Direct Payments and Risk Premia: How Fared Irish Cereal Producers Under the MacSharry Reforms?

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Abstract

The 1992 EU Common Agricultural Policy (CAP) MacSharry reforms reduced intervention prices for cereals and simultaneously introduced direct payments to producers. This paper, estimating Irish cereal producers risk attitudes and associated premiums, compares the level of the direct payment with producers' risk premia in order to see whether the introduction of the reforms was implicitly welfare enhancing for the 1993-1998 time period.

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1 Introduction

While many 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 Moro and Sckokai (1999), OudeLansink and Peerlings (1996) and Guyomard, Baudry and Carpentier (1996) for example), very little empirical investigation has explored the welfare efficiency of direct payments in the context of estimated producer risk attitudes and associated risk premium levels. Direct payments were introduced as a means of compensating producers for the anticipated decline in producer price following reductions in intervention prices.[1] However, even assuming that the direct payment just compensates for the price reduction, presenting a risk averse group of producers with a virtually riskless lump sum payment results in an additional welfare transfer to producers - a greater proportion of their income is now coming from a relatively less risky source.[2]

Given the increase in the number of studies seeking to estimate EU cereal producers' risk preferences and premium levels (see Anton and LeMouel (2002), Sckokai and Moro (2002), OudeLansink (1999) and Boyle and McQuinn (2001)) and the ongoing bugetary pressures of the existing CAP structure[3], it would seem that an *ex post* evaluation of the efficiency of the direct payment system is warranted, particularly, from a producer welfare perspective. Namely, given producer risk attitudes, did the MacSharry reforms implicitly increase producer welfare levels owing to the level of direct payments introduced?

Thus, this paper seeks to compare risk attitudes and associated risk premiums for Irish cereal producers with the level of direct payments introduced and subsequently increased under the MacSharry reforms. Producers' risk attitudes are estimated in a dual production model of price uncertainty using the highly flexible Saha (1997) mean-standard utility function (MSU). Price uncertainty is explicitly assumed to be the only source of income uncertainty confronting the producer. This assumption

[1] As a result of both sets of reforms Irish intervention prices fell by about 45 per cent between 1992 and 2000.

[2] Direct payments under Agenda 2000 are not fully decoupled - production still has to take place in order for the producer to receive the payment. However, once the decision to produce has been made, the payment is unaffected. In that sense the payment is said to be *partially* decoupled.

[3] For the year 2001, arable crops accounted for 41.5 per cent of the €44,100 million European Agricultural Guidance and Guarantee Fund (EAGGF) Guarantee Section expenditure.

is deemed appropriate in the present case as the resultant risk premiums calculated pertain only to price uncertainty and are, thus, comparable to direct payments, which were introduced to compensate for intervention price reductions. The paper uses a recently developed model of Irish grain prices (see Roche and McQuinn (2003) for more details) as the price expectations model required under the production model framework. The associated risk premium for each producer is then contrasted with the level of direct payment for the 1993-1998 time period.

The paper is laid out as follows; the next section presents the MSU. A dual production model of income uncertainty is then outlined, followed by an empirical application with a note on the price expectations model used. General results are then presented and the level of direct payments is compared with the calculated risk premiums for the entire sample. A final section offers some conclusions.

2 A Production Model under the MSU

Most structural form approaches to the estimation of risk coefficients adopt the expected utility (EU) framework that maximises a pre-defined utility function. Producers are assumed to be risk averse and the data then can only reveal the degree of risk aversion. An alternative approach involves ranking risky alternatives according to the value of a function specified over the first two moments of a producers random pay-off - the mean standard deviation (MS) approach. Typically, a compatibility or consistency between the EU and MS approach is achieved by imposing some restrictions on the risk attitudes of the producers in question.[4] However, Meyer (1987) illustrates that by imposing restrictions on the random variables in the producers' choice set rather than on their risk attitudes, a consistency can be achieved in the ranking of alternatives under both the EU and MS approach. Thus, some of the more powerful assumptions of EU analysis can then be translated in similar conditions to the MS approach without imposing the usual restrictions.

