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DYNAMIC ECONOMETRIC MODELS OF MANITOBA CROP PRODUCTION AND HYPOTHETICAL PRODUCTION IMPACTS FOR CAIS

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Abstract

This study analyzes the impact of the Canadian Agriculture Income Stabilization (CAIS) program. The study begins with a specification of dynamic crop production that decomposes static short run crop acreage allocation decisions and dynamic crop yield affects. The modelling framework accommodates risk aversion, price uncertainty, and applies recent aggregation theory to aggregate weather data. Using this framework an analytical model of the impacts of CAIS on crop production is developed. Hypothetical impacts of are simulated using an aggregate Manitoba data set. The results show that CAIS has a substantial impact on the shadow prices of both inputs and outputs. These shadow price effects resulted in a 4 percent increase in long run wheat and barley yields and a 2 percent increase for canola. CAIS has a small impact on nominal wealth but the impacts depend on the properties of producers' risk preferences. With constant relative risk aversion there is a wealth effect which in turn affects production decisions.

1.0 INTRODUCTION

Decoupling government support from agricultural production has become an important policy issue at both national and international levels. As one of the key principles of agricultural policy reform adopted by OECD Ministers in 1987, policies deemed to have minor impact on production and trade would not lead to discipline. The complexities of the decoupling issue are well described in an OECD conceptual overview (Anton, OECD 2001), and this provides the analytical framework for on-going OECD research on decoupling.

The OECD overview notes that policy can affect production and trade by the following types of mechanisms: static effects under risk neutrality; static effects under risk aversion and uncertainty; and dynamic effects related to investment decisions. Previous empirical research on the effects of policy on production and trade has focused on static effects under risk neutrality, emphasizing (e.g.) changes in expected output prices. However under risk aversion and uncertainty, price neutral policies can influence production through changes in price or revenue uncertainty and changes in exogenous income, and these effects may be substantial (e.g. Hennessy). The effects of such policies on production and trade should be most important in a dynamic setting: uncertainty and hence risk aversion are more important over a long planning horizon than in a static model, and investment decisions have long lasting effects on production.

The first phase of OECD research stemming from the overview emphasized static models under risk aversion and uncertainty (e.g. Sckokai 2002). The second phase of the research emphasized dynamic models of investment under risk aversion and uncertainty. This included an extension of the Italian static model to dynamics (Sckokai 2005) and a dynamic econometric model of Manitoba crop investment (Coyle 2005). For example, the Manitoba study estimated a dynamic model of investment in machinery and equipment aggregated over all crops, since investment data is not available for individual crops.

However the drawback of this approach, focussing on aggregate investment, is that it does not provide crop specific estimates of dynamic impacts. Thus it cannot readily be compared with estimates of static impacts on crop specific acreage demands. In addition, dynamic impacts on investment are only one component (albeit a critical component) in dynamic changes in crop production (other components are impacts of investment on yield and changes in crop acreages).

This study adopts an alternative approach to specification of dynamic crop production models: crop production is decomposed into a static short run crop acreage allocation model (conditional on yields) and dynamic crop yield models. This approach permits the estimation of short run impacts of price and programs on output (primarily impacts on acreage demands) and also long run impacts on output (long run impacts on yields and acreages).

The model is dynamic (an autoregressive distributed lag model), accommodates risk aversion and price uncertainty, and applies recent aggregation theory to aggregate weather data. In contrast, previous econometric studies of price impacts on yields have been limited to static risk neutral models, and estimated impacts have often been statistically insignificant.

This applies this methodology to a more complex econometric model of Manitoba crop production. Econometric results are used to simulate possible crop production impacts of the Canadian Agricultural Income Stabilization (CAIS) program. Crop production impacts provide a first order approximation to trade distorting impacts on CAIS, assuming effects on Canadian consumption are negligible.

A simple analytical model of the impacts of CAIS on crop production is developed here. Assuming utility maximizing behaviour under constant relative risk aversion (CRRA), the analytical model is quite different from standard models. This is especially the case in its treatment of shadow prices and changes in relative wealth. Calculations suggest that impacts of CAIS on agricultural production and trade distortions may be significant, but important caveats must qualify these hypothetical results.

2.0 MANITOBA DATA

The empirical models depend upon the following data for Manitoba (obtained primarily from the Manitoba Agriculture Yearbook 2002): annual production and harvested acres for major crops (wheat, barley, canola, oats, rye, flax), market prices for crops, operating costs and capital costs per acre for different crops, fertilizer sales, value per acre of farmland and buildings. Farm input price indexes for Western Canada were obtained for hired labour, variable crop inputs, and machinery plus equipment (Manitoba Agriculture Yearbook, Statistics Canada).

Data on government payments, premiums etc. were obtained from the following agencies or their publications: Manitoba Crop Insurance Corporation (MCIC), Net Income Stabilization Account program (NISA), Agriculture Income Disaster Assistance (AIDA), Gross Revenue Insurance program (GRIP), Western Grains Stabilization Act program (WGSA), Western Grains Transportation Act program (WGTA). Data on other programs was obtained from Agriculture Economic Statistics (Statistics Canada 26-603).

An annual crop-growth weather index GRODEX (Dyer, Narayanan and Murray) is a proxy for weather, and is available for 1961- 87 for six Manitoba weather stations, and by a similar index constructed by the Canadian Wheat Board for 1981-95 for various weather stations (sixteen have complete data over this period). This is supplemented by data on total precipitation for May-July for twenty one Manitoba weather stations over 1982-2002.

Initial wealth is defined as the sum of value of capital stock in crop production (machinery and equipment plus land and buildings) minus related debts. A perusal of historical data (not broken down for crops and livestock) on value of machinery and equipment, value of land and buildings and outstanding farm debt indicates that the value of land and buildings greatly exceeds the other two series, and the other two series largely cancel out. For example, in 2000 the value of land and buildings was $9,782 million, value of machinery and equipment was $3,315 million, and total debt was $3,735 million (Manitoba Agriculture Yearbook, 2002). The value of land and buildings in crop agriculture is a proxy for initial wealth. This is calculated from total crop acres and the value per acre of farmland and buildings for Manitoba. Initial wealth also includes various government payments for crops that are not commodity specific.

Expected market prices for crops are proxied as a one year lag on prices, p_{t-1}. Net expected prices include lagged payments minus premiums for Crop Insurance and GRIP, and also transportation costs net of subsidies. Variances and covariances of market price are calculated as a simple three year weighted average of observed prediction errors:

$$\mathrm{var}(p_t) = 0.50(p_{t-1} - E_{t-2}\, p_{t-1})^2 + 0.33\,(p_{t-2} - E_{t-3}\, p_{t-2})^2 + 0.17\,(p_{t-3} - E_{t-4}\, p_{t-3})^2$$

$$\begin{aligned}\mathrm{cov}(p_{it}p_{jt}) = {} & 0.50\,(p_{it-1} - E_{t-2}\, p_{it-1})\,(p_{jt-1} - E_{t-2}\, p_{jt-1}) + 0.33\,(p_{it-2} - E_{t-3}\, p_{it-2})\,(p_{jt-2} - E_{t-3}\, p_{jt-2}) \\ & + 0.17\,(p_{it-3} - E_{t-4}\, p_{it-3})\,(p_{jt-3} - E_{t-4}\, p_{jt-3})\end{aligned} \qquad (1)$$

$$\text{where}: E_{t-1}p_t = p_{t-1}$$

This simple naive model of expectations has been used in many studies (e.g. Chavas and Holt 1990). Moreover such models have performed better than rational expectations models in studies of Western Canadian agriculture (Coyle 1992, 1999; Mbaga and Coyle) (also see Sulewski, Spriggs and Schoney). A covariance between market prices and various government payments is defined similarly:

$$\begin{aligned}\mathrm{cov}(p_{it}G_t) = {} & 0.50(p_{it-1} - E_{t-2}\, p_{it-1})(G_{t-1} - E_{t-2}G_{t-1}) + 0.33(p_{it-2} - E_{t-3}p_{it-2})(G_{t-2} - E_{t-3}G_{t-2}) \\ & + 0.17(p_{it-3} - E_{t-4}\, p_{it-3})(G_{t-3} - E_{t-4}\, G_{t-3})\end{aligned} \qquad (2)$$

where : $G \equiv$ government payments

This is designed to capture an insurance/income smoothing effect rather than a wealth effect of government payments. So this excludes, for example, deposits to NISA savings accounts, which are contributions to wealth rather than to current expenditures. It includes withdrawals from NISA accounts, which are used to smooth expenditures.

Note that we do not adopt a truncation approach to the effect of programs on price distributions used in studies such as Chavas and Holt. These authors assumed a naive model of expectations for market prices similar to (1), and then truncated the distribution in terms of one parameter (floor price for the commodity). However in our case the truncation process seems more complex since multiple programs imply multiple truncation parameters. Moreover it is not clear that producers adopting a naive expectations process for market prices would adopt or approximate such a sophisticated truncation process for effects of programs on price distributions. Instead we assume simply that producers adopt a naive expectations process regarding market prices and how market prices covary with government payments. Moreover by modeling these as separate processes we can test if they have different impacts on investment and production decisions. Thus we can test if government programs have a separate insurance effect on investment and production.

An annual time series variance for weather is constructed for each weather station as a similar three year weighted average of prediction errors (1) in weather station data. Then the time series variance data for each weather station are combined into a weighted average variance series for Manitoba (weather stations are weighted based on crop district wheat production in 1985). This weighted average time series weather variance is used as a proxy for weather uncertainty in Manitoba in crop acreage allocation model.

3.0 ADL CROP YIELD MODELS

3.1 Introduction

Yield equations are specified here as conditional on prices but not on capital stock, i.e. capital is not included as an explanatory variable in yield equations since capital or investment data is not crop specific. Thus estimated price and program effects implicitly include indirect impacts through induced changes in capital. In other words, estimated price effects on yields include direct effects of prices on yields (at constant capital stocks) plus indirect effects on yields associated with effects of prices on capital stock.

Crop yields are modelled here as a function of expected price and price variance, initial wealth and weather. Assuming disjoint technologies (and similar capital/acreage ratios across crops), crop yield can be modelled independently of cross price effects.

An autoregressive distributed lag (ADL) crop yield model is specified for this study as follows:

$$yld_t = \beta_0 + \delta\, yld_{\tau-1} + \sum_{j=0....n_1} \beta_{1\iota}\, Ep_{t-i}/w_t + \sum_{j=0....n_2} \beta_{2\iota}\, Vp_{t-i}/(w_t)^2$$
$$+ \sum_{j=0....n_3} \beta_{3\iota}\, W^0/w_t + \gamma\, t + \sum_{j=1,2,3} \theta_j\, E\omega_{jt} + \sum_{j=1,2,3} V\omega_{jt} + e_t$$

where: $yld \equiv$ crop yield

$Ep \equiv$ expected crop price (3)

$Vp \equiv$ variance of crop output price

$W^0 \equiv$ initial wealth

$w \equiv$ crop input price index

$t \equiv$ time trend

The normalization (*Ep/w, Vp/w*^{2}*, W*^{0}*/w*) is implied by constant relative risk aversion. Here *Eω* and *Vω* are the cross section mean and variance of weather station data for three periods j: GRODEX weather index (1961-1980), CWB weather index (1981-1995), and May-July precipitation (1996-2002) (weather stations are weighted based on crop district wheat production). The choice of an ADL model and the specification of weather variables in (3) is explained in the accompanying paper (Coyle 2006).

3.2 Econometric Results for Manitoba

The unit root hypothesis was strongly rejected for yield of each crop, using standard augmented Dickey-Fuller and Phillips-Perron tests. Since *I(1)* and *I(0)* explanatory variables generally imply an *I(1)* dependent variable, this suggests that all variables are *I(0)*. Consequently the regression model is estimated in levels rather than as first differences. The model is specified as log-linear, so all coefficients can be interpreted as elasticities.

Ep includes crop specific government payments and charges (primarily related to Crop Insurance and GRIP) and an index of net transportation costs. A crop input price index is the numeraire *w*, and other input prices (hired farm labour wage, a capital price index) were statistically insignificant.

Here we discuss results for models of wheat yield, since this is the major crop. OLS estimates of a risk neutral ADL model with a four period lag in *Ep* are presented in Table A-1 of the Appendix on ADL Crop Yield Models. The sum of lag coefficients for *Ep* is +0.629 with a t-ratio of 2.68 under polynomial restrictions (Table A-1 column a). The Durbin-h statistic indicates zero autocorrelation, and the polynomial restrictions are not rejected. R^2 is .809. The statistically significant negative coefficient on lagged yield may be due to model misspecification. Results are almost identical dropping the polynomial restrictions (Table A-1 column B)

The risk averse ADL model (3) was initially specified with 8 period lags for Ep, Vp, W^0 (and fourth degree polynomials to preserve degrees of freedom). Significance of lags was tested for related models without polynomial restrictions. Results indicated these models could be reduced to four period lags, and that lagged yield was statistically insignificant.

Table A-2 reports results for the selected model under second degree polynomial restrictions. The sum of lag coefficients and corresponding t-ratios are +0.637 (2.27), -0.115 (2.45), +0.602 (3.58) for Ep, Vp, W^0 respectively for OLS (Table A-2 column A). Four of the six weather variables are statistically significant, and these have the anticipated signs. R^2 is 0.876 and the polynomial restrictions are accepted, but there is autocorrelation. Table A-2 column B reports results for a grid search maximum likelihood GLS procedure (Beach and MacKinnon). Sum of lag coefficients are similar to the OLS case: +0.572 (3.16), -0.125 (3.70), +0.567 (4.36) for Ep, Vp, W^0.

Since this and the accompanying paper are the first econometric studies of dynamic price impacts on crop yields, these results cannot be assessed relative to other studies. Nevertheless, the estimated elasticity of wheat yield with respect to initial wealth is larger than anticipated. This estimate will be critical in calculating potential impacts of CAIS on production.

The effects of omitting weather variables or economic variables from this model are illustrated in Table A-3. Omitting weather variables (and the statistically insignificant time trend), *Ep* and *Vp* are no longer statistically significant (Table A-3). R^2 for OLS estimation is 0.8763, 0.6847 and 0.6978 for the full model (A-2 column A), omission of weather (A-3 column A) and omission of economic variables (Table A-3 column B). (The corresponding adjusted R^2 is 0.7971, 0.5960 and 0.6356.) Thus the economic variables explain as much of the variation in annual wheat crop yield as does weather (and time trend).

The covariance between government payments and wheat market prices was also included in the model. This is denoted as $Cov(Gp_1)$ and is normalized by w^2. First an 8 period lag in $Cov(Gp_1)$ was added to model (3), using 4 period lags in Ep, Vp, W^0. Results are presented in Table A-4 using a fourth degree polynomial for the 8 period lag and a second degree polynomial for the 4 period lags. For OLS sum of lag coefficients and t-ratios are +0.489 (1.77), -0.065 (1.32), +0.644 (3.53), -0.0012 (1.90) for Ep, Vp, V^0, $Cov(Gp_1)$ (Table A-4 Column A). R^2 is 0.913 and the polynomial restrictions are accepted, but there is autocorrelation. For GLS, sum of lag coefficients and t-ratios are +0.415 (2.95), -0.068 (2.38), +0.463 (3.51) for Ep, Vp, W^0 and -0.0009 (2.89) for $Cov(Gp_1)$. Thus the insurance effect $Cov(Gp_1)$ has a statistically significant impact on wheat yield as anticipated, although the estimated elasticity is small here.

Rather than adopting a full general to specific testing for a preferred model given *Ep, Vp,* W^0*, Cov(*Gp_1*)*, a yield model was specified given the same lag lengths and polynomial restrictions for *Ep, Vp,* W^0 as was selected for the ADL investment equation: 2 period lag for *Ep*, 1 period lag for *Vp*, and a 4 period lag for W^0. Then testing indicated a 3 period lag for *Cov(*Gp_1*)*, in contrast to a lag for periods (3,4) in the investment model. Results are reported in Table A-5. OLS estimates and t-ratios of sum of lags are +0.531 (2.14), -0.045 (1.54), +0.432 (2.92), -0.050 (1.80) for *Ep, Vp,* W^0*, Cov(*Gp_1*)* (Table A-5 Column A). R^2 is .859 and polynomial restrictions are not rejected at the 0.05 level, but there is autocorrelation. GLS estimates correcting for autocorrelation are reported in Table A-5. GLS estimates of sum of lags are +0.383 (2.79), -0.043 (2.66), +0.263 (2.95), -0.036 (2.43) for *Ep, Vp,* W^0*, Cov(*Gp_1*).* Note that the elasticity of the insurance effect is larger in magnitude than in Table A-4.

