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**DYNAMIC ECONOMETRIC MODELS OF CROP INVESTMENT IN MANITOBA
UNDER RISK AVERSION AND UNCERTAINTY**

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NOTE BY THE SECRETARIAT

This paper provides empirical evidence of the investment effects of agricultural policy measures (particularly price support and area payments) in the Canadian province of Manitoba using aggregate time series data. This analysis is largely used for the document entitled *The Impact on Investment and Production of Different Agricultural Policy Instruments — Principal Findings* [AGR/CA/APM(2005)12/FINAL]. The author of the present paper is Barry Coyle from the University of Manitoba. The paper is issued as an unclassified document under the responsibility of the author.

TABLE OF CONTENTS

Executive Summary	4
I. Introduction	6
1. Decoupling, Risk Aversion and Dynamics.....	6
2. Canadian Programmes	7
3. Purpose of the study	7
4. Alternative Methodologies	7
II. Models of Investment in Manitoba	9
1. Data.....	9
2. ADL Investment Models	10
3. Euler Equation Investment Models	15
4. Conclusion Regarding Empirical Investment Models	15
III. Simulation of Programme Impacts for Manitoba	16
1. Policy and Wealth Effects	16
2. GRIP	17
3. Crop Insurance.....	18
4. WGSA	19
5. Western Grain Transition Payments	23
6. NISA	25
7. Conclusion Regarding the Simulated Impacts.....	29
Annex I. Main Estimations of ADL Investment Model.....	31
Annex II. Aggregation of price risk over commodities: an index number approach.....	41
Annex III. Discussion of Euler Equation Investment models.....	43
Annex IV. Parameter Assumptions for Simulations.....	49
References.....	50

EXECUTIVE SUMMARY

Earlier studies in the OECD project on decoupling estimated static models of crop production incorporating risk aversion, following the analytical framework in OECD (2001). In contrast, this paper studies primarily dynamic models of crop investment. The province of Manitoba is selected as an example of Canadian prairie crop production, and investment is defined as current expenditure on machinery and equipment for Manitoba crop production (the data are not crop specific). This is the first econometric study of dynamic crop investment decisions for Canada, and as far as can be ascertained this is the first econometric study for any country of dynamic crop investment under risk aversion. Econometric results are used to simulate impacts of agricultural programmes on crop investment.

Investment for crops is explained in terms of price effects, wealth effects, and insurance effects. The major price variables are an aggregate index of expected output prices for the six major crops and an aggregate index of price variances and covariances for these crops. A standard index approach is used to aggregate expected prices. An index approach to aggregation of price uncertainty is developed and shown to be critical to this study. Initial wealth is approximated as the value of cropland and buildings (value of machinery and equipment and of farm debts are much smaller and approximately cancel out). The insurance effect of farm programmes is measured by the covariance between government payments and market prices for crops. Econometric models use annual Manitoba data for 1960-2002.

Two alternative approaches to dynamic models of investment under risk aversion are explored: autoregressive distributed lag (ADL) models and Euler equation models. An ADL model can be viewed as an application of the general-to-specific specification of reduced form dynamic models. Given our primitive knowledge of true structural models for dynamics, econometricians now generally prefer the ADL approach to dynamic models. Euler equations are the most popular approach to specifying structural dynamic models, so these are also considered.

Econometric results for ADL investment models suggest that the expected output price index (P^{Div}), the output price variance/covariance index (VP^{Div}), wealth (W^0), and the covariance between government payments and market prices (CovGp) have statistically significant impacts on investment in the expected direction. Typical estimates of long run elasticities (sum of lag coefficients) of impacts on investment are +1.090, -0.052, +0.760, -0.074 for prices (P^{Div}), price uncertainty (VP^{Div}), wealth (W^0), and covariance between government payments and prices (CovGp), respectively. These estimates seem plausible, except the elasticity of wealth is larger than anticipated and lacks clear interpretation.

Since economists have a poor understanding of dynamic behaviour, structural Euler equation models are more likely to be seriously mis-specified than are reduced form ADL models. So less emphasis should be placed on results for Euler equation models than for ADL models. Nevertheless the simplified structure of Euler equation models facilitates testing of certain hypotheses of interest. Investment does not appear to be limited by financial constraints, and option value may have a negative impact on current investment. The latter conjecture is consistent with the modern theory of investment emphasising the evolution of information over time and irreversibility of decisions, although the estimated elasticity is quite small in this initial study.

The models estimated above measure the impacts of expected prices, price variances, wealth and insurance effects on investment and output. By specifying the effects of government programmes on these variables, an estimate of the effect of these programmes on investment and output maybe obtained. A critical assumption here is that similar expectation and response processes apply to these programme effects as in the aggregate. This approach is applied to several Canadian agricultural support programmes

in Manitoba. An alternative approach is to include programme specific variables directly in the econometric models, and so estimate programme impacts on investment and output directly. However, this approach is inappropriate for countercyclical programmes: payments from countercyclical programmes are positively correlated with low incomes, but farm level incomes are not fully specified by the explanatory variables in the econometric models, so estimators of programme impacts are biased and inconsistent.

The Gross Revenue Insurance Programme (GRIP) of 1991-95 provided payments to producers linked to crop prices. GRIP serves as a ready base of comparison for other programmes, as it was a relatively standard price support programme, and its price effects are easily simulated. The average percentage change in price due to GRIP is +12.3%. In addition there are significant insurance effects. The average long run impacts on investment are calculated as +14.6% (price effect) and +7.3% (insurance effect), for a total of 21.9%.

The Western Grains Transportation Act (WGTA) provided western crop producers substantial subsidies for transport of grain to ports for export. This programme was eliminated in 1995 and farmers received compensation in the form of "transition" payments over 1996-98. The econometric models are applied to simulate impacts of the WG Transition, ignoring secondary effects on prices (this would require a full general equilibrium model).

Elimination of WGTA led to an average decrease of 16.5% in crop prices over 1995-2002 (the smallest decrease is for canola, 5.9%), and to a 10.9% average decrease in price for 1996. Transition payments were 3.4% of wealth in 1996 and 2.4% of wealth in 1997 (0.7% over 1995-2002). The average long run impacts on investment over 1995-2002 were calculated as -19.7% (price effect) and +0.6% (wealth effect), for a total effect of -19%. Long run impacts on investment for 1996 were calculated as -13% (price effect) and +4.9% (wealth effect), for a total effect of -8.1%. These calculations for 1996 assume that farmers factor future payments into perceived wealth. Since there is ambiguity in interpreting wealth effects in the econometric model, the model was re-estimated with Transition payments treated as a separate wealth variable. This did not significantly change the results.

The Net Income Stabilization Account (NISA) programme of 1991-2002 was designed to encourage farmers to increase savings in high income years for use in low income years, thus smoothing incomes over time. The difficulties in measuring impacts of NISA on total wealth and total withdrawals make measuring the impact of this programme more difficult. NISA fund balances and withdrawals from Fund 2 are used as upper bounds on these impacts. NISA fund balances averaged 12% of wealth over 1991-2002, and withdrawals from Fund 2 show a high negative covariance with market prices and revenues, indicating the potential for significant wealth and insurance effects. Price effects seem likely to be minor, but may be important for some participants. Attempts to estimate econometric models with separate NISA-specific variables (in order to calculate NISA effects directly) were inconclusive.

Simulation results based on the results of the econometric models (without separate NISA variables) and the above information show the average impact of NISA on investment over 1991-2002 to be +6.5% (wealth effect) and +5.8% (insurance effect), for a total of 12.3%. NISA may well increase investment, but these effects, which are taken to be the upper bound, are less pronounced than for programmes such as GRIP.

DYNAMIC ECONOMETRIC MODELS OF CROP INVESTMENT IN MANITOBA UNDER RISK AVERSION AND UNCERTAINTY¹

I. Introduction

1. Decoupling, Risk Aversion and Dynamics

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 (OECD, 2001), and this provides the analytical framework for ongoing 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, emphasising (*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 emphasised static models under risk aversion and uncertainty (*e.g.* Sckokai, 2002). The second phase of the research will emphasise dynamic models of investment under risk aversion and uncertainty. This will include an extension of the Italian static model to dynamics (Sckokai, 2003).

It is important to recognize that there is very limited experience in econometric modelling of dynamic agricultural production behaviour under risk aversion and uncertainty. This has two important implications for this study. First, a study on investment under risk aversion should consider several promising approaches rather than adopting a single approach which may turn out to be inferior. The two approaches that are popular for estimation of dynamic models under risk neutrality are applied in this paper.

Second, there remains much to be learned from modelling investment under risk aversion using macro as well as micro data sets. For many countries adequate representative micro data sets are not available. For example, this is true of Canada and apparently the United States². For many countries the only option for research is to utilize macro data. This study will use province level production data.

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1. This paper was prepared by Barry T. Coyle, University of Manitoba, Canada. The author thanks James Rude and especially Jesús Antón for helpful comments.
 2. An earlier study did utilize a micro data set (ARMS) for the United States (Goodwin and Mishra), but this is not a panel data set (different farms are sampled each year, apparently without overlap). Consequently this U.S. data set is of limited value for a dynamic study.

In fact studies with micro data may not model farm-level uncertainty much more effectively than does macro data. In general prices are much more highly correlated across farms than are crop yields. For example, prices for the same commodity are quite similar across farms in the same province of Canada, whereas yields often vary substantially across farms in the same province. Thus micro data is essential to modelling farm-level crop yield uncertainty but are not necessary in modelling price uncertainty. Despite this, micro data sets are often used to model only price uncertainty even though these data sets cover only a small number of years, and a longer time horizon may be needed to model dynamic processes.

2. Canadian Programmes

Over the course of the study period, payments were provided to farmers under the Net Income Stabilization Account (NISA) and the Canadian Farm Income Programme (CFIP). NISA had been in operation for about fifteen years, and CFIP (together with the preceding Agricultural Income Disaster Assistance programme, AIDA) for seven years. The two programmes have been replaced by a new programme called the Canadian Agriculture Income Stabilization (CAIS) programme.

Both NISA and CFIP were safety net programmes providing payments generally countercyclical to farm income. NISA was a subsidized farm savings programme without insurance. A farmer deposited savings into a NISA account, and received matching government contributions to a certain limit. The account enjoyed an interest rate subsidy. The farmer was allowed to make withdrawals from this account when farm income fell below a threshold level. CFIP was an income insurance programme. When farm income fell below a threshold level, the farmer received a payment raising net farm income to a predetermined target level. The premium was highly subsidized. In practice, these programmes increased expected income and reduced negative variations in income.

This study looks also to other Canadian programmes that existed during the period under study. Programmes such as the Gross Revenue Insurance Programme (GRIP), crop insurance, the Western Grains Stabilisation Act (WGSA) and the Western Grains Transition Payments (WGTP) can potentially have affected investment decisions

3. Purpose of the study

The purpose of this research was twofold. The first objective was to estimate dynamic econometric models of investment in crop machinery and equipment for Manitoba, and also to estimate dynamic crop yield equations. These models incorporate risk aversion and uncertainty. The second objective was to apply these econometric results to simulate impacts of the major historical programs on crop investment and crop output for Manitoba. Manitoba was selected as representative of prairie crop agriculture.

This paper presents the econometric and simulation results for investment. However, results for yield and output are not released here.

4. Alternative Methodologies

Since uncertainty generally increases over the length of a planning horizon, it is obvious that uncertainty and risk aversion should be more important in dynamic models than in static models. Nevertheless, in both theoretical and econometric work, dynamic microeconomic models have generally assumed risk neutrality, and microeconomic models with risk aversion have generally assumed static equilibrium. For example, with almost no exceptions, the modern microeconomic theory of investment under uncertainty assumes risk neutrality (*e.g.* Dixit and Pindyck). Given the paucity of theoretical and econometric studies of microeconomic dynamics under risk aversion, there is no well established methodology for econometric research on dynamics under risk aversion, leaving a weak foundation for econometric research on dynamics under risk aversion. Three alternatives are considered in this section

(ADL models, Euler equation models and Optimal control models), but only the first two are retained for the empirical work.

A. ADL Models

ADL (or equivalently error correction) models express current investment decisions as a function of lagged investments and current and lagged levels of explanatory variables. An ADL model can be interpreted as a reduced form for a broad variety of structural dynamic models. This has the advantage of avoiding many specification errors in structural models. On the other hand, it is usually difficult to identify a structural model from the ADL, and structural models can be more informative in principle for policy.

B. Euler Equation Models

In both microeconomic and macroeconomic econometric studies, structural dynamic models are most commonly estimated in terms of Euler equations. In addition, theoretical studies of investment often employ related dynamic programming methods (*e.g.* the modern theory of investment under uncertainty). The success of this approach in empirical research is mixed, and most studies have assumed risk neutrality.

The Euler equation approach should also be adopted in this study, given its historical and current importance. Moreover a major advantage of this approach is that it can readily incorporate static theoretical and empirical approaches to decisions under risk aversion into a dynamic setting. Static theory is often viewed as modelling current output and variable input decisions conditional on capital stocks (and investment decisions). Then, at one level, a dynamic programming model is constructed by appending an Euler equation to a static model (*e.g.* Berndt, Fuss and Waverman). This permits an Euler equation model of investment under risk aversion to build on theoretical and empirical research on static decisions under risk aversion.

The advantages of a duality approach to econometric modelling of production decisions in static risk neutral models are well known, and recent research has extended static models to risk aversion. Chambers and Quiggin have constructed a general model of decisions under risk in a state contingent framework, but this approach seems of limited value for econometric research since the many contingent states are not observable to the econometrician. Mixed dual and primal models have been specified in terms of a standard dual cost function and a direct utility function over profits for a risk averse firm (Saha; Chavas and Holt, 1996). This approach certainly has merit for empirical research, although it loses some of the advantages of a dual approach. Dual models of static decisions under risk aversion have been constructed in a mean-variance framework (Coyle, 1992, 1999), and recently this approach has been extended to non-expected utility (Coyle, 2004). Thus the Euler equation approach to dynamics can build upon a variety of static decision models under risk aversion.

C. Optimal Control Models

An alternative approach to modelling dynamic decisions using duality was suggested by Epstein. This is based on an optimal control (rather than calculus of variations) specification of a dynamic optimization problem and the empirical model is obtained essentially by applying the envelope theorem to a Hamilton-Jacobi-Bellman (HJB) equation. This approach has some advantages over Euler equations (*e.g.* in modelling multiple investment decisions), but it fell into disrepute because it initially imposed static expectations over the planning horizon (*i.e.* the model assumed that a firm forms a dynamic plan assuming that prices will not change over the future, and when prices happen to be different in the next year, the firm forms a new dynamic plan assuming again that prices will not change over the future). In recent years there has been renewed interest in the approach among agricultural economists as studies have relaxed the assumption of static expectations (Luh and Stefanou) and incorporated elements of the modern theory of

investment under uncertainty (Chavas; Oude Lansink and Stefanou; Pietola and Myers; Stefanou). Here firms recognize that their information set will change over time, so there is an option value to waiting, and investment decisions are irreversible or at least asymmetric. Thus the approach now appears to provide a viable alternative to Euler equations. The approach has also been extended to risk aversion assuming mean-variance risk preferences (Coyle and Arnade). First priority will be given to the more popular ADL and Euler equation approaches, so dual optimal control models will not be estimated in this study.

