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A New Mean Standard Deviation Utility Function and the Behaviour Towards Risk of Specialist Irish Agricultural Producers: 1988-1997

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Abstract

This paper employs a new mean-standard deviation utility (MSU) function with more general properties than traditionally employed function in the agricultural economics literature to explore the risk preferences of Irish barley producers between 1988 and 1997. During this period EU cereal producers were subject to major policy reforms (MacSharry Reforms) that may have influenced the nature of producers' risk attitudes. Our findings support the proposed function in preference to the Linear Mean Variance (LMV) function. We find that Irish cereal producers are highly risk averse but their degree of risk aversion falls appreciably in the period before and following the MacSharry Reforms. We also find that the majority of Irish producers exhibit increasing absolute risk aversion (IARA) but the proportion displaying DARA increases substantially following the 1992 Reforms. The policy implications of our results are discussed.

1 Introduction

In this paper we apply a mean standard-deviation utility (MSU) function, which has more desirable properties than the typical function that is used in the agricultural economics literature, to determine the behaviour towards risk of specialist Irish barley producers. We examine this issue over the period 1988 to 1997. This period covers two developments in European agricultural policy that could potentially have affected producers' risk attitudes, namely, the introduction of partially decoupled direct payments under the MacSharry CAP Reforms in 1992 and the progressive liberalisation of cereal prices that has been in place now for several years and which was given an additional momentum by the MacSharry Reforms.

Our paper has three main motivations. Sandmo (1971), in a well-known paper, showed that the structure of risk preferences is of crucial importance in determining whether the introduction of, or an increase in, a lump-sum subsidy (such as the 1992 direct payments) has a positive or negative impact on the supply responsiveness of producers. Specifically, he shows that where decreasing absolute risk aversion (DARA) prevails then an increase in direct payments leads to a reduction in output. On the other hand, if increasing absolute risk aversion (IARA) holds, then an increase in direct payments reduces output, while constant absolute risk aversion (CARA) implies that changes in direct payments are unrelated to changes in output. It appears to us to be important that the MSU functions that are used in empirical applications be sufficiently flexible to allow any given data set to determine the structure of risk preferences. However, most applications in agricultural economics restrict the MSU function *a priori* to be of the CARA or DARA classes.

In the context of the growing importance of direct payments in the income and wealth of EU farmers and the debate around the decoupled nature of these direct payments, the nature of producers' attitude towards risk thus assumes a particular significance. Hence our first motivation was to establish the risk preferences of producers with a utility function that was sufficiently flexible to allow for the three different types of risk aversion, namely, DARA, IARA or CARA. Our second motivation was to establish whether the MacSharry 1992 reforms had impacted on the structure or degree of risk preferences themselves. We achieve this by exploring attitudes towards risk before and after the 1992 policy change. A final motivation was to uncover whether producers might be interested in purchasing insurance products or similar hedging instruments to any significant degree. Their preference for such products will be influenced by their attitude towards risk and its degree and also by their relationship towards risk at different levels of income.

Several studies have examined the production effects of the 1992 MacSharry and 1999 Agenda 2000 reforms of the EU Commmon Agricultural Policy (CAP) Cereal Regime, (see, for example, Moro and Sckokai (1999), OudeLansink and Peerlings (1996) and Guyomard et al (1996)). Comparatively few investigations, however, have incorporated producers' risk attitudes, (see, for example, Anton and LeMouel (2002), Sckokai and Moroa (2002), OudeLansink (1999) and Boyle and McQuinn (2001)). This study is a further application in this genre.

The paper is laid out as follows: the next section presents an integrated model of production technology and risk behaviour. The data for the empirical application to specialist Irish cereal producers are next discussed. The results of estimating the model are then presented and interpreted. A final section offers some concluding comments and proffers some implications.

2 A Production Model in the Presence of Risk

2.1 Modelling producers' risk behaviour - introducing a new MSU function

In this paper we derive a new MSU function, denoted by $V(\mu, \sigma)$, where μ is the mean of the random variable and σ is its standard deviation, by directly generating the expected utility of income given assumptions about the utility of farm income function and the distribution of farm income.[1]

A range of assumptions about both the utility function and the distribution of income is possible. In deriving our proposed function we assume a concave CRRA utility function but, crucially, we do not restrict the nature of absolute risk aversion. We also assume that farm income is distributed as a log normal. This distribution is among the more tractable and plausible distributions that could be

[1] This approach contrasts, for example, with that of Saha (1997) and Nelson and Escalante (2004) who proceed by specifying MSU functions without explicitly defining either the underlying function or the distribution of farm income.

used to model farm income. We contrast the empirical results obtained for this model with the more restrictive linear-mean-variance (LMV) MSU function, which implies the more restrictive CARA utility function and where random income is assumed to be normally distributed.

