Response to Price and Production Risk: The Case of Australian Wheat

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A model of Australian wheat grower supply response was specified under the constraints of price and yield uncertainty, risk aversion, partial adjustment, and quadratic costs. The model was solved to obtain area planted. The results of estimation indicate that risk arising from prices and climate have had a significant influence on producer decision making. The coefficient of relative risk aversion and short-run and long-run elasticities of supply with respect to price were calculated. Wheat growers' risk premium, expected at the start of the season for exposed price and yield risk, was 2.8 percent of revenue or 10.4 percent of profit as measured by producer surplus. © 2000 John Wiley & Sons, Inc. Jrl Fut Mark 20: 345–359, 2000.

INTRODUCTION

Historically, the Australian wheat industry has been exposed to high levels of price and production risk. Price per ton received by growers over the

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last five years varied between \$A154 and \$A230 and national average yields varied over the same period between 1.11 and 2.09 tons per hectare (ABARE, 1997). Price variability arises from unstable world supply, the inelastic nature of demand, and protection by governments of consumers and producers in important markets. Yield variability for Australian producers arises primarily from climatic risk and, as indicated by wide swings in national averages, tends to be highly correlated across producers. There is only limited scope for the management of either production or price risk. Other grain crops are subject to the same climatic risk and are close substitutes for wheat in both production and consumption and hence face similar price risks. The traditional form of production diversification is wool. While there has not been a futures market trading in Australia during any of the sample period, the Australian Wheat Board has undertaken some hedging on the American futures market on behalf of growers (Bond, Thompson, and Geldard, 1985). In addition, there has been some forward selling with merchants, however data are not available on the extent of this activity.

There is a growing body of published empirical work showing significant behavioral responses to volatility in state variables (see Enders (1995) for a summary of some recent empirical publications). An example of this empirical work is that by Holt and Aradhyula (1990), who relate price volatility to broiler supply using an Autoregressive Conditional Heteroscedastic (ARCH) structure in a linear supply equation. With no explicit theoretical basis, it is not clear how their coefficients should be interpreted or how other types of risk could enter their model. More recent work has been undertaken by Chavas and Holt (1990), Pope and Just (1991), and Chavas and Holt (1996). This generation of work focuses on the form of the utility function and, in particular, how risk preferences are influenced by income. This study focuses on the effects of uncertainty as measured by unanticipated changes in income. An attempt is made to capture farmers' responses to short-run volatility in prices and yield in a simple yet theoretically consistent framework that should prove useful for applications to a range of hedging and futures market problems.

The second section of this article, Theoretical Model, presents the theoretical model of the wheat market; in the third section, Empirical Results and Discussion, a constrained maximum likelihood estimation technique is used to estimate the values of the parameters in the theoretical model. The coefficient of relative risk aversion, the risk premium, and supply elasticities are then computed based on the parameter estimates. The section discusses the results with specific reference to risk response behavior. Conclusions are presented in the forth section.

THEORETICAL MODEL

It is assumed that wheat grower utility can be described using the expected value of, and the expected variation in, wheat profits. Let E(U) be expected utility, E(π) be expected profit, and k the Pratt (1964) coefficient of absolute risk aversion:

$$E(U) = E(\pi) \frac{k}{2} E[(\pi - E(\pi))^2], \qquad (1)$$

where it follows the risk premium is

$$\frac{k}{2} E[(\pi - E(\pi))^2].$$

The function is a member of a general class (E-V or mean-variance) of utility functions. General issues associated with the E-V framework are discussed in Markowitz (1959), Tsiang (1972), Anderson, Dillon, and Hardaker (1977), and elsewhere. More specifically, Meyer (1987) shows a condition that allows agent's expected utility (EU) ranking of a set of random variables to be represented by a ranking based on their means and standard deviations. This condition is, simply stated, that the choice set be composed of random variables which differ from one another only by location and scale parameters. Examples of two-parameter families of random variables are the normal and uniform distributions (for specific details see Meyer, 1987, Section II, pages 423–426). By assuming normality of the error terms in eqs. (4) and (6) below, the specification in this article meets the stated condition.