II. Models of Investment in Manitoba

1. Data

The following data for Manitoba, obtained primarily from the Manitoba Agriculture Yearbook 2002, is used: 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 was obtained for hired labour, variable crop inputs, and machinery plus equipment (Manitoba Agriculture Yearbook, Statistics Canada). Annual expenditures between 1960 and 2002 on machinery and equipment for crop production were obtained by special request from Statistics Canada, Ottawa. Published data is also available on current value of machinery and equipment for Manitoba, but this proved to be unsatisfactory for our purposes. Data on government payments and other programme details, were obtained from the following agencies or their publications: Manitoba Crop Insurance Corporation (MCIC), Net Income Stabilization Account programme (NISA), Agriculture Income Disaster Assistance (AIDA), Gross Revenue Insurance programme (GRIP), Western Grains Stabilization Act programme (WGSA), Western Grains Transportation Act programme (WGTA). Data on other programmes was obtained from Agriculture Economic Statistics (Statistics Canada 26-603).

Capital investment I is measured as current expenditures on machinery and equipment for crop production divided by the price index for machinery and equipment. Capital stock is measured with a proxy variable constructed from investment as $K_t = \sum_{s=0, \dots, S} I_{t-s} (1-\delta)^s$. In the base case $\delta = .05$ and (to conserve observations for the model) $S = 3$. This is highly correlated with proxies for capital stocks using longer lags: there is a correlation of 0.92 with K estimated using a six period lag.

Weather is measured using an annual crop-growth weather index GRODEX (Dyer, Narayanan and Murray), which 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 (16 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. 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 mil CAD 9782, value of machinery and equipment was mil CAD 3315, and total debt was mil CAD 3735 (Manitoba Agriculture Yearbook, 2002). For this reason the value of land and buildings in crop agriculture will be assumed to represent 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 assumed to be equal to 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. 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, 2005; Mbagi and Coyle) (also see Sulewski, Spriggs and Schoney).

$$\begin{aligned}
 (1) \ E_{t-1} p_t &= p_{t-1} \\
 \text{var}(p_t) &= 0.50 (p_{t-1} - E_{t-2} p_{t-1})^2 + 0.33 (p_{t-2} - E_{t-3} p_{t-2})^2 + 0.17 (p_{t-3} - E_{t-4} p_{t-3})^2 \\
 \text{cov}(p_{it} p_{jt}) &= 0.50 (p_{it-1} - E_{t-2} p_{it-1}) (p_{jt-1} - E_{t-2} p_{jt-1}) + 0.33 (p_{it-2} - E_{t-3} p_{it-2}) (p_{jt-2} - E_{t-3} p_{jt-2}) \\
 &\quad + 0.17 (p_{it-3} - E_{t-4} p_{it-3}) (p_{jt-3} - E_{t-4} p_{jt-3})
 \end{aligned}$$

A covariance between market prices and various government payments is defined similarly in equation (2), where G is various government payments. This is designed to capture an insurance/income smoothing effect rather than a wealth effect of government payments. Therefore, while this excludes 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.

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

A truncation approach to the effect of programmes on price distributions used in studies such as Chavas and Holt is not adopted here. 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). In our case the truncation process seems more complex since multiple programmes imply multiple truncation parameters. 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 programmes on price distributions. Instead it is simply assumed that producers adopt a naive expectations process regarding market prices and how market prices co-vary with government payments. By modelling these as separate processes it can be tested if they have different impacts on investment and production decisions. Thus we can test if government programmes 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. 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 models.

2. ADL Investment Models

A. Brief Literature Review

There are two general approaches to estimation of a dynamic model in econometrics. One approach is to estimate a simple specific structural dynamic model and then optionally add variables or lags, termed “specific to general approach”. Alternatively a “general to specific” approach can be followed: a general reduced form dynamic model such as an autoregressive distributed lag (ADL) model is specified, and restrictions on the model are then tested and perhaps adopted. This approach has been advocated by Hendry (Hendry 1995; Hendry and Ericsson 1991; Hendry and Krolzig 2004; Gilbert 1986) and is now generally recognized as the superior approach to estimation of dynamic models (Davidson and MacKinnon 1993; Johnston and DiNardo 1997; Greene 2003).

ADL models typically assume risk neutrality. Two recent studies of Canadian agriculture have accommodated risk aversion and output price uncertainty by including multi-period lags in expected price and price variance. These studies use industry level data for Alberta beef production (Mbagha and Coyle) and Manitoba crop production (Coyle 2001). Results support simple naive models of price expectations and uncertainty rather than rational expectations or GARCH models. Sums of lag coefficients are significant with anticipated signs for expected price and price variance, and adding price variance often increases significance of expected price.

The modern theory of investment emphasises irreversibility and the option value of waiting (Dixit and Pindyck). This implies a lumpy rather than continuous investment process, and this process is best modelled at the firm level. Because of the unavailability of micro data, an alternative approach is taken here where the impacts of option value of waiting are measured at the aggregate level.

B. Methodology

Unit root tests conducted for capital investment and price variance³ suggest that investment, the dependent variable, has a unit root (and is then trended over time), while some explanatory variables do not have unit roots. While an I(1) dependent variable can reflect a mix of I(1) and I(0) explanatory variables, this may also reflect the fact that such standard unit root tests have low power to reject the null hypothesis of a unit root. In such a situation, the standard approach is to estimate a regression model with all variables first differenced. The most appropriate method for estimation of a model with mixed I(1)/I(0) regressors may be the fully modified least squares approach of Phillips (Phillips 1995; Phillips and Durlauf, 1986). But since the standard approach of OLS estimation of the model with all variables first differenced remains valid⁴, this approach is followed in this study.

The capital investment variable for machinery and equipment in Manitoba crops cannot be decomposed into individual crops, so the capital investment decision must be modelled in terms of prices and price uncertainty for all major crops. This is impractical without some form of aggregation. Expected prices are typically aggregated over crops as a Tornqvist or Fisher price index, but theory has not provided an appropriate procedure for aggregating price uncertainty over commodities. One ad hoc approach is to calculate the variance of the aggregate price index, but this is unlikely to be satisfactory. An appropriate aggregation procedure is essential to this study and it is developed in an Annex II as an extension of standard index number theory. An aggregate price variance index VP can be defined that, under standard assumptions, correctly represents the contribution of price variances and covariances to total multiple crop revenue uncertainty (overlooking yield uncertainty). A "Tornqvist" index VP of price uncertainty is used. Price variances and covariances for individual commodities are constructed as in (1), and then these are

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3. The unit root hypothesis is not rejected at the .10 level for capital investment (current expenditures on machinery and equipment divided by the price index for machinery and equipment), capital stock and normalized cost, using augmented Dickey-Fuller and Phillips-Perron tests with and without a time trend, as encoded in Shazam 8.0. Unit root test results are ambiguous for the normalised Divisia crop price index. The unit root hypothesis is rejected at the .10 level for the normalised Divisia index for crop price variance, normalised crop input price, normalised capital price, normalised covariance between government payments and crop prices, using all tests. The unit root hypothesis is rejected at the .10 level for normalised initial wealth, using Phillips-Perron tests with trend.
 4. In a regression model with mixed I(1) and I(0) regressors, OLS estimators in levels have standard asymptotic properties for I(0) variables and I(1) variables that are a linear combination of I(0) (Sims, Stock and Watson, 1990; Stock and Watson, 1988; Park and Phillips, 1989). Estimators for coefficients of other level variables are not asymptotically normal. A common practice in the past for a regression model with mixed I(1)/I(0) variables has been to replace only I(1) variables with their first differences (*e.g.* Box and Jenkins, 1970), but this seriously mis-specifies the regression model (in effect this approach introduces omitted variables which are strongly correlated with included variables, so estimators are inconsistent).

aggregated across commodities to construct an index VP as shown in Annex II. Hereafter we will refer to this Tornqvist index as a “Divisia” index VP^{div} (although many indexes are discrete time approximations to a continuous time Divisia index).

C. Econometric Results

The three main econometric estimations in this section are presented in Tables I.1, I.2 and I.3 in Annex I. The rest of estimations can be found in the remaining tables of Annex I. The first specification is a simple risk neutral and risk averse ADL models of capital investment:

$$(3) \quad I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,..,n} \beta_{1i} P^{Div}/w_{t-i} + e_t$$

$$(4) \quad I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,..,n} \beta_{1i} P^{Div}/w_{t-i} + \sum_{i=0,..,n} \beta_{2i} VP^{Div}/w_{t-i}^2 + e_t$$

where w is the crop input price index, I is investment (current expenditures on machinery and equipment for crops divided by the price index of machinery and equipment), P^{Div} is the Divisia price index for crops, and VP^{Div} is the aggregate Divisia index of market price uncertainty for crops defined in Annex II. Since an aggregate Divisia price index is defined in theory in terms of a behavioural model based on market prices, P^{Div} is based on market prices. An alternative Divisia type index including crop specific government payments and charges was also briefly considered, but initial econometric results suggested that it was a less appropriate aggregate index. Other input prices (hired farm labour wage, a capital price index) were statistically insignificant. The normalization P/w and VP/w^2 is suggested by constant relative risk aversion (Pope 1988; Coyle 1999). All variables are specified as first differences of logarithms. Loglinearity implies that coefficient estimates of log variables can be interpreted as elasticities.

Ordinary least squares (OLS) estimates of simple first differenced log models are presented in Table I.1 of ADL Investment Models in Annex I. For the risk neutral model (Table I.1.A, *i.e.* column A of Table I.1), the sum of lag coefficients (elasticities) for the Divisia crop output price index P^{Div} is + 0.711 and is statistically significant from 0 at the .05 level (t-ratio is 2.15). For the risk averse model (Table I.1.B) using the aggregate Divisia price variance index VP^{Div} discussed in Annex II the estimated sum of lag coefficients for VP^{Div} is -0.075 with a t-ratio of 2.26. Moreover the sum of lags for P^{Div} increases to +1.150 and t-ratio 3.10. So the price variance is statistically significant with the anticipated sign, and the statistical significance of expected output price also increases. The importance of using the aggregate Divisia price variance VP^{Div} is illustrated in Table I.1.C, where VP^{Div} is replaced by VP^{BAD} , a variance of the Divisia price index. Here the sum of lags for VP^{BAD} is -0.030 with a t-ratio of only 0.94 (the sum of lags for P^{Div} is +0.761 with t-ratio 2.07). Thus this standard approach to aggregating price risk is unsuccessful in our empirical model.

Aggregate measures of price risk were also constructed from multivariate GARCH models of the six crop prices. However, resulting estimates of Equation (4) were poor: these models have lower explanatory power and the coefficient of price variance VP is statistically insignificant (Coyle, 2005).

A more general risk averse model can be specified as

$$(5) \quad I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,..,n} \beta_{1i} P^{Div}/w_{t-i} + \sum_{i=0,..,n} \beta_{2i} VP^{Div}/w_{t-i}^2 + \sum_{i=0,..,n} \beta_{3i} W^0/w_{t-i} \\ + \sum_{i=0,..,n} \beta_{4i} COVGP_{t-i}/w_{t-i}^2 + \sum_{i=0,..,n} \beta_{5i} w^L/w_{t-i} + \sum_{i=0,..,n} \beta_{6i} w^k/w_{t-i} + e_t$$

with eight period lags for all variables ($n=8$) except for the one period lag on investment (longer lags on investment were insignificant). Here w^L, w^k are hired labour wage and capital price index. Input prices were found not having statistically significant parameters and were dropped from model (5). W^0 is initial wealth. All variables except for $COVGP$ are specified as logarithms, so coefficients can be interpreted as elasticities. $COVGP$ is normalised by its 2000 (absolute) value, so its coefficient can be interpreted as an elasticity circa 2000. $COVGP$ is an index of the covariance between government payments G and market

prices p . Denote $COVGPY_t = \sum_{j=1,\dots,6} \text{cov}(G_t, p_{jt}) y_{jt}$. Then $COVGP_t = (COVGPY_t / COVGPY_0) / Y^{DIV}$ using 1985 as the constant base for $COVGPY$ and aggregate Divisia output Y^{DIV} . This Divisia type index $COVGP$ is generally negative, and can be interpreted as an aggregate insurance effect of programmes that decreases effective price uncertainty.

Empirical studies of production decisions under risk aversion typically assume that initial wealth W_0 influences decisions primarily by changing risk aversion. The stylized fact of decreasing absolute risk aversion (DARA) certainly suggests that this is an important influence of W_0 . However, an increase in W_0 will also decrease the firm's debt equity ratio, and so perhaps decrease the marginal cost of borrowing for investment. In addition current value W_0 is correlated with the constant dollar value of capital, and perhaps the latter is correlated with investment, though that was not observed in our data. Given the uncertainty in how to interpret effects of W_0 on investment in an empirical model, it is of interest to separate government payments G (that are not commodity specific) from W_0 and, in particular, separate payment variables might be defined for different programmes. However, there is insufficient data to permit a general-to-specific analysis of an ADL model with separate variables for multiple programmes. This issue will come up later when the impacts of specific programmes are assessed.

There are insufficient observations to estimate the full unrestricted model with $n=8$, so polynomial restrictions were imposed. A recommended approach to selecting the lag length and degree of polynomial is to (a) estimate unrestricted distributed lag models with long lag lengths and use simple nested tests for reducing the lag length, and (b) (given the selected lag length) use nested tests to select the degree of the polynomial (e.g. Davidson and MacKinnon, pp. 673-76; Sargan). This approach was applied and a four period lag model (6) was selected with polynomials of degree 1,1,3,1, respectively. Estimates of the selected model are reported in Tables I.2 and I.3. Standard diagnostic tests were considered⁵.

$$(6) \quad I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,1,2} \beta_{1i} P^{DIV}/w_{t-i} + \sum_{i=0,1} \beta_{2i} VP^{DIV}/w_{t-i}^2 + \sum_{i=0,\dots,4} \beta_{3i} W^0/w_{t-i} + \sum_{i=3,4} \beta_{4i} COVGP_{t-i}/w_{t-i}^2 + e_t$$

OLS estimates of model (6) (first differenced in logarithms) under polynomial restrictions are reported in Table I.2.A. The sum of lag coefficient estimates and corresponding t-ratios are +1.037 (2.91), -0.049 (1.46), +0.767 (1.95), -0.076 (1.73) for P^{DIV} , VP^{DIV} , W^0 , $COVGP$ respectively. R^2 is 0.505. Perhaps the relatively large elasticity for W_0 cannot be explained primarily in terms of DARA (the coefficient estimate is highly significant in Table I.3.C correcting for autocorrelation). The polynomial restrictions are not rejected. The lagged dependent variable is insignificant at any level. The reduction in statistical significance (t-ratios) for P^{DIV} and VP^{DIV} relative to Table I.1.B reflects the increased multicollinearity in the model.

The coefficient estimates for W_0 (current value of farmland and buildings in crops) might simply reflect a positive correlation between physical (constant dollar) capital and investment. To check this, total cropland was added to the model, with the same lag structure as for W_0 . The sum of lags for W_0 was reduced only slightly (from .767 to .630), and the sum of lag coefficients (elasticities) for total cropland was only -.00034. This indicates the relatively substantial coefficients for W_0 cannot be explained in terms of spurious correlations involving constant dollar capital.