Suppose farm income (y) has a Log Normal distribution. It is conventional to write its parameters as μ^* and σ^*, where, these are the mean and standard deviation of the Normal distribution of log y. Then, as is shown in any standard statistical textbook, the mean and variance of y are

$$\mu = e^{\mu^* + \frac{1}{2}\sigma^{*2}} \tag{1}$$

and

$$\sigma^2 = e^{2\mu^* + \sigma^{*2}} \left(e^{\sigma^{*2}} - 1 \right) \tag{2}$$

Expressing μ^* and σ^* in terms of μ and σ gives

$$\mu^* = log\ \mu - \frac{1}{2}\, log \left(1 + \frac{\sigma^2}{\mu^2} \right) \tag{3}$$

and

$$\sigma^{*2} = log \left(1 + \frac{\sigma^2}{\mu^2} \right) \tag{4}$$

We now employ the utility function

$$U(y) = 1 - \lambda e^{-\lambda\, log\, y} \quad \text{or} \quad U(z) = 1 - \lambda e^{-\lambda z}, \tag{5}$$

where $z = \log y$ and where λ is > -1. As is easily verified

$$V\left(\mu^*, \sigma^*\right) = E[U(z)] = \int U(z) f(y, \mu^*, \sigma^*) dy = 1 - \lambda e^{-\lambda\mu^* + \frac{1}{2}\lambda^2\sigma^{*2}} \tag{6}$$

Substituting (3) and (4) gives our MSU function as

$$V(\mu, \sigma) = 1 - \lambda e^{-\lambda\, log\mu + \frac{\lambda}{2}(\lambda+1)log(1+\sigma^2/\mu^2)} \tag{7}$$

It can be easily shown that V is concave provided $\mu^2 > (\lambda + 2)\sigma^2$, which implies

$$\begin{aligned} &V_\mu \geq 0 \\ &V_\sigma \leq 0 \\ &V_{\mu\mu} \leq 0 \\ &V_{\sigma\sigma} \geq 0 \\ &V_{\mu\mu}V_{\sigma\sigma} - (V_{\mu\sigma})^2 \geq 0 \end{aligned} \tag{8}$$

where, V_μ, V_σ denote the partial derivatives of V with respect to μ and σ and $V_{\mu\mu}$, $V_{\sigma\sigma}$ and $V_{\mu\sigma}$ are the corresponding second derivatives.

Given these conditions then it is well known that there is an exact correspondence between the measures of risk aversion derived from $U(y)$ and $V(\mu, \sigma)$ (see, for example, Eichner and Wagener (2005)). Specifically, the Marginal Rate of Substitution between σ and μ, that is, $S = -(V_\sigma/V_\mu)$, corresponds to the Arrow-Pratt measure of Absolute Risk Aversion (ARA). Also the partial derivative of S with respect to μ indicates the nature of ARA, specifically, if, $S_\mu < 0$ (DARA);if $S_\mu = 0$ (CARA) and if $S_\mu > 0$ (IARA). Similarly, if we increase μ and σ by the same multiple t, then the partial derivates of S with respect to t indicates the nature of relative risk aversion or RRA. Thus, $S_t < 0$ implies decreasing relative risk aversion (DRRA); $S_t = 0$ gives constant relative risk aversion (CRRA) and $S_t > 0$ gives increasing relative risk aversion (IRRA).

From (7) we have

$$V_\mu = (1 - V)\frac{\lambda}{\mu}\frac{[\mu^2 + (\lambda + 2)\sigma^2]}{\mu^2 + \sigma^2} \tag{9}$$

$$V_\sigma = -(1-V)\frac{\lambda(1+\lambda)\sigma}{\mu^2+\sigma^2} \tag{10}$$

The Arrow-Pratt measure of ARA is, thus, given by

$$S = \frac{(1+\lambda)\mu\sigma}{\mu^2+(\lambda+2)\sigma^2} \tag{11}$$

and the nature of ARA is obtained from

$$S_\mu = \frac{S}{\mu}\frac{[(\lambda+2)\sigma^2-\mu^2]}{[\mu^2+(\lambda+2)\sigma^2]} \tag{12}$$

This is a particularly interesting expression. Once income becomes large enough, decreasing absolute risk aversion or DARA must hold. However, it could be positive, implying IARA, for low mean income (or high λ and variance). Specifically, (12) will imply: CARA if $\mu = \sigma\sqrt{\lambda+2}$; IARA if $\mu < \sigma\sqrt{\lambda+2}$; and DARA if $\mu > \sigma\sqrt{\lambda+2}$.

It can also be easily shown that if σ and μ are both increased by a multiple t, then $S_t = 0$, implying that the function satisfies Constant Relative Risk Aversion or CRRA.

Thus, our proposed MSU function is capable of encompassing DARA, IARA and CARA structures and one does not have to impose a particular structure on a given dataset *a priori*. We benchmark our model against the so-called linear mean-variance MSU function defined by

$$V = \mu - \left(\frac{\lambda}{2}\right)\sigma^2 \tag{13}$$

This function, as noted by Nelson and Escalante (2004), has been widely employed in the agricultural economics literature. The function is obtained by taking the expected utility of y, where $U(y) = 1 - e^{-\lambda y}$ and y is assumed to be normally distributed. Given the latter assumption together with the fact that (13) implies CARA, it is apparent that the linear mean-variance function is highly restrictive.