This consistency property provided the basis of the following flexible utility function devised by Saha (1997) in mean standard deviation space

[4]For example, under the popular mean variance utility function for instance, producers are assumed to experience constant absolute risk aversion (CARA).

$$U(\sigma,\mu) = U(M,S) = M^{\Gamma} - S^{\Upsilon} \tag{1}$$

where Γ and Υ are parameters to be estimated and it is assumed that $\Gamma > 0$. Saha (1997) labels this the mean-standard deviation utility function or MSU. Various restrictions can be imposed on the MSU to arrive at more popular EU models. For instance, if $\Upsilon = \Gamma = 1$ is imposed, the linear U(M,S) model is obtained, if Γ is set equal to 1, then CARA attitudes are assumed. Under the MSU, α, the risk attitude measure is given by the slope of the indifference curve in mean standard deviation space

$$\alpha(M,S) = -(U_s/U_M) = (\Upsilon/\Gamma)\, M^{1-\Gamma} S^{\Upsilon-1} \tag{2}$$

The MSU exhibits

(1) Risk aversion, neutrality and risk preference corresponding to $\Upsilon > 0$, $= 0$ and < 0, respectively,

(2) Decreasing, constant and increasing absolute risk aversion as $\Gamma > 1$, $= 1$, and < 1,

(3) Decreasing, constant and increasing relative risk aversion as $\Gamma > \Upsilon$, $\Gamma = \Upsilon$, $\Gamma < \Upsilon$.

Table II[5] of Saha (1997) summarises the suite of risk attitudes, which can be accommodated within the MSU. The greater flexibility evident in this utility framework can be compared with the more restrictive structures under the traditional Arrow-Pratt measures.[6]

The dual production model adopted is one initially proposed by Coyle (1992) and expanded by OudeLansink (1999) to include area allocation. The following list

[5] p.773.

[6] Illustrated in Table I p.772 of Saha (1997).

of variables are used in the producer's decision-making process

$$\begin{aligned}
\mathbf{q} &= \text{a two dimensional vector of outputs - barley and wheat,} \\
\tilde{\mathbf{p}} &= \text{a two dimensional vector of random output prices,} \\
\mathbf{p} &= \text{a two dimensional vector of expected output prices,} \\
\mathbf{x} &= \text{a two dimensional vector of actual/planned inputs,} \\
\mathbf{n} &= \text{a two dimensional vector of input prices,} \\
\mathbf{B} &= \text{a two dimensional vector of cereal areas,} \\
\mathbf{A} &= \text{total on farm cereal area,} \\
\mathbf{D} &= \text{total cereal compensation payments (post 1992),} \\
\mathbf{I} &= \text{total producer off farm income.}
\end{aligned}$$

Random and mean income are defined as[7]

$$\tilde{\pi} = \tilde{\mathbf{p}}^T\mathbf{q} - C(\mathbf{n},\mathbf{q},\mathbf{B}) + \mathrm{D} + \mathrm{I} \tag{3}$$

$$M = \mathbf{p}^T\mathbf{q} - C(\mathbf{n},\mathbf{q},\mathbf{B}) + \mathrm{D} + \mathrm{I} \tag{4}$$

A cost function $C(\mathbf{n},\mathbf{q},\mathbf{B})$ is defined as $\mathbf{n}^T\mathbf{x}$. Irish cereal production is assumed to be non-joint in variable inputs.[8] While the presence of variable input data for both cereals is not in itself sufficient proof for non-jointness in variable inputs,[9] most barley grown in Ireland is sown in the spring, while wheat tends to be mainly a winter crop. Thus, the assumption is considered appropriate in this case. Given the underlying structure in (3), any random alternatives available to the producer are positive linear transformations of the random variable $\tilde{\mathbf{p}}$ and are thus, related to one another by location and scale parameters. As the producer's income function is linear in $\tilde{\mathbf{p}}$, consistency is ensured between expected utility and $\mathrm{U}(\sigma,\mu)$. As output

[7] Data on I, off farm income, is only available for 1998. As a result, off farm income in 1998 was regressed on a series of explanatory variables for that year. The figure for each farm for previous year was then 'backcast' using the 1998 regression results.