Table I.2.B reports OLS estimates of model (6) without the polynomial restrictions. The sum of lag coefficient estimates and corresponding t-ratios are +1.136 (2.77), -0.047 (1.40), +0.812 (1.70), -0.075

5. The Durbin h statistic could not be computed, but a grid search maximum likelihood procedure (Beach and MacKinnon) as encoded in Shazam 8.0 (White) indicated insignificant autocorrelation. Homoskedasticity was not rejected at any standard level of significance using an ARCH test due to Engle. Using the Ramsey RESET test with various powers of the predicted dependent variable, the hypothesis $Ee=0$ was not rejected at the .05 level. Sequential Chow tests did not reject parameter stability at the .05 level.

(1.66) for P^{DIV} , VP^{DIV} , W^0 , $COVGP$ respectively. R^2 is .519. These results are quite similar to estimates with polynomial restrictions imposed.

OLS estimates of model (6) excluding the lagged dependent variable I_{t-1} are reported in Table I.3. Under all polynomial restrictions the sum of lag coefficient estimates and corresponding t-ratios are +1.090 (3.40), -0.052 (1.65), +0.760 (1.96), -0.074 (1.72) for P^{DIV} , VP^{DIV} , W^0 , $COVGP$ respectively (Table I.3.A). Note that, in OLS estimates without polynomial restrictions, the second (*i.e.* fourth period) lag on $COVGP$ is significant at the .05 level by itself, with an estimate -0.060 and t-ratio 2.21 in Table I.2.B, and similarly for Table I.3.B. In the absence of any polynomial restrictions, the sum of lag coefficient estimates and corresponding t-ratios are +1.192 (3.43), -0.050 (1.54), +0.840 (1.83), -0.073 (1.66) for P^{DIV} , VP^{DIV} , W^0 , $COVGP$ respectively (Table I.3.B). For this model, a grid search maximum likelihood procedure (Beach and MacKinnon) estimated a first order autocorrelation coefficient +0.28 (t-ratio 1.87), so this model was also estimated by GLS. The sum of estimated lag coefficients are +1.114 (3.66), -0.047 (1.81), +0.996 (2.61), -0.091 (2.48) for P^{DIV} , VP^{DIV} , W^0 , $COVGP$ (Table I.3.C).

Total costs were added to model (6), as a proxy for the possible impact of financial constraints on investment. $Tcost$ is the total cost for all six crops, and it is normalised by the price of capital w^k . A one to three year lag and a first order polynomial was selected for $Tcost/w^k$, which is first differenced in logarithms. OLS estimates for model (6) plus this lag in $Tcost/w^k$ are presented in Tables I.4 and I.5 (Annex I). Results are similar to model (6). For example, using all polynomial restrictions and omitting the lagged dependent variable, the sum of lag coefficients and t-ratios are +0.968 (3.18), -0.053 (1.77), +0.731 (2.02), -0.071 (1.77) for P^{DIV} , VP^{DIV} , W^0 , $COVGP$ and +0.065 (2.16) for $TCOST$ (Table I.5.A). R^2 is .595, and there does not appear to be significant autocorrelation. Other studies of reduced form investment behaviour have also estimated that financial constraints or cash flow influence capital investment. However, there are serious criticisms of such studies. Of most importance, this relation may be spurious reflecting autocorrelated shocks to output and cash flow (Shapiro). A proper analysis apparently requires a structural model of investment behaviour (Gomes). In Annex III a structural Euler equation model is estimated, and results suggest that investment is not influenced by a financial constraint.

The next specification attempts to include a measure of change in uncertainty in the model as a proxy for effects of option value on investment. Although the modern theory of investment depends critically upon the option value of waiting as information evolves over time, empirical models of investment seldom (if ever) incorporate option value. Our standard measure of price uncertainty is a weighted average of squares of one period price differences: $Vp_t = 0.5 (p_{t-1} - p_{t-2})^2 + 0.33 (p_{t-2} - p_{t-3})^2 + 0.17 (p_{t-3} - p_{t-4})^2$. Consider a weighted average of squares of price differences s periods apart ($s > 1$) in equation (7) and its difference in (8). This latter -for various s - provides a crude measure of the reduction in price uncertainty if investment decisions are postponed for s periods. Divisia index numbers DVP^s for DVP^s were calculated by aggregating over all crop prices similarly to Annex III. Such crude proxies for the option value of waiting should be negatively related to investment, *i.e.* greater gains in information from delaying investment should have a negative impact on immediate investment. DVP^s for alternative values of s are not necessarily highly correlated (*e.g.* +0.46 for $s=1$ and $s=2$), so it is not clear that DVP^s for a single s can provide an adequate measure of reduction in uncertainty.

$$(7) Vp_t^s = 0.5 (p_{t-1} - p_{t-1-s})^2 + 0.33 (p_{t-2} - p_{t-2-s})^2 + 0.17 (p_{t-3} - p_{t-3-s})^2.$$

$$(8) DVp_t^s = Vp_t^s - Vp_t$$

This ADL model simply considers DVP^s for $s=2$ and adds a distributed lag (of first differences) to model (6). For a four period lag without polynomial restrictions, the sum of estimated lag coefficients was -0.015 and a corresponding t-ratio 1.34. Coefficients can be interpreted as elasticities circa 2000. The most significant estimate of a single coefficient was -0.005 with a t-ratio 1.54. In sum, preliminary estimates of

the ADL model do not indicate a significant effect for option value of waiting. This presumably reflects measurement errors and difficulties caused by many distributed lags in the ADL model.

3. Euler Equation Investment Models

Structural dynamic econometric models of investment are typically estimated in terms of a discrete time Euler equation model. These models explain dynamics entirely in terms of convex costs of adjustment for capital and assume risk neutrality. In spite of its popularity, this approach has generally led to poor empirical models of investment (*e.g.* Whited). In contrast, the modern theory of investment emphasises the option value of delaying investments in a setting where information evolves over time and decisions are irreversible or costly to reverse (Dixit and Pindyck). Recent empirical studies have extended the standard Euler equation model to accommodate to some extent non-convex adjustment costs, asymmetry and financial constraints. However, these studies still do not adequately represent the modern theory of investment because they seldom (if ever) incorporate changes in information over time. Moreover the assumption of risk neutrality seems particularly implausible in a dynamic model, since uncertainty is more substantial in a dynamic setting than in a static setting.

Several estimation attempts have been done following the Euler equation approach. However, since the ADL model is found to be more appropriate, the corresponding equations (9) to (29) and estimations are only discussed in Annex III.

4. Conclusion Regarding Empirical Investment Models

This study has estimated both ADL and Euler equation models of investment in machinery and equipment for Manitoba crop agriculture. ADL models are reduced forms to structural dynamic models, whereas Euler equation models are often viewed as structural models of investment. This distinction between reduced form ADL models and structural Euler equation models is valid in principle, but it does not hold up very well in the typical practice. As noted above (Annex III), at one level an Euler equation model is somewhat similar to an ADL model with lags truncated at one period, except that Euler equation models are conditional on capital stock. There are of course substantial differences in econometric specifications and methods of estimation.

In practice an Euler equation model typically does not incorporate restrictions implied by a structural model. For example, the structural dynamic optimization model (9)-(10) under risk aversion implies cross-equation symmetry restrictions on coefficients between the Euler equation (15) and output supply and factor demand equations (16). However, imposing these restrictions makes the empirical model highly nonlinear in coefficients. Consequently few empirical studies estimate an Euler equation jointly with equations such as output supply or factor demand, and even fewer impose coefficient restrictions implied by a structural model (*e.g.* Young estimates an Euler equation jointly with a cost function). In the absence of such restrictions, structural coefficients (subject to a normalisation) can still be inferred from an Euler equation, but in this respect the Euler equation does not use structure to improve estimation.

A potentially serious criticism of ADL models is the Lucas critique (econometric model estimates cannot be used to simulate changes in policy unless parameters are invariant to policy). Supporters of the “general to specific” approach to estimation of such models have argued that the Lucas critique is not critical in common practice (Banerjee, Hendry and Mizon). Although empirical Euler equation models are often intended to address this critique, it has been argued that they do not address the critique (Oliner, Rudebusch and Sichel). To a certain extent the Lucas critique applies somewhat equally to both ADL and empirical Euler equation models.

It can be concluded that “general to specific” estimation of an ADL model seems more appropriate than standard practice estimation of Euler equations. This is consistent with the dominant view in recent years among econometricians regarding “general to specific” strategies (as noted above). Nevertheless the truncation of lags in Euler equations may facilitate studies of issues such as the importance of financial constraints on investment, the importance of option value, and asymmetry of response.

Accordingly more weight should be placed on estimates from ADL models than from Euler equation models. Results for ADL models suggest that expected output price (P^{DIV}), output price variance (VP^{DIV}), initial wealth (W^0), and covariance between government payments and market prices (COVGP) have statistically significant impacts on investment. Typical estimates of long run elasticities (sum of lag coefficients) are +1.090, -0.052, +0.760, -0.074 for P^{DIV} , VP^{DIV} , W^0 , COVGP respectively (Table I.3.A in Annex I). In contrast GMM estimates of Euler equations indicate a larger elasticity for W^0 than for P^{DIV} (e.g. Table I.7), and COVGP is statistically insignificant. This result for W^0 relative to P^{DIV} is surprising and seems unreasonable. On the other hand, Euler equation results suggest that financial constraints are not binding on investment and that option value may have a negative impact on current investment. The latter is consistent with the modern theory of investment, although the estimated magnitude is quite small in this initial study.

III. Simulation of Programme Impacts for Manitoba

This section simulates the impacts of the main crops programmes in Canada. The estimated investment equations in Section II provide the main parameters. The equations and parameters needed for the simulations are summarised in Annex IV. The tables on impacts on investment are shown in this section. Program impacts on investment are simulated from estimates of the econometric models. These models estimate impacts of expected prices, price variances, wealth and insurance effects on investment. 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 investment.

In principle we should include program specific variables in the econometric models, and so estimate program impacts on investment directly. This approach is also considered briefly, but results are poor for countercyclical programs such as NISA. Apparently the difficulty is that payments from countercyclical programs are positively correlated with low incomes, but farm level incomes are not fully specified by the explanatory variables in the econometric models. Thus countercyclical program effects are significantly correlated with variables omitted from regression models, so estimators are biased and inconsistent. In contrast Western Grain Transition payments were not designed to be countercyclical. Then both approaches to simulating impacts of the Western Grain Transition program, i.e. approaches based on the above econometric models versus models with program specific variables, lead to similar simulation results.

1. Policy and Wealth Effects

The relationship between policy and wealth effects is important but difficult to address. Changes in prices due to policy can lead to changes in wealth, in particular to changes in value of cropland. Also, current decisions may reflect anticipated future wealth rather than current wealth, in somewhat the same manner as current consumption decisions may reflect permanent income rather than current income. Given lags in effective implementation of investments, producers may plan investments based on anticipated wealth (on the other hand, there is an option value to waiting). Also lending institutions may share these anticipations for future wealth to some extent, so that cost of borrowing or credit constraints may be reduced. Anticipated wealth may also reduce current risk aversion. Impacts of policy on decisions through

changes in anticipated wealth seem particularly difficult to assess⁶ and may be highly programme specific. However, these issues are likely to be of secondary importance. To the extent that expectations for future wealth (as well as price expectations) can be modelled as naive, then this may be accommodated by programme specific lags in wealth in ADL models. Alternatively leads as well as lags of programme specific wealth may be included in models.

2. GRIP

The Gross Revenue Insurance Programme (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 programme 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 programme 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 summarised as follows. 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 six crops.

Table 1. Percentage Change in Output Price due to GRIP

	wheat	Barley	canola	oats	rye	flax	total
1991	+ .4193	+ .1834	+ .0569	+ .0335	+ .3600	+ .4686	+ .2759
1992	+ .1287	+ .0373	+ .0291	+ .0026	+ .1157	+ .1051	+ .0792
1993	+ .4114	+ .1336	+ .0578	+ .0129	+ .0973	+ .2110	+ .2405
1994	+ .0338	+ .0200	+ .0029	+ .0063	+ .0208	+ .0294	+ .0235
1995	- .0093	- .0024	+ .0004	- .0014	+ .0010	+ .0129	- .0045
Avg.	+ .1968	+ .0744	+ .0340	+ .0108	+ .1190	+ .1654	+ .1229

Correlations of total indemnities with total crop market revenue, a Divisia index of market prices for the six crops, and market price for wheat over 1991-1995 are presented in Table 2. Similar correlations are presented for total indemnities minus farmer premiums. GRIP payments were very highly negatively correlated with Manitoba market revenues and prices.

6. The empirical importance of these issues is very difficult to assess. A recent USDA report summarises empirical literature on the effects of U.S. farm programmes on land values, and reports impacts ranging from 2 to 25% (Burfisher and Hopkins, p. 33). In our study, impacts of prices on decisions through changes in wealth are accommodated to some extent by the distributed lags in prices in ADL models. If prices have a substantial impact on decisions through changes in wealth, then omitting wealth variables from ADL models may lead to a significant increase in estimated coefficients for price. Omitting wealth from the ADL investment model in the corresponding Table 3A (Annex 1), the sum of lag coefficients for E_p and V_p change from +1.0903 and -0.0520 to +1.1352 and -0.0718, respectively. Thus, allowing the coefficients of price to have an indirect impact on investment through (omitted) wealth does not lead to a significant change in coefficient estimates for price. Of course such regressions are far from conclusive: programme price effects may induce changes in wealth that are not captured by correlations between total prices and total wealth.

Table 2. Correlation of GRIP payments with revenue and prices (1991-1995)

	Total Indemnities	Total Net Indemnities
Total Revenue	-.9492	-.9595
Divisia Price	-.9005	-.8846
Wheat Price	-.8813	-.8728

The estimated econometric models can be used to simulate impacts of these price supports and insurance effects on investment and output. Impacts are calculated for average levels of percentage increase in price and for 1991, when this increase was highest. This data is taken from Table 2 on price changes. Impacts of price changes are calculated using the econometric results in Annex IV. 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 Annex IV. The insurance effect indicates the impact of GRIP in truncating uncertainty and the impacts of GRIP through changes in the price variance V_p term of econometric models are not calculated. GRIP does not have a direct impact on wealth, although of course it may lead to future accumulation of wealth. Nevertheless this indirect effect of GRIP is ignored or assumed that it is implicit in our econometric models.

The calculated short-run (one year) and long run impacts of GRIP on investment in crop machinery and equipment are summarised in Table 3.

Table 3. Percentage Change in Investment due to GRIP

	Price Effect	Insurance Effect	Total Impact
Average: short run	+0.0634	0	+0.0634
long run	+0.1465	+0.0728	+0.2193
1991: short run	+0.1424	0	+0.1424
long run	+0.3289	+0.0728	+0.4017

3. Crop Insurance

Crop insurance (CI) is the longest running stabilization programme in Canada, dating back to 1939. CI programmes are designed to insure farmers against crop losses beyond their control. Individual crop yields are insured at a certain level, and payouts are received if yields fall below coverage levels. Farmer premiums are highly subsidized. Subsidy rates have varied substantially over time, and since 1996 these subsidies have risen from 1/2 to 2/3 or more of the cost of premiums in Manitoba.

