Alternative Risk Formulations in an Econometric Acreage Response Model for Northwest Wheat

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SUMMARY

This study develops an econometric acreage response model for wheat in the Pacific Northwest (Oregon, Washington, and Idaho) and tests several alternative specifications of risk and their effects on wheat acreage response. Results indicate that the existence of a significant relationship between risk and wheat acreage response depends on the way risk is measured. In general, a three year moving standard deviation of gross income per acre divided by the expected price of wheat was found to be the best measure of risk, both in terms of significance of the estimated coefficients and in terms of explanatory power of the estimated regression equation.

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ALTERNATIVE RISK FORMULATIONS IN AN ECONOMETRIC ACREAGE RESPONSE MODEL FOR NORTHWEST WINTER WHEAT

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INTRODUCTION

Recent efforts to increase the accuracy of positivistic econometric estimates of the supply responses for various agricultural commodities have focused primarily on measuring the impacts of government policies (e.g., Hoffman, 1973; Houck et al., 1976; Lidman and Bawden, 1974). There has been increasing recognition, however, that accurate evaluations of government policy impacts, as well as price elasticities of supply, may require explicit treatment of a factor which is known to interact with policy and expected price in determining production decisions; that factor is risk. Behrman (1968), McCall (1961), Nerlove (1956), Traill (1978), and others have demonstrated that output under uncertainty tends to be smaller than when conditions are more certain, and that supply models which exclude risk likely will have biased estimated price elasticities.

Just (1974), on the other hand, found that risk is not an important element of models of crops subject to government production control programs or of highly aggregated models. Nevertheless, the a priori belief that risk should play an important role in management decisions of even regulated crops and the encouraging results of modelling risk at disaggregated levels have stimulated efforts to capture the impacts of this risk component more accurately.

There is, however, no consensus on how risk variables should be formulated in econometric acreage response studies. Behrman (1968) separated the effects of price and yield risk. He formulated the first as a standard deviation of crop price over the preceding three years relative to the standard deviation of an alternative crop price over the same production period. The latter is a simple three-year standard deviation of crop yield. Just (1974) combined yield and price risks into a single income risk formulation and assumed individual decision-makers formed risk expectations by geometrically weighting past observations of price and yield. Traill (1978) experimented with several formulations of risk in a national onion model. He considered a polynomial lag of the difference between actual market and expected prices, a simple moving standard deviation of price, and a moving probability distribution of expected price. He found that at national levels the supply response was more sensitive to length of lag than to the actual formulation of risk and that the length of lag is relatively short. Traill concluded that farmers may hold stable perceptions of risk and merely adjust them slightly in light of recent information. Hence, much of the risk element, i.e., the long term effects, would be captured by the intercept term. In this case, the actual risk coefficients could be expected to be small, but important.

This study represents an attempt to measure the effect of risk on a regional model of a wheat, regulated crop. The regulation component is treated with government policy parameters and risk is represented as a moving standard deviation of gross incomes per acre. Varying lengths of lag are examined, as well as differential impacts of risk at different price levels and with positive versus negative variations in gross income. Estimated coefficients of traditional policy variables--Houck et al.'s (1976) effective wheat support price and acreage diversion--and of the expected price of wheat are then compared with previous national and regional estimates.

THE MODEL

The hypothesized wheat supply response model is based on Winter and Whittaker's (1979) model, which adds a risk component to the Houck et al. (1976) analysis of governmental policy impacts on wheat acreage. Planted wheat acreages are specified as a function of expected wheat price, government-sponsored wheat support prices, diversion payments, and the expected variation in gross returns of the crop (risk).

The dependent variable, acreage of wheat planted, is theoretically only a proxy for producers' planned output. Behrman (1968) and Penn (1973) have cautioned against assuming a perfect correlation between planted acreage and planned output. But the constraints of data availability, especially at highly aggregated levels, make planted acreage an attractive surrogate. In fact, it is likely that alternative estimates of planned output would include a much larger error factor than does the planted acreage variable, particularly if attitudinal surveys are necessary and the model is at a state, regional, or national level.

The expected price of wheat in this study is simply the market price of wheat lagged one year, a common proxy which has been shown to be a good approximation of expected price and a better predictor of output than many alternative expected price formulations (Lin, 1978; Nerlove, 1958). The lagged price of wheat accounts for the most recent revision of long-run price expectation, the long-run expectation being captured by the intercept (Nerlove, 1956). The anticipated sign on the estimated lagged wheat price coefficient is positive, i.e., an increase in the expected wheat price will elicit an increase in planted wheat acreage.