Appendix A on ADL crop yields also reports econometric results for barley, canola and oats. For barley, results for risk neutral models and simple risk averse models (with *Ep, Vp,* W^0) were similar to wheat (Tables A-6 and A-7), but *Cov(*Gp_2*)* was not statistically significant. For canola, a period 3 lag on *Ep* and a period 4 lag on *Vp* were statistically significant, instead of distributed lags (Table A-8). W^0 and *Cov(*Gp_3*)* were statistically insignificant. For oats, only a period 9 lag on *Ep* was statistically significant (Tables A-9 and A-10).

We cannot have much confidence in the estimated elasticities of yield with respect to wealth, which will be critical in simulating impacts of CAIS. There are no comparable studies, since other studies of price effects on yields have estimated static risk neutral models (and generally obtained insignificant results). A dynamic ADL model of Manitoba investment in crop machinery and equipment under risk aversion led to similar magnitudes for estimated elasticity of investment with respect to wealth (Coyle 2005).

Nevertheless we suspect that this study overestimates elasticity of wheat and barley yields with respect to wealth. An elasticity of +0.20 lies within the 95 % confidence interval about the reported estimates of +0.4322 and +0.4395 for wheat and barley yields (Table A-5 column A and A-7 column A). This may well be more realistic than the reported estimates.

We also briefly considered ADL yield models conditional on capital stocks. The difficulty is that investment/capital data is not crop specific, unlike yields. Yield models for individual crops can be estimated conditional on total capital per acre, but this assumes that capital per acre is approximately equal across crops. This is presumably an unrealistic assumption, although similar capital requirements per acre are often assumed in MDA budget forecasts for different crops.

An alternative approach is to specify a yield model, using a Tornqvist index1 for yields, as conditional on capital per acre. Here yield is defined as a Tornqvist index of crop outputs relative to crop acres. Although this approach does not impose identical capital requirements per acre over crops, there is a loss of information in aggregating outputs over crops. Correlations between the Tornqvist index for yields and individual crop yields are relatively high (+0.919, +0.954, +0.925 for wheat, barley and oats, respectively). Nevertheless, econometric results for such an ADL model are weaker than (e.g.) an ADL wheat yield model. This presumably reflects errors in aggregation of outputs.

4.0 CROP ACREAGE ALLOCATION MODEL

4.1 Introduction

Our econometric results for ADL yield models suggest that crop yields are essentially independent of price expectations within the current year. Indeed this assumption underlies most empirical studies of crop acreage demands.

Here we focus on the allocation of total crop acres among individual crops, i.e. we model crop acreage decisions conditional on total crop acres. This necessarily implies that crop acreage demand depends on cross prices, but this is likely to be the case even if we do not adopt such a specification, since total cropland is largely quasi-fixed. Moreover there are two major advantages to such a specification: adding-up restrictions lead to more precise estimates (Coyle 1993), and the model presumably is largely static rather than dynamic (predetermined yields and total crop acres reflect dynamic processes). In addition by specifying a static model, we can more readily interpret measured impacts of *Ep* and *Vp* as a response to specified expectations, i.e. we do not have the complex identification problem of disentangling lags in expectations and lags in response as in dynamic ADL models.

A four crop acreage allocation model is considered: wheat, barley, canola and other crops (oats, rye, flax). Crop acreage allocation models generally treat price uncertainty in an *ad hoc* manner in the case of multiple crops and ignore yield uncertainty in the case of aggregate (macro) data. Since these are serious misspecifications, we attempt to deal with them here. Construction of a time series representative farm level weather variance VW is discussed in the accompanying paper (Coyle 2006). Price uncertainty and predetermined yields lead to a relatively simple Tornqvist type index of aggregate variance of revenues per acre VR_4 for other crops and revenue per acre covariances $VR_{14}, VR_{24}, VR_{34}$ between the

1 Tornqvist index is a discrete approximation to a continuous Divisia index which is a weighted sum of the growth rates of the various components The weights are shares of total value. With a Tornqvist index the growth rates are defined as the difference in natural logarithms of successive observations of the components and the weights are equal to the mean of the factor shares of the components in the corresponding pair of years.

three commodities and other crops. These are included in the acreage demand model along with revenue per acre variances and covariances between the three commodities, constructed similarly to (1).

4.2 Econometric Results for Manitoba

The crop acreage allocation model is specified most generally as:

$$
\begin{aligned}
sz_{it} = {} & \beta_{ias} + \sum_{j=1,2,3} \beta_{ij}\, Er_{jt}/Er_{4t} + \beta_{i4}\, Er_{4t} + \gamma_{i1}\, Vr_{1t}/(Er_{4t})^2 + \gamma_{i2}\, Vr_{2t}/(Er_{4t})^2 \\
& + \gamma_{i3}\, Vr_{3t}/(Er_{4t})^2 + \gamma_{i12}\, Vr_{12t}/(Er_{4t})^2 + \gamma_{i13}\, Vr_{13t}/(Er_{4t})^2 + \gamma_{i23}\, Vr_{23t}/(Er_{4t})^2 \\
& + \gamma_{i4}\, VR_{4t}/(Er_{4t})^2 + \gamma_{i14}\, VR_{14t}/(Er_{4t})^2 + \gamma_{i24}\, VR_{24t}/(Er_{4t})^2 + \gamma_{i34}\, VR_{34t}/(Er_{4t})^2 \\
& + \theta_{i0}\, W_{0t}/Er_{4t} + \theta_{i1}\, w_{1t}/Er_{4t} + \theta_{i2}\, w_{2t}/Er_{4t} + + \theta_{i3}\, Z_{t-1} + \theta_{i4}\, (Z_t - Z_{t-1}) \\
& + \theta_{i5}\, d_t + \theta_{i6}\, t + {}_{i1}\, VW_{1t} + {}_{i2}\, VW_{2t} + {}_{i3}\, VW_{3t} + e_t \qquad i = 1,.,4
\end{aligned}
\tag{4}
$$

The dependent variables are crop acreage shares sz_i $(= z_i/Z)$. Here Er_i denotes expected revenues per acre $(r = p\ yld)$ for wheat, barley and canola, r_4 is a Tornqvist revenue per acre index for other crops, Vr_{ij} is variance or covariance in revenues per acre for wheat, barley and canola, VR_4 is the aggregate variance in revenues per acre index for other crops, VR_{i4} is the aggregate covariance index for crop *i* (wheat, barley, canola) with other crops 4. The normalization of Vr by $(Er_4)^2$ corresponds to constant relative risk aversion (CRRA). Under CRRA, the separate term Er_4 should drop out of the model. W_0 is initial wealth, w_1 and w_2 are farm price indexes for variable crop inputs and hired labour, Z is total crop acreage, d is a dummy variable for the 1970 LIFT program, and t is a time trend. Change in total crop acreage Z_t-Z_{t-1} is included in the model since acreage demands may depend on lags in adjustment of the overall crop rotation (Coyle 1993). VW_i is the time series representative farm level weather variance constructed from GRODEX weather station index data (1961-1984), CWB weather station index data (1985- 1995), and annual May-July weather station precipitation data (1996-2002), respectively. Management data on operating costs per acre for different crops (Manitoba Agricultural Yearbook) was also considered but did not improve the empirical model. All price variables (except for covariances, which are negative as well as positive) are specified as logarithms. Under risk neutrality the model reduces to:

$$
\begin{aligned}
sz_{it} = {} & \beta_i + \sum_{j=1,2,3} \beta_{ij}\, Er_{jt}/Er_4 + \beta_{i4}\, Er_4 + \theta_{i1}\, w_{1t}/Er_4 + \theta_{i2}\, w_{2t}/Er_4 \\
& + \theta_{i3}\, Z_{t-1} + \theta_{i4} (Z_t - Z_{t-1}) + \theta_{i5}\, d_t + \theta_{i6}\, t + e_t \qquad i = 1,.,4\,.
\end{aligned}
\tag{5}
$$

A standard approach is to estimate crop share equations as a system of seemingly unrelated regressions (SUR). However we emphasize single equation methods of estimation for share equations. The primary reason is that, in system methods, estimates of one equation are generally biased and inconsistent under specification errors in any other equation (except for identical regressors, which is unlikely in a final model omitting insignificant variables). So, single equation methods are more robust.

Unit root hypotheses were tested using standard augmented Dickey-Fuller and Phillips-Perron tests, and an alternative test due to Kwiatkowski et.al. Results were ambiguous for the key variables crop shares. So it was decided to consider Durbin-Watson (DW) statistics from OLS estimates of equations in levels. DW statistics are often recommended as an additional indicator of unit roots. Indeed in simple unit root models the DW statistic converges asymptotically to 0 (Phillips 1986). OLS estimates in levels generally indicated DW statistics of approximately 2 for wheat and canola share equations, but lower DW statistics for barley equations. Consequently it was concluded that wheat and canola share equations can be estimated in levels, but as a precaution barley share equations should be estimated with all variables in first differences.

Single equation OLS estimates for wheat acreage share equations are presented in Table B-1 of the Appendix B on crop acreage demands. OLS estimates of the risk averse model (4), omitting insignificant variables of no direct interest here, are presented in Table B-2 column B. Coefficient estimates and t-ratios are +0.105 (2.04), -0.016 (2.35), and -0.158 (3.65) for Er_1, Vr_1 and W_0, respectively, and + 0.016 (2.67) for VW_2. Opposite coefficient signs for W_0 and VW_2 are expected, assuming decreasing absolute risk aversion (DARA). Then an increase in W_0 leads to a reduction in risk aversion, whereas an increase in VW_2 implies an increase in risk. The negative coefficient of the LIFT dummy in the wheat equation is expected (payments were made for every 1969 wheat acre taken out of production in 1970). The Durbin-Watson statistic 1.83 indicates low autocorrelation. R^2 is 0.8677.

The own price elasticity for wheat is similar to results in an earlier study of Manitoba under risk neutrality (Coyle 1993). The elasticity of response with respect to wealth is similar to results for winter wheat (-0.448) in a multi-crop model for Ontario (von Massow and Weersink), but the own price (expected revenue) elasticity of response is much larger than here (+0.911). Given differences in relative importance of wheat in the two provinces, perhaps these results suggest that the elasticity of response in wealth is overestimated here for Manitoba.

OLS estimates of the corresponding risk neutral equation are presented in Table B-1 column B. The DW statistic falls to 0.95 and the R^2 falls to 0.6513. OLS estimates of the corresponding risk averse model, but omitting weather uncertainty, are presented in Table B-1column C. R^2 falls to 0.8170 (adjusted R^2 falls from .8063 to .7579) and DW statistic is 1.79 .

Single equation estimates for barley acreage share are presented in Table B-2. All variables are first differenced. The final model deleting various insignificant coefficients is presented in Table B-2 column A. The estimated coefficient of Er_2 is +0.102 (2.58), but Vr_2 and W_0 are insignificant at +0.0001 (0.02) and +0.0067 (0.30). R^2 is .380 and the DW statistic is 1.45. When this equation is estimated in levels, R^2 is .548 and DW statistic is 1.12. A grid search maximum likelihood procedure is used to estimate first order autocorrelation, and GLS is applied (Table B-2: column B). However coefficients of Vr_2 and W_0 remain statistically insignificant. OLS estimates of the corresponding model under risk neutrality and omitting weather uncertainty are also presented in Table B-2 column C.

Single equation OLS estimates of canola acreage share equations are presented in Table B-3. The final model deleting various insignificant coefficients is presented in Table B-3 column A. OLS coefficient estimates are +0.099 (6.67), -0.0052 (2.43), -0.018 (0.91) and -0.0056 (2.49) for Er_3, Vr_3, W_0 and VW_2, respectively. R^2 is .9848 and DW statistic is 2.02. For the corresponding risk neutral model, R^2 is .9689 and DW statistic is 1.76 (Table 3B). For the corresponding model omitting weather uncertainty, R^2 is 0.9803 and DW statistic is 2.05.

Then the three acreage share equations for wheat, barley and canola are estimated as a system by SUR. The wheat and canola equations are specified as levels, but the barley equation is specified in first differences for all variables. SUR estimates for the above final models are presented in Table B-4. Any contemporaneous covariance in disturbances for this specification of the model is weak, since the hypothesis of a diagonal covariance matrix is not rejected at the .01 level using the Breusch-Pagan LM test. Coefficient estimates are similar to the above single equation results. SUR estimates of the corresponding risk neutral model and the model omitting weather uncertainty are presented in Tables B-5 and B-6, respectively. Reciprocity (symmetry) conditions implied by static competitive profit maximization and predetermined yields (Coyle 1993) are rejected.

CRRA implies that the separate term Er_4 can be dropped from each equation. The corresponding joint hypothesis is $\beta_{i4} = 0 \; i=1,2,3$. This hypothesis is not rejected at the .05 level for the SUR model. Thus the hypothesis of CRRA is not rejected. This has important implications for analyzing the impacts of CAIS on crop production.

5.0 SIMULATION OF PROGRAM IMPACTS FOR MANITOBA

5.1 Introduction

Program impacts on output are simulated from estimates of the econometric models above. These models estimate impacts of expected prices, price variances, wealth and insurance effects on investment and output. Here we calculate effects of programs on these variables. A critical assumption here is that similar expectation and response processes apply to these program effects as in the aggregate. Then we combine these assumed program impacts on variables with econometric results in order to simulate program impacts on output. Crops 1, 2, 3 are wheat barley and canola respectively.

5.2 Parameter Assumptions

Long run and short run impacts on yields are estimated from ADL yield equations. However an ADL model is a reduced form dynamic model that does not identify separate lags in response from lags in expectations. This can be a serious problem in estimating response along a dynamic path, but it does not seem so serious in estimating long run response, where it is not so important to identify the source of lags. In inferring short-run responses from our estimates, we assume that lags in expectations follow our simple naive model, so that short run lags in response are identified.

First consider wheat yields. The relevant long-run elasticities (sum of lag coefficients) from Table A- 5 columns A and B in Appendix A are as follows:

Table 1 Long Run Elasticity of Wheat Yield

	Table A-5 Column A	Table A-5 Column B
Expected Output Price (Ep_1)	+0.5306	+0.3834
Variance of Output Price (Vp_1)	-0.0451	-0.0426
Initial Wealth (W_0)	+0.4322	+0.2626

We will use estimates from Table A- 5 Column A (OLS, excluding the insignificant lagged dependent variable). The short run (current period) elasticities are:

Table 2 Short Run Elasticity of Wheat Yield

	Table A-5 Column A	Table A-5 Column B
Expected Output Price (Ep_1)	+0.1768	+0.1278
Variance of Output Price (Vp_1)	-0.0407	-0.0434
Initial Wealth (W_0)	+0.2923	+0.1730
Insurance Effect ($CovGp_1$)	-0.0309	-0.0304

For barley yields, the relevant long-run elasticities from Table A- 7columns A and B Appendix A:

Table 3 Long Run Elasticity of Barley Yield

	Table A-7 Column A	Table A-7 Column B
Expected Output Price (Ep_2)	+0.5310	+0.5202
Variance of Output Price (Vp_2)	-0.1363	-0.1451
Initial Wealth (W_0)	0.4395	+0.4105
Insurance Effect ($CovGp_2$)	0	0

We will use estimates from Table 7A (OLS, excluding the insignificant lagged dependent variable). The short run (current period) elasticities are:

Table 4 Short Run Elasticity of Barley Yield

	Table A-7 Column A	Table A-7 Column B
Expected Output Price (Ep_2)	+0.1768	+0.1029
Variance of Output Price (Vp_2)	-0.0407	-0.0473
Initial Wealth (W_0)	+0.2923	+0.2534
Insurance Effect ($CovGp_2$)	0	0

For canola yields, the relevant long run elasticities from Table 8AB in the Appendix are (initial wealth is statistically insignificant):

Table 5 Long Run Elasticity of Canola Yield

	Table A-8 Column A	Table A-8 Column B
Expected Output Price (Ep_3)	+0.3203	+0.3143
Variance of Output Price (Vp_3)	-0.0520	-0.0541
Initial Wealth (W_0)	0	0
Insurance Effect ($CovGp_3$)	0	0

We will use estimates from Table 8A (OLS, excluding the insignificant lagged dependent variable).
The short run (current period) elasticities are all zero.

Crop acreage demands are specified as static models providing only a current period (short run) response conditional on yields rather than a dynamic response. Coefficient estimates are from the following Tables in the Appendix B on crop acreage demands: Table B-1 column A for wheat, Table B-2 column B for barley, and Table B-3 column A for canola. Elasticities are calculated based on average shares over 1960-2002 (*$sz_1 = 0.465$, $sz_2 = 0.175$, $sz_3 = 0.118$, versus $sz_1 = 0.445$, $sz_2 = 0.136$, $sz_3 = 0.265$* in 2000).