[8] Seperate production functions in variable inputs are assumed for both cereals.

[9] Fertiliser spread on barley could technically end up in a wheat field in which case, it would belong in the wheat production function.

prices by assumption are the only source of uncertainty, then, the standard deviation of the producer's random income is

$$S = \left(\mathbf{q}^T \mathbf{Vp} \mathbf{q}\right)^{\frac{1}{2}} \tag{5}$$

$\mathbf{Vp}$ is the (symmetric, positive definite) covariance matrix of output prices. Following Saha (1997), and using (4) and (5), the MSU can be represented as follows

$$U^* \left(\mathbf{p}, \mathbf{n}, \mathbf{Vp}, \mathbf{B}\right) = \max_{\mathbf{q}} U \left(\mathbf{p}^T \mathbf{q} - C\left(\mathbf{n}, \mathbf{q}, \mathbf{B}\right), \left(\mathbf{q}^T \mathbf{Vp} \mathbf{q}\right)^{\frac{1}{2}}\right) \tag{6}$$

The first order condition is given by

$$U_M \left(\mathbf{p} - C_{\mathbf{q}}\left(\mathbf{n}, \mathbf{q}, \mathbf{B}\right)\right) + U_S \mathbf{Vp} \mathbf{q} = 0 \tag{7}$$

which can be rearranged as

$$\mathbf{p} - C_{\mathbf{q}}\left(\mathbf{n}, \mathbf{q}, \mathbf{B}\right) = -\frac{U_S}{U_M} \mathbf{Vp} \mathbf{q} \tag{8}$$

$U^* \left(\mathbf{p}, \mathbf{n}, \mathbf{Vp}, \mathbf{B}\right)$ is the indirect utility function corresponding to $U\left(M, S\right)$. The standard price equal marginal cost result of perfect certainty is achieved if either U_S is zero, or, if price variances and covariances are zero. Optimal output supplies $\left(\mathbf{q}^*\right)$ are attained by solving (8) in terms of $\mathbf{q}$. Thus, $\mathbf{q}^*$ will now be a function of input prices $\mathbf{n}$, output price variances $\mathbf{Vp}$ and the area vector $\mathbf{B}$

$$\mathbf{q}^* = \mathbf{q}\left(\mathbf{n}, \mathbf{p}, \mathbf{Vp}, \mathbf{B}\right) \tag{9}$$

It can be shown that conditional input demand equations $\left(\mathbf{x}\right)$ are obtained by differentiating either $U^* \left(\mathbf{p}, \mathbf{n}, \mathbf{Vp}, \mathbf{B}\right)$ or $C\left(\mathbf{n}, \mathbf{q}^*, \mathbf{B}\right)$ with respect to input prices $\mathbf{n}$[10]

[10] See Coyle (1992).

$$\mathbf{x}(\mathbf{p},\mathbf{n},\mathbf{Vp},\mathbf{B}) = -\frac{\partial U^*(\mathbf{p},\mathbf{n},\mathbf{Vp},\mathbf{B})}{\partial \mathbf{n}} = \frac{\partial C(\mathbf{n},\mathbf{q}^*,\mathbf{B})}{\partial \mathbf{n}} = \mathbf{x}(\mathbf{n},\mathbf{q}^*,\mathbf{B}) \tag{10}$$

Area allocation equations are obtained in an analogous fashion to OudeLansink (1999), where total land area is fixed in the short run. These may be equivalently obtained by partially diferentiating either, the indirect utility function or the cost function

$$\frac{\partial U^*(\mathbf{p},\mathbf{n},\mathbf{Vp},\mathbf{B})}{\partial b_1} = \frac{\partial U^*(\mathbf{p},\mathbf{n},\mathbf{Vp},\mathbf{B})}{\partial b_2} \text{ subject to } \sum_{i=1}^{2} b_i = \mathbf{A}, \text{ or}$$
$$\frac{\partial C(\mathbf{n},\mathbf{q}^*,\mathbf{B})}{\partial b_1} = \frac{\partial C(\mathbf{n},\mathbf{q}^*,\mathbf{B})}{\partial b_2} \text{ subject to } \sum_{i=1}^{2} b_i = \mathbf{A} \tag{11}$$

The next section presents the functional form used to estimate the system outlined in (9), (10) and (11).