Unlike other E-V formulations, the version above is Constant Absolute Risk Aversion (CARA), which is more desirable than the quadratic utility specification discussed by Anderson et al. (1977), which implies Increasing Absolute Risk Aversion (IARA). CARA, from Pratt (1964), means $-U''/U'$ does not change with increments or decrements to expected profit. CARA occurs in our specification since it implies that the risk premium $[(k/2)\, E(\pi - E(\pi))^2]$ (measuring the cost of risk) is independent of fixed cost or initial wealth.

Production occurs within a one year cycle so that producers maximize utility in period t conditional on information available in period $t-1$:

$$E_{t-1}(U_t) = E_{t-1}(\pi_t) \frac{k}{2} E_{t-1}[(\pi_t - E_{t-1}(\pi_t))^2]. \qquad (2)$$

Using p_t to denote price in period t, q_t to denote actual production, and

$E_{t-1}(q_t)$ to denote production planned for period t in period $t-1$, profits are defined as:

$$\pi_t = p_t q_t - a - bE_{t-1}(q_t) - c[E_{t-1}(q_t)]^2. \tag{3}$$

Thus, wheat profits are revenue minus costs where costs are assumed to be a quadratic function of planned production. There are two simplifications in this treatment of costs. First, it is assumed the majority of variable costs are incurred prior to the realization of yield with land preparation and sowing and subsequent yield fluctuations do not influence costs at all. That is, low or high yielding crops are assumed to incur the same costs of harvest. (This is plausible in the Australian context since weedy or sparse crops are often expensive to harvest and often "bumper" crops harvest very evenly.) This simplifying assumption means the farmers' decision reflects costs that are independent of yield risk.[1] The second simplification in eq. (3) is the quadratic form which can be viewed as a second order approximation to any higher order differentiable cost function.

Prices are assumed to follow a naive expectations model with an exponential error term:

$$p_t = p_{t-1}(1 + u_{1t}) \tag{4}$$

where u_{1t} is a random variable with zero mean and variance σ_1^2. Realised production is:

$$q_t = A_t Y_t \tag{5}$$

where A_t, the decision variable, is area sown in t-1 resulting in the area of wheat in t, and Y_t is yield per hectare. Yield is also stochastic and has a similar functional form to price. However, the equation incorporates a trend, t, to capture improvements in crop husbandry over time:

$$Y_t = (d + \beta_1 t)(1 + u_{2t}) \tag{6}$$

where u_{2t} is a random variable with zero mean and variance σ_2^2. Hence, remembering that A_t is known in period $t-1$, planned production is:

$$E_{t-1}(q_t) = A_t(d + \beta_1 t) \tag{7}$$

The specification of the error term in eqs. (4) and (6) if flexible in the

[1]Pope and Chavas (1994) provide an excellent discussion of ex ante versus ex post treatments of costs.

Just and Pope (1978) sense and it also avoids the introduction of the undesirable incentive to "like risk" imposed by, for example, an exponential error term.[2]

As the Australian wheat industry contributes less than three percent of world supply, growers are price takers both individually and collectively. Hence u_{1t} and u_{2t} are assumed to be independent so that $E_{t-1}(u_{1t}u_{2t}) = 0$. Profits can be rewritten as:

$$\pi_t = p_{t-1}(1 + u_{1t})q_t - a - bE_{t-1}(q_t) - c[E_{t-1}(q_t)]^2 \qquad (8)$$

and, conditional expected profits are:

$$E_{t-1}(\pi_t) = p_{t-1}E_{t-1}(q_t) - a - bE_{t-1}(q_t) - c[E_{t-1}(q_t)]^2 \qquad (9)$$