Table 6 Short run elasticities of acreage demands, conditional on yields

	Wheat Acres Share		Barley Acres Share		Canola Acres Share	
	Coef	elasticity	Coef	elasticity	Coef	elasticity
Er_1	+0.1051	+0.2257	-0.0542	-0.3092	0	0
Er_2	+0.0866	+0.1860	+0.1246	+0.7108	-0.1339	-1.1347
Er_3	-0.0702	-0.1508	-0.0236	-0.1346	+0.0993	+0.8415
Vr_1	-0.0156	-0.0335	0	0	0	0
Vr_2	0	0	0	0	-0.0115	-0.0975
Vr_3	0	0	0	0	-.0052	-.0441
Vr_{12}	+0.00005	+0.0001	0	0	0	0
Vr_{24}	0	0	-0.0072	-0.0411	0	0
W_0	-0.1577	-0.3388	0	0	0	0
VW_2	+.0156	+.0335	0	0	-0.0056	-0.0475

In the long run, changes in acreage demands also depend on long run changes in yields. For example, a LR elasticity of +0.53 for Ep_1 on wheat yield implies that a 1% increase in Ep_1 induces a 1.53% increase in Er_1 in the LR, so the LR elasticities of acreage demands with respect to Ep_1 are 1.53 times larger than the impacts of Er_1 reported in the previous table. The change in wheat yield also induces a change in revenue uncertainty per acre Vr_1 and Vr_{1i}, but these effects on acreage demands are minor (given small elasticities of demand with respect to Vr) and are ignored here. Similarly a LR elasticity of -0.045 for Vp_1 on wheat yield implies that a 1% increase in Vp_1 induces a 0.045% decrease in Er_1 in the LR, and this is incorporated into LR impacts of Vp_1 on demands. This impact on wheat yield also induces a change in LR Vr_1 and Vr_{1i}, but these effects are minor and therefore are ignored here (e.g. approximating Vr_1 as $yld_1{}^2 Vp_1$, then a 1% increase in Vp_1 induces a 0.998% LR increase in Vr_1). Since initial wealth W_0 induces statistically significant long run changes in both wheat and barley yields, these effects multiplied by effects of Er on acreage are incorporated into long run impacts of W_0 on acreage. LR impacts on acreage demands are summarized in the following table.

Table 7 Elasticities of Acreage Demands in Short and Long Run

	Wheat Acres Share		Barley Acres Share		Canola Acres Share	
	SR	LR	SR	LR	SR	LR
Ep_1	+0.2257	+.3455	-0.3092	-.4733	0	0
Ep_2	+0.1860	+.2848	+0.7108	+1.0882	-1.1347	-1.7372
Ep_3	-0.1508	-.1991	-0.1346	-.1777	+0.8415	+1.1110
Vp_1	-0.0335	-.0437	0	+.0139	0	0
Vp_2	0	-.0254	0	-.1253	-0.0975	+.0572
Vp_3	0	+.0078	0	+.0070	-.0441	-.0879
Vp_{12}	+0.0001	+.0001	0	0	0	0
Vp_{24}	0	0	-0.0411	-0.0411	0	0
W_0	-0.3388	-.1595	0	+.1788	0	-.0588
VW_2	+.0335	+.0335	0	0	-0.0475	-0.0475

These estimates for yield and acreage response can be combined into long run and short run impacts on output (conditional on total crop acres). Long run impacts on output are the sum of long run impacts on yields plus long run impacts on acres. Short run impacts on output are the sum of short run impacts on yields plus short run impacts on acres.

Table 8 Output Elasticities Short and Long Run

	Wheat Output		Barley Output		Canola Output	
	SR	LR	SR	LR	SR	LR
Ep_1	+.4025	+.8761	-.3092	-.4733	0	0
Ep_2	+.1860	+.2848	+.8876	+1.6192	-1.1347	-1.7372
Ep_3	-.1508	-.1991	-.1346	-.1777	+.8415	+1.4303
Vp_1	-.0742	-.0888	0	+.0139	0	0
Vp_2	0	-.0254	-.0407	-.2616	-.0975	+.0572
Vp_3	0	+.0078	0	+.0070	0	-.1399
Vp_{12}	+.0001	+.0001	0	0	0	0
Vp_{24}	0	0	-.0411	-.0411	0	0
W_0	-.0405	+.2727	+.2923	+.6183	0	-.0588
CovGp	-.0309	-.0503	0	0	0	0
VW_2	+.0335	+.0335	0	0	-.0475	-.0475

5.3 GRIP

Here we briefly consider one relatively standard program that can serve as a benchmark for the current program. The Gross Revenue Insurance Program (GRIP) was introduced in 1989 and provided payments to Manitoba crop producers from 1991 to 1995. GRIP was the major source of government payments to Manitoba crop producers over this period. The program was intended to stabilize partially gross income from crops. In Manitoba and all provinces except Saskatchewan, revenue insurance was crop specific. A farmer would pay a premium to insure gross revenues of a crop at a certain level, and he would receive an indemnity when area revenues fell below the coverage level. Premiums were subsidized (typically producers would pay 1/3 of the insurance premium). Target revenue depended on long-run yields and a 15 year moving average price. The crop specific nature of the insurance program presumably encouraged a shift in production from low risk crops to high risk crops.

Annual GRIP data was obtained for 1991-1995 from the Farm Financial Branch of AAFC. Data relevant to our simulation of GRIP is summarized as follows.

Table 9 Percentage Change in Output Price due to GRIP

	wheat	barley	canola	oats	rye	flax	total
1991	0.419	0.183	0.057	0.034	0.360	0.469	0.276
1992	0.129	0.037	0.029	0.003	0.116	0.105	0.079
1993	0.411	0.134	0.058	0.013	0.097	0.211	0.241
1994	0.034	0.020	0.003	0.006	0.021	0.029	0.024
1995	-0.009	-0.002	0.000	-0.001	0.001	0.013	-0.005
avg.	0.197	0.074	0.034	0.011	0.119	0.165	0.123

These figures are the percentage change in farm level crop output prices due to GRIP (indemnities minus farmer premiums) relative to market prices. Total is an index of these changes for all 6 crops.

Correlations of total indemnities with total crop market revenue, a Divisia index of market prices for the 6 crops, and market price for wheat over 1991-1995 are presented in the following table. Similar correlations are presented for total indemnities minus farmer premiums.

Table 10 Correlations of total indemnities (1991-1995)

	Total Indemnities	Total Net Indemnities
Total Revenue	-0.9492	-0.9595
Divisia Price	-0.9005	-0.8846
Wheat Price	-0.8813	-0.8728

Thus GRIP payments were very highly negatively correlated with Manitoba market revenues and prices.

Our econometric models can be used to simulate impacts of these price supports and insurance effects on output. We consider impacts for average levels of percentage increase in price and for 1991, when this increase was highest. This data is taken from table 8 with respect to price changes. Impacts of price changes are calculated using summary tables of econometric results in section 3.1 above. Since GRIP was the major source of insurance effects for Manitoba crop producers over this period, the insurance impact of GRIP (or of removing GRIP) is approximated by coefficients of insurance effect in the previous section. The insurance effect indicates the impact of GRIP in truncating uncertainty (so we do not also calculate impacts of GRIP through changes in the price variance Vp term of econometric models). GRIP does not have a direct impact on wealth, although of course it may lead to future accumulation of wealth. Nevertheless we ignore this indirect effect of GRIP, or we assume that it is implicit in our econometric models.

The calculated short-run (one year) and long-run impacts of GRIP on crop outputs are summarized in the following tables.

Table 11 Percentage Change in Wheat Output due to GRIP

		price effect	insurance effect	total impact
Average	short run	0.0879	0.0309	0.1188
	long run	0.1868	0.0503	0.2371
1991	short run	0.1943	0.0309	0.2252
	long run	0.4083	0.0503	0.4586

Table 12 Percentage Change in Barley Output due to GRIP

		price effect	insurance effect	total impact
Average	short run	0.0006	0	0.0006
	long run	0.0213	0	0.0213
1991	short run	0.0255	0	0.0255
	long run	0.0885	0	0.0885

Table 13 Percentage Change in Canola Output due to GRIP

		price effect	insurance effect	total impact
Average	short run	-0.0558	0	-0.0558
	long run	-0.0486	0	-0.0486
1991	short run	-0.1602	0	-0.1602
	long run	-0.2372	0	-0.2372

6.0 CAIS

6.1 Introduction

The Canadian Agricultural Income Stabilization (CAIS) Program was approved in late 2003 and is now Canada's single safety net program, in place of the Net Income Stabilization Account (NISA), Canadian Farm Income Plan (CFIP) and related provincial programs. The program structure can be summarized briefly as follows. A production margin is intended to reflect revenues and expenses that are directly related to production for the firm. A reference production margin is an average of the five previous production margins for the firm, excluding the high and low margins. If the production margin for the firm is less than the reference margin for the year, then claims generally are triggered for the difference. Claim payments are financed from producer deposits and government contributions, and shares vary with the difference relative to the reference margin as follows: producer deposits cover 1/2 of 0 to 15% loss, 3/10 of 15 to 30% loss, and 1/5 of loss greater than 30 % (claims are in principle limited by available producer deposits, which must be a minimum of 14 % of the current reference production margin). The government finances 60% of negative production margins. By supporting incomes during low income years, CAIS may influence production decisions.

CAIS payouts are based on farm specific losses relative to reference margin rather than on a regional measure of loss (as in the Western Grain Stabilization Act). Current payouts to a farm depend on the difference between (farm) reference margin and current profits for the particular farm. Presumably this creates moral hazard problems. From the viewpoint of the current paper, this also substantially changes shadow prices for outputs and inputs, and so in theory this has a substantial impact on production.

Most analytical studies of farm program impacts under risk focus on analyzing a formal specification of an expected utility (or mean variance) maximization problem with integral versions of truncated distributions for price or yield. Much has been accomplished in this way (e.g. Chavas and Holt; Hennessy). However relations between shadow prices and risk under various insurance type programs have often been obscured (two exceptions in the context of crop insurance are Quiggin, Karagiannis and Stanton; Ramaswami).

In contrast, this study focuses directly on moments of distributions of net returns under programs. Simple statistics, essentially only expectations operators, are adequate for our purposes. This approach clarifies CAIS program impacts on shadow prices under risk. Calculated changes in moments translate into production impacts using homogeneity conditions under constant relative risk aversion and estimates of an econometric model.

We adopt the following notation. For a firm in a particular year t, let p denote a vector of market prices for commodities $i = 1,.,n$, y is a vector of corresponding output levels, w is a vector of market prices for inputs $j=1,.,m$, and x is a vector of corresponding input levels. Let $\pi = p\, y - w\, x$ denote farm level profits or market income. We will assume that this represents production margin, which is a reasonable approximation for our purposes. Let π^T denote the reference production margin. This is the average of the production margins over the previous 5 years, deleting the high and low, and denote this as $\pi^T = \Sigma_s\, \pi_{t-s}\ /s$. Total payouts under CAIS to the farm are $\pi^T - \pi$ if $\pi^T > \pi \geq 0$ (assuming adequate producer deposits are available). Let θ denote the government share of total payouts in this case ($\theta = 0$ if there is no trigger, i.e. $\pi^T - \pi \leq 0$ and $\pi \geq 0$)). Let W_0 denote initial wealth of the farm. Then end of year wealth for the farm can be defined as:

$$W_F = W_0 + \pi + \theta(\pi^T - \pi) \qquad (6)$$

Suppose for the moment that all prices and yields are known with certainty, so that π is known for any given production decisions y,x. Assume that, in the absence of CAIS, the firm maximizes profits: $max_{(x,y)\varepsilon T}\ \pi = py - wx$ or equivalently $max_y\ \pi = py - C(w,y)$ where $C(w,y)$ is the firm's dual cost function. At the profit maximizing output levels y^* the corresponding first order conditions are

$$\partial\pi(y^*)/\partial y = p - \partial C(w,y^*)/y = 0 \qquad (7)$$

In the presence of CAIS, the firm will consider the impact of its decisions y,x on government payments as well as farm profits π. In addition, if the firm expects CAIS to continue into the future, it will consider the impact of current decisions on future reference margins and in turn on future government payouts. The function $\theta = \theta(\pi^T,\pi)$ is continuous in π and is defined by CAIS rules. Then the firm's maximization decision for year t can be represented as follows (assuming for simplicity that current π is neither high nor low π over each of the next 5 years):

$$\max_y \psi = W_0 + \pi + \theta(\pi^T - \pi) + \sum_{s=1,.,5} \theta_{t+s}\pi/3(1+r)^s \qquad (8)$$
$$\text{where: r} \equiv \text{firm's intertemporal discount rate}$$

Equation (8) also assumes that current π and future π are independent, i.e. the production model is static rather than dynamic. The first order conditions for a solution to this problem are

$$\begin{aligned}\partial\psi/\partial y &= (1-\theta)\,\partial\pi/\partial y + \sum_{s=1,.,5}\theta_{t+s}\,\partial\pi/\partial y\,/\,3(1+r)^s + \partial\theta/\partial\pi\ \partial\pi/\partial y\,(\pi^T - \pi)\\ &\quad + \sum_{s=1,.,5}\partial\theta_{t+s}/\partial\pi\ \ \partial\pi/\partial y\ \pi/3(1+r)^s\\ &= 0\end{aligned} \qquad (8a)$$

At the profit maximizing level y^* in the absence of CAIS, $\partial\pi(y^*)/\partial y = 0$. Then y^* also solves the first order conditions (8) under CAIS. This implies that, in the absence of uncertainty, CAIS would not influence production decisions.

The first order conditions (8a) also imply effective or shadow prices for outputs and inputs under CAIS. Substituting $p - \partial C(w,y)/\partial y$ *for* $\partial\pi/\partial y$ *in (8),* $\partial\psi/\partial y_i = \gamma\, p_i - \gamma\, \partial C(w,y)/\partial y_i = \gamma\, p_i - \partial C(\gamma w,y)/\partial y_i$ by linear homogeneity of $C(w,y)$ in w, where $\gamma \equiv (1\text{-}\theta) + \Sigma_{s=1,.,5}\ \theta_{t+s}$ /

$3(1+r)^s$ + $\partial\theta/\partial\pi\ (\pi^T - \pi)$ + $\Sigma_{s=1,.,5}\ \partial\theta_{t+s}/\partial\pi\ \ \pi\ /3(1+r)^s$. Then effective prices under CAIS can be defined as

$$
\begin{aligned}
p_i^* &= \gamma\, p_i \\
w_j^* &= \gamma\, w_j
\end{aligned}
\qquad (9)
$$

Note that $Ep^* / w^* = Ep / w$, i.e. relative shadow prices are unchanged by CAIS. So decisions are unchanged by CAIS in the absence of uncertainty.

These shadow prices have a simple explanation. Assume an exogenous $\theta > 0$ and $\pi^T - \pi > 0$, and ignore impacts on future reference margins. Then an increase in market income $\Delta\pi$ leads to a decrease in government payments $\theta\,\Delta\pi$, i.e. the net effect is $(1 - \theta)\ \Delta\pi$. Thus the net effect on total income (farm income plus government payments) is $(1 - \theta)$ for a \$1 increase in revenue and also $(1 - \theta)$ for a \$1 decrease in cost. So the shadow prices for outputs and inputs are $(1 - \theta)\ p$ and $(1 - \theta)\ w$.

6.2 Modeling Impacts of CAIS under Price Uncertainty and Risk Aversion

The contribution of current period production decisions to farm wealth under CAIS assuming certainty is summarized essentially by the above function ψ (8). If the farmer suspects that CAIS will not exist in the future, then current production margin π does not create an investment in future reference margins π^T, and then ψ reduces to

$$
\begin{aligned}
\psi' &= W_0 + \pi + \theta(\pi^T - \pi) \\
&= W_0 + (1-\theta)\ \pi + \theta\pi^T
\end{aligned}
\qquad (10)
$$

We now recognize that output price and in turn π and θ are random variables, but we ignore uncertainty in y,w,x (and W_0). We assume that producer deposits are sufficient to cover loss claims under CAIS, and we do not address any impacts of opportunity costs for farmer deposits on production incentives.

Truncation effects of support programs under risk are often analyzed as in Chavas and Holt (e.g. von Massow and Weersink). This approach provides analytical evaluations of truncations of price moments for simple price support programs which place a floor under market price. However CAIS is a more complex program, so these simple evaluations of truncations presumably do not apply here. Nevertheless effects of CAIS (or at least our idealized version of CAIS) on price distributions can be analyzed by simple statistics plus synthetic data, as we will show.