2.1 Symmetric Generalized McFadden Cost Function

The Diewert and Wales (1987) cost function is used as the relevant proxi for underlying producer technology. The form builds on work developed by McFadden and Lau and allows for the imposition of curvature properties with relative ease. Given the assumption of non-jointness in variable inputs, a separate cost function is specified and estimated for both wheat and barley. The cost function for wheat is given as

$$\begin{aligned} C(q_2,\mathbf{n},\mathbf{z}) &= h(\mathbf{n})\,q_2 + \sum_{i=1}^{2} s_{ii} n_i q_2 + \sum_{i=1}^{2} s_i n_i + \sum_{i=1}^{2} s_{it} n_i z_i q_2 \\ &+ s_{z_i}\left(\sum_{i=1}^{2} \epsilon_i n_i\right) z_i + s_{qq}\left(\sum_{i=1}^{2} \theta_i n_i\right) q_2^2 + s_{z_i z_i}\left(\sum_{i=1}^{2} \omega_i n_i\right) t_i^2 q_2 \end{aligned} \tag{12}$$

where $\mathbf{z}$ is a vector with z_1 = a time trend and $z_2 = b_2$ = wheat area (the second component in the $\mathbf{B}$ vector),[11] θ, ϵ and ω are vectors of parameter values pre-selected

[11] In the case of barley, $z_2 = b_1$ (the first component in the $\mathbf{B}$ vector).

by the researcher. The parameters s are the only ones estimated. The function $h(\mathbf{n})$ is defined as

$$h(\mathbf{n}) = \frac{1}{2}\left(\mathbf{n}^T\mathbf{L}\mathbf{n}\right)\left[v^T\mathbf{n}\right]^{-1} \tag{13}$$

where $\mathbf{L} = \mathbf{L}^T = [l_{ij}]$ is a 2 x 2 negative, semidefinite, symmetric matrix and $v^T = [v_1, v_2] > 0^T$ is a vector of non-negative constants, not all equal to zero and $\mathbf{s}$ is a matrix of parameters to be estimated. Under these set of restrictions, $h(\mathbf{n})$ can be shown to be globally concave. As terms involving $\mathbf{s}$ are linear in input prices, they do not appear in the Hessian matrix of C. Thus, $\nabla^2_{nn}C(\mathbf{n}, q, \mathbf{z}) = \nabla^2_{\mathbf{nn}}h(\mathbf{n})\,q$. Therefore, if the estimated $\mathbf{L}$ matrix is negative semidefinite, then the cost function C given by (13) is globally concave in input prices. Negative semi-definiteness can be imposed in various ways. In this instance, the Wiley, Schmidt and Bramble (1973) technique is adopted, with $\mathbf{L}$ being set equal to $-\mathbf{E}\mathbf{E}^T$ where $\mathbf{E}^T = [e_{ij}]$ is an upper triangular matrix. Consequently, the $\mathbf{L}$ matrix in (13) can be shown to be equal to[12]

$$\mathbf{L} = -\begin{bmatrix} e_{11}^2 & e_{11}e_{21} \\ e_{11}e_{21} & (e_{21}^2 + e_{22}^2) \end{bmatrix} = e_{11}^2 \begin{bmatrix} -1 & 1 \\ 1 & -1 \end{bmatrix}$$

where e_{11} is now the parameter to be estimated.

3 Empirical Model

Given the cost function (12), the first order equation, given by (7), may be written as[13]

[12] See Barnett and Zhou (2000) for details.