However, due to its widespread use in the literature it is useful to contrast the implications of using this MSU function as against our less restrictive alternative.

The marginal rate of substitution (S) for the linear mean-variance model is simply

$$S = \lambda\sigma \tag{14}$$

2.2 The producers' choice problem

In this section of our paper we follow the seminal contribution of Sandmo (1971)

Irish specialist cereal producers are assumed to maximise the expected utility of income with income defined as follows

$$y = pq - c(q) - f + d + o \tag{15}$$

where: y is annual income; p and q are prices and outputs respectively; $c(q)$ are the allocable costs to barley; f denotes non-allocatable fixed costs; d are Direct Payments; and o denotes 'other' income which includes income from other farming activity, off-farm household income and income deriving from non-farm assets.

We assume that the only source of randomness that affects the optimal level of (non-stochastic) q, derives from random prices p. These assumptions imply that

$$\begin{aligned} \mu &= \mu_p - c(q) - f + d + o \\ \sigma &= \sigma_p q \end{aligned} \tag{16}$$

where: μ_p and σ_p denote the mean and standard deviation of the random barley price.

Producers are assumed to maximise $V(\mu, \sigma)$ where μ and σ are given by (16) by choosing the level of q. It can easily be shown that the first-order condition for a maximum is given by

$$\mu_p - c'(q) = S\sigma_p \tag{17}$$

where: $c'(q)$ denotes the marginal cost of production.

We substitute (11) for S in (17) to obtain the central equation of our model. We contrast the results obtained from this model with the results generated by the traditional linear mean-variance model which is obtained by substituting (14) for S in (17).

We complete the model by specifying cost and production functions that are assumed to be of the Cobb-Douglas form. Thus, the complete empirical model is given by

$$\begin{aligned}
&(i) \quad \log q = \log(a_0) + \sum_{j=1}^{5} a_j \log(x_j) + \epsilon \qquad (18)\\
&(ii) \quad \log c = (1/a_1)\log(1/a_0) + \log(w_1) + (1/a_1)\log(q) + \\
&\sum_{j=1}^{5}(a_1/a_{j+1})\log(x_{j+1}) + \xi \\
&(iii) \quad \log(\mu_p - (1/a_1)(c(q)/q)) = \log(S\sigma_p) + \eta
\end{aligned}$$

where: the x's are production inputs, x_1 = allocatable (variable) inputs (fertilisers, crop protection and other inputs), x_2 = (quasi-fixed) labour inputs, x_3 = (quasi-fixed) capital inputs, x_4 = (quasi-fixed) land area and x_5 = an indicator of soil quality; the a's are production/cost parameters to be estimated; w_1 denotes the price of allocatable (variable) inputs and ϵ, ξ and η are error terms.

2.3 Some useful comparative statics

While Sandmo (1971) has derived the comparative statics in the general case, it is useful to derive some of the key comparative statics results for our particular system. The key results are obtained by totally differentiating the first-order condition in (18)(iii). Denoting the partial derivatives of (18)(iii) by $f(\ldots)$ some of the key comparative static results may be shown to be obtained as follows

$$
\begin{aligned}
&(i) \quad f_{\mu} = -\sigma_p \left[\frac{S}{\mu}\right]\left[\frac{(\lambda+2)\sigma^2-\mu^2}{\mu^2+(\lambda+2)\sigma^2}\right] \\
&(ii) \quad f_{\mu_p} = 1-\sigma_p \left[\frac{Sq}{\mu}\right]\left[\frac{(\lambda+2)\sigma^2-\mu^2}{\mu^2+(\lambda+2)\sigma^2}\right] \\
&(iii) \; f_{\sigma_p} = -2S\left[\frac{\mu^2}{\mu^2+(\lambda+2)\sigma^2}\right] \\
&(iv) \quad f_q = -c''(q)-\sigma_p\left[\frac{S}{\mu q}\right]\left[\frac{((\lambda+2)\sigma^2-\mu^2)(q(p-c'(q))-\mu)}{\mu^2+(\lambda+2)\sigma^2}\right] \\
&(v) \quad \frac{\delta q}{\delta \mu_p} = -\frac{f_{\mu_p}}{f_q} \\
&(vi) \quad \frac{\delta q}{\delta \mu} = -\frac{f_{\mu}}{f_q} \\
&(vii) \; \frac{\delta q}{\delta \sigma_p} = -\frac{f_{\sigma_p}}{f_q}
\end{aligned} \tag{19}
$$

where: $c''(q)$ denotes the second derivative of cost with respect to output.