The expected value of (10) is:

$$E\psi' = W_0 + (1\text{-}E\theta)\, E\pi + \mathrm{cov}(\text{-}\theta,\pi) + E\theta\, \pi^T \tag{11}$$

Then the marginal impacts of output decisions y on $E\psi'$ are:

$$\partial E\psi'/\partial y_i = (1\text{-}E\theta)\, \partial E\pi/\partial y_i + \partial E\theta/\partial y_i\, (\pi^T\text{-}E\pi) \text{ - } \partial\, \mathrm{cov}(\theta,\pi)/\partial y_i \tag{12}$$

If θ and π are related approximately by a linear model $\theta = \alpha_0 + \alpha_1\, \pi + \alpha_2\, Z + e$ where $Ee=0$, then $E\theta = \alpha_0 + \alpha_1\, E\pi + \alpha_2\, EZ$ and in turn $\partial E\theta/\partial y_i = \partial E\theta/\partial E\pi\; \partial E\pi/\partial y_i$. $\partial E\theta/\partial E\pi = \alpha_1$ and $\partial E\pi/\partial y_i = Ep_i - \partial C(w,y)/\partial y_i$.

Furthermore it might be tempting to evaluate $\partial cov(\theta,\pi)/\partial y$ as $\partial cov(\theta,\pi)/\partial\pi\; \partial\pi/\partial y$ and approximate $\partial\pi/\partial y$ as $Ep - \partial C(w,y)/\partial y$. In this case (12) would imply that CAIS would not change relative expected prices Ep and w. However this argument would be incorrect.

$\partial cov(\theta,\pi)/\partial y$ can be analyzed as follows. First, by definition,

$$\begin{aligned}
\mathrm{cov}(\theta,\pi) &\equiv E\{(\theta - E\theta)(py\text{-}wx\ \text{ - } E(py\text{-}wx))\} \\
&= E\{(\theta - E\theta)(py \text{ - } E(py))\} && \text{if } wx \text{ nonstochastic} \\
&\equiv \mathrm{cov}(\theta, R) && R \equiv p\ y \\
&\equiv E\{(\theta - E\theta)(\textstyle\sum_i p_i y_i - \sum_i E(p_i y_i))\} \\
&= E\{(\theta - E\theta)(\textstyle\sum_i (p_i - Ep_i) y_i)\}\ (\ = E_i\, \mathrm{cov}(\theta, p_i y_i)) && \text{if } y \text{ nonstochastic} \\
&= \textstyle\sum_i E\{(\theta - E\theta)(p_i - Ep_i)\} y_i && \text{if } y \text{ nonstochastic} \\
&\equiv \textstyle\sum_i \mathrm{cov}(\theta, p_i)\ y_i
\end{aligned} \tag{13}$$

Note that the above argument does not make any assumptions regarding the form (e.g. linearity) of the relation between θ and π.

The argument can be summarized in words as follows. As perceived by the firm, θ and π are random variables. We assume that the cost component wx of π is known to the firm, so the covariance $cov(\theta,\pi)$ of the random variables θ and π (as perceived by the firm) reduces to the covariance $cov(\theta,R)$ between θ and revenues R. Assuming for simplicity that y is non-stochastic, $cov(\theta,R)$ reduces to $\Sigma_i\, cov(\theta,p_i)\, y_i$. In sum,

$$\begin{aligned}
\mathrm{cov}(\theta,\pi) &= \mathrm{cov}(\theta,R) && \text{if } wx \text{ non-stochastic} \\
&= \textstyle\sum_i \mathrm{cov}(\theta,p_i)\, y_i && \text{if } y \text{ non-stochastic}
\end{aligned} \tag{13a}$$

We will need to evaluate $cov(\theta,p_i)$ or more precisely $\partial cov(\theta,R)/\partial y_i$, which requires an evaluation of $cov(\theta,R)$. Assume that the firm's subjective probability distribution for θ can be approximated by a linear regression model $\theta = \alpha_0 + \alpha_1 R + \alpha_2 Z + e$ relating current θ to current R, where Z denotes other variables conditioning the firm's subjective probability distribution for θ, and e is a disturbance. Current θ and current R are uncertain (not known with certainty) to the firm and hence both are stochastic in the firm's subjective distribution. Z is essentially information known with certainty that conditions the firm's subjective distribution for θ (Z is essentially current costs and lagged θ and π). Then Z is non-stochastic, so stochastic current R will not covary with Z in terms of this subjective probability distribution. Then $cov(\theta,R)$ is evaluated as follows in terms of the firm's subjective probability distribution.

$$
\begin{aligned}
\mathrm{cov}(\theta,R) &\equiv E\{(\theta - E\theta)(R\text{-}ER)\} \\
&= E\{(\alpha_1(R\text{-}ER)(R\text{-}ER)\} + E\{(\alpha_2(Z-EZ)(R-ER)\} && \text{if } \mathrm{cov}(e,R)=0 \\
&= \alpha_1 \,\mathrm{var}(R) + \alpha_2 \,\mathrm{cov}(Z,R) \\
&= \alpha_1 \, y^T \, Vp \, y + \alpha_2 \sum_j \mathrm{cov}(Z,p_j)\, y_j && \text{if } y \text{ is non-stochastic} \quad (13b) \\
&= \alpha_1 \, y^{\mathrm{T}} \, Vp \, y && \text{if } z \text{ is non-stochastic}
\end{aligned}
$$

where: Vp ≡ output price covariance matrix

Comparing (13a) and (13b),

$$
\begin{aligned}
\mathrm{cov}(\theta,\pi) &= \sum_i \mathrm{cov}(\theta,p_i)\, y_i \\
&= \sum_i (\alpha_1 \sum_j (Vp_{ij} \; y_j)\, y_i
\end{aligned}
\qquad (14)
$$

so

$$
\mathrm{cov}(\theta,p_i) = \alpha_1 \sum_j \mathrm{cov}(p_i,p_j)\, y_j \qquad (15)
$$

Since all $cov(p_i,p_j) \geq 0$ for Manitoba and $\alpha_1 \leq 0$ by CAIS rules, then $cov(\theta,p_i) \leq 0$. So

$$
\begin{aligned}
\partial\, \mathrm{cov}(\theta,\pi)/\partial y_i &= 2\,\alpha_1 \sum_j \mathrm{cov}(p_i p_j)\, y_j && by\,(14) \\
&= 2\,\mathrm{cov}(\theta,p_i) && by\,(15)
\end{aligned}
\qquad (16)
$$

This is our key result for calculating impacts of CAIS on relative expected prices Ep/w and Ep_i/Ep_j . Note that α_1 can be calculated from (13b) as

$$\alpha_1 = \mathrm{cov}(\theta,\pi) / y^T\ Vp\ \ y \qquad (17)$$
given estimates of Vp and $\mathrm{cov}(\theta,\pi)$

The subjective $cov(\theta,\pi)$ (π^T known with certainty) may be better calculated from data as $cov(\theta,\pi-\pi^T)$, which can be decomposed in data to $cov(\theta,\pi) - cov(\theta,\pi^T)$. Thus it is not necessary to estimate α_1 by econometrics. Indeed regression is inappropriate since the covariance of the joint subjective distribution for θ and π reflects a nonlinear identity corresponding to CAIS rules. Also an appropriate regression model may have nonlinear errors in variables.

One difficulty with the above analysis (13 b)-(17) is the assumption of a linear relation between θ and π or $\pi-\pi^T$, whereas CAIS rules define a nonlinear relation. In order to see the complications implied by nonlinearity, consider a quadratic relation $\theta = \alpha_0 + \alpha_1 \pi + \alpha_2 \pi^2$. Then wx non-stochastic implies $cov(\theta,\pi) = \alpha_1 var(R) + \alpha_2 cov(R^2,R)$, so α_1 and α_2 cannot both be calculated from this equation. However we will conclude in the next section that the linear approximation appears to be adequate for a particular Manitoba data set.

Note that, since Ep_i is non-stochastic, $cov(\theta,p_i) \equiv E\{(\theta-E\theta)(p_i-Ep_i)\} = Ep_i E\{(\theta-E\theta)(p_i/Ep_i-1)\} \neq Ep_i E\{(\theta-E\theta)(p_i-1)\}$ unless $Ep_i=1$. Thus $cov(\theta,p_i)$ is not directly proportional to Ep_i (although scaling the random variable p_i by λ multiplies Ep_i and $cov(\theta,p_i)$ by λ). So relative expected output prices Ep_i / Ep_j are changed (to at least some extent) by CAIS.

Under our assumptions, (16) implies that (12) can be approximated as

$$\partial E\psi' / \partial y_i = \tau \cdot Ep_i - \tau \cdot \partial C(w,y)/\partial y_i - 2\,\mathrm{cov}(\theta, p_i) \qquad (18)$$
$$\text{where: } \tau = (1 - E\theta) + \partial E\theta / \partial E\pi\, (\pi^T - E\pi)$$

Thus the increase in $E\psi'$ for a marginal increase in y_i, excluding the marginal cost $\partial C/\partial y_i$, is $\tau Ep_i - 2\, cov(\theta,p_i)$. In turn shadow expected returns for a marginal increase in y_i are $\tau Ep_i - 2\, cov(\theta,p_i)$. Also note that $\tau \cdot \partial C(w,y)/\partial y_i = \partial C(\tau w,y)/\partial y_i$ since the cost function is linear homogeneous in prices. Thus our assumptions imply that effective expected prices and effective expected initial wealth under CAIS are approximately

$$Ep_i{}^* = \tau \cdot Ep_i - 2\ \mathrm{cov}(\theta, p_i) \qquad (19)$$
$$w_j{}^* = \tau \cdot w_j$$
$$EW_0{}^* = W_0 + E\theta\, \pi^T$$

where: $\mathrm{cov}(\theta,\mathrm{p_i})$ are not directly proportional to $\mathrm{Ep_i}$

The difference between effective expected output and input prices under CAIS and expected market prices can be explained somewhat intuitively as follows. First, an increase in output and hence income leads to a decrease in government payments both directly (through a change in π) and indirectly (through an induced change in θ). This tends to reduce uniformly shadow prices of outputs and inputs as indicated by the terms τEp_i and τw_j, $\tau = (1-E\theta) + \partial E\theta/\partial E\pi\ (\pi^T - E\pi)$. These effects by themselves do not change relative prices.

Second, consider the covariance effect. Costs are assumed to be non-stochastic, so the covariance between θ and income reduces to a covariance between θ and market revenue py. Since this covariance is independent of costs, it does not imply a transformation of input prices. This covariance does imply a transformation of output prices.

The expected value of investment in future CAIS payouts can be approximated as

$$\sum\nolimits_{s=1,..,5} E\theta_{t+s}\ E\pi / 5(1+r)^s + \sum\nolimits_{s=1,..,5} \text{cov}(\theta_{t+s}, \pi)/5(1+r)^s \qquad (20)$$

Here the covariance between government share of payouts in the future (θ_{t+s}) and current profits π is positive (an increase in current production margin leads to an increase in future reference margin without influencing future production margins, so θ is likely to increase in the future). So $cov(\theta_{t+s},\pi) > 0$. We will simplify (20) to

$$\begin{aligned}&\sum\nolimits_{s=1,..,5} E\theta_F\ E\pi / 5(1+r)^s + \sum\nolimits_{s=1,..,5} \text{cov}(\theta_F, \pi)/5(1+r)^s \\ &\quad \text{where: } \theta_F \equiv \text{scalar approximation to } \theta \text{ in the future } (\theta_{t+1}, ., \theta_{t+5})\end{aligned} \qquad (21)$$

The marginal impact of y on (21) is approximated as:

$$\{E\theta_F\ \partial E\pi / \partial y_i + \partial E\theta_F / \partial E\pi\ \partial E\pi / \partial y_i + 2\,\text{cov}(\theta_F, p_i)\} \quad \sum\nolimits_{s=1,..,5} 1/5(1+r)^s \qquad (22)$$

Assuming that farmers expect CAIS to continue into the future and factor investment in future CAIS payouts into their decisions, then the effective expected prices under CAIS are modified from (19) to

$$\begin{aligned}&Ep_i^{**} = (\tau+\gamma)\, Ep_i - 2\,\text{cov}(\theta, p_i) + 2\,\text{cov}(\theta_F, p_i) \sum\nolimits_{s=1,..,5} 1/5(1+r)^s \\ &\quad w_j^{**} = (\tau+\gamma)\, w_j \\ &\quad \text{where: } \gamma \equiv \{E\theta_F + \partial E\theta_F / \partial E\pi\} \sum\nolimits_{s=1,..,5} 1/5(1+r)^s\end{aligned} \qquad (23)$$

Next consider how CAIS modifies income uncertainty. The variance of ψ' (8) is:

$$
\begin{aligned}
Var\,\psi' &= Var\{(1\text{-}\theta)\pi\} + Var(\theta\pi^T) + 2\,Cov\{(1\text{-}\theta)\pi, \theta\pi^T\} \\
&= (1\text{-}E\theta)^2\ Var\pi + Var(1\text{-}\theta)\ E\pi^2 + Var(1\text{-}\theta)\ Var\pi \\
&\quad \text{-}2(1\text{-}E\theta)E\pi\ \mathrm{cov}(1\text{-}\theta,\pi)\text{ - }\mathrm{cov}(1\text{-}\theta,\pi)^2 + \mathrm{cov}((1\text{-}\theta)^2,\pi^2) \\
&\quad + Var\theta\,(\pi^T)^2 + 2\,Cov\{(1\text{-}\theta)\pi,\theta\}\pi^T \\
&= (1\text{-}E\theta)^2\ y^T Vp\ y + Var\theta\ E\pi^2 + Var\theta\ y^T Vp\ y \\
&\quad + 2(1\text{-}E\theta)E\pi\ \mathrm{cov}(\theta,\pi)\text{ - }\mathrm{cov}(\theta,\pi)^2 + \mathrm{cov}((1\text{-}\theta)^2,\pi^2) \\
&\quad + Var\theta\,(\pi^T)^2 + 2\,Cov(\theta,\pi)\pi^T\text{-}2\,Cov(\theta\pi,\theta)\pi^T \\
&= \{(1\text{-}E\theta)^2 + Var\theta\}\ y^T Vp\ y + \{E\pi^2 + (\pi^T)^2\}\ Var\theta \\
&\quad + 2(1\text{-}E\theta)E\pi\ \mathrm{cov}(\theta,\pi)\text{ - }\mathrm{cov}(\theta,\pi)^2 + \mathrm{cov}((1\text{-}\theta)^2,\pi^2) \\
&\quad + 2\,Cov(\theta,\pi)\pi^T\text{ - }2\,Cov(\theta\pi,\theta)\pi^T
\end{aligned}
\qquad (24)
$$

where : $\mathrm{Vp} \equiv$ covariance matrix of market prices p

If the firm could predict θ with certainty, then (24) would imply $Var\ \psi' = (1\text{-}E\theta)^2\ Var\ \pi$, i.e. the variance measure of uncertainty under CAIS would be less than the variance of market income. In principle, an effective price covariance matrix *Vp** under CAIS can be obtained by twice differentiating (24) with respect to *y*.

Alternatively, we can define two components of the effective variance of ψ' (24) as follows:

$$Vp^* = \{(1\text{-}E\theta)^2 + Var\theta\}\ Vp \qquad (25\text{-A})$$

$$\varnothing \equiv 2(1 - E\theta)E\pi\ \mathrm{cov}(\theta,\pi) - \mathrm{cov}(\theta,\pi)^2 + \mathrm{cov}((1\text{-}\theta)^2,\pi^2) + 2\,Cov(\theta,\pi)\pi^T\text{ - }2\,Cov(\theta\pi,\theta)\pi^T \qquad (25\text{-B})$$

(these include 7 of the 8 terms in the last expression for (24)). Note that (25a) are not necessarily less than market price uncertainty *Vp*, due to uncertainty about θ. However it seems very unlikely that CAIS will increase the variance measures of price uncertainty. *Vp** and $\varnothing$ can be substituted into the estimated econometric model in place of market price uncertainty *Vp* and covariance between government payments and market prices *cov(G,p)*, respectively.