[13] The empirical derivations for the output supply, input demand and area allocation follow that in OudeLansink (1999), albeit using a different functional form. OudeLansink uses the normalised quadratic.

$$\Gamma M^{\Gamma-1}\Bigg[p_1 - h\left(\mathbf{n}\right) - \sum_{i=1}^{2} s_{ii}n_i - \sum_{i=1}^{2}\sum_{j=1}^{2} s_{iz_j}n_i z_j - 2\sum_{i=1}^{2} s_{q_1q_1}\theta_i n_i q_1 - \sum_{i=1}^{2}\sum_{j=1}^{2} s_{z_jz_j}\omega_i n_i z_j^2\Bigg] - \Upsilon S^{\Upsilon-1}\Bigg[\sum_{j=1}^{2} q_j V p_{1j}\Bigg] = 0 \quad \text{for } i = 1\ldots 3 \tag{14}$$

Re-expressing this in closed-form solution for q_1 results in

$$q_1 = \frac{1}{\Gamma M^{\Gamma-1} 2\sum_{i=1}^{2} s_{q_1q_1}\theta_i n_i + \Upsilon S^{\Upsilon-1} V p_{11}} \Gamma M^{\Gamma-1}\Bigg[p_1 - h\left(\mathbf{n}\right) - \sum_{i=1}^{2} s_{ii}n_i - \sum_{i=1}^{2}\sum_{j=1}^{2} s_{iz_j}n_i z_j - 2\sum_{i=1}^{2} s_{q_1q_1}\theta_i n_i q_1 - \sum_{i=1}^{2}\sum_{j=1}^{2} s_{z_jz_j}\omega_i n_i z_j^2\Bigg] - \Upsilon S^{\Upsilon-1} q_2 V p_{12} \tag{15}$$

The input demand equations are obtained by applying Shephard's lemma to (12)

$$\begin{aligned} x_i &= \frac{1}{2}\Bigg\{2\left(\mathbf{Ln}\right)\left[\upsilon^T\mathbf{n}\right]^{-1} - \left[\upsilon^T\mathbf{n}\right]^{-2}\upsilon\left(\mathbf{n}^T\mathbf{Ln}\right)\Bigg\}q + s_{ii}q_1 + s_i + \sum_{j=1}^{2} s_{iz_j}z_j q_1 \\ &\quad + \sum_{j=1}^{2} s_{z_j}\epsilon_i z_j + s_{q_1q_1}\theta_i q_1^2 + \sum_{j=1}^{2} s_{z_jz_j}\omega_i z_j^2 q_1 \quad \forall \;\; i \end{aligned} \tag{16}$$

The area allocation equations corresponding to (11) are obtained by taking the partial derivatives of the cost function with respect to the elements of the $\mathbf{B}$ matrix for the wheat and barley cost functions. This results in barley output and parameters of the barley cost function appearing in the area allocation for wheat and vice versa. These parameters and variables are denoted by $\{*\}$ in the following area allocation for wheat

$$b_2 = \frac{1}{2\sum_{i=1}^{2} s_{z_2z_2}\omega_i n_i q_1 + 2\sum_{i=1}^{2} s^*_{z_2z_2}\omega^*_i n_i q^*_1}\Bigg[\left(\sum_{i=1}^{2} s^*_{iz_2} n_i q^*_1 - \sum_{i=1}^{2} s_{iz_2} n_i q_1\right) + \left(\sum_{i=1}^{2} s^*_{z_2}\epsilon^*_i n_i - \sum_{i=1}^{2} s_{z_2}\epsilon_i n_i\right) + 2\sum_{i=1}^{2} s^*_{z_2z_2}\omega^*_i n_i q^*_1 \mathbf{A}\Bigg] \tag{17}$$

The next section discusses the data used in the study and in particular, the price expectations model proposed by Roche and McQuinn (2003) is briefly discussed.

3.1 Data and Price Expectations

The data used for the analysis is obtained from the National Farm Survey (NFS) conducted by Teagasc[14].[15] An unbalanced panel, from 1984 to 1998, comprising data on specialist cereal producers who simultaneously plant both barley and wheat, was compiled.[16] The production system given by (15), (16) and (17) is estimated for wheat and barley. The two variable input items used in the analysis are nitrogenous fertiliser and 'other' inputs. Note, that the 'other' inputs item contains all other variable inputs. The prices for these input items are considered to be non-stochastic and known to producers in advance of the input application decision. In addition to output prices, input prices are also assumed to be constant across space and variable only through time. The prices used for these input items are national aggregate price indices and are from the Irish Central Statistics Office (CSO). In total, seven equations are estimated.[17] All estimations are conducted using the nonlinear three-stage systems estimator in SAS/ETS.[18]

One of the key features of the model employed in this paper is the price expectations model that is used to generate the expected mean, variance and covariance of Irish grain prices. The model used is that presented in Roche and McQuinn (2003).