The sign of the supply response effects (expressions $(19(v)-19(vii))$ depend on the signs of the partial derivatives. With the exception of the partial derivative of the standard deviation of price $(19(iii))$, which is unambiguously negative in sign, the signs of the remaining derivatives depend on the structure of risk aversion. It is apparent that their signs depend on the relative magnitudes of μ^2 and $(\lambda+2)\sigma^2$ and, as noted previously, if $\mu^2 < \sigma^2(\lambda+2)$ IARA is implied, while DARA holds if $\mu^2 > \sigma^2(\lambda+2)$. If $\mu^2 = \sigma^2(\lambda+2)$, then CARA holds and the expressions are considerably simplified. If IARA holds we can see that f_μ will be negative, as will f_q, provided $-c''(q)$ exceeds the second expression in $(19(iv))$ which will be positive under IARA. It follows then that an increase in expected income brought about by an increase in decoupled direct payments will cause output to fall, *ceteris paribus*, under IARA but to increase under DARA. Irrespective of whether DARA or IARA holds, it is apparent that an increase in the standard deviation of price will always lead to a fall in output, *ceteris paribus.* It is apparent, however, that the responsiveness of output will be lower under DARA than IARA. It is also evident that output responds positively to a change in expected price under both DARA and IARA. It is interesting to note, however, that we cannot determine whether the degree of output-price responsiveness will be higher or lower under DARA or IARA since the numerator and denominator of expression (19(v)) will be affected in opposite directions.

3 Data Generation

The data used for the analysis are obtained from the Irish National Farm Survey (NFS) conducted by Teagasc[2]. The survey is conducted annually and is based on a representative sample of Irish farms and is used as the Irish input into the Farm Accounts Data Network of the EU.[3]

The data available for our analysis spanned the years 1988 to 1997. As farms can be surveyed over several years, we decided to exploit this feature by constructing balanced-panel data sets. We formed separate panels for the period 1988-1992, comprising 63 producers and for the period 1993-1997, comprising 41 producers. This allows us to estimate our model for the period before and after the MacSharry reforms. The selected panels were confined to specialist barley producers. Barley production is the single largest grain produced in Irish agriculture (between 1997 and 2002, the total area of barley accounted for, on average, 65 per cent of total cereals grown).[4]

The prices for all input items are considered to be non-stochastic and known to producers in advance of the input-application decision. Both output and input prices are assumed to be constant across space and variable only through time. This assumption is particularly valid in an Irish context as most cereal production is located within a relatively small geographical area. The prices used for inputs are national aggregate price indices as produced by the Irish Central Statistics Office.

[2]The Irish Agriculture and Food Development Authority

[3]For more on the NFS see Heavey et al. (1998).

[4]The definition 'specialist' is based on the EU's farm classification system as described in Commission Decision 85/377/EEC. A producer is deemed to be a specialist cereal producer if more than two thirds of the standard gross margin on the farm comes from cereal production.

3.1 Generating the expected value and variance of Irish barley prices

A novel feature of the empirical model that we estimate is the price expectations model used to generate the expected mean, variance and covariance of Irish grain prices.[5] We draw on the model of grain prices proposed by Roche and McQuinn (2003). Only a brief outline of their model will be given here.

Most studies of price uncertainty follow the Chavas and Holt (1990) model. Roche and McQuinn (2003) hypothesise a long-run relationship between Irish and UK grain prices and model expected variances and covariances within an ARCH framework. They explicitly test the forecasting performance of their model against the Chavas and Holt approach using standard forecast statistics (the mean squared error (MSE) and the mean absolute deviation (MAD)) as well as the recent test of "superior predictive ability" (SPA) produced by Hansen (2001). In all cases, the Roche and McQuinn (2003) model outperforms the Chavas-Holt specification. The model in its error correction form is expressed as

$$\begin{aligned} \triangle\mu_{p_t}^{bir} &= \alpha_0 + \alpha_1(p_{t-1}^{bir} - \beta_1 p_{t-1}^{buk}) + \alpha_2 \triangle p_{t-1}^{bir} \\ &+ \sum_{i=0}^{1} \alpha_{i+3} \triangle p_{t-i}^{buk} + u_t \\ &[u_t] \sim N(0, h_t) \end{aligned} \tag{20}$$

The series p^{bir} is the price of Irish feed barley, p^{buk} is the price of British feed barley and u is a stochastic error terms. The conditional variance of Irish barley prices, h_t is estimated as an autoregressive conditional heteroscedasticity (ARCH(1,1)) model.

The use of the results from the ARCH model have a number of attractive features. First, the error correction model captures the dynamic nature of price transmissions between the mean Irish cereal prices and that of its largest grain trading partner. Second, allowing for ARCH errors has been shown to improve the efficiency of the results achieved in such a transmission framework (see Bollerslev et al. (1992) for example). The model is estimated on a rolling basis and forecasts are

[5] Pagan and Ullah (1988), amongst others, raise issues regarding both estimation and inference when using estimated results for expected prices and variances.

generated for the sample period 1984-1998. We refer to this model as the "rational price expectations" (RPE) model.