CI programmes have influenced both expected returns and the variability of returns for crop producers. Changes in expected returns can be approximated from aggregate (Manitoba level) data. However, it is difficult to measure changes in variability of returns from aggregate data. The difficulty is that insurance is specific to yields at a regional level, but yield losses are far from perfectly contemporaneously correlated across regions. As a result, time series data on Manitoba level yields substantially underestimate yield variation or uncertainty at the farm level, and it is difficult to measure impacts of CI on farm level uncertainty from such data. Without a measure of CI impacts on farm level yield uncertainty, insurance effects for CI cannot be reasonably estimated. For example, a simple regression of crop output on CI payments with aggregate data will show a negative impact of CI payments on crop output, since CI payments will be positively correlated with omitted adverse weather conditions at the micro level and CI payments do not fully compensate for these conditions.

CI data aggregated at the Manitoba level was received from the Manitoba Crop Insurance Corporation for 1960-2002. This data was used to calculate impacts of CI on expected returns but not on uncertainty of returns. Table 4 shows the average impact of CI on net output prices over 1960-2002. Change in price for a crop is calculated as total payments to farmers minus farmer premiums, divided by crop output. CI provides payments to insured cropland, so strictly speaking it is a subsidy to cropland rather than to output per se. However, it is simpler to model this as a subsidy to output, and this is appropriate in our econometric models to the extent that yields are largely predetermined in the short run.

Table 4. Percentage Change in Output Price due to Crop Insurance

	Wheat	Barley	Canola	Oats	Rye	Flax	Total
Average	+0.0106	+0.0067	+0.0125	+0.0097	+0.0089	+0.0146	+0.0098

These figures are the percentage change in farm level crop output prices due to CI (indemnities minus farmer premiums) relative to market prices. Total is an index of these changes for all six crops. The average impact on prices is small: approximately 1%. The change in the total price index varies from a low of -.0137 (1990) to a high of +.0915 (1988).

Correlations between total CI net payments (total indemnities - total farmer premiums) and yields, and between percentage change in aggregate price due to CI and yields, are presented in Table 5. CI payments are somewhat negatively correlated with aggregate Manitoba yields. However, it is not clear that such correlations provide an accurate measure of CI insurance effects at the farm level.

Table 5. Correlation of Crop Insurance payments with yields (1960-2002)

	CI Net Payments	% Change in Total Price
Wheat Yield	-0.174	-0.212
Barley Yield	-0.016	-0.039
Canola Yield	-0.072	-0.086
Oats Yield	-0.073	-0.097
Rye Yield	-0.025	-0.064
Flax Yield	-0.073	-0.078

Since these calculated impacts on output price are so small, and there is not an available measure of insurance effects, we do not simulate impacts on investment. The impacts of the price effects are approximately 10% of the impacts of price effects reported in the above tables for GRIP.

4. WGS

The WGS (Western Grains Stabilization Act) of 1978-1991 essentially guaranteed net cash flow for Prairie crop agriculture at a level not below the previous five year average. The total value of the WGS payout to all farmers in a given year was determined by the difference between the net cash flow of the western region for the current year with the average of the preceding five years. Then the individual farmer's share of this payout was determined by comparing the farmer's contributions to the programme (levies) in the current and previous two years with total levies of all farmers over the same period. One third of programme costs were paid by farmers, and participation in the programme was optional. A maximum was placed on the farm level of crop receipts eligible for WGS payouts (CAD 25 000 in 1978 to CAD 60 000 in 1991).

Although the WGSA did not provide payments specific to commodities, it is possible to infer effective prices for commodities under the programme. One approach is based on viewing production as an investment in share of WGS payouts over a three year period. A simpler approach more in keeping with our models of naive price expectations is as follows. For farmers participating in the programme with crop receipts below the maximum eligible amount, define an effective representative crop price at year t as

$$(32) \quad p_t = (1 + \alpha/2) p_t - \text{prem}_t$$

where p is market price, α is an average Western Canada mark-up calculated as $\alpha = \text{WGS gross payouts} / \text{market payments}$ for all eligible crops, and $\text{prem} = \text{producer premiums} / \text{market payments}$ for all eligible crops. Approximately half of crop receipts in the programme were for farms with receipts above the maximum eligible amount, so for these farms WGSA payouts do not influence effective prices at the margin. So WGSA effective average prices can be approximated as above.

In turn, since premiums are largely known with certainty, the variance of the effective price p^{eff} is approximated as $E(p^{\text{eff}} - E p^{\text{eff}})^2 = (1 + \alpha/2)^2 E(p - E p)^2$. Thus effective price variances and covariances are related to market price variances/covariances V_p as

$$(33) \quad V p_t = (1 + \alpha/2)^2 V p_t .$$

WGSA data was collected from Western Grain Stabilization Annual Reports from Agriculture Canada. Table 6 presents calculations of α and an expected α defined as a 3 period weighted lag of observed α 's: $E \alpha_t = 0.5 \alpha_{t-1} + 0.33 \alpha_{t-2} + 0.17 \alpha_{t-3}$. Table 6 also presents the percentage change in price variances and covariances due to WGSA, calculated as $(1 + \alpha/2)^2$ from (33). The column with heading DEp is an aggregate index (over six crops) of the percentage changes in prices due to WGSA, calculated from (32). Percentage changes in individual prices are almost identical for all six crops.

Alternatively producers may not formulate effective price impacts for WGSA. Then WGSA payments should be viewed as contributing directly to wealth and creating insurance effects. WGSA payments were included in initial wealth W_0 for estimation of econometric models. WGSA net payments (payments - farmer premiums) as a percentage of initial wealth are indicated in the last column Table 6. Thus WGSA payments are generally a small fraction of wealth, so the WGSA is unlikely to have large wealth effects on decisions.

Table 6. Some data on the WGSA Programme

	A	E α	DVp	DEp	WGSA Net Payments/Initial Wealth W ₀
1978	.0218		.0219	.0109	.0068
1979	.0522		.0529	.0261	.0157
1980	0		0	-.00003	-.0035
1981	0	.0210	0	-.00003	-.0028
1982	0	.0088	0	-.00003	-.0027
1983	0	0	0	-.00003	-.0033
1984	.0284	0	.0286	.0142	.0080
1985	.0636	.0142	.0646	.0318	.0238
1986	.1694	.0412	.1766	.0847	.0472
1987	.3268	.1105	.3535	.1634	.0889
1988	.1542	.2301	.1602	.0771	.0427
1989	.0311	.2138	.0313	.0154	.0017
1990	0	.1220	0	-.0001	-.0073
1991	.0305	.0308	.0365	.0152	.0031
Avg.	.0627	.0725	.0657	.0313	.0156

Correlations between WGSA payments and market revenues and prices are indicated in Table 7. Correlations with percentage change in effective prices due to WGSA are indicated in the second column.

Table 7. Correlations of WGSA payments with revenue and prices (1978-1991)

	WGSA Payouts	DEp
Total Revenue	-2743	-.3689
Divisia Price	-.4185	-.4427
Wheat Price	-.4599	-.4882

However, there are sufficient observations to isolate a WGSA insurance effect using the methodology adopted in the econometric model.

First, the relation between insurance effects and total income uncertainty should be clarified. Assuming that uncertainty is limited to programme payments G and commodity prices p , then the variance measure of income uncertainty $V(G+\pi)$ can be expressed as

$$(33) \text{Var}(G+py) = \text{Var}(G) + y^T V_p y + 2 \text{Cov}(G,p) y$$

where output levels y are assumed to be nonstochastic. More generally (assuming y and p are distributed independently) $\text{Var}(py) = y^T V_p y + p^T V_y p + \sum_i \sum_j V_{p_{ij}} V_{y_{ij}}$, but output uncertainty V_y cannot be measured from output data y aggregated over agents (since weather is not perfectly contemporaneously correlated across farms). Therefore only a measure of (33) is calculated.

Variance measures of income uncertainty (33) are calculated under alternative scenarios for government payments G , assuming that historical data on y and price covariances V_p remain constant

under different scenarios. Of course this only provides (at best) a first order approximation to impacts of programmes on income uncertainty. Nevertheless these results may be somewhat suggestive about the magnitudes of insurance effects.

Table I.9 in Annex I reports measures of income uncertainty (33) over 1982-1991 under two scenarios: (1) payments G are limited to the WGSA and (2) payments G include WGSA plus miscellaneous programmes operating at the time. Since WGSA premiums were largely known with certainty, they are not subtracted from payments in application of (33). The covariance between payments and income, $Cov(Gp)y$, is always negative for WGSA only, as is expected for a countercyclical programme. WGSA is also the primary influence on $Cov(Gp)y$ when miscellaneous programmes are included. The average reductions in our measure of income uncertainty due to WGSA are 16.42% and 18.20% for the two scenarios, respectively. The maximum reductions in income uncertainty due to the WGSA are calculated as 45.15% in 1989, and 42.8% in 1987.

In order to specify a WGSA specific wealth variable in econometric models, WGSA net payments were separated out of initial wealth W_0 . Final models for investment, yield and acreage demand were re-specified with two wealth variables, one for WGSA net payments and the other for W_0 less WGSA payments. A four period lag on WGSA net payments was added to final ADL investment models, but WGSA payments were statistically insignificant. A similar lag was statistically significant in some ADL yield equations, but with a negative sign. So this variable does not seem to be measuring a wealth effect. A possible explanation is that WGSA payments do not fully compensate for low net cash flows, and low net cash flows are not fully measured in the rest of the model. Then estimated coefficients for WGSA payments may largely reflect the impact of an omitted variable, low net cash flow. WGSA payments were statistically significant in the canola acreage equation, with an elasticity of +0.0020 and a t-ratio of 2.70. However, WGSA payments were not statistically significant in other acreage demands. In sum, including a separate WGSA payment variable in econometric models does not seem to elucidate effects of WGSA on decisions.

The impacts of WGSA on investment are now calculated based on the alternative scenarios of price effects and of wealth/insurance effects. Under the first scenario, the calculated short run (one year) and long run impacts of WGSA on investment in crop machinery and equipment are summarised in Table 8. Results are presented under average conditions and for 1987, when WGSA presumably had the largest impact on effective prices.

Table 8. Percentage Change in Investment due to WGSA: Scenario 1

	Price Effect	Price Variance Effect	Total Impact
Average: short run	+0.0162	-0.0018	+0.0144
long run	+0.0373	-0.0033	+0.0340
1987: short run	+0.0843	-0.0098	+0.0745
long run	+0.1948	-0.0257	+0.1691

Next consider the second scenario. Here the WGSA has wealth and insurance effects. Impacts on investment are calculated from econometric models without WGSA specific variables. The calculated impacts are presented in table 9.

Table 9. Percentage Change in Investment due to WGSA: Scenario 2

	Wealth Effect	Insurance Effect	Total Impact
Average: short run	+0.0096	0	+0.0096
long run	+0.0131	+0.0120	+0.0251
1987: short run	+0.0547	0	+0.0547
long run	+0.0747	+0.0312	+0.1059

Wealth and insurance variables were statistically insignificant in yield and acreage equations for canola.

5. Western Grain Transition Payments

The federal government has a long history of regulating rail rates and subsidising rail transportation of grains for farmers. These subsidies have been substantial and have been criticised as export subsidies. With the demise of the Western Grains Transportation Act (WGTA) in 1995, such subsidies to crop producers/railways were eliminated. Annual subsidy payments to railways to hold down prices for farmers averaged CAD 705 million across the prairies from 1985-1995, but farmers only received compensation of CAD 1.6 billion for termination of the subsidies. Given relatively high farm incomes for 1995-96, this increase in wealth was probably retained in large part for several years.

In the econometric models, estimated costs to farmers of moving grains from a mid-prairie point to export position (Thunder Bay, although similar results hold for Vancouver) are subtracted from product price. CAD /tonne charges to farmers for transport from 1991 to 2001 were 11.07, 11.98, 13.73, 14.72, 31.23, 33.46, 34.13, 33.81, 32.46, 33.10, 34.43 (the Crow rate through 1982 was 5.09) (Canadian Grain Commission). From 1994 to 1995, farmer transport costs increased by CAD 16.51/tonne, from CAD 14.72 to CAD 31.23, with elimination of the WGTA. Assuming as an approximation that elimination of the WGTA increased farmer transport costs by CAD 16.51/tonne each year, the percentage change in farm prices due to elimination of the WGTA is calculated as in Table 10. Here, 16.51 is divided by farm price including market price, net payouts from GRIP and Crop Insurance, and transport costs.

Table 10. Percentage Change in Output Price due to Elimination of WGTA

	Wheat	Barley	Canola	Oats	Rye	Flax	Total
1995	.1252	.2456	.0527	.2469	.2416	.0685	.1697
1996	.0958	.1272	.0491	.1474	.1669	.0607	.1088
1997	.1299	.1972	.0478	.1603	.1406	.0554	.1430
1998	.1458	.2350	.0492	.1809	.1346	.0548	.1629
1999	.1476	.2653	.0514	.2094	.2703	.0591	.1785
2000	.1855	.2756	.0834	.2829	.4846	.0925	.2181
2001	.1956	.2230	.0794	.2765	.4635	.0791	.2077
2002	.1447	.1678	.0591	.1053	.1970	.0555	.1351
avg.	.1463	.2171	.0590	.2012	.2624	.0657	.1655

The compensation received by Manitoba crop producers for elimination of WGTA is indicated in Table 11. For example, 1995-96 Transition payments of CAD 147 million are designated as contributing to

initial wealth for 1996. Table 11 also indicates the ratio of this compensation relative to other wealth, defined as the value of cropland plus payments from non-commodity specific programmes).

Table 11. Compensation under the WGTA

	Transition Payments (CAD 1 000)	Ratio to Wealth
1996	147040	.0344
1997	109263	.0239
1998	516	.0001

The impacts of the WG Transition on investment are now calculated based on the above price and wealth effects. The calculated short run (one year) and long run impacts of the Transition on investment in crop machinery and equipment are summarised in table 12 (estimation 1). Results are presented under average conditions and for 1996, the first year of Transition payments to farmers. The wealth effect for 1996 is calculated under two scenarios: (a) change in wealth is limited to current year payments, and (b) change in wealth includes payments for all three years (1996-98). The second scenario assumes that in 1996 farmers correctly forecast future payments and include them in perceived wealth for 1996. This seems a reasonable assumption.

Table 12. Percentage Change in Investment due to WG Transition, estimation 1

	Price Effect	Wealth Effect		Total Impact	
avg.: short run	-0.0854	+0.0046		-0.0808	
long run	-0.1973	+0.0062		-0.1911	
1996: short run	-0.562	(a)+.0212	(b)+.0359	(a)-.0354	(b)-.0203
long run	-0.1297	(a)+.0289	(b)+.0491	(a)-.1008	(b)-.0806

Since there is some ambiguity in interpreting the initial wealth variable W_0 , models are re-estimated treating WG Transition payments and other wealth as separate variables (first differenced). A four period lag on both wealth variables (normalised by a variable crop input price index) was specified within the final ADL investment model, and the lag on WG payments was then reduced to one. OLS estimates and t-ratios for sum of lag coefficients were +0.8666 (2.87), -0.0508 (1.72), -0.0691 (1.75), +0.6000 (1.68) for $E_p, V_p, Cov(G_p), W_0$ minus Transition payments, respectively, and +0.0311 (2.28) for Transition payments. Although this last elasticity is small in comparison to elasticities for initial wealth, it can be attributed entirely to Transition payments. Thus Transition payments are estimated to increase investment by 3.11% in 1996. This is very similar to the wealth effect on investment calculated above.