[14]The Irish Agriculture and Food Development Authority.

[15]For more on the NFS see Heavey, Roche and Burke (1998).

[16]The sample totaled 913 observations.

[17]Two output supply equations, four input demand equations and one area allocation equation (wheat).

[18]Programs are available from the authors upon request.

Most studies of price uncertainty within a dual production setting use the Chavas and Holt (1990) model of expected mean, variance and covariances.[19] 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 (1990) approach using standard forecast statistics (the mean squared error (MSE) and the mean absolute deviation (MAD)) as well as the recently developed test of superior predictive ability (SPA) by Hansen (2001).[20] In all cases, the Roche and McQuinn (2003) model outpreforms that of Chavas and Holt (1990). The model is summarised in Equation (3) of Roche and McQuinn (2003) and is presented in its linear autoregressive distributed lag form as[21]

$$\begin{aligned} p_t^{wir} &= f\left(1, t, p_{t-1}^{wir}, p_{t-2}^{wir}, p_{t-1}^{wuk}, p_{t-2}^{wuk}, e_{t-1}, e_{t-2}, u_t^w\right) \\ p_t^{bir} &= f\left(1, t, p_{t-1}^{bir}, p_{t-2}^{bir}, p_{t-1}^{buk}, p_{t-2}^{buk}, e_{t-1}, e_{t-2}, u_t^b\right) \end{aligned} \tag{18}$$

$$\begin{bmatrix} u_t^w \\ u_t^b \end{bmatrix} = [u_t] \sim MN\left(0, \mathbf{H}_t\right)$$

The series p^{wir} is the price of MCA[22] adjusted Irish feed wheat, p^{wuk} is the price of British feed wheat, e is the punt/sterling exchange rate, p^{bir} is the price of MCA adjusted Irish feed barley, p^{buk} is the price of British feed barley, t is a trend term and the u^i are stochastic error terms. The covariance matrix of Irish grain prices, $\mathbf{H}_t$, is estimated following Baba, Engle, Kraft and Kroner (1991) and Flavin and Wickens (2001) using the following MVARCH(1,1) model

$$\mathbf{H}_t = \mathbf{A}'\mathbf{A} + \mathbf{B}'\left(u_{t-1}u_{t-1}\right)\mathbf{B} \tag{19}$$

The use of the results from the MVARCH model has a number of attractive features. First, the error correction model captures the dynamic nature of price

[19] For a more detailed discussion on this point see McQuinn (2003).

[20] The SPA tests for the best standardized forecasting performance relative to a benchmark model.

[21] The model proposed in Equation (3) of Roche and McQuinn (2003) is for the *growth rates* in prices, whereas the model expressed in (18) is for price *levels.*

[22] Monetary Compensation Amount.

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, Chou and Kroner (1992) for example). The model is estimated on a rolling basis and forecasts are generated for the sample period 1984-1998.

The parameters outlined in (12) are given the following values: $\omega = 1, \epsilon = 1$ and $\theta = 1$. For the parameter υ, the same approach is taken as that outlined in the footnote to p.54 of Diewert and Wales (1987). This implies that υ_i is measured in the units of input i. Therefore, υ_i is chosen to be equal to the average amount of input i ($\overline{x}_i$) as this ensures invariant elasticity estimates.

4 Estimation Results

As the maximum number of parameters to be estimated in any one equation outnumbered the total number of exogenous variables, instrumental variables were required. The instruments chosen were the price of phosphates fertiliser and the price of crop protection. The total list of parameter estimates are contained in Table 1. (Insert Table 1 here.)