While some empirical evidence (see, for example, Chavas (2000)) suggests that certain producers may adhere to simpler price-expectations' mechanisms such as naïve expectations, we favour the rational expectations' model, that is, price expectations are, on average, informed exclusively by the underlying data-generating process of domestic grain prices. The close relationship between UK and Irish grain prices revealed by the Roche and McQuinn (2003) model is a common feature of supply-response analysis in an Irish context (see the FAPRI-Ireland models of the Irish agricultural sector, for example, Binfield et al. (2000) and Binfield et al. (2001)). The price-taking assumption which is routinely invoked in analyses of a small open agri-economy such as Ireland's is not accommodated within a naïve-price-expectations' model. However, as a sensitivity check on our results, we also employ the following hybrid of the naïve expectations' and Chavas-Holt (CH) approaches

$$\mu_{p_t}^{bir} = \psi_0 + \psi_1 p_{t-1}^{bir} + e_t \tag{21}$$

$$Var\left(p_t\right) = 0.5\ \left(e_{t-1}^2\right) + 0.33\ \left(e_{t-2}^2\right) + 0.17\ \left(e_{t-3}^2\right) \tag{22}$$

We label this model the "C&H naïve price expectations' model" (CHNPE).

In Table 1 (insert Table 1 here) we provide summary statistics of the variables used in the analysis.

Some of the more interesting features of these data are the substantial increase that is revealed for the standard deviation of barley prices between the first and second periods of our analysis and the emergence of direct payments as a significant component of expected farm income subsequent to the MacSharry Reforms.

4 Estimation and Results

4.1 Specification tests

Our preferred model specification is given by equation (18), with μ_p and σ_p derived from equation (20) and (21), that is, our RPE specification and the Marginal Rate of

Substitution between μ and σ, that is S, given by equation (11). We benchmark this model against two alternatives. First, we replace the expected value and standard deviation of prices by equations (21) and (22), that is by the CHNPE specification. Second, we replace the Marginal Rate of Substitution between μ and σ by that implied by the linear mean-variance model (LMV).

The system (18) is modified in the estimation by the addition of farm-specific intercepts to capture possible time-invariant efficiency differences across farms and we also include a time dummy variable to capture inter-temporal efficiency changes that may be common to all farms over the period of estimation. Thus, the parameters in (18) of our preferred and alternative specifications are estimated as a fixed-effects-panel-data model. Because of the possible endogeneity of right-hand-side variables, we also employ an instrumental variables estimator, using the price of phosphates and plant protection products as instruments. Finally, to enhance the efficiency properties of our estimates we allow for contemporaneous correlation in the error terms of the three equations in (18).[6] The system (18) is estimated for two different time-periods (1988-1992), (1993-1997) using a balanced panel in each case.[7]

Before presenting the results for our preferred specification, we first report the results of the specification tests that we conducted in respect of the nature of price expectations (that is, RPE vs. CHNPE) and the specification of the MSU function (that is, our proposed new MSU function vs. the LMV specification). As the alternative models cannot be obtained by parametric restrictions on the preferred specification we need to employ a non-nested specification test. An appropriate test for this exercise is the overlapping version of the Vuong (1989) test. This is essentially a likelihood-ratio test, suitably normalised. A test value close to zero would suggest that there was no significant difference between the preferred specification and the proposed alternatives. The results are given in Table 2 (insert Table 2 here). The test results imply strong support for our preferred specification so we now proceed to discuss the detailed results obtained for this model.

[6] We use the non-linear three-stage systems estimator in WinRATS 6.00. The relevant programs are available from the authors upon request.

[7] We could have obtained estimates for a balanced panel over the full period, 1988-1997, but there were only 16 producers available.

4.2 Estimation results - all farms

The coefficient estimates of the system in (18) are shown in Table 3 (insert Table 3 here). In terms of the standard input production parameters, all of the coefficients are significant at the one per cent level. All of the production inputs, bar labour, have the correct positive signs and are of a reasonable magnitude. This result is common across both time-periods. The negative sign on labour could arise for a number of reasons. First, the basic data on the labour input refer to labour that is available for use on farms as opposed to what is actually employed. Thus the negative sign may suggest that there is more (family) labour available than is actually required on specialist Irish cereal farms. This may be reflected in the higher than average off-farm employment activity among specialist cereal producers. A second and related explanation of the negative coefficient could be due to the fact that cereal production is a highly mechanised activity and there is extensive use of contract services on cereal farms.

An F-test on the coefficients of the farm-specific dummy variables rejects the null hypothesis of joint non-significance. In Table 4 (insert Table 4 here) we report the results of miss-specification tests performed on the system. Both specification tests are for the null of autocorrelation and heteroscedasticity in a fixed effects model and are taken from Chapter 10 of Verbeek (2000). From the table it would appear that the error processes in all estimations are 'well-behaved'.