Based on these revised wealth effects, the impacts of the WG Transition circa 1996 are calculated as in Table 13.

Table 13. Percentage Change in Investment due to WG Transition, estimation 2

	Price Effect	Wealth Effect	Total Impact
1996: short run	-0.0562	+0.0311	-0.0251
long run	-0.1297	+0.0311	-0.0986

6. NISA

The stated purpose of the Net Income Stabilization Account (NISA) programme of 1991-2002 was to encourage farmers to save more funds in high income years for use in low income years, so as to smooth incomes over time. This is similar to the objective of the WGSA. Unlike WGSA, each participant has an individual NISA account. NISA was originally designed for grain and oilseed farms but was later extended to other sectors.

A farmer can contribute up to 2 or 3% (more recently 3) of eligible net sales as savings to a Fund 1 NISA account, and the federal and provincial governments generally make a matching contribution to a Fund 2 NISA account for the individual. Matching funds earn a competitive interest rate, but the farmer's own deposits receive a 3% interest bonus paid by government. In addition a farmer can contribute up to an additional 20% of eligible net sales to his Fund 1 NISA account. Such additional farmer contributions are not matched by government, but these additional contributions do receive the 3% interest bonus from government. Farmer contributions are not tax deductible. All interest from both accounts is accumulated in Fund 2. There is a maximum net eligible sales of CAD 250 000 that can qualify for matching funds. According to unpublished AAFC data for NISA in 1995, 54.5% of NISA participants across Canada exceeded these maximum net eligible sales. 9% of participants made contributions of less than 3% of net eligible sales (Rude).

In years of declining income, farmers can withdraw funds from their accounts in amounts determined by either one of two trigger mechanisms. A stabilization trigger permits withdrawals based essentially on average incomes for the last five years. A second trigger permits withdrawals based on a minimum income threshold. There is ambiguity about classifying NISA as a countercyclical programme. Savings accumulate during high income years, and incomes are smoothed during low income years. This is referred to as a countercyclical programme.

Table 14 summarises the most important available data on NISA for our purposes. Data on fund balances and participants was obtained from NISA (balances as of March 31), and data on withdrawals from Fund 2 is from Agriculture Economics Statistics (Statistics Canada cat. no. 21-603). All data is in CAD 1 000. Fund ratio is the ratio of NISA balances (Funds 1 and 2) to other measures included in initial wealth W_0 , primarily the value of cropland and buildings. The average fund ratio over this period was .07785, *i.e.* 7.785%. According to this data, withdrawals from Fund 2 are often approximately 10-20% of balances in Fund 2 (withdrawals from Fund 1 are less). There are many anecdotes about farmers using NISA solely as a savings programme for retirement. This data suggests that there are also many farmers using NISA to smooth income over time, as was intended for the programme.

In principle correct modelling of impacts of NISA on production requires a dynamic farm household model (Chen). However, this is not practical using data currently available for Canada. The framework used corresponds to a dynamic production model where production decisions are separable from consumption decisions. In addition our econometric models have been limited to crops, whereas NISA includes livestock as well as crops.

Given access to NISA, a producer can be viewed as solving a stochastic optimal control problem. The producer's choice variables relate to production, investment, savings in and outside of NISA, and (contingent on events) withdrawals from NISA and other savings accounts. Such a model implies that savings/wealth decisions and withdrawal/smoothing decisions are in effect joint aspects of the complex dynamic "plan" (accommodating option value) formulated by the firm in response to access to NISA. In other words, such a model would generally be mis-specified by assuming a recursive structure where NISA related savings decisions are made independently or separable from income smoothing decisions. More generally, NISA related savings and withdrawal decisions are jointly related to production decisions. So in

order to model the impact of NISA on production, modelling the impact of both savings and withdrawal decisions on production is required.

Table 14. Balance and withdrawals from NISA funds

	Fund 1	Fund 2	Fund 2	Fund
Participants	Balance CAD	Balance CAD	Withdrawals CAD	Ratio
1991	18 791	42 137*	323	.01320
1992	17 808	36 981	62 278	.02202
1993	16 746	57 394	24 046	.03309
1994	17 947	89 990	4 152	.04829
1995	18 720	131 882	4 465	.06975
1996	18 775	167 475	14 714	.08082
1997	19 061	199 769	20 065	.09221
1998	19 314	214 043	29 678	.09668
1999	19 551	230 457	58 267	.11312
2000	19 508	248 525	61 628	.10734
2001	20 393	271 718	58 642	.13636
2002	20 530	228 330	77 725	.12127

* For 1991, this figure refers to the total for Fund 1 and Fund 2.

As noted above, the contributions of NISA to savings/wealth can be classified as procyclical (since contributions occur when incomes are high) whereas withdrawals from savings are countercyclical. This does not diminish or negate the importance of withdrawals. NISA related contributions to wealth (whether procyclical or countercyclical) generally increase investment and production (assuming decreasing absolute risk aversion). The income smoothing effect of withdrawals when incomes are low is an insurance effect that also presumably increases investment and production. In addition, NISA also presumably leads to price effects that influence investment and production decisions. This is because increases in market revenues generally permit farmers to increase NISA savings and subsidies, as discussed above. These price effects are relatively simple to calculate, in contrast to NISA savings and income smoothing effects. These price effects can be substantial for certain participants, but they are minor for most participants.

In response to the NISA programme, a participant will place savings in his Fund 2 and receive matching funds in his Fund 1. Since this savings is so heavily subsidized, a participant will presumably divert (substitute) some outside savings into the NISA programme, *i.e.* savings in Fund 1 plus Fund 2 will overestimate the impact of NISA on total savings. Furthermore NISA with its matching funds may lead to a decrease in individual savings efforts, *i.e.* matching savings in Fund 1 is not necessarily a lower (or upper) bound on the impact of NISA on total savings. Presumably NISA leads to some increase in total savings (NISA plus other accounts).

In response to the NISA programme, the above data suggests that many participants will withdraw some funds from NISA accounts when incomes are low. Since savings in Fund 2 receive no interest subsidies, withdrawals from these funds presumably substitute in part for withdrawals from other savings in the absence of NISA. So withdrawals from Fund 1 plus Fund 2 presumably overestimate the impact of NISA on total withdrawals. Given the substantial interest rate subsidies to savings in Fund 1, participants will be reluctant to withdraw savings from this fund. So withdrawals from Fund 1 may underestimate the impact of NISA on total withdrawals (these withdrawals are negligible). Presumably NISA leads to some increase in total withdrawals.

In addition, NISA presumably has impacts on effective output prices for at least some producers, *i.e.* participants with eligible net sales of less than the maximum of CAD 250 000 but contributing at least 3% of net sales to NISA. For these participants, market revenue constrains NISA subsidies. Then a CAD 100 increase in revenue permits the farmer to make an additional CAD 3 contribution to his Fund 1 NISA account. There are three possible NISA subsidy effects associated with this CAD 100 increase in market revenue for this producer. First, government provides a matching CAD 3 to his Fund 2 NISA account. Second, government provides a 3% interest bonus on the additional savings in Fund 1. This bonus above competitive market/discount rates implies that an additional CAD 1 in Fund 1 has a present value of CAD 1.344 in terms of interest over a 10 year horizon ($1.03^{10} = 1.344$), and a present value of CAD 1.159 in terms of interest over a five year horizon ($1.03^5 = 1.159$). Thus the additional CAD 3 in Fund 1 leads to a present value subsidy of CAD 1.032 ($3 \times .344$) over a ten year savings horizon and a present value subsidy of CAD 0.477 ($3 \times .159$) over a five year savings horizon. Third, suppose this individual is also contributing an additional 20% to NISA. Presumably this is a relatively small number of participants. This participant can contribute up to an additional CAD 20 (20% of market revenue) to his Fund 1 NISA account. There are no matching government funds, but these contributions do qualify for the 3% interest bonus. Then this additional CAD 20 in Fund 1 leads to a present value subsidy of CAD 10.32 ($20 \times .344$) over a ten year savings horizon and a present value subsidy of CAD 4.77 over a five year savings horizon.

In sum, an additional CAD 100 of market revenue may lead to the following maximum additional subsidies under these scenarios: CAD 14.352 ($3 + 1.032 + 10.32$) over a ten year savings horizon and CAD 8.247 ($3 + .477 + 4.77$) over a five-year savings horizon. These NISA subsidies associated with market revenue suggest, under conditions noted above, the following increases in effective (shadow) output prices relative to market prices.

Table 15. Increase (percentage) in output prices due to NISA

Savings Horizon	Matching Funds	Interest Subsidy A	Interest Subsidy B	Total
Ten years	.03	.01032	.1032	.14352
Five years	.03	.00477	.0477	.08247

Interest subsidy A relates to the savings (3% of additional revenues) that qualify for matching contributions, and interest subsidy B relates to the savings (20% of additional revenues) not qualifying for matching contributions. Changes in effective prices due to matching funds and interest subsidy A apply to participants with net sales below the eligible maximum of CAD 250 000 and contributing 3% or more of their sales to NISA. This may be approximately 40% of participants circa 1995. Changes in effective prices due to interest subsidy B only apply to participants with net sales below the eligible maximum and contributing an additional 20% of sales to NISA accounts. Presumably this is a much smaller number of participants. Note that there is also some uncertainty associated with NISA impacts on effective prices. Although the matching contribution and interest rate subsidy rules are known with certainty, market revenue and in turn NISA subsidies associated with given output levels are uncertain. However, this is presumably minor.

Withdrawals from Fund 2 are used as an approximate upper bound on the impact of NISA on total withdrawals (from NISA and other funds). This is justified simply as follows: withdrawals from NISA funds presumably overestimate the impact of NISA on total withdrawals, and withdrawals from Fund 1 are negligible. Consequently withdrawals from Fund 2 were included in calculations of insurance effects. Here insurance impacts of NISA are modelled in terms of annual withdrawals from Fund 2. Impacts of these withdrawals on income uncertainty are calculated as in (33) for WGSA above. Table I.10 in Annex I reports measures of income uncertainty under NISA and other programmes. The table covers 1995-2002 (results are not obtained for earlier NISA years because our approximations to variances and covariances

related to NISA withdrawals involve four year lags). Estimated covariances between NISA withdrawals and Manitoba income, $Cov(Gp)_y$, are always negative, as is expected for a programme that is countercyclical. When other programmes are considered in addition to NISA, covariances remain negative except in 2002.

Table 16 summarises total covariance between government expenditures G and market revenues, $Cov(Gp)_y$, and covariances between G and individual prices, $Cov(Gp)_i$. G_A includes miscellaneous programme payments, AIDA payments and NISA withdrawals from Fund 2. G_B excludes NISA. The mean is over 1995-2002. Units are CAD 1 000. As anticipated, NISA withdrawals contribute substantially to a negative covariance between G and market revenues and prices: in all cases the covariance is substantially more negative due to NISA withdrawals. The changes are also calculated in the table as $(G_A - G_B)/G_B$.

Table 16. Covariance of G with Revenues and Prices

	Market	Wheat	Barley	Canola
	Revenue	Price	Price	Price
	$Cov(Gp)_y$	$Cov(Gp)_1$	$Cov(Gp)_2$	$Cov(Gp)_3$
G_A: 2000	-22676	-1739892	-305244	-7980387
2001	-10479	-1088610	+221201	-5021256
mean	-4664	-415730	+100412	-2036600
G_B: 2000	-17473	-1346426	-205870	-6187956
2001	-7967	-843308	+243397	-3873253
mean	-2595	-170740	+121830	-1430700
Percentage change in Covariance				
2000	.2978	.2922	.4827	.2897
2001	.3153	.2909	.0912	.2964
mean	.7973	1.8038	.1758	.4235

The impacts of NISA on investment are now calculated based on the above scenarios of price effects, wealth effects and insurance effects. Calculated short run (one year) and long run impacts of NISA on investment in crop machinery and equipment are summarised in table 17. Results are presented under average conditions and for 2001. Price A refers to the output price effect under a five-year savings horizon, and price B refers to output price effect under a ten year savings horizon. Sum0 refers to the case of zero price effects. Since price effects do not apply to the majority of producers, impacts representative for Manitoba are most likely to be approximated by Sum0.

Table 17. Percentage Change in Investment due to NISA

	PriceA	PriceB	Wealth	Insurance	Sum0	SumA	SumB
avg.: short run	.0426	.0741	.0479	0	.0479	.0905	.1220
long run	.0983	.1711	.0654	.0580	.1234	.2247	.2945
2001: short run	.0426	.0741	.0839	0	.0839	.1266	.1580
long run	.0983	.1711	.1146	.0230	.1376	.2359	.3087

Given these ambiguities, and ambiguities about price effects, it is also attempted to estimate directly the impact of NISA on investment, yield and acreages. The final ADL investment equation (6) is re-specified as follows:

$$(30) I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,1,2} \beta_{1i} P^{Div}/w_{t-i} + \sum_{i=0,1} \beta_{2i} VP^{Div}/w_{t-i}^2 + \sum_{i=0,..,4} \beta_{3i} W^0/w_{t-i} + \sum_{i=3,4} \beta_{4i} COVGP/w_{t-i}^2 + \sum_{i=0,1,2} \alpha_{1i} D^{NISA} P^{Div}/w_{t-i} + \sum_{i=0,..,4} \alpha_{3i} W^{NISA}/w_{t-i} + \sum_{i=3,4} \alpha_{4i} With^{NISA}/w_{t-i} + e_t$$

with polynomials as before. D^{NISA} is a dummy variable for NISA (= 1.0 for 1991-2002). Since the percentage change in output price due to NISA is calculated as a constant for all years, the terms $D^{NISA} P^{Div}/w_{t-i}$ in (30) capture in principle this impact on investment. W^{NISA} is savings in Fund 1 plus Fund 2 NISA accounts, and $With^{NISA}$ is withdrawals from Fund 2. W^0 and $COVGP$ were recalculated so as to exclude effects of NISA. All variables (excluding D^{NISA}) are in logs, and all variables are first differenced.

There are two somewhat plausible interpretations of the terms W^{NISA} and $With^{NISA}$ in the investment equation: they measure the true impacts of NISA on total savings and withdrawals with error, and they represent a “reduced form” for investment conditional on these variables. By the equivalence of linear measurement error models and linear simultaneous equations models, both these interpretations are valid. In both cases OLS estimators are inconsistent. Moreover in the first case, standard instrumental variables (IV) procedures will not lead to consistent estimators of coefficients in the investment equation for true impacts of NISA on total savings and withdrawals, since W^{NISA} and $With^{NISA}$ include biased measurement errors. Regarding the second approach, there is difficulty in finding adequate instruments to avoid the weak instruments problem. These considerations illustrate the difficulty in estimating a model with NISA specific variables such as (30).