In total, 70 per cent of parameters are significant at the 5 per cent level, while 60 per cent are significant at the 1 per cent. This compares favourably with results elsewhere (OudeLansink (1999) had 41 per cent of parameter estimates at the 5 per cent level in his unrestricted model and 28 per cent significant in a restricted model). The risk coefficients, as well as parameter tests are further summarised in Table 2. (Insert Table 2 here)

The results presented in Table 2 may be compared with Table II of Saha (1997). Irish cereal producers are risk averse, as both α and Υ are clearly positive. Further tests suggest that the MSU utility function specification constitutes a better representation of producers' risk attitudes than the more restrictive specifications of CARA used in other studies.[23] The associated risk premium[24] with this utility function specification is €3,767 - 12 per cent of average annual income, suggesting Irish producers would be prepared to accept a sure income level somewhat below

[23]Such as Coyle (1992) and OudeLansink (1999).

[24]Calculated as E(π) - $\hat{\pi}$, where $\hat{\pi}$ is the certainty equivalent of income given by $M^{\Gamma} - S^{\Upsilon}$.

its expected level in order to forego price uncertainty.

The χ^2 test demonstrates that the linear U(M,S) model nested within the non-linear specification can be rejected as an adequate representation of producer's attitudes at the 1 per cent level. Similarly, the first t-test of $\Gamma = 1$ is clearly rejected at the same significance level, suggesting that producers in the sample do not display constant absolute risk aversion. This would generally conform to *a priori* expectations about economic agents i.e. one would expect less affluent producers to be more risk averse than relatively more affluent producers. Therefore, the rejection of the null hypothesis of CARA is not altogether surprising. The final t-test rejects the null hypothesis that producers possess constant relative risk aversion. Under CRRA, producers would exhibit a constant level of risk aversion to the same proportional risk i.e. as a proportion of income.

From the results presented in Table 2, it appears that Irish producers exhibit DARA and DRRA. This initial result of DARA is in line with *a priori* expectations i.e. as producers experience increased income, one expects them to become less risk averse. Table 1 of Saha, Shumway and Talpaz (1994) illustrates that in many international studies CARA has been rejected in favour of DARA. Thus, the result of DARA has considerable support in the literature.

The finding of DRRA, that is, a declining level of risk aversion to the same proportional risk, is also noteworthy, as Saha, Shumway and Talpaz (1994) note, studies on relative risk aversion frequently yield ambiguous results. In particular, most studies reveal either CRRA or DRRA. The importance of the flexibility of the utility function adopted is underlined by the finding of DRRA. As noted by Saha (1997), "most prior studies have not investigated whether the nature of relative risk aversion preference differs according to income levels".[25]

4.1 Direct Payments and Producer Risk Premiums

The risk premium that is estimated may be interpreted as the cost of producer price uncertainty. It is the amount of uncertain income the producer will forego in order to have a certain income level $\hat{\pi}$. Assuming that the level of direct payment just compensates the producer for the subsequent price decline under MacSharry,

[25]Elasticity results are available from the authors upon request.

then risk-averse producers are implicitly in receipt of a welfare transfer - an uncertain amount of income has been replaced by a vitually certain direct payment D. Producers could receive a direct payment less in total than the income decline experienced due to the price drop and still remain in a welfare neutral state.

Table 3 (insert Table 3 here) presents a comparison of the average estimated risk premium and direct payment received per farmer for each year between 1993 and 1998. To standardise the comparison, all values are on a € per hectare basis. On average, Irish cereal producers received direct payments worth over €124 per hectare between 1993 and 1998 (column 5). For the same period, the average risk premium on a per hectare basis was €92 (column 4). Accordingly, column 6 reports an adjusted direct payment per hectare - D*. This amount scales the actual direct payment by the ratio (1-(ρ/E(π))), thus generating the level of direct payment required to leave producers in a welfare-neutral state, given their risk attitudes. Comparing columns 5 and 6 suggests that, on average, Irish cereal producers received welfare transfers of approximately €16 per hectare throughout the 1990s relative to their pre-MacSharry position.

5 Concluding Comments

This paper has estimated risk attitudes and premiums, in the presence of output price uncertainty, for Irish cereal producers using the flexible MSU framework. Expected Irish prices are generated using the recently developed Roche and McQuinn (2003) MVARCH model of Irish grain prices. From the empirical results, it may be observed that Irish producers are risk averse and display both DARA and DRRA.