If farmers factor investment in future CAIS payouts into their decision making, then the measure of uncertainty under CAIS is more complex. The variance of investment in future CAIS payouts can be approximated as:

$$\sum_{s=1,..,5} \mathrm{var}(\theta_F\ \pi) / [5(1+r)^s]^2$$

or

$$\upsilon\ \mathrm{var}(\theta_F\ \pi) \tag{26}$$

where : $\mathrm{var}(\theta_F\ \pi) = (E\pi)^2\ \mathrm{var}\,\theta_F + (E\theta_F)^2\ \mathrm{var}\,\pi\ \text{-2}\quad E\theta_F\ E\pi\ \mathrm{cov}(\theta_F, \pi)$
$\quad - \mathrm{cov}(\theta_F, \pi)^2 + \mathrm{cov}(\theta_F{}^2, \pi^2)$

The covariance between ψ' and value of investment is approximated as:

$$\mathrm{cov}\{(1\text{-}\theta)\pi + \theta\pi^T, \theta_F\pi\} \sum_{s=1,..,5} 1/5(1+r)^s \tag{27}$$

where : $\mathrm{cov}\{(1\text{-}\theta)\pi + \theta\pi^T, \theta_F\pi\} = \mathrm{cov}(\pi, \theta_F\pi)\text{-}\mathrm{cov}(\theta\pi, \theta_F\pi) + \mathrm{cov}(\theta, \theta_F\pi)\quad \pi^T$

Then (26) and twice (27) should be added to the variance expression (24). In principle the effective price covariance matrix can be derived from second derivatives with respect to y of this altered variance expression. Alternatively, investment in future CAIS payments alters the components (25) of the effective price covariance matrix to approximately

$$Vp^* = \{(1\text{-}E\theta)^2 + Var\,\theta + (E\theta_F)^2\ \upsilon\}\quad Vp \tag{28}$$

6.3 Estimation of CAIS Specific Parameters of Response

Time series data on Manitoba level farm incomes over 1960- 2002 are used to construct a hypothetical time series of government shares of payouts under CAIS, by combining this data with CAIS rules for payouts. Production margin is approximated as crop and livestock market receipts minus operating expenses after rebates and minus depreciation charges, per farm (Manitoba Agriculture Yearbook 2002, pp. 209-210). Reference margin is the average of the five previous production margins, excluding high and low. Percentage loss measures a negative difference between production and reference margin, relative to (absolute value) of reference margin. Government payouts for a year is calculated under the following assumptions: government share of payouts for a negative margin is 0.60, and government share of payouts from 0 margin to a (positive) reference margin is decomposed as 0.50 for 0-15 % loss, 0.70 for 15-30 % loss, and 0.80 above 30 % loss relative to reference margin. An average government share of total payouts θ is calculated for each year.

Results for this exercise are summarized in Table 1 of Appendix C. Margins are measured in $ per farm. The time series mean and variance for these estimates of θ (government share of CAIS payouts) are 0.3907 and 0.0882, respectively. The time series correlation between θ and production margin π is -0.6299, and the correlation between θ and market revenue per farm R is -0.1418.

Alternatively $E\theta$ (expected government share of payouts) can be approximated as follows. If income for a representative farm does not show significant upward trend over time, then CAIS rules imply payouts in approximately half the years. In the Manitoba data set under CAIS rules, payouts would have occurred in 24 of 37 years, so this result is higher than expected. The lowest government share of payouts is 50% (for losses less than 15% relative to reference margin). So 0 payout and 50% payout average to a 25% expected payout, i.e. $E\theta = 0.25$. So we have two scenarios for $E\theta$: (A) $E\theta = 0.3907$ from Manitoba data, and (B) $E\theta = 0.25$.

In order to assess the magnitude of $cov(\theta,p)$ or more precisely $\partial cov(\theta,\pi)/\partial y_i$ relative to p for this data set, we consider two approaches. First, note that (12') $cov(\theta,R) = \Sigma_i\, cov(\theta,p_i)\, y_i$ assuming only output prices are stochastic, and market revenues are $R \equiv \Sigma_i\, p_i\, y_i$. For the data set, $cov(\theta,R) = -1833.0$ and average $R = 61539$, so $cov(\theta,R)$ is 2.98% of average R. Since all $cov(\theta,p_i) \leq 0$, magnitudes of $cov(\theta,p)$ may roughly approximate 3% of p. By (16) and (19), the CAIS effects $\partial cov(\theta,\pi)/\partial y$ imply an approximate 6% increase in effective output prices p relative to input prices w.

Second, consider the earlier discussion (13)-(16). Assume that current θ is related to current R by a linear regression model $\theta = \alpha_0 + \alpha_1 R + \alpha_2 Z + e$ where Z denotes other variables (essentially current reference margin and costs, lagged θ and R) explaining the firm's subjective probability distribution for θ. Then (13b) $cov(\theta,\pi) = \alpha_1\, y^T\, Vp\, y$ assuming y is non-stochastic. This permits calculation of α_1 as $\alpha_1 = cov(\theta,\pi) \,/\, y^T\, Vp\, y$. The price covariance matrix Vp is measured over 1966-2002 Manitoba market price data for the six main crops and livestock and hogs, and y is defined as the average outputs of these commodities over the period. Then $y^T\, Vp\, y = .24591 \text{ X } 10^9$. For the data set, $cov(\theta,\pi) = -1153.3$ and $cov(\theta,\pi\text{-}\pi^T) = -1300.1$. The latter may be a more appropriate measure of $cov(\theta,\pi)$ for the subjective distribution conditional on knowledge of reference margin π^T, but the difference is relatively minor. Using the latter, α_1 is calculated as $\alpha_1 = -0.5287\text{x } 10^{-5}$.

By (15) and (16), $cov(\theta,p_i) = \alpha_1\, \Sigma_j\, cov(p_i,p_j)\, y_j$ and $\partial cov(\theta,\pi)/\partial y_i = 2\, cov(\theta,p_i)$. The following table reports the estimated $cov(\theta,p_i)$ for the six major crops and also for cattle and hogs (units are \$/tonne for crops and \$/animal for cattle and hogs). The table also reports the percentage change in output prices related to $\partial cov(\theta,\pi)/\partial y_i$, i.e. $-2cov(\theta,p_i)/p_i$. All calculations are based on average Manitoba data over 1966-2002.

Table 14 Covariance of share of government payments and price

	cov(θ,pi)	-2cov(θ,pi)/pi
Wheat	-3.35	0.0528
Barley	-2.53	0.0548
Canola	-6.82	0.0533
Oats	-2.7	0.0542
Flax	-2.51	0.0561
Rye	-6.21	0.0502
Cattle	-17.31	0.0607
Hogs	-2.15	0.0425

These calculations suggest that CAIS would increase price ratios Ep/w by 5-6% for all crops. This would tend to increase yields. On the other hand, the output price ratios Ep_i/Ep_j for different crops are essentially unchanged by CAIS, so CAIS would have no direct impact on crop acreage shares through changes in relative output prices.

The essentially unchanged output price ratios p_i/p_j under CAIS reflect the similar terms $\Sigma_j\ cov(p_i,p_j)\ y_j\ /\ p_i$ for all crops and livestock in this data set. These terms can be expressed as $E[\varphi(p_i\text{-}Ep_i)]/p_i$ where $\varphi \equiv \Sigma_j\ (p_j\text{-}Ep_j)\ y_j$. Apparently this similarity reflects the particular historical conditions in Manitoba over 1966-2002, i.e. these terms need not be so similar.

Note that the results in this table are similar to the 6% estimate of change in Ep/w obtained by the first method, although the two methods use somewhat different assumptions. The first method approximated $cov(\theta,p_i)$ without assuming a linear relation between θ and π, whereas the second method estimated $cov(\theta,p_i)$ assuming a linear relation between θ and π (or $\pi\text{-}\pi^T$). CAIS rules imply a nonlinear relation between θ and $\pi\text{-}\pi^T$ (and θ was calculated accordingly in this data set), but the similarity of results for the two methods suggest that the linear approximation in the second method has been acceptable for our purposes.

Assuming a linear approximation as above between θ and π, then $\partial E\theta/\partial E\pi$ can be approximated as α_1 = -0.5287 X 10^{-5}. This provides an estimate of $\partial E\theta/\partial E\pi$ in τ for the shadow price equations (19). Then under scenario A ($E\theta$ = 0.3907), $\tau = (1\text{-}E\theta) + \partial E\theta/\partial E\pi\ (\pi^T\text{-}E\pi)$ can be approximated as τ = 0.6093 - 0.000005287 $(\pi^T\text{-}E\pi)$. τ = 0.6136 using mean data for $\pi^T\text{-}\pi$ = -827 over the data set. Under scenario B ($E\theta$ = .25), τ = 0.75 - 0.000005287 $(\pi^T\text{-}E\pi)$ = 0.7543 using mean data for $\pi^T\text{-}\pi$.

The impact of CAIS on effective expected prices and expected wealth can then be approximated as follows under the two scenarios (A) $E\theta$ = 0.3907 from Manitoba data, and (B) $E\theta$ = 0.25.

$$Ep_i* = .6136\ Ep_i\ -2\,\text{cov}(\theta,p_i)$$
$$w_j* = .6136\ w_j \qquad (29\text{-A})$$
$$EW_0* = W_0 + .3907\ \pi^T$$

$$Ep_i* = .7543\ Ep_i\ -\ 2\ \text{cov}(\theta,p_i)$$
$$w_j* = .7543\ w_j \qquad (29\text{-B})$$
$$EW_0* = W_0 + .25\ \pi^T$$

Where the covariances *cov(θ,p$_i$)* are indicated in table 12.

The components (25) of the effective variance of ψ' are calculated as follows. Cov(θ,π) is approximated as -1153 from the data set (correlation is -0.6299). Similarly cov($(1-\theta)^2,\pi^2$) and Cov(θπ,θ) are measured from the data set as 16261000 and 0.82088, respectively (correlations are .3913 and .0013). Then (25) are measured as:

$$Vp* = (.609^2 + .088)\ Vp \qquad \text{under (A)}\ E\theta = 0.3907$$
$$= 0.459\ Vp \qquad (30\text{-A})$$
$$= 0.651 \qquad \text{under (B)}\ E\theta = 0.25$$

$$\varnothing \equiv\ 2(.609)(-11543)\ E\pi - (-1153)^2 + 16261000$$
$$+ 2(-1153)\pi^T - 2(.821)\ \pi^T \qquad \text{under (A)} \qquad (30\text{-B})$$
$$= -1404.354\ E\pi\ -2307.642\ \pi^T + 1493159.1$$

Thus effective price uncertainty *Vp** under CAIS are substantially less than market *Vp*, as approximated here. ∅ = -4173712.5 at average data over the data set. The remaining term in *varψ'* (24) is $\{E\pi^2 + (\pi^T)^2\}\ Var\theta$ = 1478272.3 at average data, which is smaller in magnitude than the estimate of ∅. *Vp** and ∅ might be substituted into the estimated econometric model in place of market price uncertainty *Vp* and covariance between government payments and market prices *cov(G,p)*, respectively.

6.4 Hypothetical Impacts of CAIS on Prices, Wealth and Price Risk

Short run and long run crop production impacts of CAIS are simulated from our Manitoba data set and econometric estimates of acreage and yield response for Manitoba. Simulations are conducted using average time series data for Manitoba over 1966- 2002 and also using 2002 data.

Calculation of impacts of CAIS on effective relative expected prices *Ep/w* is simple, given above results. From (20) and an above table of estimates for changes in *p*, percentage changes in effective relative expected prices Ep/w due to CAIS are (under scenarios (A) $E\theta$ = 0.3907 and (B) $E\theta$ = 0.75)

$$(Ep_i^*/w^* - Ep_i/w) / Ep_i/w$$

$$= \{[Ep_i/w - 2\,\mathrm{cov}(\theta, p_i)/.6136w] - Ep_i/w\} / Ep_i/w \qquad \text{under (A) } E\theta = 0.3907$$

$$= -2\,\mathrm{cov}(\theta, p_i) / .6136\, Ep_i$$

$$\begin{aligned}
&= .0860\,(A) \quad = .0700\,(B) \quad \text{wheat} \\
&= .0893\,(A) \quad = .0726\,(B) \quad \text{barley} \\
&= .0869\,(A) \quad = .0707\,(B) \quad \text{canola} \\
&= .0883\,(A) \quad = .0718\,(B) \quad \text{oats} \\
&= .0914\,(A) \quad = .0744\,(B) \quad \text{flax} \\
&= .0818\,(A) \quad = .0665\,(B) \quad \text{rye}
\end{aligned} \qquad (31)$$

Thus CAIS would lead to an approximately 8 to 9% increase in effective relative expected prices Ep_i/w for all crops under scenario (A) $E\theta$ = 0.3907, or to an approximately 7% increase under scenario (B) Eθ = 0.25.

However further calculations of impact on relative wealth and price uncertainty is not straightforward. A brief summary of homogeneity conditions and appropriate normalizations under alternative risk preferences is required in order to clarify these calculations. A reason for confusion in normalization is that CAIS changes all expected effective prices, i.e. CAIS changes shadow prices for inputs as well as for outputs.

Constant relative risk aversion (CRRA) is a benchmark assumption in the theory of decision making under risk. Moreover the CRRA hypothesis was tested and not rejected for the above Manitoba acreage demands model. It is difficult to test the expected utility (or mean variance) hypothesis within our framework.

Assuming CRRA and mean-variance or expected utility maximizing behaviour, the zero homogeneity condition on output supplies is $y(\lambda Ep,\lambda w,\lambda^2 Vp,\lambda W_0) = y(Ep,w,Vp,W_0)$ for any $\lambda>0$ (Pope 1988; Coyle 1999), and here we select $\lambda = 1/w_i$ for normalization, i.e. the crop index input price index is the numeraire price. Any vectors (Ep, w, Vp, W_0) that are proportional as in the homogeneity condition lead to the same decision y, and it is variation in these proportions that leads to changes in decisions. Similarly under CAIS and CRRA the identical homogeneity condition is $y(\lambda Ep^*,\lambda w^*,\lambda^2 Vp^*,\lambda W_0^*) = y(Ep^*,w^*,Vp^*,W_0^*)$ for any $\lambda>0$, and we can select either $\lambda = 1/w_i$ or $\lambda = 1/w_i^*$. In comparing market and CAIS variables, it is convenient to select $\lambda = 1/w_i$ to normalize market variables (Ep,w,Vp,W_0)

and $\lambda = 1/w_i^*$ (or any other expected price under CAIS) to normalize CAIS variables (Ep^*,w^*,Vp^*,W_0^*).

On the other hand, consider the more restrictive assumption of constant absolute risk aversion (CARA). Under expected utility maximizing behavior and CARA, the zero homogeneity condition on output supplies is $y(\lambda Ep,\lambda w,\lambda Vp) = y(Ep,w,Vp)$ for any $\lambda>0$ (Coyle 1992), and again we select $\lambda = 1/w_i$ (initial wealth does not influence decisions under CARA). Here Vp is normalized by w_i rather than by w_i^2 as under CRRA. The same homogeneity condition under CAIS is $y(\lambda Ep^*,\lambda w^*,\lambda Vp^*) = y(Ep^*,w^*,Vp^*)$ for any $\lambda>0$, and we can select $\lambda = 1/w_i^*$. Note that, in more general models of risk preferences than CRRA or CARA, there generally does not exist any appropriate homogeneity condition or normalization for Ep,w,Vp,W_0 or counterparts under CAIS (Chavas and Pope).

The main impact of CAIS on effective (Ep, w, Vp, W_0) is an approximately equi-proportional decrease in effective Ep,w (essentially because an increase in profits implies a reduction in government payouts) and a small increase in nominal wealth. This implies in effect a substantial increase in relative wealth. For example, under the homogeneity condition for CRRA and utility maximization, a 40 percent decrease in Ep,w (and decrease in Vp to $.6^2$ of initial levels) and a constant W_0 would imply the same decisions as 0 change in Ep,w (and Vp) and a 67 percent increase in W_0: $y(.6Ep,.6w,.6^2Vp,W_0) = y(Ep,w,Vp,1/.6\, W_0)$ $(\lambda=1/.6)$. Thus, in effect, CAIS has a substantial impact on relative wealth through substantial reductions in effective prices Ep,w.

Changes in relative wealth W_0/w due to changes in effective input prices w are calculated as follows. Under scenario (A) $E\theta$ = 0.3907, effective input prices decrease by 39%, so normalized wealth increases by 63%. Under scenario (B) $E\theta$ = 0.25, effective input prices decrease by 25%, so normalized wealth increases by 33%.

The impact of CAIS on relative price variances depends on the form of risk aversion, as can be seen from the different homogeneity conditions under CRRA and CARA. Assuming CRRA, our normalization Ep/w_i (and w/w_i) of expected prices implies the following normalization of price variances and covariances: $Vp / (w_i)^2$ where Vp are market price variances and covariances. Similarly our normalization Ep^*/w_i^* under CAIS implies $Vp^* / (w_i^*)^2$. Under scenario (A) $E\theta$ = 0.3907, Vp^* = .459 Vp and w^* = 0.6136 w using (25), (29) and mean data for the data set. Under scenario (B) $E\theta$ = 0.25, Vp^* = .651 Vp and w^* = .7543 w. Then under CRRA the percentage change in relative price uncertainty due to CAIS is:

$$(Vp^*/w_i^{*2} - Vp/w_i^2) / Vp/w_i^2 = .459/ .6136^2 - 1$$
$$= + 0.2191 \quad \text{(A)} \qquad (32)$$
$$= + 0.1441 \quad \text{(B)}$$

i.e. CAIS leads to a 22 or 14% increase in relative price uncertainty, assuming CRRA.

