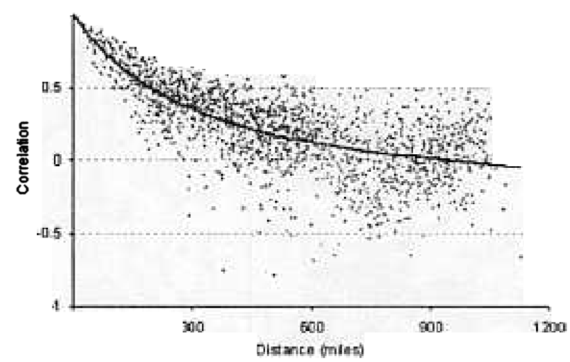

Figure 4 Soybean Yield Correlation,Adair, IA

Table 1. Estimation of Correlation Tendency

	Corn			Soybeans		Wheat		
	Adair	Lancaster	Redwood	Adair	Marshall	Whitman	Chouteau	Sumner
b_1	0.230	-10.001	-0.167	-0.300	-0.317	0.046	-0.128	0.053
	(1.447)*	(-0.147)	(-0.306)	(-0.536)	(-0.638)	(1.097)	(-1.751)**	(0.942)
b_2	0.0003	0.0064	0.0005	0.0023	0.0025	0.0002	0.00005	0.0005
	(0.170)	(0.170)	(0.253)	(0.325)	(0.314)	(0.116)	(0.132)	(0.166)
b_3	1.549	0.3354	1.237	1.064	1.088	1.875	1.905	1.793
	(1.339)*	(2.242)**	(1.547)*	(1.560)*	(1.557)*	(1.092)	(1.439)*	(1.456)*

Note: The numbers in parentheses are t-statistics, and ** means significant at 5% while * means significant at 10%.