Given these risk attitudes, Irish cereal producers could have received direct payments worth approximately 88 per cent of the value of the actual payments over the 1993-1998 time period and still have been in a welfare-neutral position vis-à-vis CAP reform.

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 'complete' risk free payments has obvious implications for the provision of market provided risk management tools such as derivatives or insurance markets. Studies by Meuwissen, Huirne and Hardaker (1999a) and Meuwissen, Huirne and Hardaker (1999b) have emphasised the removal of institutionalised income support as a pre-requisite for the successful adoption of these products.

Finally, the welfare impacts of reform are worth highlighting 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, a welfare-neutral result can be achieved by offering direct payments which are less in value than the expected decline in incomes brought about by further reductions in intervention prices.

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Table 1: MSU-type Cost Function Estimates for
Irish Wheat and Barley Producers under Price Uncertainty

	Wheat		Barley	
Parameter	Estimate	T-Value	Estimate	T-Value
e_{11}	11.57	4.67	31.23	5.91
s_{11}	51.52	7.23	5.51	1.70
s_{22}	0.87	0.11	-43.41	-12.18
s_1	-2446.23	-2.87	876.13	2.54
s_2	-2926.11	-3.47	1052.23	3.45
s_{1z_1}	1.26	1.30	0.65	1.33
s_{2z_1}	2.18	2.19	0.77	1.52
s_{1z_2}	0.02	1.39	-0.03	-4.39
s_{2z_2}	-0.005	-2.05	-0.031	-3.59
s_{z_1}	-120.58	-1.32	-115.72	-2.39
s_{z_2}	69.31	16.07	71.19	16.11
s_{qq}	-0.03	-4.87	0.042	12.37
$s_{z_1z_1}$	-0.04	-0.65	0.012	0.59
$s_{z_2z_2}$	0.000006	10.69	-0.0001	-10.76
Γ	8.78	50.77	8.78	50.77
Υ	5.42	44.23	5.42	44.23

Sample size = 913 observations.
Note: Diagnostic tests for heteroscedasticity and autocorrelation were conducted using the Baltagi (2001) LM test. This tests for individual and time effects in the error term and is is asymptotically distributed as χ^2_2. Residuals from (15) * 2, (16) * 2 were all tested under the null of homoscedastic errors. At the one per cent level, the null was not rejected in any of the seven cases.

Table 2: Estimated Risk Attitudes
and Parameters for Irish Cereal Producers

Parameter/Test	Description	Estimate	Std. Error/P-Values
Γ		8.78	(0.000)
Υ		5.42	(0.000)
α*	$(\Gamma/\Upsilon)\, \mathrm{M}^{1-\Upsilon}\mathrm{S}^{\Gamma-1}$	13493.99	(0.000)
$\mathrm{H_O}$: $\Gamma = \Upsilon = 1$**	Linear U(M,S) model	4435.3	(0.000)
$\mathrm{H_O}$: $\Gamma = 1$***	CARA Attitudes	50.71	(0.000)
$\mathrm{H_O}$: $(\Upsilon - \Gamma) = 0$***	CRRA Attitudes	48.15	(0.000)

Note: * denotes evaluated at the sample mean. ** denotes Asymptotic $\chi^2(2)$ test statistic, p-value in parentheses. *** denotes Asymptotic t-test statistic, p-value in parentheses.

Table 3: Comparison of Irish Cereal Producers'
Average Risk Premia and Direct Payment Levels (€ per hectare)

Year	E(π)	$\hat{\pi}$	ρ	D	D*
1993	867.54	801.96	65.57	76.27	70.50
1994	973.04	910.70	62.34	106.77	99.93
1995	804.54	711.16	93.38	140.83	124.48
1996	854.35	775.40	78.96	140.83	127.82
1997	702.71	579.61	123.10	140.83	116.16
1998	668.68	543.02	126.66	140.83	114.37
Average	811.81	720.31	91.51	124.39	108.89

Note: E(π) = expected income, $\hat{\pi}$ = certainty equivalent of income = $M^{\Gamma} - S^{\Upsilon}$, ρ = risk premium, D = direct payment, D* = direct payment adjusted downwards given the magnitude of producers' risk premia i.e DP* = DP(1-(ρ/E(π))).