The government policy variables used are those of Houck et al. (1976). The wheat support price proxy variable is the wheat support rate weighted by the percent of wheat acreage eligible for this payment. Following the argument of Houck et al., this variable will have the same effect on acreage as an increase in farm income. Hence, its expected sign is also positive. Similarly, the diversion price proxy is the wheat diversion price weighted by the percent of wheat acreage eligible for this payment. The government-sponsored voluntary wheat diversion program fostered land withdrawal as an alternative to production and, as such, a negative coefficient is hypothesized. As the price of this production alternative increases, a shift from wheat production to diversion could be expected, other factors being equal.[1/]

In contrast to Winter and Whittaker's (1979) formulation of risk, which was a simple three-year moving standard deviation of incomes, the formulation used in this study is based on Just's (1977), Ryan's (1977) and Traill's (1978) contentions that the supply response to income

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[1/] See Houck et al. (1976) pages 7 to 10 for a complete discussion of and justification of these two policy variables.

variation will be attenuated when prices are relatively high.[2/] Thus, the risk parameter is calculated as a moving standard deviation of gross income per acre divided by expected price. Because Traill found that results appeared to be responsive to the length of lag used in calculating the standard deviation, two, three, and five-year lag models are examined

$$\mathrm{RISK}_t = \left[\sum_{i=t-j}^{t-1} (\mathrm{GI}_i - \overline{\mathrm{GI}})^2/(j-2) \right]^{\frac{1}{2}} / \mathrm{PW}_{t-1} \quad \text{for } j = 3, 4, 6 \quad (1)$$

where

RISK_t = measure of risk in year t;

GI_i = gross income in year i;

$\overline{\mathrm{GI}}$ = mean gross income during the years t - j to t - 1;

PW_{t-1} = the price of wheat in year t - 1.

It is conceivable that producers react differently to positive deviations in gross income than to negative deviations. A risk averse attitude implicitly assumes that any variation, either positive or negative deviations in gross income from the mean, will shift the supply curve leftward. The magnitude of the shift, however, may differ for negative versus positive deviations. To test this hypothesis, the risk variable was broken into two parts; the first consists of a semi-standard deviation of positive gross income deviations relative to the expected price of wheat and the second a semi-standard deviation of negative gross income deviations relative to the expected wheat price. These variables were

[2/] Ryan also measured risk as income variability relative to the prices of alternative crops, but previous research suggests that Northwest winter wheat has no economically viable substitutes (Winter and Whittaker, 1979).

calculated for three and five year lags (Equation 2).

$$RISKP_t = \left[\sum_{i=t-j}^{t-1} (GI_i - \overline{GI})^2/(j - 2) \right]^{\frac{1}{2}} / PW_{t-1} \quad \text{for } j = 4, 6 \text{ and } GI_i \geq \overline{GI}$$

$$RISKN_t = \left[\sum_{i=t-j}^{t-1} (GI_i - \overline{GI})^2/(j - 2) \right]^{\frac{1}{2}} / PW_{t-1} \quad \text{for } j = 4, 6 \text{ and } GI_i \leq \overline{GI} \tag{2}$$

where

$RISKP_t$ = measure of risk in year t for positive deviations;

$RISKN_t$ = measure of risk in year t for negative deviations; and

all other variables are as defined above.

Assuming that agricultural producers are risk averse (see, for example, Just, 1974, and Lin, 1977), the coefficient of the risk variable in equation 1 is expected to be negative. An increase in the variability of gross income per acre relative to expected price is hypothesized to induce a decrease in wheat acreage planted. Again, assuming risk averse behavior by farmers, the coefficients on the risk variables in equations 2 will both be negative. If negative deviations in gross income cause larger supply shifts than positive deviations, the estimated coefficient for the second equation 2 will be greater (in absolute value) than that for the first.

Intercept shifters are included in the model to account for major differences among the three states in mean planted wheat acreages. Table 1 shows that Washington is the largest producer in the region, followed by

Table 1. Area planted in wheat 1,000 acres

	1977	1976	1975	1974	1973	1972	1971	1970	1969	1968	1967	1966	1965	1964
Washington	3,120	3,285	3,108	3,280	3,345	2,746	2,416	2,413	2,890	2,875	3,002	2,473	2,545	2,094
Oregon	1,278	1,370	1,310	1,278	1,100	918	809	745	834	1,045	1,080	824	942	809
Idaho	1,330	1,575	1,480	1,550	1,200	1,034	1,046	1,001	1,166	1,338	1,398	1,089	1,201	1,238

SOURCE: USDA Agricultural Statistics, USGPO, Washington. Various issues.