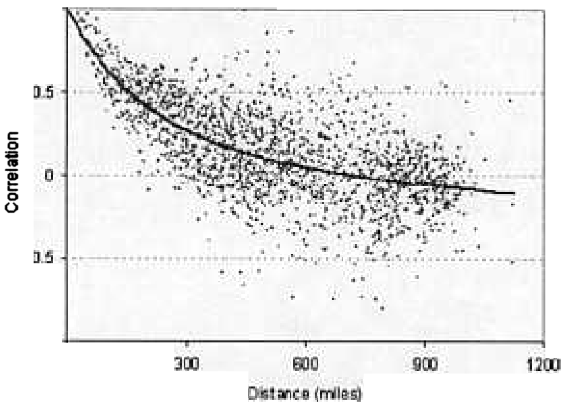

Figure 5 SoybeanYield Correlation,Marshall, IA

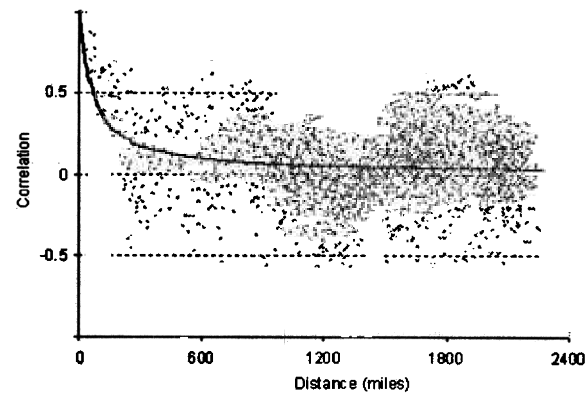

Figure 6 WheatYield Correlation,Whitman, WA

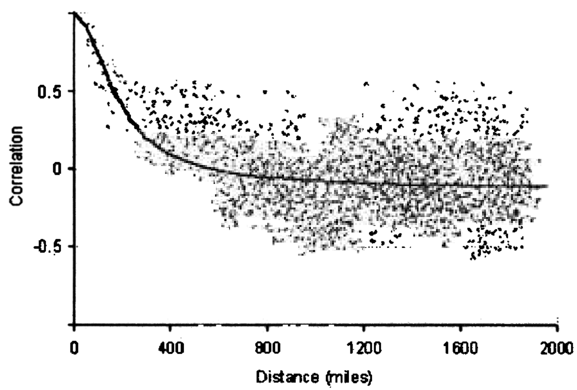

Figure 7Wheat Yield Correlation,Chouteau, MT

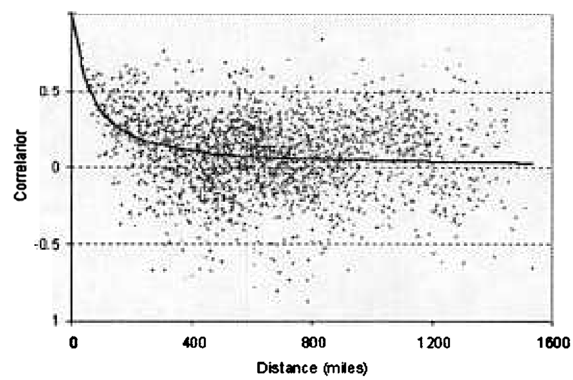

Figure 8 Wheat Yield Correlation,Sumner, KS

The tendency is either approaches to zero or even a negative number for all crops and reference counties selected except for corn-Adair, whose converging point is 0.23. One reason to explain the positive converging point is that there are not many observations beyond 1000 miles from Adair. The negative converging point does not strictly conform with the a.n.c. theorem, however, it is better to serve the purpose of having $\lim_{N\to\infty} \sum_{t=1}^{N} a_t$ bounded from above.

2. Converging speed

A practical concern is how "fast" the sequence converges to zero. The converging speed will suggest how large a geographical area has to be included in the pool in order to reduce the risk faced by the insurance company to an acceptable level. An alternative way to ask this questions is for a given geographical region, how much the pooled risk would be reduced compared to individual risk.

The last line in equation (9) indicates if all individual yields are identically distributed, $\frac{1}{N} + \frac{2(N-1)}{N^2}\sum_{t=1}^{N-1} a_t$ is the upper bound of the proportion that variance of the average loss account for the variance of the individual loss. This proportion, denoted by α, is a critical number for risk pooling. If α is significantly smaller than 1, we can claim the risk pooling is effective, and a low premium loading can be acceptable by both insureds and insurers using the Chebyshev's inequality. Using the estimated correlation ρ based on each reference county as an approximation to *a*, the α levels are estimated and reported in table 2.

Table 2. Estimation of Risk Pooling Effect

	Corn			Soybeans		Wheat		
County	Adair	Lancaster	Redwood	Adair	Marshall	Whitman	Chouteau	Sumner
N	2591	2591	2591	2000	2000	2660	2660	2660
α	0.75	0.78	0.61	0.43	0.29	0.12	-0.13	0.18
$\sqrt{\alpha}$	0.87	0.88	0.78	0.66	0.54	0.35	N/A*	0.42

* Because α is negative, the squared root is not applied.

When the whole country is included in the pool, the variance of the average loss is less than 18% of the variance of the individual loss for wheat, less than 43% for soybeans, and at most 78% for corn.

3. Premium loading

Since the premium loading for each insured is $L = k\sigma\sqrt{\alpha}$ from the Chebyshev's Inequality. If the insurer's critical probability of ruin is 10%, $k = 3.16$, the σ can be calculated by equation (A2) in the appendix. Given the average standard deviations of exponential error term (ε_t) are 0.24, 0.20, and 0.33, and the variance of the log residuals (δ^2) are 0.10, 0.06, and 0.15, standard deviation of yield (σ) are 23%, 19%, and 30% of their mean level for wheat, soybeans, and corn respectively. This indicates pooling all individual losses can reduce the risk to 30% or below. The premium loading are then calculated as 28%, 36% and 76%.