From the independence of u_{1t} and u_{2t} it follows than p_t and q_t are independent. Thus, the conditional expectation of utility is:

$$E_{t-1}(U_t) = E_{t-1}(q_t)[p_{t-1} - b] - [E_{t-1}(q_t)]^2 \left[c + \frac{k}{2} p_{t-1}^2(\sigma_1^2 + \sigma_2^2 + \sigma_1^2\sigma_2^2)\right] - a \qquad (10)$$

which is maximized with respect to the decision variable, area planted:

$$\frac{\partial E_{t-1}(U_t)}{\partial A_t} = (d + \beta_1 t)(p_{t-1} - b) - 2A_t(d + \beta_1 t)^2 \left[c + \frac{k}{2} p_{t-1}^2(\sigma_1^2 + \sigma_2^2 + \sigma_1^2\sigma_2^2)\right] = 0 \qquad (11)$$

This is solved for the desired area, $\hat{A}_t$:

$$\hat{A}_t = \frac{p_{t-1} - b}{[2c + kp_{t-1}^2(\sigma_1^2 + \sigma_2^2 + \sigma_1^2\sigma_2^2)](d + \beta_1 t)}. \qquad (12)$$

Desired area, $\hat{A}_t$, is the acreage that would maximize utility. There is a second and independent component to "desired area" coming from the incorporation of marginal land into agricultural production over the sample period. This has occurred as a result of government sponsored land clearance programs and gradual upgrading of grazing pasture to wheat land following fertilizer applications. This is modeled as an additional

[2] If the error term were of the form e^u, then an increase in uncertainty (in either prices or production) would result in a positive increase in expected revenue creating an incentive to like-risk. We thank an anonymous referree for pointing this out.

element to the desired area by adding a time trend to the right hand-side of (12). This time trend should capture the increased availability and productivity of suitable wheat land to farmers:

$$\hat{A}_t = \frac{p_{t-1} - b}{[2c + kp_{t-1}^2(\sigma_1^2 + \sigma_2^2 + \sigma_1^2\sigma_2^2)](d + \beta_1 t)} + \beta_2 t. \quad (12')$$

The parameter β_2 is expected to be positive and significantly different from zero.

Assuming area sown follows a partial adjustment mechanism:

$$A_t - A_{t-1} = (1 - g)[\hat{A}_t - A_{t-1}] + \varepsilon_t \quad (13)$$

where $0 < g < 1$, $E(\varepsilon_t) = 0$ and $E[\varepsilon_1^2 = \sigma_\varepsilon^2$ (Nerlove, 1958). Area planted in period t is then given by:

$$A_t = gA_{t-1} + (1-g)\,\hat{A}_t + \varepsilon_t \quad (14)$$

where, $\hat{A}_t = f(p_{t-1}, t)$, is given by eq. (12′). The results of the estimation of eq. (14) are presented in the next section. The distribution of the error term, ε_t, in eq. (14) is assumed normal with mean zero and variance σ_ε^2.

EMPIRICAL RESULTS AND DISCUSSION

Data for this study were obtained from the Australian Bureau of Agricultural and Resource Economics (ABARE, 1997) for prices received by Australian growers, yield per hectare, and area planted for the period 1955–56 to 1996–97. Prices were deflated to 1989–1990 values using the Australian Consumer Price Index (CPI).

A weighted constrained maximum likelihood estimation (bootstrapped) was used to estimate eq. (14). Prior to estimation, a test for ARCH(1) errors in the price equation was conducted. The residuals $\hat{u}_{1t}$ were computed from eq. (4), $\hat{u}_{1t} = p_t/p_{t-1} - 1$. The sample autocorrelation function and Ljung-Box-Pierce Q-statistic of $\hat{u}_{1t}$ were computed. No noticeable patterns appeared in the correlogram plot and the resulting Q-statistic (12 lags) was 2.7 with a probability value of 0.997. Moreover, the Lagrange multiplier (LM) for ARCH(1) errors was computed from the estimated eq. (15):

$$\hat{u}_{1t}^2 = \alpha_o + \alpha_1\hat{u}_{1,t-1}^2 + \varepsilon_t. \quad (15)$$