Idaho and then Oregon. Because the binary shift variables are interpreted as the change in the constant if the observation is from Washington (BW) or Idaho (BI) both estimated coefficients should be positive.

Economic theory suggests that prices of alternative crops should be important determinants of planted wheat acreage. However, previous work by Hoffman (1973) and Winter and Whittaker (1979) suggests that wheat has no economically viable production alternatives.[3/] Therefore, no prices of production alternatives appear in the model.

A slope shifting term also was included in the model (as suggested by Winter and Whittaker). This shifter accounts for the interaction of the expected price of wheat with the Washington intercept shifter. This variable was included because the price elasticity in Washington is different from that in the rest of the region because of the relatively low cost of growing wheat in this state (Winter and Whittaker, 1979).

The issue of whether it is theoretically proper to deflate prices in supply response is an ongoing one and likely to remain so. Many other supply response studies have used deflated values: Adams and Behrman (1976), Houck et al. (1976), Lidman and Bawden (1974), Lin (1977), Nerlove (1956). On purely pragmatic grounds, this model was more stable and more of the variables dictated by theory were statistically significant when prices were deflated. Therefore, wheat price, wheat support price, and wheat diversion payments all were deflated by a farm producers' price index in this research.

[3/] When wheat acreage response is broken into substate regions, Moe (1979) has found significant production alternatives, but they are not significant when the aggregate units are multi-state regions. (Grass seed and potatoes are significant substitutes in certain regions of Oregon and peas, alfalfa, and sugarbeets are substitutes in various regions of Washington.)

In summary, the empirical model in this research is as follows:

$$AWP_{it} = f(PW_{i,t-1}, EWSP_t, WD_t, RISK_{it}, BW, BI, PWBW), \qquad (3)$$

where

AWP_{it} = acreage of wheat planted in state i in year t, in thousands of acre

$PW_{i,t-1}$ = average annual price of wheat for state i in year t-1, per bushel, deflated by an index of prices received by farmers;

$EWSP_t$ = wheat price support rate, weighted by percent of wheat acreage eligible for this government payment in year t, in dollars per bushel deflated by an index of prices received by farmers;

WD_t = wheat diversion payment rate, weighted by the percent of wheat acreage eligible for this payment in year t, in dollars per bushel deflated by an index of prices received by farmers;

$RISK_{it}$ = risk in state i in year t as defined either by equation 1 or equations 2;

BW = a binary intercept shift variable for Washington (BW = 1 if the observation is from Washington, and BW = 0 otherwise);

BI = a binary intercept shift variable for Idaho (BI = 1 if the observation is from Idaho and BI = 0 otherwise);

PWBW = $PW_{i,t-1}$ times BW.

EMPIRICAL RESULTS

The empirical application of the above model is to wheat in the Pacific Northwest (Oregon, Washington, Idaho) for the 1964-1977 period. The 1964 starting point was selected because 1964 marked the beginning of several fundamental changes in government policies directly affecting wheat production. Wheat marketing quotas and penalties were terminated, mandatory controls were dropped in favor of direct payment inducement, and loan levels were reduced drastically to $1.25 per bushel (Just, 1974; Houck et. al., 1976; Lidman and Bawden, 1979). Voluntary allotments continued, but prices became less stable. These changes were of such import to wheat production decisions that it was felt the accuracy of a response model might depend to some degree on recognition of these changes and separate treatment of the production period after the 1963 changes.

Because separate treatment of the 1964-1977 period severely limits the size of the data base, a pooled cross-sectional time-series approach was utilized (Bancroft and Whittaker 1977). The pooled element implicitly assumes that the relationship of the dependent variable, acres of wheat planted, with the independent variables is consistent throughout the region. In a linear model, this means that a given change in any independent variable would induce a change of equal magnitude in wheat acreage in each of the three states. Since considerable differences in wheat acreages among the three states exist, such a response is not reasonable. Therefore, a double logarithmic functional form which implicitly enforces the more reasonable assumption that interstate acreage elasticities are equal was used in this research.

The ordinary least squares (OLS) estimates of the supply response parameters are summarized in Table 2. Equations 2.1, 2.2, and 2.3 contain risk variables calculated using equation 1 with three, five, and two lagged observations, respectively. Equations 2.4 and 2.5 contain risk variables calculated using equations 2 with three and five lagged observations, respectively.