The highly significant estimate of the risk coefficient, λ, indicates that specialist Irish cereal producers are risk averse. It is apparent that the estimate declines considerably between the earlier and later period, that is, from 32764 for the 1988-1992 period to 1108 for the 1993-1997 time period. Thus over the full period 1988-97 Irish barley producers appear to have become much less sensitive to a given degree of price risk.

4.3 Estimation results for small ($<$ 50 acres) and large ($>$ 50 acres) farms

We also estimated the system in (18) for small farms ($<$50 acres) and large farms ($>$50 acres). Our primary interest was to ascertain whether there were significant differences in the attitudes towards risk between small scale and large-scale producers. The results are presented in Table 5 (insert Table 5 here).

The production coefficients are fairly similar across both farm categories and relatively stable for the two time periods. However, there is a substantial difference in the risk parameter for the first-period estimates. Large farms have a risk parameter that is only 54 per cent of the magnitude of smaller farms. However, in the second period the magnitude of this parameter has fallen substantially for both categories and the risk parameter is seen to be of a very similar magnitude for both farm-size groups.

4.4 Some key comparative static results

Following our derivations in Section 2.3, we now consider the key comparative statics of our estimated model. For ease of interpretation, we present the comparative static findings in terms of elasticities, which are estimated at the sample means. The findings are given in Table 6 (insert Table 6 here).

The elasticity of the marginal rate of substitution (S) with respect to changes in expected farm income reveals that specialist Irish cereal producers exhibit IARA in both periods, although the magnitude of the elasticity declines between the first period and the second. It's also interesting to note that we do not find any difference in the elasticity values between large and small farms.

While DARA is often preferred as an assumed risk structure, the finding of IARA is not uncommon in the literature, see, for example, Wolf and Pohlman (1983) and Abdulkadri et al. (2003). From a policy perspective, the finding of IARA suggests that the receipt post 1992 of direct payments by producers would have had a negative impact on the production of barley. This is revealed by our estimates of the elasticity of output with respect to changes in the level of direct payments. These are zero in the first period as there were no DPs but in the second period the elasticities as estimated at the sample means are fairly large (ranging from -0.13 to -0.19), bearing in mind that the direct payments are denominated in nominal terms. Larger farms would also appear to have relatively lower elasticities than smaller farms.

The elasticity of output with respect to the standard deviation of the barley price is seen to increase significantly from the first to the second period. The elasticity is found to be much bigger for smaller farms in both periods, most notably in the 1988-92 period. The main reason for the increase in this elasticity in the second period is because of the substantial increase in the estimated standard deviation of barley prices. Perhaps a more useful way to represent the impact of this variable would be in terms of the percentage impact on output of a change in the standard deviation of prices. For the case of all farms, the latter coefficient is estimated at -0.26 in the first period and -0.15 for the period 1993 to 1997. This result reflects our finding that the estimated λ coefficient in our utility function declines substantially between the first and the second period.

In respect of the elasticity of supply response, we observe a substantial increase in the coefficient between the first and second periods. As with the standard-deviation elasticity, we find that the estimates are bigger for smaller-scale farms.

4.5 Distribution of comparative static results

In the previous section we presented key comparative static effects that were evaluated at the sample means. These estimates will vary across the sample since μ and σ will vary due to differences in output in the case of σ and due to output, production costs and direct payments in the case of μ. As we noted in Section 2.3, the key supply response effects are determined by whether or not the structure of risk preferences are characterised by DARA or IARA. Thus to illustrate the nature of the potential differences that are likely to arise across our sample of farms, we present in Figures 1 and 2 (insert Figures 1 and 2 here) the distribution of elasticities of S with respect to expected income that are generated from our estimates based on all farms[8] for the two periods of our analysis.

These results indicate that a significant proportion of specialist barley producers exhibit DARA in both periods. However, the percentage of the sample displaying DARA increases significantly from about 24 per cent in the period 1988-92 to 34 per cent in the period 1993-97. As noted already, the shift in risk preferences is reflected in a decline in the mean elasticity, but perhaps the more notable statistic is the decline in the median value from 0.61 to 0.31.

[8]The distributions for large and small farms may be obtained from the authors upon request.

5 Concluding Comments

This paper has estimated the risk attitudes of specialist Irish barley producers using a new mean standard deviation utility (MSU) function over the period 1988-1997. We split our data analysis into two periods: 1988-92 and 1993-97, which coincide with introduction of the MacSharry Reforms in 1992.

Our analysis rejects a naïve specification of the first and second moments of expected barley prices that has been used in similar studies. Also, our proposed MSU function is preferred in our particular application over the traditional Linear Mean Variance MSU function that is often employed in the agricultural economics literature.

As to the estimated risk preferences of specialist Irish barley producers, we find a number of interesting results. Irish producers are found to be highly risk averse in relation to barley prices but the intensity of risk aversion declines appreciably pre and post the MacSharry 1992 Reforms. We also found that relative large farms (>50 acres) were much less sensitive to risk compared with small farms (<50 acres) in the first period but both groups displayed a similar, and much lower, risk sensitivity in the period 1993-1997.