OLS estimates of (30) led to the following estimates of sum of lag coefficients for NISA specific variables (no polynomial restrictions on these variables): +1.4673 (1.24), -0.471 (1.82) for W^{NISA} , $With^{NISA}$, and +1.2347 (1.70) for DUM^{NISA} . The negative coefficient on $With^{NISA}$ is not surprising: although this variable smoothes income which should increase investment, this variable is positively correlated with low incomes which are not completely specified in the regression model. It is difficult to find an instrument that is highly correlated with NISA withdrawals yet is uncorrelated with incomes. Estimated coefficients for W^{NISA} and DUM^{NISA} have the anticipated signs but are not statistically significant. Note that IV will not correct this problem: IV leads to consistent estimators but with larger standard errors than for OLS. These ambiguous results are unchanged under alternative specifications of (30).

7. Conclusion Regarding the Simulated Impacts

Estimates of our econometric models incorporating price, price variance, wealth and insurance effects have been applied to simulate impacts of the major government programmes for crop agriculture over the last thirty years. These include GRIP, Crop Insurance, WGSA, WGTA, and NISA. Apparently this is the first study to attempt to simulate impacts of the WG Transition and NISA from an econometric model. Indeed this may be the first study to attempt to simulate impacts on investment from an econometric model for any of these programmes.

GRIP was a relatively standard price support programme, and its price effects are easily simulated. The average percentage change in price due to GRIP is calculated here as +12.3%. In addition, there are significant insurance effects. The average long run impacts on investment are calculated as +14.6% (price effect) and +7.3% (insurance effect), for a total of 21.9%.

Crop Insurance impacts are difficult to measure. Price effects are minor, but data is inadequate to measure farm level yield uncertainty and the effective reduction/compensation for yield uncertainty. There

was some suggestion from econometric results how full compensation for remaining uncompensated losses might influence acreage demands and investment.

Although the WGSA did not provide commodity specific payments, it is possible to infer price effects. The change in prices is calculated as +3.1% over the programme, and the corresponding estimated change in long run investment is +3.7%. In the absence of price effects, percentage increase in wealth is calculated as +1.6%. The average percentage change in investment is estimated as +1.3% (wealth effect) plus +1.2% (insurance), for a total of +2.5%.

Historically Western crop producers have received substantial subsidies for transport. The econometric models are used to simulate impacts of the WG Transition; ignoring secondary effects on prices (this would require a full general equilibrium model). This Transition essentially removed transport subsidies while providing only partial compensation in terms of interim payments. Elimination of WGTA led to an average decrease of 16.5% in crop prices over 1995-2002 (the smallest decrease is for canola, 5.9%), and to a 10.9% average decrease in price for 1996. Transition payments were 3.4% of wealth in 1996 and 2.4% of wealth in 1997 (0.7% over 1995-2002). The average long run impacts on investment over 1995-2002 for this Transition were calculated as -19.7% (price effect) and +0.6% (wealth effect), for a total effect of -19%. Long run impacts on investment for 1996 were calculated as -13% (price effect) and +4.9% (wealth effect), for a total effect of -8.1%. These calculations for 1996 assume that farmers factor future payments into perceived wealth. Since there is ambiguity in interpreting wealth effects in the econometric model, the model was re-estimated with Transition payments treated as a separate wealth variable. This led to similar simulated wealth impacts as for the original econometric model without separate Transition specific variables.

Impacts of NISA are much more difficult to simulate with an econometric model than are impacts of GRIP, WGSA and WG Transition payments. This is largely due to difficulties in measuring impacts of NISA on total wealth and total withdrawals. NISA fund balances and withdrawals from Fund 2 are used as upper bounds on these impacts. NISA fund balances average 12% of wealth over 1991-2002, and withdrawals from Fund 2 show a high negative covariance with market prices and revenues. Thus NISA may have significant wealth and insurance effects. There can also be substantial price effects for some participants, but price effects seem likely to be minor for most participants. Attempts to estimate econometric models with separate NISA specific variables (in order to calculate NISA effects directly) were inconclusive. Simulation results based on econometric models (without separate NISA variables) and the above assumptions can be summarised as follows. The average impact of NISA on investment over 1991-2002 is calculated as +6.5% (wealth effect) and +5.8% (insurance effect), for a total of 12.3%. Interpreting these effects as upper bounds, it is concluded that NISA may well increase investment, but these effects are less pronounced than for programmes such as GRIP.

*Annex I.***Main Estimations of ADL Investment Model****Table I.1. ADL Investment**

OLS (first differences)

		A		B		C	
	lag	coef	t-stat	coef	t-stat	coef	t-stat
I	1	.0478	0.32	-.0473	0.31	.0348	0.22
P ^{Div}	0	.2751	1.95	.4418	2.80	.3116	1.91
P ^{Div}	1	.2369	2.16	.3832	3.10	.2537	2.07
P ^{Div}	2	.1987	1.34	.3246	2.14	.1958	1.28
VP ^{Div}	0			-.0310	1.44		
VP ^{Div}	1			-.0440	2.14		
VP ^{Bad}	0					-.0186	0.87
VP ^{Bad}	1					-.0114	0.46
t		-.0029	1.29	-.0033	1.51	-.0030	1.32
constant		.0865	1.55	.1044	1.94	.0899	1.58
Sum of lag coefficients:							
P ^{Div}		.7107	2.15	1.1496	3.10	.7611	2.07
VP ^{Div}				-.0750	2.26		
VP ^{Bad}						-.0301	0.94
R ²		.1920		.3050		.2154	
rho		.132		.141		.107	
F - test of polynomial restrictions:							
		F (1,35) = .450		F (1,33) = .282		F(1,33) = .713	
		Prob = .507		Prob = .599		Prob = .404	

Table I.2. ADL Investment

OLS (first differences)

		A		B	
	lag	Coef	t-stat	coef	t-stat
I	1	.0573	0.36	.0493	0.27
P ^{Div}	0	.3865	2.57	.4941	2.39
P ^{Div}	1	.3457	2.91	.2326	1.31
P ^{Div}	2	.3050	2.11	.4098	2.03
VP ^{Div}	0	-.0214	1.06	-.0260	1.21
VP ^{Div}	1	-.0272	1.26	-.0210	0.90
W ₀	0	.5670	1.80	.6243	1.86
W ₀	1	.1679	2.05	.1586	0.48
W ₀	2	.0209	1.58	.0201	1.46
W ₀	3	.0064	0.52	.0063	0.49
W ₀	4	.0047	0.39	.0029	0.23
Cov Gp	3	-.0166	0.60	-.0141	0.49
Cov Gp	4	-.0591	2.18	-.0607	2.18
t		.0004	0.15	.0004	0.14
Constant		-.0163	0.23	-.0158	0.22
Sum of lag coefficients:					
P ^{Div}		1.0372	2.91	1.1364	2.77
VP ^{Div}		-.0487	1.46	-.0470	1.40
W ₀		.7669	1.95	.8122	1.70
Cov Gp		-.0758	1.73	-.0748	1.66
R ²		.5047		.5191	
rho		.178		.153	
F - test of polynomial restrictions:					
			F (2,26) = .388	no restrictions	
			Prob = .682		

Table I.3. DL Investment

(first differences)

		A (OLS)		B (OLS)		C (auto)	
lag		coef	t-stat	coef	t-stat	coef	t-stat
P ^{Div}	0	.4053	2.91	.5162	2.78	.4174	2.87
P ^{Div}	1	.3634	3.40	.2446	1.45	.2648	2.03
P ^{Div}	2	.3216	2.37	.4311	2.36	.4319	2.98
VP ^{Div}	0	-.0234	1.22	-.0276	1.36	-.0229	1.46
VP ^{Div}	1	-.0286	1.37	-.0221	0.98	-.0243	1.46
W ₀	0	.5645	1.82	.6153	1.87	.6973	2.84
W ₀	1	.1637	2.05	.1962	0.68	.2658	1.16
W ₀	2	.0191	1.58	.0186	1.51	.0202	1.91
W ₀	3	.0073	0.61	.0072	0.59	.0081	0.79
W ₀	4	.0050	0.41	.0032	0.26	.0051	0.53
Cov Gp	3	-.0154	0.56	-.0128	0.46	-.0261	1.21
Cov Gp	4	-.0584	2.20	-.0600	2.21	-.0647	3.06
t		.0002	0.08	.0003	0.12	.0010	0.40
Constant		-.0110	0.17	-.0144	0.20	-.0381	0.53
Sum of lag coefficients:							
P ^{Div}		1.0903	3.40	1.1920	3.43	1.1141	3.66
VP ^{Div}		-.0520	1.65	-.0497	1.54	-.0473	1.81
W ₀		.7595	1.96	.8405	1.83	.9966	2.61
Cov Gp		-.0738	1.72	-.0728	1.66	-.0908	2.48
R ²		.5024		.5178		.5442	
rho		.208		.179		.078	
DW		1.58		1.64		1.83	
F - test of polynomial restrictions:							
		F (2, 27) = .430 Prob = .654		no restrictions		no restrictions	

Table I.4. ADL Investment

OLS (first differences)

		A		B	
	lag	Coef	t-stat	Coef	t-stat
I	1	.0009	0.01	-.0029	0.01
P ^{Div}	0	.3497	2.43	.4222	2.07
P ^{Div}	1	.3223	2.85	.2557	1.41
P ^{Div}	2	.2948	2.17	.3603	1.80
VP ^{Div}	0	-.0319	1.55	-.0319	1.39
VP ^{Div}	1	-.0209	1.02	-.0168	0.71
W ₀	0	.5290	1.78	.5824	1.58
W ₀	1	.1589	2.07	.1902	0.52
W ₀	2	.0238	1.86	.0231	1.70
W ₀	3	.0111	0.95	.0112	0.88
W ₀	4	.0085	0.73	.0074	0.58
CovGp	3	-.0130	0.50	-.0113	0.40
CovGp	4	-.0579	2.28	-.0589	2.18
t		.0016	0.67	.0018	0.66
Tcost	1	.0062	0.39	.0059	0.30
Tcost	2	.0217	2.12	.0252	1.32
Tcost	3	.0373	2.26	.0322	1.42
constant		-.0633	0.90	-.0669	0.87
Sum of lag coefficients					
P ^{Div}		.9669	2.85	1.0382	2.58
VP ^{Div}		-.0528	1.63	-.0487	1.32
W ₀		.7313	1.98	.8143	1.75
CovGp		-.0710	1.72	-.0702	1.60
Tcost		.0652	2.12	.0634	1.86
R ²		.5952		.6008	
rho		.060		.065	
F-test of polynomial restriction					
		F(3,23) = .1069		no restrictions	
		Prob = .955			

Table I.5. DL Investment

OLS (first differences)

		A		B	
	lag	Coef	t-stat	Coef	t-stat
P ^{Div}	0	.3500	2.65	.4212	2.22
P ^{Div}	1	.3226	3.18	.2547	1.53
P ^{Div}	2	.2951	2.32	.3593	1.96
VP ^{Div}	0	-.0320	1.70	-.0317	1.53
VP ^{Div}	1	-.0209	1.06	-.0167	0.75
Wo	0	.5290	1.81	.5836	1.66
Wo	1	.1588	2.13	.1876	0.60
Wo	2	.0237	2.06	.0232	1.90
Wo	3	.0111	0.99	.0111	0.93
Wo	4	.0085	0.74	.0073	0.60
CovGp	3	-.0130	0.51	-.0114	0.42
CovGp	4	-.0579	2.33	-.0589	2.24
t		.0016	0.70	.0018	0.68
Tcost	1	.0062	0.42	.0060	0.31
Tcost	2	.0217	2.16	.0253	1.38
Tcost	3	.0373	2.37	.0321	1.55
constant		-.0632	0.94	-.0670	0.89
Sum of lag coefficients					
P ^{Div}		.9677	3.18	1.0352	3.00
VP ^{Div}		-.0529	1.77	-.0485	1.47
Wo		.7312	2.02	.8130	1.82
CovGp		-.0709	1.77	-.0703	1.66
Tcost		.0652	2.16	.0633	1.91
R ²		.5952		.6008	
rho		.061		.064	
DW		1.87		1.86	
F-test of polynomial restrictions					
			F(3,24) = .1115	no restrictions	
			Prob = .952		

Table I.6. Euler Equations for Investment
GMM (first differences)

		A		B	
	lag	Coef	t-ratio	Coef	t-ratio
I	1	-.1347	1.23	-.2280	2.30
P ^{Div}	1	.0281	0.72	.1839	2.23
P ^{Div}	2	-.0428	1.13	-.1127	1.19
V ^{P^{Div}}	0	--		-.0081	2.12
V ^{P^{Div}}	1	--		.0035	0.87
W ₀	0	--		.1174	0.29
W ₀	1	--		.6976	1.75
K	0	2.9174	5.56	3.4457	13.20
K	1	-1.8574	4.95	-2.1894	11.01
t		-.0019	0.91	-.0014	0.86
constant		.0544	1.18	.0269	0.75
R ²		.6358		.6586	
rho		-.029		-.124	
J-test of over identifying restrictions:					
		$\chi^2(6) = 1.587$		$\chi^2(8) = 2.701$	

Table I.7. Euler Equation for Investment
GMM (first differences)

		A		B	
	lag	coef	t-ratio	coef	t-ratio
I	1	-.1735	1.63	--	
P ^{Div}	0	.2027	3.08	.1894	3.14
VP ^{Div}	0	-.0096	3.85	-.0091	4.35
W ₀	0	.1372	0.48	.2470	1.01
W ₀	1	.6341	1.64	.2903	0.95
K	0	2.9875	6.26	2.4821	7.58
K	1	-1.9402	5.83	-1.7539	5.44
t		-.0016	0.81	-.0025	1.46
Constant		.0344	0.72	.0555	1.34
R ²		.6613		.6507	
rho		-.073		-.068	
DW		--		2.06	

J-test of over identifying restrictions:

$\chi^2(9) = 5.119$	$\chi^2(10) = 5.598$
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Table I.8. Euler Equation for Investment

GMM (first differences)

		A		B	
	lag	Coef	t-ratio	coef	t-ratio
I	1	-.1189	1.28	--	
P ^{Div}	0	.1357	1.96	.1217	1.84
VP ^{Div}	0	-.0071	2.80	-.0070	2.82
W ₀	0	.4055	1.74	.4973	1.96
W ₀	1	.5796	1.56	.4040	1.19
K	0	3.1437	9.70	2.9483	11.14
K	1	-2.1632	7.95	-2.1747	8.01
OV	0	-.0034	0.50	-.0035	0.46
OV	1	-.0029	1.78	-.0034	2.06
t		-.0001	0.54	-.0020	1.25
constant		.0058	0.14	.0272	0.74
R ²		.6624		.6583	
rho		-.025		-.018	
DW		--		1.93	
J-test of over identifying restrictions:					
		$\chi^2(11) = 8.718$		$\chi^2(12) = 9.489$	