All the signs of the coefficients conform with _a priori_ expectations, with the exception of the risk estimates in equations 2.2, 2.4, and 2.5. However, none of these estimated coefficients is significant at 5 percent (the risk coefficient for positive deviations in equation 2.4 is, however, significant at 10 percent). With few exceptions, the remainder of the estimated coefficients are statistically significant at the 5 percent level, and the coefficients are relatively stable with changes in the formulation of risk. The acreage response elasticity cannot really be calculated because wheat price enters the risk variable in a nonlinear manner (Ryan, 1977), as well as serving both as an independent variable and as a deflator for the risk variable. However, if the effect of a price change on acreage that comes from a change in the numerator of the risk variable is ignored, the acreage response elasticity (for Oregon and Idaho) is the difference between the estimated coefficients for wheat price and risk.[4/]

4/ This fact can be simply demonstrated. Assume the estimated model is

$$\ln AWP = a+b(\ln PW)+ c\ln(RISK/PW). \qquad (4)$$

Taking the antilog of (4) yields

$$AWP = a'PW^{b}(RISK/PW)^{c} = a'PW^{b-c}\ RISK^{c}, \qquad (5)$$

where $a' = e^{a}$. Differentiating equation 5 with respect to the price of wheat results in

$$\partial AWP/\partial PW = (b-c)a'PW^{b-c-1}RISK^{c} = (b-c)\ AWP/PW \qquad (6)$$

If one multiplies equation 6 by PW/AWP to transform it into an elasticity, the result is b-C.

Table 2. Risk Response Model Coefficients and t Values

Equation	Constant	PW	EWSP	WD	RISK3[1/]	RISK5[2/]	RISK2[3/]	PWBW	BW	BI	R^2
2.1	6.652	.361	.323	-.384	-.043			-.227	1.142	.231	
	(76.450)	(4.952)	(2.287)	(5.413)	(1.734)			(2.232)	(18.137)	(6.251)	.983
2.2	6.513	.303	.359	-.331		.040*		-.186	1.128	.250	.964
	(62.797)	(3.849)	(2.486)	(3.437)		(.844)		(1.768)	(17.451)	(6.303)	
2.3	6.579	.333	.371	-.386			-.006*	-.204	1.131	.239	.964
	(87.262)	(4.513)	(2.583)	(5.228)			(.504)	(1.946)	(17.379)	(6.276)	
					RISKP3[1/]	RISKN3[1/]					
2.4	6.605	.399	.371	-.392	-.049	.019*		-.217	1.141	.231	.968
	(75.852)	(4.873)	(2.567)	(5.271)	(1.960)	(.697)		(2.141)	(18.102)	(6.261)	
					RISKP5[2/]	RISKN5[2/]					
2.5	6.455	.171*	.335	-.242	.049*	.035*		-.170*	1.127	.261	.967
	(63.587)	(1.541)	(2.371)	(2.240)	(1.325)	(.878)		(1.643)	(17.603)	(6.583)	

*Not significant at $P < .10$

[1/] Three year lag
[2/] Five year lag
[3/] Two year lag

The acreage response elasticity for Washington is calculated as the elasticity for Oregon and Washington plus the coefficient on the price elasticity shifter variable for Washington (PWBW). These "approximate" elasticities are presented in Table 3 for each equation in Table 2. With the exception of the elasticity calculated from equation 2.5, the Oregon-Idaho acreage response elasticities are generally in line with Nerlove's (1956) national estimates of .34 to .48, Houck et al.'s (1976) national estimate of .39 and Winter and Whittaker's (1979) Oregon-Idaho estimate of .376. The Washington acreage response elasticities are somewhat smaller than those of Oregon and Idaho, and three are similar in magnitude to Winter and Whittaker's (1979) Washington estimate of .219. This intra-regional difference in price elasticity was expected because Washington wheat acreage is much greater than the other two states and because the opportunity cost of growing wheat in Washington is lower relative to Oregon and Idaho. The lower opportunity cost in Washington reflects the condition that alternative crops are not as economically attractive as in Oregon and Idaho. Again, the Washington elasticity calculation from equation 2.5 is out of line (in fact it is negative) with the other estimates. The low elasticity estimates from equation 2.5 are likely attributable to multicollinearity, as evidenced by a relatively high R^2, and several (four) non-significant coefficients for this equation.

As mentioned above, the acreage response elasticities presented in Table 3 are somewhat biased since they do not include the price component embodied in the numerator of the risk variable. Estimates of elasticities that include this component cannot be calculated, but the direction of the bias can be indicated. If the change in price tends to increase the

Table 3. Acreage Response Elasticities with Respect to the Price of Wheat

Equation from which elasticity was derived	Oregon and Idaho	Washington
2.1	.404	.177
2.2	.263	.077
2.3	.339	.135
2.4	.429	.212
2.5	.087	-.08

variation in gross income, the estimates in Table 3 are biased upwards, i.e., the percent change in acreage will be less than that indicated by the Table 3 elasticity estimates. If the price change decreases the variation in gross income, the Table 3 estimates are biased downwards (Ryan, 1977).