$LM = T^*R^2$ where T is sample size and R^2 is the estimated coefficient of determination from eq. (15). The probability value of the calculated

LM test is 0.93393, and thus fail to reject the null hypothesis (H_o: $\alpha_1 = 0$) at the five per cent significance level. These results indicate that the assumption of $E_{t-1}(u_{1t}^2) = \sigma_1^2$ is reasonable. The residuals in the yield equation, $\hat{u}_{2t}$, reflected changes in climate. It has been argued in the literature that the southern oscillation weather index is introducing some predictable volatility (Watson, Zinyowera, and Moss, 1996). Thus, a test for ARCH was conducted on $\hat{u}_{2t}$. Estimation of eq. (6) by OLS was carried out to obtain estimates of d and β_1 and the residuals were tested. The LM test for ARCH(1) errors yielded a probability value of 0.79038, thus it was reasonable to assume that $E_{t-1}(u_{2t}^2) = \sigma_2^2$.

From the estimated residuals of eq. (4) an estimate of σ_1^2 is,

$$\hat{\sigma}_1^2 = \sum_{t=2}^{T} \hat{u}_{1t}^2/(T - 1) \tag{16}$$

From eq. (6)

$$\hat{u}_{2t} = \frac{Y_t}{(\hat{d} + \hat{\beta}_1 t)} - 1 \tag{17}$$

then,

$$\hat{\sigma}_2^2 = \sum_{t=1}^{T} \hat{u}_{2t}^2/T \tag{18}$$

Therefore, estimates for d, β_i, g, b, c, and k are needed. The algorithm used to estimate these parameters is as follows:

1. Starting values for d, β_1, c, k from Simmons and Rambaldi (1997), b=100 from a grid search, $\beta_2 = 0.002$ from a linearized version of eq. (12′), and g = 0.5.
2. Compute $\hat{\varepsilon}_t$, from eq. (14).
3. Generate with replacement 5000 bootstrap samples of size m = 42 (i.e., sample size).
4. Obtain a new set of estimates by a weighted constrained maximum likelihood estimation (bootstrapped), where the constraints are on the parameters' space as follows:

   $0 < g < 1$;

   d, β_i, and $c \geq 0$ and

   $1e-16 < k < 0.1$.

TABLE I

Bootstrap estimates of the parameters

Parameter	*Initial Value*	*Bootstrap Estimates (5000 Bootstrap Samples)*
g	0.5	0.6539**
		(0.03499)
b	100	100.001**
		(0.01676)
d	1	0.9891**
		(0.0188)
β_1	0.0124	3.8e-06
		(0.0001)
β_2	0.002	0.01**
		(0.0008)
c	0.0117	0.00066*†
		(0.0005)
k	9.633e-8	1.604e-06**
		(1.605e-07)
σ_1^2	0.03918	0.0392
σ_2^2	0.03494	0.0973
Log-Likelihood	−9.3091853	−8.34847
R^2		0.84

Note: Standard Errors appear in brackets; **Significant at the 5%; *Significant at the 10%, *†p-value = 0.17.

The word "constrained" stems from the fact that bounds on the parameter spaces were set for the estimation. The constraints were not binding for any of the estimates as can be observed for the histograms presented below.

The weights are Poisson pseudo-numbers with expected value equal to the number of observations, which is asymptotically equivalent to the simple random sampling with replacement.[3] The advantage in generating re-sampled parameter estimates and their covariance is that histograms can be produced for each of the parameters in the model. In this case, the shape of the distribution of $\hat{k}$ is of practical importance.

5. Compute $\hat{\sigma}_2^2$ using eq. (18).
6. Iterate 2 to 5 until convergence (for these data, the procedure converged in six iterations).

Table I presents mean values of the bootstrap estimates and goodness-of-fit measures. Figures 1 to 7 show the histograms based on the

[3]The estimation was performed with the CML module in GAUSS System, Aptech.