This result may seem counterintuitive, but it follows from the homogeneity/normalization condition under CRRA and utility maximization. In nominal terms Vp^* = .459 Vp (under A), which represents a reduction in uncertainty under CAIS. However CAIS also reduces effective w (w^* = .6136 w). Under the CRRA normalization Vp/w^2 and $Vp^*/(w^*)^2$ and the Manitoba data set, the net effect of CAIS is an increase in relative uncertainty Vp.

On the other hand, consider the more restrictive assumption of CARA. In this case, an appropriate normalization for price variances and covariances is Vp/w and similarly Vp^*/w^*. Then under CARA the percentage change in relative price uncertainty due to CAIS is

$$
\begin{aligned}
(Vp^*/w_i^* - Vp/w_i) / Vp/w_i &= .459/.6136 - 1 \\
&= -0.2519 \quad \text{(A)} \\
&= -0.1369 \quad \text{(B)}
\end{aligned}
\tag{32a}
$$

i.e. CAIS leads to a 25 or 14 % decrease in relative price uncertainty, as defined here assuming CARA.

The change in expected wealth due to CAIS can be evaluated as follows. An appropriate normalization for initial wealth W_0 (corresponding to our previous normalizations) is W_0/w_i under CRRA, and similarly expected wealth $EW_0^* = W_0 + 0.391\,\pi^T$ can be normalized as EW_0^* / w_i^*, where $w_i^* = .6136\, w_i$ at average τ. Then the percentage change in relative wealth due to CAIS is

$$
\begin{aligned}
&(EW_0^* / w_i^* - W_0 / w_i) / W_0/w_i \\
&\quad = \{1/.6136\,[W_0/w + .391\,\pi^T/w] - W_0/w\} / W_0/w \\
&\quad = 1/.6136\,[1 + .391\,\pi^T/W_0] - 1 \\
&\quad = 0.6297 + 0.6372\,\pi^T/W_0 \qquad \text{(A)} \\
&\quad = 0.3257 + 0.3314\,\pi_T\,/W_0 \qquad \text{(B)}
\end{aligned}
\tag{33}
$$

as defined here. Note that the difference in w^* and w (w^* = 0.6136 w or 0.75 w)) implies a substantial increase in normalized expected wealth, given this normalization. Here $w^* < w$ implies an increase in normalized expected wealth under CAIS irrespective of π^T, as indicated by the substantial term +0.6297 or +0.3257 above. This is much larger than the effect of π^T relative to W_0. This large change in relative wealth can have a substantial impact on production decisions assuming CRRA. On the other hand, under CARA decisions are independent of initial wealth irrespective of normalization, so there is no impact on decisions under CARA.

Above we specified expected wealth under CAIS as $EW_0 * = W_0 + E\theta\, \pi^T$. Denote this as scenario I. This scenario assumes that the producer perceives only current year government payments as adding to his stock of wealth. This may be a reasonable assumption if CAIS is uncertain to continue.

Consider an opposite scenario II where the producer perceives an infinite flow of possible CAIS payouts as adding to his stock of wealth. Approximate the present expected value of CAIS benefits as $E\theta\, \pi^T / r$, where r is an inter-temporal discount rate. Assuming r = 0.10, the above formula is modified as

$$
\begin{aligned}
(EW_0^* / w_i^* - W_0/w_i) / W_0/w_i &= 0.6297 + 6.372\, \pi^T / W_0 \quad \text{(A)} \\
&= 0.3257 + 3.314\, \pi^T / W_0 \quad \text{(B)}
\end{aligned}
\tag{34}
$$

The mean reference margin calculated over the data set is π^T = \$2,455.7 and the 2002 level is \$6,722.5. This is close to the maximum reference margin calculated for the data set (\$7,325 for 1979). The mean initial wealth W_0 over the data set is \$95,816.0 per farm, and the 2002 level is \$190,188.6. Reference margins are 0.0256 and 0.0353 of initial wealth for average data and 2002, respectively. Then the percentage changes in normalized expected wealth under scenario I are 0.6457 (A) and 0.3342 (B) for average data and 0.6522 (A) and 0.3373 (B) for 2002.

$\varnothing$ is calculated using mean data as above, and $\varnothing / w_i^{*2}$ is calculated using 2000 level of crop input price index w. $\varnothing / w_i^{*2}$ is then divided by the 2000 normalizations for $cov(G,p_i)$ used in econometric models. The resulting normalized level of $\varnothing / w_i^{*2}$ is 6.090 for use in wheat equations. The covariance between historical government payments and market prices, $cov(G,p_i)$ was statistically significant for wheat but not other crops in econometric models. The short and long run elasticities for wheat output with respect to $cov(G,p_i)$ were estimated as -0.0309 and -0.0503. Combining these elasticities with the normalized level of $\varnothing$, this insurance effect of CAIS on short run and long run wheat output is calculated as +0.1882 and +0.3063, respectively. However it is unclear whether this complex measure $\varnothing$ of truncated uncertainty under CAIS is comparable to the historical measure $cov(G,p_i)$ used in the econometric models. If $\varnothing$ is not adequately comparable to historical $cov(Gp_i)$, then we cannot adequately measure insurance effects of CAIS from estimates of econometric models. So we will exclude $\varnothing$ from calculations.

6.5 Hypothetical Simulation Results for CAIS in Manitoba

Long run and short run impacts of CAIS on crop production are now estimated from the earlier tables on long run and short run output elasticities and the above calculations. This assumes that CAIS can be simulated adequately by plugging calculated impacts of CAIS on relative expected prices, price uncertainty and wealth into the econometric model based on historical data. This may be reasonable in the case of a transparent program such as GRIP.

However the impacts of CAIS, as modelled here using economic theory of decision making under risk, presumably are far from transparent for a typical firm (and perhaps for a typical economist). If CAIS would remain in place for many years then perhaps firm behaviour would come to approximate the economic model, but this is unlikely to happen. Nevertheless there does not appear to be an alternative dynamic approach based on econometrics, since there is insufficient data to estimate a CAIS specific time series model.

We omit CAIS insurance related effects since these do not seem comparable to the historical situation. Effects related to investment in future reference margins are also omitted, since firms presumably view CAIS as transitory. Including these effects would presumably increase the simulated impacts of CAIS on production.

In the above discussion of impacts of CAIS on variables (Ep, w, Vp, W_0), we emphasized impacts of CAIS on (in effect) the normalization ($Ep/w_i, w/w_i, Vp/(w_i)^2, W/w_i$) consistent with CRRA and utility maximization. The econometric models were also specified and estimated imposing this normalization. So there is no contradiction in applying the theoretical analysis to the estimated econometric model.

We first consider impacts of CAIS based on estimates of the crop acreage allocation model conditional on yields. The econometric model specified acreage share equations for wheat, barley, canola and a fourth aggregate crop (primarily oats, also flax and rye), so share equations were estimated for the first three crops. Market price covariances *Vp* were calculated for the 1966-2002 Manitoba data. Using this *Vp* and average output price and quantity data over the period, CAIS would happen to have essentially zero effect on relative output prices p_i/p_j . In this case CAIS would not have a relative price effect on acreage shares (conditional on yield).

The impact of CAIS on price uncertainty varies here with normalization appropriate under CRRA (Vp/w^2) and CARA (Vp/w). Results also vary with the scenario for expected government share of payments: (A) $E\theta$ = .3907 using Manitoba data and (B) $E\theta$ = .25. Given these different normalizations, relative price uncertainty is calculated to decrease by 25% (A) or 14% (B) under CARA but to increase by 22% (A) or 14% (B) under CRRA. CRRA is considered a less restrictive assumption than CARA, and the CRRA hypothesis was not rejected for the acreage demand model. Due to relatively small elasticities of response to uncertainty for wheat and barley (-0.0334, -0.0411), this difference between CRRA and CARA scenarios has little impact on acreage shares. However, estimated elasticities are larger for canola (a sum of -.1416), so these two scenarios have a larger impact on canola acres.

Under CRRA, the appropriate normalization for wealth (corresponding to the above normalizations) is W_0/w and (essentially) $W_0{}^*/w^*$. Although CAIS has a relatively small impact on nominal wealth W_0, it leads to a large reduction in shadow prices for inputs. This implies a surprisingly large increase in relative wealth: 64.6% (Eθ = 0.3907) or 33.4% ($E\theta$ = 0.25). Initial wealth does not influence decisions under CARA. In the econometric model under CRRA, the estimated elasticity of wheat acres share with respect to wealth is -0.3388 (scenario A), and wealth is statistically insignificant in equations for barley and canola. This estimated elasticity for wheat seems high, so we also consider an alternative elasticity -.15 (scenario B).

Simulations of hypothetical impacts of CAIS on wheat, barley and canola acreage demands (conditional on yields and total crop acres) under average Manitoba data over 1966-2002 are presented in the following table. Here scenario A is $E\theta$ = 0.3907 and econometric estimates of wealth effects. Scenario B is $E\theta$ = 0.25 and elasticity of wheat acres share with respect to wealth is -0.15. This table presents relative changes in acres in decimal form. For example, the number +0.0084 in the table corresponds to a +0.84% increase in crop acres (for wheat). Except for wheat acres under CRRA, hypothetical impacts of CAIS are relatively small. Since the sum of shares is identically 1, in principle off setting impacts are implied for the fourth crop, which is primarily oats.

Table 15 Short Run Impacts of CAIS (average data)

	Wheat Acres		Barley Acres		Canola Acres	
	(A)	(B)	(A)	(B)	(A)	(B)
CRRA	-0.2262	-0.0549	-0.009	-0.0059	-0.031	-0.0204
CARA	0.0084	-0.0455	0.0103	0.0056	0.0357	0.0194

Next we consider hypothetical long run impacts of CAIS on crop yields. As in the acreage demands, these calculations are based on average data for the 1966-2002 Manitoba data set. CAIS leads to 8-9% increases in relative expected prices *Ep/w* assuming (A) $E\theta$ = 0.3907, and to 7% increases assuming (B) $E\theta$ =0 .25. Long run elasticities of yield with respect to price are estimated as +0.5306, +0.5310, +0.3203 for wheat, barley, canola, respectively. These lead to 4.56%, 4.73% and 2.79% increases in yield for wheat, barley and canola in scenario A and to 3.71%, 3.86% and 2.26% increases in yield in scenario B. The effects of price risk are quite small for wheat and canola, and somewhat larger for barley.

Under CRRA, CAIS has a substantial impact on yields through changes in relative wealth. Although CAIS has only a minor effect on nominal wealth, a large reduction in effective prices implies a large increase in normalized wealth: 65% under scenario A and 33% under scenario B. Moreover the estimated elasticities of wheat and barley yields with respect to wealth under CRRA are substantial: +0.4322 and +0.4395. This implies a 28% increase in wheat and barley yields through wealth effects (scenario A). There are no wealth effects under CARA.

There are no comparable studies estimating dynamic yield response to wealth, but we suspect that this study overestimates elasticity of wheat and barley yields with respect to wealth. An elasticity of +0.20 lies within the 95% confidence interval about these estimates, and this may well be more realistic than the reported estimates. This alternative elasticity implies a 13% increase in long run wheat and barley yield through wealth effects under CRRA (scenario B).

Hypothetical long run impacts of CAIS on yields are presented in the following table. Scenario A is $E\theta$ = 0.3907 and the econometric estimates of wheat and barley yield equations. Scenario B is $E\theta$ = 0.25 and elasticity of wheat and barley yields are +0.20. Results are in decimal form.

Table 16 Long Run Impacts of CAIS on Yields (average data)

		Price	Price Risk	Wealth	Total
Wheat	CRRA A	+0.0456	-0.0099	+0.0668	+0.0974
	CRRA B	+0.0371	-0.0065	+0.0668	+0.0974
	CARA A	+0.0456	+0.0114	-	+0.0570
	CARA B	+0.0371	+0.0062	-	+0.0433
Barley	CRRA A	+0.0473	-0.0299	+0.2839	+0.3013
	CRRA B	+0.0386	-0.019	+0.0668	+0.0858
	CARA A	+0.0473	+0.0343	-	+0.0816
	CARA B	+0.0386	+0.0187	-	+0.0573
Canola	CRRA A	+0.0279	-0.0114	-	+0.0165
	CRRA B	+0.0226	-0.0075	-	+0.0151
	CARA A	+0.0279	+0.0131	-	+0.0410
	CARA B	+0.0226	+0.0071	-	+0.0297

Induced long run changes in yields also have a feedback effect on acreage demands. These feedback effects of CAIS on acreages are indicated in the following table. These secondary impacts of CAIS on acreage demands are obtained by multiplying impacts in the above table by short run elasticities of acreage demand with respect to *Er* (expected revenues per acre) (or equivalently by differences between long run and short run acreage responses in an earlier table). These effects are generally quite small. The exception is wealth effects on wheat and barley acreages under CRRA, scenario A, where impacts are +5.4%.

Table 17 Acreage Feedback Effects of CAIS Impacts on Long Run Yields

		Price	Price Risk	Wealth	Total
Wheat	CRRA A	+0.0078	-0.0017	+0.0475	+0.0536
	CRRA B	+0.0063	-0.0011	+0.0114	+0.0166
	CARA A	+0.0078	+0.0019	-	+0.0097
	CARA B	+0.0063	+0.0010	-	+0.0073
Barley	CRRA A	+0.0081	-0.0017	+0.0475	+0.0539
	CRRA B	+0.0066	-0.0033	+0.0114	+0.0147
	CARA A	+0.0081	+0.0019	-	+0.0100
	CARA B	+0.0066	+0.0032	-	+0.0098
Canola	CRRA A	-0.0093	+0.0033	-	-0.0060
	CRRA B	-0.0075	+0.0025	-	-0.0050
	CARA A	-0.0093	-0.0038	-	-0.0131
	CARA B	-0.0075	-0.0024	-	-0.0099

Hypothetical long run impacts of CAIS on wheat, barley and canola output are presented in the following table. This is the sum of impacts in the previous three tables. Scenario A is $E\theta$ = 0.3907 and econometric estimates of yield and acreage equations. Scenario B is $E\theta$ = .25, elasticity of wheat and barley yields with respect to initial wealth is +0.20, and elasticity of wheat acres with respect to wealth is -0.15. Scenario B may be more realistic. Results are in decimal form. In scenario B under CRRA, CAIS is calculated to have the following impacts on long run output: +5.91% for wheat, +9.46% for barley, and -1.03% for canola. In scenario B under CARA, impacts on long run output are +0.51% for wheat, +7.27% for barley, and +0.04 % for canola.

Table 18 Long Run Impacts of CAIS (average data)

	Wheat Acres		Barley Acres		Canola Acres	
	(A)	(B)	(A)	(B)	(A)	(B)
CRRA	+0.1423	+0.0591	+0.3462	+0.0946	-0.0205	-0.0103
CARA	+0.0751	+0.0051	+0.1019	+0.0727	+0.0636	+0.0004

7.0 Extensions to Revenue Insurance Programs

The theoretical analysis developed here presumably has implications beyond CAIS. In the U.S., farmers have been encouraged to rely more on subsidized insurance programs and less on direct government subsidies. Revenue insurance rather than income insurance programs are common. The above theoretical analysis can easily be extended to revenue insurance programs, provided that an increase in current revenue for the individual farm can lead to a reduction in payouts to the farm, as is usually the case.

First assume that output levels y are non-stochastic, as in the above CAIS model. Let $R \equiv p\ y$ and $C(w,y)$ denote revenue and cost for the firm, and R^T is the insured level of revenue. Premiums can be viewed as embedded in the cost function. Assume a whole farm insurance program (the model can also easily accommodate crop specific programs, which are more common). So profits plus insurance payments are $\psi = R - C + \theta\,(R^T\text{-}R) = (1\text{-}\theta)\cdot R - C + \theta\, R^T$, where θ is the net share of payouts received by the firm. Most simply, $\theta = 1$ *if* $R^T\text{-}R > 0$ and $\theta = 0$ *if* $R^T\text{-}R < 0$. Since R is stochastic, θ is also stochastic. Then taking expectations $E\psi = (1\text{-}E\theta)\ ER - C - cov(\theta,R) + E\theta\ R^T$ where $ER = Ep\ y$ assuming y is non-stochastic. Then $\partial E\psi/\partial y = (1\text{-}E\theta)\ Ep - \partial C(w,y)/\partial y - \partial E\theta/\partial y\ \ (R^T\text{-}R) - \partial cov(\theta,R)/\partial y$. Under the insurance program, the expected shadow price for output *is* $Ep_i^* = (1\text{-}E\theta)\ Ep_i - \partial E\theta/\partial y_i\ (R^T\text{-}R) - \partial cov(\theta,R)/\partial y_i$ whereas the shadow price for inputs is not influenced by the program: $w^* = w$. Here it is quite possible that $Ep^*/w^* < Ep/w$, in contrast to CAIS. The analysis of components of Ep^* is similar to CAIS. Thus such revenue insurance programs may have very different impacts on production than does CAIS, and may well lead to decreases in production. Apparently the effects of revenue insurance programs on shadow prices under risk have not been adequately recognized in the directly related literature (e.g. Babcock and Hennessey; Hennessey, Babcock and Hayes; Wu and Adams).