**Table I.9. Variance Measures of Income Uncertainty with WGSA (CAD 1 000)
(Manitoba)**

A. WGSA Only (no other programmes in G)				
	Var (G)	Cov (Gp)y	Var (py)	Var (G + py)
1982	659.356	-2584.373	42129.36	37619.97
1983	291.218	-1440.437	21082.84	18493.19
1984	0.000	0.000	35540.28	35540.28
1995	652.580	-621.962	30159.85	29568.51
1986	1764.512	-6775.796	38659.90	26872.82
1987	3088.015	-11942.36	48591.16	27794.46
1988	7739.023	-4572.704	12476.23	11069.84
1989	12933.69	-33856.92	121321.0	66540.90
1990	12477.30	-17680.93	166808.9	143924.3
1991	6597.857	-1069.104	106294.7	110754.5
B. WGSA + miscellaneous programmes (= G)				
	Var (G)	Cov (Gp)y	Var (py)	Var (G + py)
1982	702.323	-3354.585	42129.36	36122.52
1983	362.542	-1277.338	21082.84	18890.71
1984	42.966	232.668	35540.28	36048.59
1985	667.188	-309.034	30159.85	30208.97
1986	1764.512	-6775.796	38659.90	26872.82
1987	3088.015	-11942.36	48591.16	27794.46
1988	7739.023	-4572.704	12476.23	11069.84
1989	12933.69	-33856.92	121321.0	66540.90
1990	8478.285	-33506.11	166808.9	108274.9
1991	11025.59	3711.184	106294.7	124742.7

Table I.10. Variance Measures of Income Uncertainty with NISA (CAD 1 000) (Manitoba)

A. NISA Only (no other programmes in G)				
	Var (G)	Cov (Gp)y	Var (py)	Var (G + py)
1995	1332.774	-1403.176	43807.08	42333.50
1996	379.140	-1929.969	130211.3	126730.5
1997	119.834	-1924.544	67163.73	63434.48
1998	48.997	-881.418	41086.76	39372.92
1999	73.510	-824.467	16159.47	14584.05
2000	444.028	-5202.981	62475.29	52513.35
2001	291.077	-2512.297	24981.63	20248.11
2002	147.132	-1876.990	59407.28	55800.43
B. NISA + AIDA (= G)				
	Var (G)	Cov (Gp)y	Var (py)	Var (G + py)
1995	1332.774	-1403.176	43807.08	42333.50
1996	379.140	-1929.969	130211.3	126730.5
1997	119.834	-1924.544	67163.73	63434.48
1998	48.997	-881.418	41086.76	39372.92
1999	73.510	-824.467	16159.47	14584.05
2000	2539.701	-12343.42	62475.29	40328.14
2001	1688.286	-6127.864	24981.63	14414.18
2002	2231.241	4504.338	59407.28	70647.20
C. NISA + miscellaneous programmes (= G)				
	Var (G)	Cov (Gp)y	Var (py)	Var (G + py)
1995	6640.746	-93.018	43807.08	50261.79
1996	2461.866	-1267.093	130211.3	130139.0
1997	126.305	-1921.912	67163.73	63446.22
1998	63.359	-985.203	41086.76	39179.71
1999	66.865	-833.028	16159.47	14560.28
2000	4048.236	-15536.32	62475.29	35450.89
2001	3558.807	-6863.952	24981.63	14812.53
2002	2008.653	-5442.639	59407.28	50530.66
D. NISA + AIDA + miscellaneous programmes (= G)				
	Var (G)	Cov (Gp)y	Var (py)	Var (G + py)
1995	6640.746	-93.018	43807.08	50261.79
1996	2461.866	-1267.093	130211.3	130139.0
1997	126.305	-1921.912	67163.73	63446.22
1998	63.359	-985.203	41086.76	39179.71
1999	66.865	-833.028	16159.47	14560.28
2000	8719.982	-22676.76	62475.29	25841.75
2001	6769.921	-10479.52	24981.63	10792.51
2002	4650.870	938.688	59407.28	65935.53

Annex II.

Aggregation of price risk over commodities: an index number approach

Index number methods are commonly used to aggregate prices over commodities. For example, assuming a Translog multioutput cost function and competitive profit maximization, then a Tornqvist index of outputs Y_1/Y_0 is interpreted as an aggregate output index for two time periods 0,1. In turn a corresponding implicit index of output prices P_1/P_0 is defined by the identity $P_1/P_0 = Y_1/Y_0 / \sum_j p_{j1} y_{j1} / \sum_j p_{j0} y_{j0}$, and can be interpreted as a ratio of aggregate price (Diewert 1976, 1982).

However, an index number procedure apparently has not been presented for aggregation of price uncertainty (or, more precisely, price risk as defined by Knight) over commodities. Since this involves price covariances as well as variances, this is a higher dimensional problem than is aggregation of prices. Common approaches to this problem are (a) to reduce the dimension of the problem by omitting many covariances and variances or (b) to calculate a variance of an aggregate price index. The second approach is commonly applied (*e.g.* Coyle 1992, 1999), but it has no basis in theory.

Ignoring output uncertainty, price uncertainty and output levels jointly contribute to revenue uncertainty as $VR_t = y_t^T Vp_t y_t$, where Vp is the price covariance matrix and y is a vector of output levels. An appropriate aggregation procedure for price uncertainty Vp will preserve the contribution of Vp to revenue uncertainty while controlling for effects of output levels y .

The following “Laspeyres” index is the most obvious approach:

$$(A-1) \quad VP_1/VP_0 = y_0^T Vp_1 y_0 / y_0^T Vp_0 y_0$$

using base period weightings y_0 throughout the index. However, this generally misrepresents the contribution of price uncertainty Vp to revenue uncertainty VR , just as aggregation with standard Laspeyres indexes generally loses the economic meaning of the subaggregates (Diewert). There seems to be no applications or references even to this approach to aggregation of uncertainty in the literature.

In the spirit of index number theory, an aggregate uncertainty index VP and an aggregate output index Y should satisfy the following identity:

$$(A-2) \quad VP_1/VP_0 (Y_1/Y_0)^2 = y_1^T Vp_1 y_1 / y_0^T Vp_0 y_0.$$

Thus a price uncertainty index VP_1/VP_0 can be correctly calculated from (A-2) given an appropriate aggregate output index Y_1/Y_0 . This connects aggregation of price uncertainty to the formulation of aggregate index numbers for multiple outputs.

Assuming risk aversion, index numbers for multiple outputs cannot be calculated from profit maximization, and in general index numbers depend on knowledge of risk preferences or the corresponding dual utility function (Chambers 1983). Nevertheless, it can be shown that a Tornqvist-like aggregate output quantity index is appropriate assuming a Translog cost function and constant returns to

scale disjoint technologies, which are common assumptions in index number theory. Then a “Tornqvist” index for aggregation of price risk can be calculated from (A-2) as

$$(A-3) \quad VP_1/VP_0 = (y_1^T Vp_1 y_1 / y_0^T Vp_0 y_0) / (Y_1/Y_0)^2$$

using a Tornqvist output quantity index. This approach can be extended to aggregation of higher moments of price uncertainty across commodities.

This approach can also be used to partition price uncertainty into aggregate subgroups. For example, variance of total revenue can be partitioned in terms of two commodity groups A and B as $VR = y_A^T Vp_A y_A + y_B^T Vp_B y_B + 2 y_A^T Vp_{AB} y_B$. Matrices Vp_A and Vp_B can be aggregated similarly to the above. A Laspeyres index analogous to (A-1) could be defined for Vp_{AB} . However, a more appropriate approach is to define a “Tornqvist” index for Vp_{AB} as the following simple generalization of (A-3):

$$(A-4) \quad VP_{AB1}/VP_{AB0} = (y_{A1}^T Vp_{AB1} y_{B1} / y_{A0}^T Vp_{AB0} y_{B0}) / (Y_{A1}/Y_{A0}) (Y_{B1}/Y_{B0}).$$

In crop acreage models where yields are predetermined, much simpler aggregation procedures may be acceptable. Since yields are predetermined, then in effect the exogenous measure of uncertainty faced by firms are variances and covariances of revenues per acre, *i.e.* uncertainty is represented by the revenue per acre covariance matrix $Vr = yld^T Vp yld$ for a predetermined yield vector yld . Then this uncertainty translates into total revenue uncertainty for the firm as $VR = z^T Vr z$ where z is a vector of crop acreage allocations. Therefore the contribution of Vr_t to VR_t apparently can be indexed simply as VR_t/Z_t where Z is total crop acres, *i.e.* the index number aggregation of Vr apparently does not require an index number for outputs y . Uncertainty for two groups of commodities A and B can be aggregated similarly, where the aggregate covariance is $VR_{ABt}/(Z_{At}Z_{Bt})$ and Z_A, Z_B are total crop acres in groups A and B.

Index numbers of the contribution of price uncertainty Vp (measured as (1) assuming naïve expectations) for all six crops to total revenue uncertainty VR were calculated, and correlations are reported in table II.1. A Laspeyres index is calculated from (A- 1), and is labelled as STLASPVP. Other indexes (VPTORN, VPFISH, VPLASP) are calculated from (A-2) using Tornqvist, Fisher and Laspeyres index numbers respectively for total output Y_t/Y_0 (there is a constant base at 1975). A variance of a Tornqvist index price P^{Div} is also presented (VPDIVBAD). The index number calculations based on (A-1) and (A-3) are similar and highly correlated, as often seems to happen with Laspeyres and more theoretically appropriate index numbers. On the other hand the variance of P^{Torn} , is a particularly poor index number for price uncertainty, is very poorly correlated with the other index numbers.

Table II.1. Correlation Matrix for Aggregate Indexes of Price Uncertainty for Manitoba Crops

VPTORN	1.0000				
VPFISH	0.98458	1.0000			
VPLASP	0.98713	0.99974	1.0000		
STLASPVP	0.96783	0.98975	0.98728	1.0000	
VPDIVBAD	0.51290	0.49302	0.49595	0.46563	1.0000
	VPTORN	VPFISH	VPLASP	STLASPVP	VPDIVBAD

Accompanying this study, an index number approach to aggregation of price risk over commodities is developed in detail by Coyle (2005). This paper measures price risk VP in terms of rational expectations/GARCH models as well as naïve expectations. The paper concludes that naïve expectations models (1) are more appropriate for this study.

Annex III.

Discussion of Euler Equation Investment models

Estimations of Euler Investment equations are presented in Tables I.6 to I.8 in Annex I.

A. Reduced Form Euler Equations in Relation to ADL Models

Assuming mean-variance risk preferences define a short run dual indirect utility function conditional on capital stock K and change in capital stock I as in Coyle (1999) and in equation (9). The coefficient of risk aversion function is $\alpha(.) = \alpha(W_0 + E_p y - w x - w^k K, y^T V_p y)$ and the production function with adjustment costs is $y(.) = f(x, K, I)$. Given discrete time data net investment I_t at time t is approximated as $I_t = K_{t+1} - K_t$ rather than $K_t - K_{t-1}$.

$$(9) U^*(E_p, V_p, W_0, w, w^k, K, I) = \max_x [W_0 + E_p y(.) - w x - w^k K - \alpha(.) / 2 y(.)^T V_p y(.)]$$

The discrete time calculus of variations problem is defined in equation (10), subject to initial conditions. Assuming an interior solution for K , the first order condition $\delta L / \delta K_t = 0$ implies Euler equation (11) at t conditional on information at t :

$$(10) \max_K L = \sum_{t=0, \dots, \tau} U^*(E_{pt}, V_{pt}, W_t, w_t, w_t^k, K_t, I_t) / (1+r)^t$$

$$(11) \delta U^*(.) / \delta K_t - \delta U^*(.) / \delta I_t + \delta U^*(.) / \delta I_{t-1} (1+r) = 0.$$

This Euler equation generally specifies current investment I_t as a function of lagged investment I_{t-1} and of current variables $(E_p, V_p, W, w, w^k, K)_t$ plus a one period lag in these variables. Then the reduced form equation for I_t from the Euler equation can be approximated in loglinear form as

$$(12) I_t = \beta_0 + \beta_1 I_{t-1} + \sum_s \beta_{1s} z_t^s + \sum_s \beta_{2s} z_{t-1}^s + e_t$$

where $z = (E_p/w, V_p/w^2, W/w, w^l/w, w^k/w, K, t)$. In the empirical model, all variables are specified as first differences of logarithms. This specification clearly nests the Euler equation within an ADL investment equation conditional on capital. An ADL model has a more general lag structure.

The alternative approximation $I_t = K_t - K_{t-1}$ is more common in the empirical literature. This implies that $\delta L / \delta K_t = 0$ specifies current I_t as a function of I_{t+1} and current and future z , conditional on information at time t . (A lagged dependent variable I_{t-1} is sometimes appended to this model in an ad hoc manner in order to incorporate lags in adjustment.

The choice between the approximations $I_t = K_t - K_{t-1}$ and $I_t = K_{t+1} - K_t$ with discrete data seems essentially *ad hoc*. We adopt the latter approximation primarily because it directly specifies current investment in terms of a lag on investment and other variables in rather than leads, and so makes the relation between Euler equations and ADL models more transparent. In addition, econometric estimation is less complex than in the standard case as will be discussed in the next section.

In principle a longer lag structure for the approximation can be obtained by specifying a longer lag structure in adjustment costs, *i.e.* by specifying $y(.)$ as $y_t = f(x_t, K_t, K_{t-1}, \dots, K_{t-s})$ rather than as $y_t = f(x_t, K_t, I_t)$.

However, such a specification of adjustment costs would be uncommon and is difficult to justify. This specification presumably would not actually explain the longer lags in ADL models.

Input prices w^L and w^k were insignificant in estimates of model (12), as in the ADL investment models. So the selected model is

$$(13) \quad I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,1} \beta_{1i} P^{DIV}/w_{t-i} + \sum_{i=0,1} \beta_{2i} VP^{DIV}/w_{t-i}^2 + \sum_{i=0,1} \beta_{3i} W^0/w_{t-i} + \sum_{i=0,1} \beta_{4i} COVGP/w_{t-i}^2 + \sum_{i=0,1} \beta_{5i} K_{t-i} + e_t.$$

This model was estimated by OLS with data as first differences in logarithms. In comparison with Table I.2.B, t-ratios for all sum of lag coefficients are higher for the ADL model, in spite of higher multicollinearity. Coefficient estimates for P^{DIV} and $COVGP$ are considerably different between these models. The difference in results for $COVGP$ is not surprising, since only a three and four period lag is included in the final ADL model. Thus, due to shorter lag lengths, an Euler equation type model does not adequately approximate sum of lag coefficient estimates of the ADL model.

B. Structural Euler Equations

A normalised quadratic flexible functional form is assumed for the above dual indirect utility function (9). Then the derivatives of the dual relevant to the Euler equation are

$$(14) \quad \begin{aligned} \delta U^*(.)/\delta K_t &= \alpha_{0k} + \alpha_{1k} P^{DIV}/w_t + \alpha_{2k} VP^{DIV}/w_t^2 + \alpha_{3k} W^0/w_t \\ &\quad + \alpha_{4k} w^L/w_t + \alpha_{5k} w^k/w_t + \alpha_{6k} K_t + \alpha_{7k} I_t + \alpha_{8k} t \\ \delta U^*(.)/\delta I_t &= \alpha_{0l} + \alpha_{1l} P^{DIV}/w_t + \alpha_{2l} VP^{DIV}/w_t^2 + \alpha_{3l} W^0/w_t \\ &\quad + \alpha_{4l} w^L/w_t + \alpha_{5l} w^k/w_t + \alpha_{6l} K_t + \alpha_{7l} I_t + \alpha_{8l} t \end{aligned}$$

in terms of coefficients α . Substituting (14) into the Euler equation (11),

$$(15) \quad I_t (\alpha_{7l} - \alpha_{7k}) = (\alpha_{0k} + \alpha_{0l} r) + \sum_i (\alpha_{ik} - \alpha_{il}) z_{it} + \sum_i \alpha_{il} z_{it-1} (1+r) + \alpha_{7l} I_{t-1} (1+r)$$

where $z = (P^{DIV}/w, VP^{DIV}/w^2, W^0/w, w^L/w, w^k/w, K)$. Given a normalization $\alpha_{7l} - \alpha_{7k} = 1$, the Euler equation (15) is identified and linear in coefficients α . Variables in the Euler equation are conditional on information available at time t .