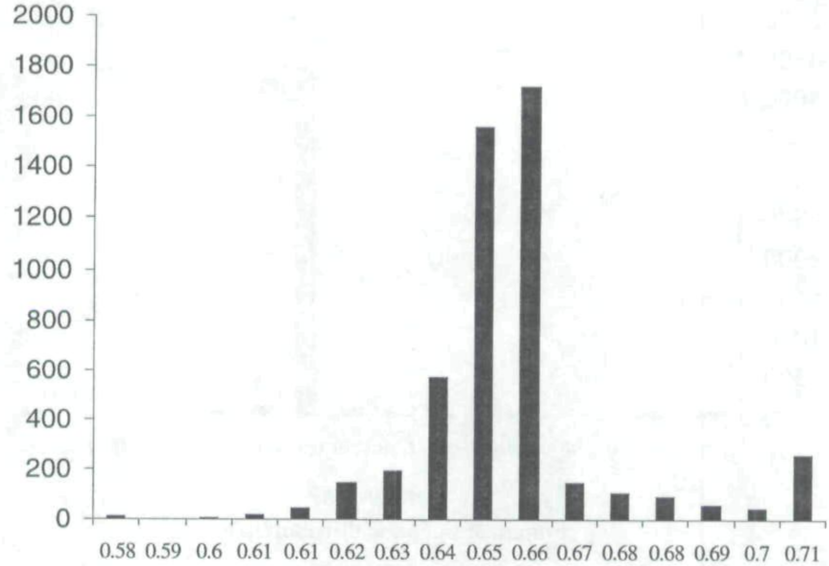

FIGURE 1
Empirical sampling distribution of "ĝ".

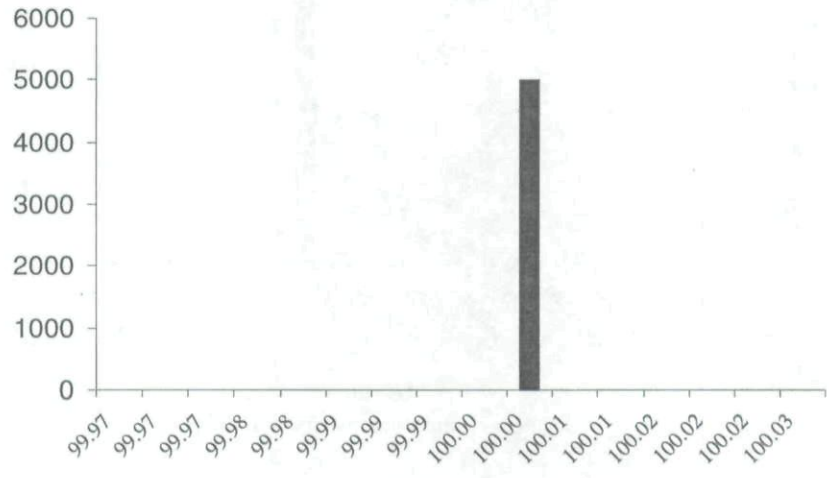

FIGURE 2
Empirical sampling distribution of "$\hat{b}$".

bootstrap samples (empirical sampling distributions). All point estimates are statistically significant except for the estimate of β_1 and c which are only marginally significant. To assess goodness-of-fit, a conventional R^2 was computed as the square of the correlation between A_t and predicted A_t. The value of R^2 is 0.84. Visual inspection of the histograms seems to