Perhaps our most interesting findings concern the nature of risk preferences (that is the substitution of expected payoff for risk) in relation to the level of farm income. We find evidence, at the sample means, of increasing absolute risk aversion, or IARA, for both periods but there is a decline in the IARA elasticity pre and post reforms. This essential finding is reflected in the elasticities for large and small farms. The distribution of elasticity values reveals that about a quarter of the sample exhibit DARA in the first period but about a third do so in the second period. This underlines the apparently significant shift in the structure of risk preferences that we find in the period before and after the MacSharry Reforms.

The most important policy implication of our results concerns the finding of IARA as this implies that an increase in DPs will lead to a reduction in output, whereas DARA suggests the opposite effect.

The MacSharry Reforms introduced partially-decoupled direct payments so these did not constitute a completely risk-free source of income. However, the June 2003 medium-term review (MTR) of the CAP potentially removes any element of risk associated with direct payments following the full decoupling of payments from 2005. Producers will be in receipt of single-farm payments without any production obligation. The receipt by producers of risk-free payments has important implications for the provision of market-provided risk-management tools such as derivatives or insurance markets. Studies by Meuwissen et al (1999a) and Meuwissen et al (1999b) have emphasised the removal of institutionalised income support as a prerequisite for the successful adoption of these products. Our results may have relevance in this context.

Finally, the welfare impacts of reform are worth investigating given the impending budgetary pressures of EU enlargement. Ackrill (2003), for instance, demonstrates that if the EU is to operate under existing budgetary levels under enlargement, it may not be able to offer more than 25 per cent of current payment levels to new member states. The risk-averse nature of most agricultural producers suggests, that under policy reform, where, prices and incomes may have become more risky, the receipt of 'riskless' direct payments can reduce the cost of risk to producers relative to what it would have been if the reform had not actually taken place.

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Table 1: Mean Values of Variables Used in Production System

Variable	1988-1992	1993-1997	1988-1997
μ	30,701	23,216	30551.9
μ_p	128	107.9	117.9
σ_p	1.7	6.9	4.3
q	160.2	133	167.7
x_1	29.2	29.3	44.9
x_2	96.8	66.9	59.4
x_3	227	187	226.8
x_4	61.4	55.8	66.7
x_5	1.5	1.7	5.5
w_1	100.7	112.8	106.7
DP	NA	7,618	NA
N	315	205	160

Note: DP is total direct payments per producer. DP, μ and μ_p are in euros with μ_p in euros per tonne. q is in tonnes and x_4 is in acres. x_1 - x_3 is in index form where quantity = total expenditure / relevant input price index. w_1 is in index form with 1990 = 100 and x_5 is from a 1-5 index where 1 denotes optimal soil quality.

Table 2: Vuong Specification Tests for Models of the Risk Behaviour of Irish Barley Producers

	1988-1992	1993-1997
Test		
Preferred Model v CHNPE	96.572*	27.311*
Preferred Model v LMV	160.434*	110.044*

Note: Actual scores are reported, * denotes significance at the 1 per cent level. The Vuong test is normally distributed asymptotically.

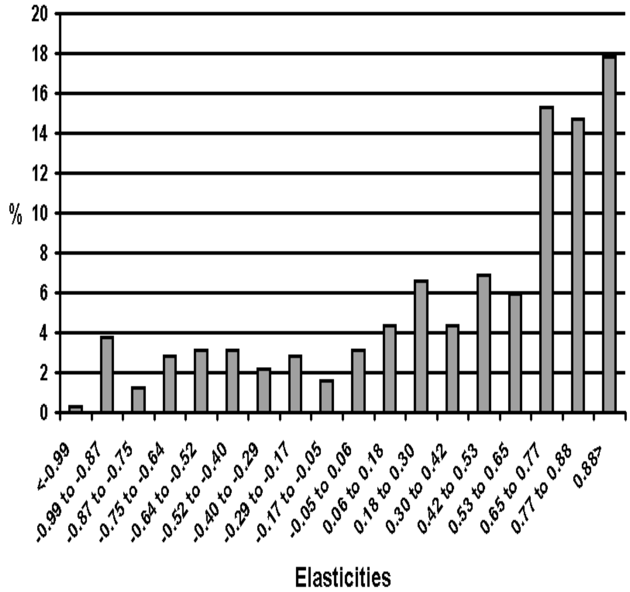

Figure 1: Distribution of Elasticties of 'S' with respect to Expected Income on Irish Specialist Barley Farms, 1988-92