Second, assume that output levels y are stochastic and (for simplicity) output prices p are non-stochastic. The cost function can be defined as $C(w,m_y,m_\omega) = min_x\ wx$ *s.t.* $T(x,m_y,m_\omega) = 0$ where m_y are moments of output distribution and m_ω are moments of weather distribution (e.g. Pope and Chavas). Assume $C = C(w,m_y,m_\omega) + e^c$ where $Ee^c = 0$ (for simplicity). Then $E\psi = (1\text{-}E\theta)\ \ ER - EC - cov(\theta,R) + E\theta\ R^T$ where $ER = p\ Ey,\ EC = C(w,m_y,m_\omega)$, and in turn $\partial E\psi/\partial Ey = (1\text{-}E\theta)\ p - \partial C(w,m_y,m_\theta)/\partial Ey - \partial E\theta/\partial Ey\ (R^T\text{-}R) - \partial cov(\theta,R)/\partial Ey$. The expected shadow price for output (Ey_i) *is* $Ep_i^* = (1\text{-}E\theta)\ p_i - \partial E\theta/\partial Ey_i\ (R^T\text{-}R) - \partial cov(\theta,R)/\partial Ey_i$ whereas the shadow price for inputs is not changed by the program: $w^* = w$. So again it is quite possible that $Ep^*/w^* < Ep/w$, in contrast to CAIS.

Calculation of components of Ep^* is more complex when y is stochastic. Assume a linear relation between θ and R: $\theta = \alpha_0 + \alpha_1\ R + \alpha_2 Z + e\ (Ee=0)$, so $\partial E\theta/\partial Ey_i = \alpha_1\, p_i$. Then $cov(\theta,R) = \alpha_1\ \ varR$, where $varR = p^T\ Vy\ p$. So $\partial cov(\theta,R)/\partial Ey_i = \alpha_1\ \partial varR/\partial Ey_i = \alpha_1\ \Sigma_j\ \partial cov(y_i y_j)/\partial Ey_i\ p_i\ p_j$, and α_1 can be calculated as $\alpha_1 = cov(\theta,R)\ /\ varR$. Calculation of

$\partial cov(\theta,R)/\partial Ey_i$ apparently requires some econometrics. A multioutput stochastic transformation function $T(x,m_y,m_\omega) = 0$ (e.g., a multioutput generalization of a Just-Pope production function) can be estimated. Then perhaps all $\partial varR/\partial Ey_i$ can be calculated as shadow prices for the minimization problem min $_x p^T\ Vy\ p$ s.t. $T(x,Ey,Vy,.,m_\omega) = 0,\ Ey_i = Ey_i^*\ i=1,.,N$ using the estimated function $T(.)$ and reference p, Ey.

In crop insurance programs, payments for yield losses typically are conditional on acreage enrolments. Let x_{LA} denote acreages enrolled in CI, and let x_{LB} denote other acreages. Let yld_I denote insured levels of yield and p_A corresponding prices, and ρ is premium per acre. Define the cost relation $C = C(w,m_y,m_\omega,x_{LA},x_{LB}) + e^C$ where w excludes cropland market prices w_L. Then $\psi = p\ (y_A+y_B) - C - w_L\ (x_{LA}+x_{LB}) - \rho\ x_{LA} + \theta\ p_A\ (yld_I - y_A/x_{LA})\ x_{LA}$. Taking expectations and differentiating with respect to Ey_A and x_{LA} yields shadow expected prices for output and land under CI. These differ from market prices and imply that CI may decrease output. Early theoretical literature argued that CI would increase crop output, but more recent literature has recognized the ambiguity (Ramaswami; Chambers and Quiggin). Empirical studies suggest that U.S. CI may increase acreages and reduce inputs per acre, but apparently there are no empirical studies of net impacts on output (Glauber).

8.0 CONCLUSION

Econometric models of crop supply response typically estimate crop acreage demands. Such models presumably are adequate for measuring short run production response, but they are inadequate for the long run if there are significant long run yield responses to price.

The first part of this study estimates a dynamic econometric model of crop yield response to price for Manitoba over 1960- 2002. Previous studies of crop yield response to price have ignored dynamics and risk aversion, and generally have not obtained statistically significant results. Here crop yields are specified as distributed lags of expected output price, output price variance, initial wealth and program insurance effects. These are normalized by a crop input price index as implied by constant relative risk aversion (CRRA) and mean variance or expected utility maximization. The model also includes measures of average weather conditions and heterogeneity of weather across Manitoba. Estimates of yield response to price are statistically significant, especially for the major crop, wheat. There are no comparable studies estimating dynamic yield response to economic variables, but we suspect that this study overestimates wheat and barley yield response to wealth. So a lower estimate is emphasized in simulating CAIS.

A static econometric model of Manitoba crop acreage demands conditional on yields is also estimated. Crop acreages depend on expected revenues per acre, revenue variances and covariances per acre (assuming only price uncertainty), initial wealth, program insurance effects, and a variance of weather indexes serving as a proxy for weather/yield uncertainty. The model adopts a normalization implied by CRRA, and the CRRA hypothesis is not rejected.

A comparison of results for the two models indicates that in this case long run yield responses to price appear quite significant relative to short run acreage responses. This suggests that acreage demand models may well underestimate long run crop production response to price.

The second part of this study develops a simple analytical model of crop production response to CAIS under risk aversion and output price uncertainty. Assuming CAIS has negligible impacts on domestic consumption impacts on production may approximate impacts on trade, i.e. trade distorting impacts. Hypothetical impacts of CAIS are simulated using an aggregate Manitoba data set over 1966-2002, which is prior to CAIS.

Since increases in current income reduce current payouts under CAIS, CAIS has a substantial impact on shadow prices for outputs and inputs. Over the data set, effective prices for outputs increase by approximately 8% relative to effective input prices. Using estimates of the econometric model, this implies 4-5% increases in long run yields for wheat and barley and 3% for canola. CAIS has a small impact on nominal wealth. Nevertheless impacts on production depend critically upon properties of risk preferences. Under the production homogeneity condition implied by CRRA, the substantial change in shadow prices relative to nominal wealth implies, in effect, a substantial increase in relative wealth and in turn a substantial impact on production.

Under more plausible assumptions, our simulation of hypothetical long run impacts of CAIS on annual crop production over our 1966-2002 Manitoba data set are as follows. Under CRRA: +6 % for wheat, +9 % for barley, and -1 % for canola. Under CARA: +5 % for wheat, +7% for barley, and 0% for canola. These calculations primarily reflect impacts on long run yields, of increases in shadow prices of outputs, relative to inputs and (in the case of CRRA) increases in normalized wealth. Given estimates of the yield equation for canola, these impacts are negligible for canola. These impacts for wheat and barley are less substantial than simulated impacts for the historical Gross Revenue Insurance Program, which provided large and transparent subsidies to crop output prices.

Two important caveats to these results are in order. First, this study plugs calculated hypothetical impacts of CAIS on relative expected prices, price uncertainty and wealth into an econometric model based on historical data. This is less reasonable for a complex program such as CAIS than for a transparent program such as GRIP. Nevertheless there does not appear to be an alternative dynamic approach based on econometrics, since there is insufficient data to estimate a CAIS specific time series model.

Second, firm behaviour under CAIS is analyzed using economic theory of decision making under risk. However this standard theory may be too complex to approximate behaviour under CAIS, especially if the program is transitory. This is suggested by various counterintuitive results that we have derived under CRRA and utility maximization.

Given these caveats, and the use of a synthetic data base, these results certainly do not imply that CAIS has significant trade distorting impacts on Canadian crop production. On the other hand, the analysis should further raise the possibility that CAIS is not a decoupled farm program.

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ECONOMETRIC RESULTS FOR MANITOBA

Appendix A

ADL Yield Models

Table A - 1. ADL Wheat Yield OLS

	lag	A Coef	A t-stat	B coef	B t-stat
YLD1	1	-.3310	2.35	-.3078	2.07
Ep1	0	-.0844	0.74	-.0308	0.22
Ep1	1	.1307	2.18	.0639	0.51
Ep1	2	.2358	3.46	.2498	1.94
Ep1	3	.2309	3.78	.2815	2.16
Ep1	4	.1160	1.11	.0742	0.59
T		.0164	3.53	.0163	3.41
EW1		.4212	3.01	.4437	3.03
EW2		.7141	4.50	.7370	4.46
EW3		.6510	3.34	.6788	3.35
VW1		.0125	1.21	.0126	1.19
VW2		-.0680	2.68	-.0665	2.55
VW3		.3110	2.02	.3133	1.98
Constant		-4.1933	1.13	-4.2523	3.66
Sum of lag coefficients					
Ep1		.6290	2.70	.6385	2.64
R^2		.8096		.8139	
rho		.068		.089	
F-test of polynomial restrictions					
		F(2,28) = .3216 prob = .7276		no restrictions	

Table A - 2. ADL Wheat Yield

		A	OLS	B	auto
	lag	Coef	t-ratio	coef	t-ratio
Ep1	0	.1219	1.23	.0829	1.18
Ep1	1	.2497	3.20	.2412	4.68
Ep1	2	.2525	2.99	.2569	4.75
Ep1	3	.1303	1.92	.1302	3.15
Ep1	4	-.1168	1.04	-.1391	1.72
Vp1	0	-.0554	2.72	-.0592	4.00
Vp1	1	-.0140	1.44	-.0158	2.50
Vp1	2	.0022	0.20	.0013	0.18
Vp1	3	-.0068	0.72	-.0078	1.26
Vp1	4	-.0409	2.09	-.0431	3.10
W_o	0	.4248	3.47	.4004	4.20
W_o	1	.1762	3.49	.1663	4.27
W_o	2	.0240	2.10	.0229	2.84
W_o	3	-.0317	2.56	-.0300	3.11
W_o	4	.0091	1.04	.0077	1.33
t		.0052	0.56	.0091	1.63
EW1		.5337	3.21	.5218	4.25
EW2		.5825	3.41	.5274	4.43
EW3		.6488	3.32	.5836	4.17
VW1		-.0002	0.02	.0012	0.14
VW2		-.0715	2.99	-.0764	4.61
VW3		..0926	0.55	-.0085	0.07
Constant		-11.826	4.63	-11.012	5.50

(continued)

Table A - 2. DL Wheat Yield (continued)

	A	OLS	B	auto
	Coef	t-ratio	coef	t-ratio
Sum of lag coefficients				
Ep1	.6376	2.27	.5722	3.16
Vp1	-.1150	2.45	-.1246	3.70
W_o	.6025	3.58	.5673	4.36
R^2	.8763		.8879	
rho	-.238		-.084	
DW	2.46		2.16	
F-test of polynomial restrictions				
	F(6,19) = 1.09		no restrictions	
	prob = .398			

Table A - 3. ADL Wheat Yield OLS

		A		B	
	lag	Coef	t-stat	coef	t-stat
Ep1	0	-.0241	0.21	--	
Ep1	1	.1058	1.52	--	
Ep1	2	.1292	1.49	--	
Ep1	3	.0461	0.69	--	
Ep1	4	-.1436	1.13	--	
Vp1	0	-.0442	2.25	--	
Vp1	1	-.0001	0.01	--	
Vp1	2	.0186	1.42	--	
Vp1	3	.0121	1.16	--	
Vp1	4	-.0198	0.93	--	
W_o	0	.2856	3.17	--	
W_o	1	.1368	3.70	--	
W_o	2	.0384	3.35	--	
W_o	3	-.0094	0.86	--	
W_o	4	-.0069	0.63	--	
t		--		.0083	1.75
EW1		--		.5015	3.18
EW2		--		.7554	4.10
EW3		--		.6819	3.06
VW1		--		.0169	1.41
VW2		--		-.0530	1.88
VW3		--		.1362	0.86
constant		-6.9194	2.98	-1.1020	5.99

(continued)

Table A - 3. ADL Wheat Yield OLS (continued)

	A		B	
lag	Coef	t-stat	coef	t-stat
Sum of lag coefficients:				
Ep1	.1134	0.60	--	
Vp1	-.0335	0.89	--	
W_o	.4446	3.71	--	
R^2	.6847		.6978	
rho	-.102		-.059	
DW	2.19		1.94	
F-test of polynomial restrictions:				
	F(6,26) = 2.218			
	prob = .073			

Table A-4. DL Wheat Yield

		A OLS		B auto	
	lag	Coef	t-ratio	coef	t-ratio
Ep1	0	.0716	0.72	.0282	0.50
Ep1	1	.2129	2.71	.2011	4.96
Ep1	2	.2260	2.67	.2285	5.51
Ep1	3	.1109	1.70	.1105	3.83
Ep1	4	-.1324	1.17	-.1529	2.31
Vp1	0	-.0495	2.50	-.0515	4.53
Vp1	1	-.0091	0.89	-.0137	2.78
Vp1	2	.0091	0.81	.0052	1.02
Vp1	3	.0051	0.50	.0053	1.03
Vp1	4	-.0210	0.91	-.0134	0.98
W_o	0	.4501	3.42	.3256	3.41
W_o	1	.1866	3.42	.1335	3.39
W_o	2	.0259	2.24	.0170	2.47
W_o	3	-.0320	2.56	-.0238	2.74
W_o	4	.0129	1.50	.0109	2.56
t		-.0111	0.95	-.0007	0.13
CovGP1	0	-.0002	0.83	-.0002	1.63
CovGP1	1	-.0002	2.24	-.0002	4.55
CovGP1	2	-.0001	1.30	-.0001	2.10
CovGP1	3	-.0000	0.02	.0001	0.83
CovGP1	4	.0000	0.22	.0001	1.60
CovGP1	5	-.0001	0.54	.0000	0.61
CovGP1	6	-.0002	1.90	-.0001	1.90
CovGP1	7	-.0003	2.46	-.0002	3.86
CovGP1	8	-.0001	0.47	-.0002	1.26

(continued)

Table A-4. DL Wheat Yield (continued)

	AOLS		B auto	
lag	Coef	t-ratio	coef	t-ratio
EW1	.6068	3.61	.6759	6.51
EW2	.7574	3.96	.8132	7.25
EW3	.8571	4.07	.8290	6.84
VW1	.0018	0.19	.0116	1.66
VW2	-.0577	2.01	-.0616	3.38
VW3	.3041	1.57	.2034	1.57
Constant	-11.845	4.28	-9.1922	4.43
Sum of lag coefficients:				
Ep1	.4891	1.77	.4153	2.95
Vp1	-.0654	1.32	-.0682	2.38
W_o	.6437	3.53	.4632	3.51
CovGp1	-.0012	1.89	-.0009	2.89
R^2	.9132		.9376	
rho	-.402		-.204	
DW	2.78		2.38	
F-test of polynomial restrictions:				
	F(10,10) = .925			
	prob = .548			

Table A - 5. DL Wheat Yield

		A OLS		B Auto	
	lag	Coef	t-stat	coef	t-stat
Ep1	0	.0686	0.63	-.0045	0.07
Ep1	1	.1768	2.14	.1278	2.79
Ep1	2	.2851	2.31	.2601	3.62
VP1	0	-.0407	1.71	-.0434	2.93
VP1	1	-.0044	0.15	.0009	0.05
W_o	0	.2923	2.76	.1730	2.75
W_o	1	.1262	2.80	.0763	2.78
W_o	2	.0233	1.90	.0160	1.98
W_o	3	-.0165	1.57	-.0077	1.25
W_o	4	.0069	0.75	.0050	0.91
t		-.0013	0.15	.0047	1.02
Cov Gp1	0	-.0309	1.38	-.0304	2.01
Cov Gp1	1	-.0014	0.11	-.0105	1.32
Cov Gp1	2	.0022	0.15	.0008	0.09
Cov Gp1	3	-.0202	1.00	.0038	0.29
EW1		.4140	2.22	.3643	3.08
EW2		.5603	2.90	.5382	4.53
EW3		.6784	3.07	.6365	4.41
VW1		.0049	0.48	.0090	1.17
VW2		-.0679	2.33	-.0918	4.45
VW3		.1656	0.94	-.0418	0.36
Constant		-9.1030	3.75	-6.0877	4.06

Continued

Table A - 5. DL Wheat Yield (Continued) OLS

Sum of lag coefficients:	Coef	t-stat	coef	t-stat
Ep1	.5306	2.14	.3834	2.79
VP1	-.0451	1.54	-.0426	2.66
W_o	.4322	2.92	.2626	2.95
Cov Gp1	-.0503	1.80	-.0363	2.43
R^2	.8592		.8881	
rho	-.207		-.104	
DW	2.40		2.16	
F-test of polynomial restrictions:				
				F (4,20) = 2.566
				Prob = .072