These premium loadings for wheat and soybeans are reasonable, while too high for corn. The main reason for this inefficient risk pooling is that the corn producing region in the Mid-West is small relative to soybeans and wheat. There will not be many participant for corn yield insurance beyond that region even though the insurance is offered nationwide. However, in reality, the insurance company is not restricted to any single crop, selling multiple crops and serving larger area will reduce the premium loading in general.

In summary, we expect the existence of a private reinsurance market with minimum government involvement, where the risks are traded at the national level (or international level) at which yield losses for various crops and regions are hardly correlated. In addition, the yield risks can also be pooled over time because annual yield losses are believed independent. In this case, the premium loading will be reduced by the square root of the number of periods. Thus, multi-year contracts may also help to facilitate the insurance market.

Conclusion

We have reviewed the statistical foundations of insurance in this article, and demonstrated that a necessary condition for effective risk pooling is not the independency, but a weaker condition, asymptotically nonpositive correlation for the loss variable. While the former condition does not hold for agricultural yield, which leads to the common belief that private market for crop insurance is doomed to fail (Miranda and Glauber,1997), we conduct empirical tests for major US grain yields for the latter condition.

The results indicate the possibility of private agricultural insurance markets, if the insurance/reinsurance markets can cover the whole country. Our contention is not that the effect of government intervention is necessarily negative, but that the possibility of a private agricultural insurance and reinsurance market is not precluded by the statistical nature of agricultural production, and this possibility should not be dismissed out of hand.

These results may help alleviate the government's pressure of assuming total responsibility of providing crop insurance. Further research in feasibility investigation of new market products is needed to diversify the crop insurance portfolio.

Appendix

Using a quadratic trend model with a lognormal error term, $lnY_t = \alpha + \beta t + \gamma t^2 + e_t$, where Y_t are random yields, α, β and γ are parameters, e_t are iid random error terms, and assuming $e_t \sim N(0, \delta^2)$, the mean and variance of log yield are $E(lnY_t) = \alpha + \beta t + \gamma t^2$, and $Var(lnY_t) = \delta^2$, respectively.

First, we introduce the procedure to calculate correlations between the yield levels of any two counties in the following:

Since $Y_t = e^{\alpha+\beta t+\gamma t^2} e^{e_t}$, we can denote the trend by $\mu_t = e^{\alpha+\beta t+\gamma t^2}$ and the exponential error by ε_t. Using superscripts indicating a particular county, $Y_t^1 = \mu_t^1 \varepsilon_t^1$, $Y_t^2 = \mu_t^2 \varepsilon_t^2$ the correlation between county 1 and 2 are:

$$
\begin{aligned}
Corr(Y_t^1, Y_t^2) &= Corr(\mu_t^1 \varepsilon_t^1, \mu_t^2 \varepsilon_t^2) \\
&= \frac{E(\mu_t^1 \varepsilon_t^1 - \mu_t^1)(\mu_t^2 \varepsilon_t^2 - \mu_t^2)}{\sqrt{(\mu_t^1 \varepsilon_t^1 - \mu_t^1)^2}\sqrt{(\mu_t^2 \varepsilon_t^2 - \mu_t^2)^2}} \\
&= \frac{E(\varepsilon_t^1 - 1)(\varepsilon_t^2 - 1)}{\sqrt{(\varepsilon_t^1 - 1)^2}\sqrt{(\varepsilon_t^2 - 1)^2}} \\
&= Corr(\varepsilon_t^1, \varepsilon_t^2) \\
&= Corr(e^{e_t^1}, e^{e_t^2})
\end{aligned}
$$

Second, we introduce the way to express yield standard deviation as a proportion of yield

$$
\begin{aligned}
E(Y_t) &= e^{\mu_t} E(\varepsilon_t) = e^{\mu_t} e^{\delta^2/2} \\
Std(Y_t) &= e^{\mu_t} Std(\varepsilon_t) = E(Y_t) e^{-\delta^2/2} Std(\varepsilon_t)
\end{aligned}
$$

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