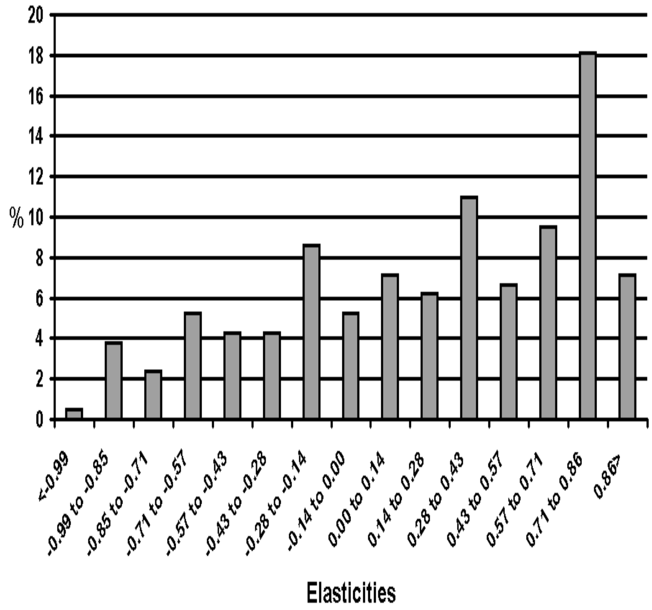

Figure 2: Distribution of Elasticties of 'S' with respect to Expected Income on Irish Specialist Barley Farms, 1993-97

Table 3: Production and Risk Parameter Estimates for Specialist Irish Barley Producers Assuming Rational Price Expectations

	1988-1992	1993-1997
Parameter		
a_0	1.203	1.155
	(0.192)	(0.295)
a_1	0.403	0.526
	(0.021)	(0.033)
a_2	-0.096	-0.113
	(0.006)	(0.006)
a_3	0.339	0.313
	(0.027)	(0.035)
a_4	0.479	0.450
	(0.032)	(0.046)
a_5	0.263	0.393
	(0.056)	(0.125)
Risk		
λ	32764.355	1107.480
	(3072.97)	(147.01)
Tests		
CRS	37.114	18.072
	(0.000)	(0.000)
N	315	205

Note: Standard errors are in parentheses for coefficient estimates, p-values are reported for tests. CRS = constant returns to scale, test = $H_o : a_1 + a_2 + a_3 = 1$, Dummy variable results (both individual and time) are surpressed but are available, upon request, from the authors.

Table 4: Miss-Specification Tests for Estimated Equations

	1988-1992	1993-1997
Heteroscedasticity		
Equation 18(i)	0.105	0.189
Equation 18(ii)	0.447	0.326
Equation 18(iii)	0.203	0.098
Autocorrelation		
Equation 18(i)	0.131	0.215
Equation 18(ii)	0.644	0.426
Equation 18(iii)	0.148	0.112

Note: P-values are reported for the heteroscedasticity and autocorrelation tests. Both tests for heteroscedasticity and autocorrelation are χ^2 tests.

Table 5: Production and Risk Parameter Estimates for Specialist Irish Barley Producers Assuming Rational Price Expectations

	1988-1992		1993-1997	
	(< 50 acres)	(> 50 acres)	(< 50 acres)	(> 50 acres)
Parameter				
a_0	1.185	1.272	0.932	0.694
	(0.267)	(0.363)	(0.255)	(0.251)
a_1	0.404	0.389	0.562	0.537
	(0.027)	(0.031)	(0.055)	(0.044)
a_2	-0.094	-0.093	-0.132	-0.116
	(0.007)	(0.008)	(0.012)	(0.008)
a_3	0.319	0.339	0.393	0.371
	(0.041)	(0.037)	(0.051)	(0.058)
a_4	0.481	0.485	0.539	0.474
	(0.041)	(0.050)	(0.539)	(0.474)
a_5	0.271	0.222	0.625	0.156
	(0.089)	(0.062)	(0.279)	(0.032)
Risk				
λ	42084.3	22850.9	1197.1	1187.7
	(4719.5)	(3376.3)	(223.36)	(206.76)
Tests				
CRS	42.354	15.185	28.789	35.775
	(0.000)	(0.000)	(0.000)	(0.000)
N	155	160	100	105

Note: Standard errors are in parentheses for coefficient estimates, p-values are reported for tests. CRS = constant returns to scale, test = $H_o : a_1 + a_2 + a_3 = 1$, Dummy variable results (both individual and time) are surpressed but are available, upon request, from the authors.

Table 6: Elasticities Estimated at the Sample Mean

	All farms	Small farms (< 50 acres)	Large farms (> 50 acres)
S wrt μ			
1988-92	0.44	0.36	0.36
1993-97	0.27	0.30	0.30
Output wrt expected DPs			
1988-92	0.00	0.00	0.00
1993-97	-0.13	-0.19	-0.15
Output wrt expected σ_p			
1988-92	-0.44	-0.68	-0.44
1993-97	-1.07	-1.26	-1.11
Output wrt expected μ_p			
1988-92	0.49	0.64	0.55
1993-97	1.19	1.38	1.22

Note: DPs are direct payments and *wrt* is "with respect to".