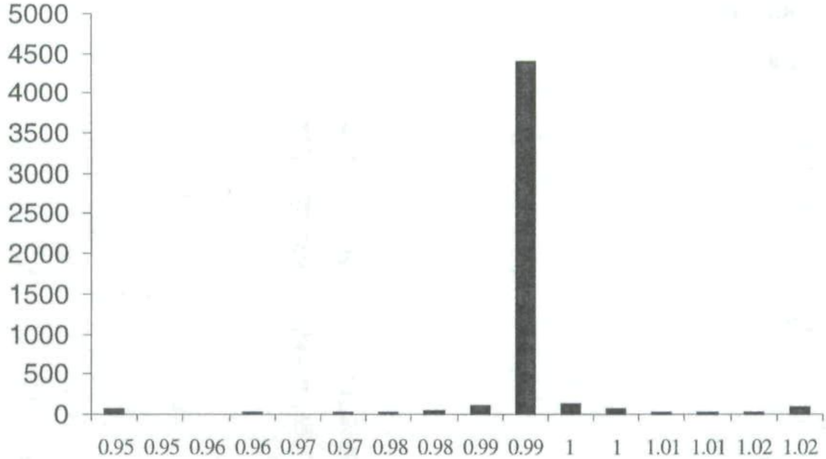

FIGURE 3
Empirical sampling distribution of "$\hat{d}$".

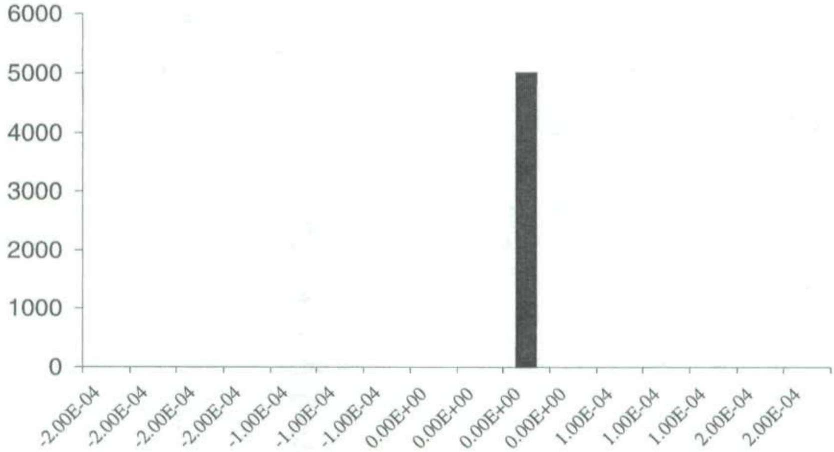

FIGURE 4
Empirical sampling distribution of "$\hat{\beta}_1$".

indicate that the empirical sampling distributions of $\hat{g}$, $\hat{\beta}_2$, and $\hat{k}$ show some excess kurtosis, and for the case of $\hat{k}$, some skewness. The empirical distribution for $\hat{k}$ is roughly within the range 1.26e-06 to 1.86e-06 with a mean value of 1.60e-06, and slightly negatively skewed (see Fig. 6). Because the measurements of risk aversion are dependent on k this is of practical relevance (details below).

The short- and long-run own-price elasticities were calculated as point estimates at the means based on partial derivatives of area with

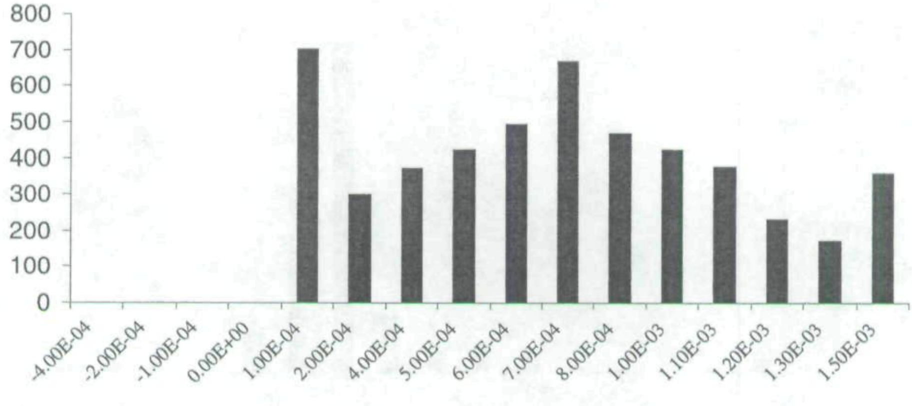

FIGURE 5
Empirical sampling distribution of "$\hat{\beta}_2$".