Table A - 6. ADL Barley Yield OLS

		A		B	
	lag	coef	t-stat	coef	t-stat
YLD 2	1	.0294	0.19	.0768	0.48
Ep2	0	.0765	0.69	.1760	1.26
Ep2	1	.0767	1.36	-.0721	0.50
Ep2	2	.0861	1.21	.1498	1.07
Ep2	3	.1049	1.70	.1490	1.03
Ep2	4	.1330	1.21	.0771	0.57
t		.0215	3.39	.0206	3.20
EW1		.5758	3.51	.6065	3.61
EW2		.7873	4.07	.8276	4.17
EW3		.7685	3.15	.7814	3.17
VW1		.0124	0.97	.0135	1.04
VW2		-.0208	0.71	-.0231	0.77
VW3		.1192	0.68	.0152	0.84
Constant		-3.2493	3.14	-3.2743	3.13
Sum of lag coefficients:					
Ep2		.4772	2.06	.4798	2.05
R^2		.8148		.8239	
rho		-.003		.022	
F - test of polynomial restrictions:					
		F (2,28) = .727			no restrictions
		Prob = .492			

Table A - 7. DL Barley Yield

	lag	A OLS coef	A OLS t-stat	B Auto coef	B Auto t-stat
Ep2	0	..1112	1.14	.1029	1.41
	1	.1056	1.44	.1023	1.94
	2	.1031	1.13	.1029	1.56
	3	.1037	1.37	.1046	1.96
	4	.1075	0.97	.1076	1.31
VP2	0	-.0436	1.75	-.0473	2.58
	1	-.0138	0.96	-.0141	1.36
	2	-.0056	0.44	-.0050	0.55
	3	-.0191	1.45	-.0199	2.10
	4	-.0541	2.16	-.0588	3.12
W_o	0	.2757	2.30	.2534	2.80
	1	.1322	2.58	.1247	3.20
	2	.0383	2.79	.0390	3.72
	3	-.0061	0.53	.0035	0.41
	4	-.0007	0.07	.0030	0.40
t		.0154	1.45	.0181	2.49
EW1		.3473	1.69	.3152	1.94
EW2		.4645	1.99	.4156	2.27
EW3		.3576	1.24	.2981	1.30
VW1		.0006	0.05	.0004	0.04
VW2		-.0438	1.71	-.0500	2.65
VW		.1221	0.82	.0682	0.59
Constant		-8.5936	3.60	-8.0972	4.38
Sum of lag coefficients:					
Ep2		.5310	1.80	.5202	2.44
Vp2		-.1363	1.87	-.1451	2.71
W_o		.4395	2.61	.4105	3.22
R^2		.8982		.9013	
rho		-.136		-.016	
DW		2.27		2.03	
F - test of polynomial restrictions:					
		F (6,19) = 1.790			
		Prob = .155			

Table A-8. DL Canola Yield

		A	OLS	B	auto
	lag	coef	t-stat	coef	t-stat
Ep3	3	.3203	2.82	.3143	3.67
Vp3	4	-.0520	3.29	-.0541	4.82
W_o	0	.0533	0.54	.0263	0.35
W_o	1	.0233	0.55	.0123	0.37
W_o	2	.0039	0.34	.0026	0.28
W_o	3	-.0049	0.54	-.0029	0.42
W_o	4	-.0032	0.37	-.0041	0.64
EW1		.3240	2.25	.3247	2.93
EW2		.4862	2.88	.4821	3.69
EW3		.5244	2.64	.5254	3.32
VW1		.0072	0.73	.0088	1.07
VW2		-.0670	2.91	-.0773	4.36
VW3		-.0214	0.17	-.0916	0.93
t		.0103	1.66	.0129	2.90
Constant		-3.7550	2.07	-3.2104	2.34
Sum of lag coefficients					
W_o		.0724	0.52	.0342	0.32
R^2		.8516		.8579	
rho		-.157		-.054	
DW		2.31		2.10	
F-test of polynomial restrictions					
		F(2,27) = 1.172			
		prob = .325			

Table A - 9. DL Oats Yield OLS

	lag	A coef	A t-stat	B coef	B t-stat
Ep4	0	-.0211	0.26	-.0429	0.41
Ep4	1	-.0531	0.80	.0353	0.26
Ep4	2	.1326	0.21	-.0835	0.56
Ep4	3	.0578	0.82	.1838	1.18
Ep4	4	.0468	0.62	-.0281	0.19
Ep4	5	.0014	0.02	.0575	0.36
Ep4	6	-.0325	0.52	.0365	0.22
Ep4	7	-.0154	0.28	-.0488	0.31
Ep4	8	.0557	1.40	.0079	0.05
Ep4	9	.1159	3.13	.1461	2.93
Ep4	10	.0024	0.08	-.0085	0.27
t		.0066	1.46	.0054	1.10
EW1		.3963	2.55	.3455	1.84
EW2		.6757	3.91	.6439	3.16
EW3		.7239	3.18	.7387	2.62
VW1		.0168	1.63	.0129	1.12
VW2		-.0511	2.00	-.0470	1.66
VW3		.1012	0.64	.1104	0.60
Constant		-1.6607	0.90	-1.9453	0.94
Sum of lag coefficients:					
Ep4		.1712	0.40	.2553	0.53
R^2		.8562		.8675	
rho		-.063		-.017	
DW		2.12		2.03	
F - test of polynomial restrictions:					
		F (5,23) = .3927 Prob =.849		no restrictions	

Table A-10.

DL OatsYield

	lag	A OLS coef	A OLS t-stat	B auto coef	B auto t-stat
Ep4	9	.0965	2.47	.0742	2.72
t		.0048	1.00	.0071	2.39
EW1		.0362	0.23	.1380	1.18
EW2		.3276	1.81	.4063	3.05
EW3		.3797	1.72	.5148	3.00
VW1		.0141	1.16	.0165	1.64
VW2		-.0794	2.84	-.0815	4.08
VW3		.0347	0.22	-.0822	0.65
Constant		-.8992	7.65	-.9520	11.41
R^2		.7555		.7831	
rho		-.2799		.059	
DW		2.40		1.65	

Appendix B.

Crop Acreage Allocation Models

Table B - 1. Wheat Acres Demand

	A		B		C	
	coef	t-stat	coef	t-stat	coef	t-stat
Er1	.1051	2.04	.1313	1.90	.1366	2.48
Er2	.0866	1.85	.1183	1.93	.1301	2.61
Er3	-.0702	2.13	-.0950	2.14	-.0552	1.56
Er4	.0409	1.60	-.0340	1.09	.0219	0.83
Vr1	-.0156	2.35	------		-.0209	2.96
Vr2	.0001	2.32	------		.0001	4.08
W1	.2582	4.07	.0263	0.53	.2375	3.50
Wo	-.1577	3.65	------		-.1581	3.40
TSVW1	.0009	0.10	------		------	
TSVW2	.0156	2.67	------		------	
TSVW3	-.0168	0.61	------		------	
t	.0025	1.07	.0018	0.82	.0034	1.75
DLIFT	-.1673	4.67	-.1853	3.57	-.1664	4.18
Constant	2.1707	3.75	-.2300	0.90	1.7928	3.05
R^2	.8677		.6513		.8170	
rho	.077		.500		.094	
DW	1.83		0.95		1.79	

Table B - 2. Barley Acres Demand
(first differences)

	A OLS		B auto		C OLS		D OLS	
	coef	t-stat	Coef	t-stat	coef	t-stat	coef	t-stat
Er1	-.0517	1.49	-.0542	2.22	.0370	1.19	-.0472	1.44
Er2	.1024	2.57	.1246	4.14	.0862	2.78	.1050	3.13
Er3	-.0291	1.29	-.0236	1.51	-.0305	1.49	-.0279	1.29
Vr2	.0001	0.02	.0007	0.19	------		-.0003	0.08
Vr4	-.0068	1.30	-.0072	1.94	------		-.0065	1.36
Z	-.1692	1.98	-.1293	2.06	-.0987	1.41	-.1541	1.91
Wo	.0065	0.30	.0007	0.04	------		.0017	0.08
TSVW1	-.0070	0.81	-.0023	0.31	------		------	
TSVW2	.0027	0.54	.0032	0.81	------		------	
TSVW3	.0063	0.23	-.0146	0.72	------		------	
t	-.0003	0.77	-.0001	0.13	-.0003	0.76	-.0003	0.74
DLIFT	.0454	2.12	.0395	2.75	.0400	2.05	.0417	2.04
Constant	.0062		.0005	0.05	.0050	0.58	.0058	0.62
R^2	.3800		.4253		.3079		.3556	
rho	.212		-.064		.268		.265	
DW	1.45		1.99		1.39		1.35	

Table B - 3. Canola Acres Demand
OLS

	A		B		C	
	coef	t-stat	coef	t-stat	coef	t-stat
Er1	.0312	1.12	.0157	0.52	.0132	0.55
Er2	-.1339	6.01	-.1735	6.44	-.1402	6.15
Er3	.0993	6.67	.0901	4.58	.1081	7.00
Er4	.0191	1.53	.0133	0.92	.0270	2.16
Vr2	-.0115	4.49	------		-.0117	4.80
Vr3	-.0052	2.43	------		-.0063	2.92
W1	-.0980	2.02	-.0108	0.25	-.0579	1.56
W2	.0866	2.34	.0168	0.43	.0564	1.84
Wo	-.0184	0.91	------		-.0140	0.66
TSVW1	.0049	1.04	------		------	
TSVW2	-.0056	2.49	------		------	
TSVW3	-.0089	0.66	------		------	
t	.0071	6.43	.0057	4.87	.0061	6.81
DLIFT	-.0124	0.80	-.0191	0.87	-.0063	0.38
Constant	-.0656	0.21	.1772	1.10	.0888	0.35
R^2	.9848		.9602		.9803	
rho	-.036		.155		-.049	
DW	2.02		1.68		2.05	

Table B-4. Acres Demand SUR

	Wheat		Barley (first differences)		Canola	
	coef	t-stat	coef	t-stat	coef	t-stat
Er1	.0931	1.89	-.0466	1.39	.0257	.096
Er2	.0877	1.94	.0973	2.54	-.1321	6.18
Er3	-.0804	2.52	-.0267	1.24	.0973	6.80
Er4	.0227	0.93	--		.0134	1.12
Vr1	-.0158	2.50	--		--	
Vr2	--		.0019	0.42	-.0119	4.82
Vr3	--		--		-.0054	2.68
Vr4	--		-.0087	1.79	--	
Vr12	.0001	2.35	--		--	
W1	.2138	3.51	--		-.1081	2.33
W2	--		--		.0831	2.35
Z	--		-.2318	2.93	--	
Wo	-.1229	3.00	.0045	.022	-.0078	0.41
TSVW1	.0025	0.30	-.0042	0.50	.0053	1.18
TSVW2	.0150	2.67	.0022	0.44	-.0056	2.62
TSVW3	-.0223	.085	.0055	0.21	-.0108	0.84
t	.0026	1.15	-.0003	0.78	.0072	6.82
DLIFT	-.1624	4.63	.0546	2.63	-.0101	0.67
Constant	1.7735	3.22	.0069	0.70	-.1728	0.57
R^2	.8624		.3562		.9844	
rho	.174		.161		-.007	
DW	1.62		1.52		1.96	
BP-LM test for diagonal covariance matrix: $\chi2(3) = 9.052$						

Table B-5. Acres Demand SUR

	Wheat		Barley (first differences)		Canola	
	coef	t-stat	coef	t-stat	coef	t-stat
Er1	.1091	1.64	-.0343	1.14	.0165	0.56
Er2	.1358	2.32	.0677	2.26	-.1745	6.59
Er3	-.0972	2.27	-.0318	1.60	.0891	4.61
Er4	-.0419	1.42	--		.0143	1.00
W1	.0395	0.85	--		-.0167	0.42
W2	--		--		.0233	0.64
Z	--		-.1133	1.71	--	
t	.0017	0.83	-.0003	0.87	.0056	4.92
DLIFT	-.1662	3.29	.0487	2.54	-.0192	0.89
constant	-.2002	0.82	.0062	0.71	.1575	1.01
R^2	.2931		.2932		.9601	
rho	.511		.241		.155	
DW	0.92		1.44		1.68	
BP-LM test for diagonal covariance matrix: $\chi2(3)$ = 8.942						

Table B-6 Acres Demand SUR

	Wheat		Barley (first differences)		Canola	
	coef	t-stat	coef	t-stat	coef	t-stat
Er1	.1262	2.35	-.0453	1.42	.0102	.044
Er2	.1338	2.76	.0944	2.91	-.1399	6.44
Er3	-.0618	1.80	-.0304	1.47	.1058	7.16
Er4	.0063	0.25	--		.0221	1.84
Vr1	-.0200	3.01	--		--	
Vr2	--		.0008	0.21	-.0123	5.33
Vr3	--		--		-.0062	3.07
Vr4	--		-.0078	1.72	--	
Vr12	.0001	3.95	--		--	
W1	.1942	2.94	--		-.0754	2.14
W2	--		--		.0564	1.96
Z	--		-.2070	2.75	--	
Wo	-.1235	2.75	-.0024	.012	.0001	0.01
t	.0031	1.65	-.0003	0.88	.0060	6.97
DLIFT	-.1596	4.07	.0494	2.48	-.0028	0.17
Constant	1.3667	2.41	.0076	0.81	-.0821	0.33
R^2	.8122		.3389		.9797	
rho	.167		.215		-.008	
DW	1.63		1.43		1.98	
BP-LM test for diagonal covariance matrix: $\chi2(3)$ = 7.696						

Appendix C.

Hypothetical CAIS Data for Manitoba, 1966-2002

	producer margin	reference margin	Relative Loss	average government share of payments	government payments
1966	3748.7	2371.6	0.000	0.000	0.000
1967	5658.9	2854.2	0.000	0.000	0.000
1968	2572.9	3324.1	-0.225	0.567	425.900
1969	2534.1	3324.1	-0.238	0.574	453.480
1970	2374.3	3259.5	-0.272	0.590	885.200
1971	2572.8	2951.9	-0.128	0.500	189.540
1972	4680.2	2560	0.000	0.000	0.000
1973	6799.7	2560	0.000	0.000	0.000
1974	9774.4	3262.4	0.000	0.000	0.000
1975	10007.7	4684.2	0.000	0.000	0.000
1976	5233.0	7084.8	-0.261	0.585	1083.310
1977	3038.0	7269	-0.582	0.697	2949.020
1978	6967.7	7269	-0.041	0.500	150.630
1979	6098.4	7325	-0.167	0.520	637.850
1980	5300.0	6099.7	-0.131	0.500	399.850
1981	2343.6	5543.8	-0.577	0.696	2227.350
1982	3959.0	4812.1	-0.177	0.531	452.970
1983	3103.4	5119.1	-0.394	0.648	1306.140
1984	4000.0	4120.8	-0.029	0.500	60.400
1985	2125.4	3687.5	-0.424	0.658	1027.840
1986	4572.7	3135.3	0.000	0.000	0.000
1987	-3332.1	3135.3	-2.063	0.668	4320.220
1988	-4085.5	3076.3	-2.328	0.660	4726.760
1989	-7969.6	931.1	-9.559	0.615	5473.950
1990	-2769.2	-1764.1	-0.569	0.600	1661.540
1991	-5135.0	-3395.6	-0.512	0.600	3080.990
1992	-7921.6	-4184.2	-0.893	0.600	4752.940
1993	-3162.1	-5713.9	0.000	0.600	1897.230
1994	-5320.0	-5406.2	0.000	0.600	3192.000
1995	202.4	-4539	0.000	0.000	0.000
1996	7259.2	-4539	0.000	0.000	0.000
1997	11708.3	-2759.9	0.000	0.000	0.000
1998	6835.4	1433.2	0.000	0.000	0.000
1999	1623.9	4765.7	-0.659	0.709	2275.510
2000	-45.5	5239.5	-1.009	0.739	3905.580
2001	11295.1	5239.5	0.000	0.000	0.000
2002	24809.5	6722.5	0.000	0.000	0.000

	MEAN	ST. DEV
producer margin	3282.7	6164.1
reference margin	2455.7	3847.4
Relative Loss	-0.574	1.6077
average government share of payments	0.39073	0.29706
government payments	1284.8	1654.6

Correlation Matrix for Crop and Animal Prices

wheat	1							
barley	0.91573	1						
canola	0.84541	0.83948	1					
oats	0.76591	0.89189	0.80742	1				
flax	0.83867	0.81947	0.78893	0.77262	1			
rye	0.87765	0.89176	0.85913	0.8041	0.81292	1		
cattle	0.58088	0.63555	0.67731	0.7385	0.48716	0.49894	1	
hogs	0.64508	0.62883	0.7655	0.66467	0.59798	0.55613	0.78262	1