Both the government policy parameter estimates--the wheat support price and the acreage diversion payments--are significant in all five equations. The values of the coefficients in each equation were higher than the coefficients for the same variables estimated by Winter and Whittaker (1979). The wheat support price coefficient ranges from .323 to .371, suggesting that the government-sponsored floor price for wheat has a large effect on acreage planted. A strong inverse relationship (-.242 to -.392) was estimated between wheat acreage and government acreage diversion payments, indicating that the diversion programs have successfully induced producers in the region to reduce wheat acreage planted. It was not possible with the data in this study to investigate whether producers removed marginal lands from production or compensated for foregone acreage with more intensive usage of other inputs in the production mix, so it is not possible to judge whether the acreage diversion scheme was effectively supply output restrictive. However, the coefficient estimates do suggest that wheat acreage was reduced as a result of diversion payments.

Neither the two-year nor the five-year risk coefficients were significant though both had the expected inverse relationship to acreage. The three-year risk coefficient was significant at greater than the 5 percent level and the estimate of -.043 is similar to that estimated by Lin (1977) for Kansas wheat and Winter and Whittaker (1979) for Northwestern wheat.

An attempt to measure the hypothesized differential impact of positive revenue variations versus negative revenue variations on producers' risk involved the calculation of the standard deviation of those revenue periods above the three (or five)-year mean and the standard deviation of those revenue periods below. Of the four risk coefficients estimated, three had signs opposite to those predicted and were non-significant. Only the three-year positive revenue variance was significant, and it did have the expected negative coefficients.

PREDICTIVE CAPABILITIES

For a further comparison of alternative measures of risk, each of the equations in Table 2 was tested for its ability to predict regional planted wheat acreage during the years used for parameter estimation (1964 to 1977). The results are presented in Table 4.

The most accurate predictions are provided by the model where risk is separated into positive and negative deviations and the values of gross income from the previous three years are used to calculate the risk coefficient. The other four equations have predictive capabilities that are very comparable to each other. The equations where risk is measured using gross income from the previous five years (equations 2.2 and 2.5) have the smallest variation in their prediction errors.

Table 4 suggests little concerning measurement of risk using a single standard deviation versus separating deviations in gross incomes into positive and negative deviations and using two risk variables. The table does suggest, however, that (at least in this case) incorporating three years of lagged gross income values into the risk variable gives the most accurate predictions (equations 2.1 and 2.4), but calculating risk from five lagged years of gross income results in the least variation in prediction errors. Care should be used in generalizing these results, however, since the differences in the means and variances of the prediction errors are small.

Table 4. Prediction Capabilities of the Models

Equation from which predictions were made	Average percent prediction error	Standard deviation of the percent prediction error
2.1	5.56	5.33
2.2	5.85	4.78
2.3	5.74	5.39
2.4	4.99	5.88
2.5	5.57	4.45

CONCLUSIONS

Government regulation of wheat production via support payments and diversion programs has substantially impacted production decision making. Establishment of minimum wheat prices has reduced the level of risk to the producer and diversion policies have generated shifts in the Northwest region's output mix. Welfare questions of who gains from the induced shifts and at what cost to whom is not within the scope of this study, but should be addressed before any conclusions are drawn regarding an optimum wheat policy.

Market risk also exerted significant influence on production trends after 1964. The statistical significance of the risk to wheat price ratio supported Ryan's (1977) and Traill's (1978) beliefs that a producer perceives crop riskiness as lower when the price of the crop is relatively high. However, this risk factor was sensitive to adjustments in lag. While Traill found a two-year lag was best for onions, in Northwestern wheat a three-year lag seems to be most significant (Winter and Whittaker). In any case, the appropriate risk lag seems to be low, capturing only a producer's short run adjustment of his long run perception of risk. Thus, the concern that risk coefficients tend to be lower than theory would suggest may be unsound. Given that long-run expectations of risk are found in the intercept term, the short run revisions in light of recent information may indeed be slight; the riskiness of wheat probably has not changed dramatically in the last 14 years. Furthermore, government stabilization programs, upward trends in yields per acre from technological change, and a lack of an attractive alternative to winter wheat in this region may render risk a minor consideration. Nevertheless, inclusion of a risk element in a supply response model of a regulated crop at a regional level can provide the model with a better fit and more accurate estimation of price elasticities of supply and responses to government policies.

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