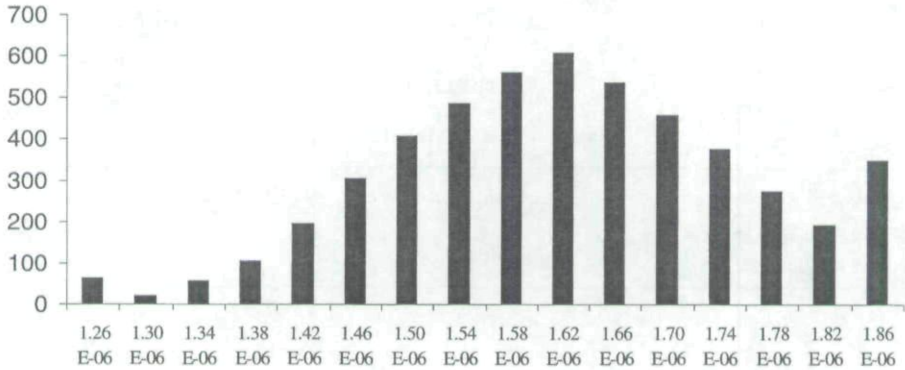

FIGURE 6
Empirical sampling distribution of "$\hat{k}$."

respect to price from eqs. (12) and (14). Table II presents the short- and long-run elasticity of supply with respect to price, 0.42 and 1.21, respectively. These values are slightly lower than expected from previous studies, however within the relevant range (Vincent, Dixon, and Powell, 1980; McKay, Lawrence, and Vlastuin, 1982; Fisher and Wall, 1990; Johnson, Powell, and Dixon, 1990.)

Empirical estimates of the coefficient of absolute risk aversion, k, are not useful on their own. To bring the degree of risk aversion into context, the risk premium, expressed as a percent of expected profits, and the coefficient of relative risk aversion were calculated. Pratt (1964, p. 124) defines the risk premium (π) "such that he would be indifferent between receiving a risk z and receiving the non-random amount $E(z)$ −

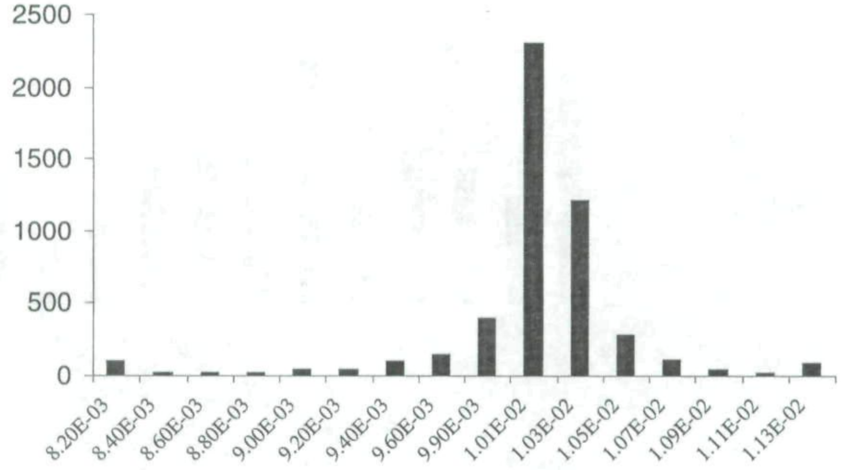

FIGURE 7
Empirical sampling distribution of "ĉ."

TABLE II

Elasticities and Measures of Risk

Elasticity of Supply With Respect to Wheat Price	*Risk Premium* $\hat{R} = \frac{\hat{k}}{2} p_{t-1}^2 q_t^2 [\hat{\sigma}_1^2 + \hat{\sigma}_2^2 + \hat{\sigma}_1^2\hat{\sigma}_2^2]$	*Risk Aversion*
Short-Run: 0.42	10.27% of expected profits	Coefficient of absolute risk aversion ($\hat{k}$): 1.604e-06
Long-Run: 1.21		Coefficient of relative risk aversion ($\pi \times \hat{k}$): 0.1049

π . . . " Based on this, the risk premium was calculated in our study as the difference between expected profit and certainty equivalent profit. The risk premium, R, is:

$$R = \frac{k}{2} p_{t-1}^2 [E_{t-1}(q_t)^2 (\sigma_1^2 + \sigma_2^2 + \sigma_1^2\sigma_2^2)] \tag{19}$$

which is evident from eq. (10).

Expected profit is the area between the price line and $\hat{A}_t$ which can be interpreted in this context as the marginal cost curve. They can simply be calculated as expected revenue minus costs, by substituting the estimates from Table I. The risk premium is 10.37 percent of expected profits or 2.76 percent of expected revenue.

The coefficient of relative risk aversion is a "unit-free" measure of risk aversion that allows comparisons between groups and between results from different studies (Pratt, 1964). It is measured as $-\pi \times (U'/U")$ or, in this study, as $\pi \times k$. Using the measure of expected profit, the coefficient of relative risk aversion is 0.105, which is considerably lower than values around 0.7 obtained in previous work. Previously measured coefficients of relative risk aversion range widely with Chavas and Holt (1996) reviewing previous work and finding estimates between 0 and 7. The measures of risk computed in this article appear in Table II. The low level of risk aversion found is consistent with Bond and Wonder's 1980 conclusion that Australian wheat growers are not very risk averse. Bond and Wonder questioned farmers about responses to hypothetical scenarios using the ELCE technique described in Anderson, Dillon, and Hardaker (1977).

Finally, given the standard deviation of the sampling distribution, an interval estimate for k can be constructed that would result in an interval estimate for the coefficient of relative risk aversion. A simple rule of approximation to the sampling error is $\pm 2 \times$ standard error. Applying this simple rule, a 95% Interval Estimate for k is given by [1.58E-06, 1.63E-06], resulting in an interval for the Coefficient of Risk aversion of 0.103 to 0.106.

While the results show that farmers are significantly risk averse in their response to price and climate volatility, the risk premium seems small since it amounts to only a few dollars for each ton of wheat produced, however, it is 10.37% of profits. A willingness to pay 10 percent of profits is clearly economically significant. The seemingly small value of the coefficient of relative risk aversion compared to a relatively "large" risk premium reflects the way the risk premium is measured. That is, the risk aversion coefficient is multiplied by the variance of profit which is relatively large.

The results of the study show that farmer's willingness to pay for certainty is significant when compared to their expected profits. This indicates that improved management of risk, either through education and extension or through industry or government policy, has the potential to significantly improve welfare. Clearly, this is a matter of evaluating benefits and costs associated with different proposals however knowing, from studies such as this one, that there are potential benefits is very important. This is particularly so for programs in developing countries where small-holders are likely to be considerably more averse to risk than farmers in Australia.

CONCLUSIONS

This study examined the effect of agent response to price and climate volatility on producer decision making in a theoretically consistent framework. Considering theory and method separately, neither the theoretical model nor the empirical approach used in the study are new from a conceptual standpoint. The E-V framework has been used extensively by most "branches" of economics and its development in this study to incorporate production and price risk separately in an expectations framework is, in the broader scheme of things, minor. Similarly, there is a growing body of published empirical work showing significant behavioral responses to volatility in state variables. However, this study represents an improvement over previous approaches of assessment of theoretically important measures such as the coefficients of absolute and relative risk aversion because it links derived theory and appropriate method.

The results shows that farmers are significantly risk averse in their response to price and climate volatility. The risk premium might seem small because it amounts to only a few dollars for each ton of wheat produced, however, at 10.37% of profits, it is an economically significant result. Certainly, if the results of the analysis are accepted, they indicate that risk is an integral consideration in the understanding of farmers and their management systems.

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