Alternatively, all coefficients α of (15) can be identified by estimating the Euler equation jointly with output supply and factor demand equations

$$(16) \quad \begin{aligned} y &= \delta U^*(.)/\delta Ep / \delta U^*(.)/\delta W^0 \\ -x &= \delta U^*(.)/\delta w / \delta U^*(.)/\delta W^0 \end{aligned}$$

(Coyle 1999). However, the cross-equation (symmetry) restrictions on coefficients between the Euler equation and (16) imply a model that is highly nonlinear in coefficients and hence difficult to estimate. In principle equations (16) can be estimated using a Divisia index for crop output y and fertilizer sales (tonnes) x , but this approach is not followed here.

Euler equation (15) can be expressed equivalently in the simpler form (assuming a constant discount rate r , as in most studies)

$$(17) \quad I_t = \beta_0 + \beta_1 I_{t-1} + \sum_i \beta_{1i} z_t^i + \sum_i \beta_{2i} z_{t-1}^i + e_t$$

where $\beta_0 = (\alpha_{0k} + \alpha_{0l} r) / (\alpha_{7l} - \alpha_{7k})$, $\beta_1 = \alpha_{7l} (1+r) / (\alpha_{7l} - \alpha_{7k})$, $\beta_{1i} = (\alpha_{ik} + \alpha_{il} r) / (\alpha_{7l} - \alpha_{7k})$, $\beta_{2i} = \alpha_{il} (1+r) / (\alpha_{7l} - \alpha_{7k})$. This equation is linear in coefficients and can identify the structural coefficients α of the Euler

equation, given a normalization such as $\alpha_{3l} - \alpha_{3k} = 1$. Variables are not logged. Instead variables are normalised to 1.0 for 2000, so coefficients can be interpreted as elasticities circa 2000. In the empirical model all variables are specified as first differences. Note that model (17) is very similar to the reduced form approximations (including K) in the previous section, except that variables are not in logarithmic form.

Generalized method of moments (GMM) is the most common method of estimation for structural Euler equations, in spite of poor small sample properties for such estimators. In principle GMM provides the optimal instrumental variable estimator under heteroscedasticity or autocorrelation. So only GMM estimates are reported. The weighting matrix was constructed using a heteroskedastic-consistent covariance matrix (White 1980) as encoded in Shazam 8.0. A Bartlett weighting matrix was also considered, but the estimated matrix was singular so estimation was unsuccessful. This is not a serious problem since serial correlation of the disturbance appears to be minor. Price variables and lagged capital were treated as exogenous. Additional one period lags on all exogenous variables in the equation were used as additional instruments. Hired labor wage w^l and capital price index w^k were statistically insignificant, so they were deleted from the model. Note that model estimation is less complex than under the assumption $I_t = K_t - K_{t-1}$.

Alternatively, consider the standard case where net investment is approximated as $I_t = K_{t+1} - K_t$. Then (10) implies the following Euler equation at time t:

$$(18) \delta U^*(.)_t / \delta K_t - \delta U^*(.)_t / \delta I_t + \delta U^*(.)_{t+1} / \delta I_{t+1} (1+r) = 0$$

conditional on information available at time t. Substituting (14) into (18),

$$(19) I_t (\alpha_{7l} + \alpha_{7k}) = -(\alpha_{0k} + \alpha_{0l} r) - \sum_i (\alpha_{ik} - \alpha_{il}) z_{it} + \sum_i \alpha_{il} z_{it+1} / (1+r) + \alpha_{7l} I_{t+1} / (1+r)$$

Assuming a constant discount rate r, this equation can be simplified to

$$(20) I_t = \beta_0 + \beta_1 I_{t+1} + \sum_i \beta_{1i} z_{it} + \sum_i \beta_{2i} z_{it+1} + e_t$$

where $\beta_0 = -(\alpha_{0k} + \alpha_{0l} r) / (\alpha_{7l} + \alpha_{7k})$, $\beta_1 = \alpha_{7l} / (\alpha_{7l} + \alpha_{7k}) (1+r)$, $\beta_{1i} = -(\alpha_{ik} + \alpha_{il} r) / (\alpha_{7l} + \alpha_{7k})$, $\beta_{2i} = \alpha_{il} / (\alpha_{7l} + \alpha_{7k}) (1+r)$. This equation is linear in coefficients and can identify the structural coefficients of the Euler equation, given a normalization such as $\alpha_{7l} + \alpha_{7k} = 1$.

The Euler equation (20) is conditional on information at time t, so planned I_{t+1} and anticipated Z_{t+1} are not observed. This substantially complicates estimation, in comparison to (17). I_{t+1} and K_{t+1} can be modelled assuming rational expectations, but it is not entirely clear how to model price elements of Z_{t+1} (e.g. as naïve or rational expectations). In any case, GMM estimation is much more complex in the case of (20) than in (17).

Table I.6 reports GMM estimates of simple Euler equations under risk neutrality and risk aversion:

$$(21) I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,1} \beta_{1i} P^{Div}/w_{t-i} + \sum_{i=0,1} \beta_{4i} K_{t-i} + e_t$$

$$(22) I_t = \beta_0 + \delta I_{t-1} + \sum_{i=0,1} \beta_{1i} P^{Div}/w_{t-i} + \sum_{i=0,1} \beta_{2i} VP^{Div}/w_{t-i}^2 + \sum_{i=0,1} \beta_{3i} W^0/w_{t-i} + \sum_{i=0,1} \beta_{4i} K_{t-i} + e_t$$

This model assumes $I_t = K_{t+1} - K_t$, and this model is conditional on information available at time t. So GMM estimation is relatively simple here.

First consider results for the risk neutral model (Table I.6.A, i.e. column A of Table I.6). The overidentifying restrictions implied by moment conditions for orthogonality of instruments and disturbance

were not rejected using Hansen's J-test for GMM models. However, estimated coefficients of the expected output price Divisia index P^{DIV} were statistically insignificant (and with a negative sign). Estimated coefficients and t-ratios are +0.028 (0.72) and -0.043 (1.13) for current and lagged P^{DIV} . R^2 is .636.

GMM estimates of the above risk averse model are reported in Table I.6.B. Current period P^{DIV} and VP^{DIV} are significant at the .05 level with anticipated signs: coefficient estimates and t-ratios are +0.184 (2.23) and -0.008 (2.12) for P^{DIV} and VP^{DIV} . Lagged initial wealth has the largest elasticity: +0.698 (1.75). Other coefficients for P^{DIV} , VP^{DIV} , W^0 are insignificant. The insignificance of current wealth is not surprising: according to theory the coefficient of risk aversion depends on initial (prior) wealth rather than current wealth. R^2 is .659. Over-identifying restrictions for valid instruments are not rejected using Hansen's J test. The statistical significance of price variance VP^{DIV} implies rejection of the risk neutral model.

Table I.7 reports GMM results for model (22) deleting insignificant coefficients for P^{DIV} and VP^{DIV} . In 7A with lagged investment, estimated coefficients and t-ratios are +0.203 (3.08), -.010 (3.85) for P^{DIV} and VP^{DIV} , and +0.137 (0.48), +0.634 (1.64) for current and lagged W^0 . R^2 is .661, and over-identifying restrictions are accepted. In 7B excluding lagged investment, estimated coefficients are +0.189 (3.14), -0.009 (4.35) for P^{DIV} and VP^{DIV} , and +0.247 (1.01) and +0.290 ((0.95) for W^0 . The Durbin-Watson statistic is 2.06, indicating zero autocorrelation. R^2 is .651, and over-identifying restrictions are accepted. In this case initial wealth W^0 has a higher elasticity than does output price P^{DIV} , in contrast to results for ADL investment above. However, W^0 is less statistically significant than is P^{DIV} .

Current and one period lag on COVGP, covariance between government payments and market prices, were insignificant in the Euler equation. This is not surprising, since only third and fourth period lags were significant in the ADL investment model.

Financial constraints were incorporated into the Euler equation model. Euler equation studies typically specify financial constraints as exogenous and related to internal finance, so an exogenous financial constraint is appended to the above dynamic model as follows:

$$(23) U^*(Ep, Vp, W_0, w, w^k, K, I, VC) = \max_x W_0 + Ep y(.) - w x - w^k I - \alpha(.) / 2 y(.)^T Vp y(.)$$

$$\text{s.t. } w x \# VC$$

$$(24) \max_{K, VC} \sum_{t=0, \dots, T} U^*(Ep_t, Vp_t, W_t, w_t, w_t^k, K_t, I_t, VC_t) / (1+r)^t$$

$$\text{s.t. } VC_t + w^k I_t \# TC_t$$

where VC is variable cost and TC is total cost. This implies an Euler equation

$$(25) \delta U^*(.) / \delta K_t - \delta U^*(.) / \delta I_t + \delta U^*(.)_{t-1} / \delta I_{t-1} (1+r) - \lambda_t w^k_t + \lambda_{t-1} w^k_{t-1} (1+r) = 0$$

where λ_t is the shadow price of the financial constraint t in the dynamic problem (and a first order condition for VC). The difficulty in estimation is to model λ . This is typically done in an ad hoc manner by relating λ to financial variables that may influence the constraint (e.g. Hubbard and Kashyap 1992; Hubbard, Kashyap and Whited, 1995). These approaches have been criticized as *ad hoc*.

In order to adopt a less ad hoc approach, the shadow prices for the financial constraints in problems (23) and (24) may be seen as related. For example these shadow prices would be identical if they can be reduced to a single constraint $w x + W^k I \# TC$, i.e. variable input and capital expenses draw on the same underlying constraint. In this case, the Euler equation can be specified by substituting the shadow price $\delta U^*(.) / \delta VC$ for λ in the Euler equation:

$$(26) \delta U^*(.) / \delta K_t - \delta U^*(.) / \delta I_t + \delta U^*(.)_{t-1} / \delta I_{t-1} (1+r) - \delta U^*(.) / \delta VC w^k_t + \delta U^*(.)_{t-1} / \delta VC w^k_{t-1} (1+r) = 0 .$$

This can be estimated jointly with the following envelope relations for the dual $U^*(.)$ (Coyle 1990, 1999):

$$(27) \quad y = \delta U^*(.)/\delta E_p / \delta U^*(.)/\delta W_0$$

$$x_1 = -\delta U^*(.)/\delta w_1 / \{\delta U^*(.)/\delta W_0 + \delta U^*(.)/\delta VC\}$$

for Divisia aggregate crop output and fertilizer input. The null hypothesis of no financial constraint can be specified as $\delta U^*(.)/\delta VC = 0$ and VC can be excluded from the dual $U^*(.)$ in (26)-(27). However, (26)-(27) imposing cross-restrictions for $\delta U^*(.)/\delta W_0$ are highly nonlinear in coefficients and could not be estimated here.

Alternatively, the Euler equation (26) was estimated independently of (27). In the absence of financial constraints both Lagrange multipliers λ_t and λ_{t-1} equal 0 in the Euler equation (25), *i.e.* $H_0: \lambda_t = \delta U^*(.)/\delta VC_t = 0$ and $\lambda_{t-1} = \delta U^*(.)/\delta VC_{t-1} = 0$ in (26). The hypothesis H_0 was tested using Euler equation (26).

The hypothesis of no financial constraints was tested for OLS estimates by standard methods, and also for GMM. Testing linear restrictions in GMM is problematic. In one approach, a likelihood ratio test statistic for GMM is simply the difference between the minimized criterion function (using the same instruments and weighting matrix) evaluated at restricted and unrestricted estimates, and this is asymptotically distributed as χ^2 . However, this test assumes that an optimal weighting matrix, leading to asymptotically efficient GMM estimators, has been selected (Newey and West 1987). In a second approach when GMM estimates minimize a nonoptimal criterion function, restrictions are tested as a Wald test. The second approach is encoded in Shazam 8.0. A further difficulty in testing the hypothesis is that the Euler equation (26) involves many coefficients and low degrees of freedom. This can be a particularly serious problem for GMM.

Given these difficulties in testing financial constraints for GMM estimation, the focus was on testing results for OLS. For simplicity we test the hypothesis H_0 that all coefficients of $\lambda_t = \delta U^*(.)/\delta VC_t$ and $\lambda_{t-1} = \delta U^*(.)/\delta VC_{t-1}$ are equal to 0 in the Euler equation (26) at t . Acceptance of H^* certainly implies acceptance of H_0 . A direct test of H_0 could be constructed similarly to Coyle (1990).

Test results for OLS estimates of (26) can be summarised as follows. Denote λ_t and λ_{t-1} in the Euler equation at t as λ_1 and λ_2 . Then (a) $H_0: \lambda_1=0$ and $\lambda_2=0$ $F(12,14) = 0.934$, $\text{prob} = 0.542$, (b) $H_0: \lambda_1=0$ $F(6,14) = 1.156$, $\text{prob} = 0.377$, and $H_0: \lambda_2=0$ $F(6,14) = 1.165$, $\text{prob} = 0.377$. Somewhat similar results hold for GMM. Thus the hypothesis of no financial constraints on investment is not rejected at any level of significance. This result supports the conclusion of Gomes (2001).

Option value was proxied as both a Divisia index DVp^s of reduction in crop price uncertainty (described above) and more precisely as the related reduction in uncertainty of profits. Reduction in uncertainty of profits is measured as $y^{\text{DIV}^2} DVp^{\text{sDIV}}$ using Divisia indexes for y and DVp , or alternatively as $Y_t [VP_t^s - VP_t]$ where VP is a 6×6 crop price covariance matrix and Y is a 6×1 vector of crop outputs. All measures are normalised by w^2 assuming CRRA. Suppose that there is a marginal cost of current investment that is linear in such a measure of option value (OV), *i.e.* marginal cost is θ OV where θ is a parameter to be estimated. Then the maximization problem and Euler equation can be modified as

$$(28) \quad \max_K \sum_{t=0, \dots, \tau} \{U^*(E_p, V_p, W_t, w_t, w^k, K_t, I_t) - \theta OV_t I_t\} / (1+r)^t$$

$$(29) \quad \delta U^*(.)/\delta K_t - \delta U^*(.)/\delta I_t + \delta U^*(.)/\delta I_{t-1} (1+r) + \theta OV_t - \theta OV_{t-1} (1+r) = 0.$$

Thus OV_t and OV_{t-1} are appended to the Euler equation (17)/(12). The focus was made on OV measured as $Y^{\text{DIV}^2} DVp^{\text{sDIV}}$ using Divisia indexes for y and DVp , and $s=2$. Denote this measure of option value as OV^* .

Table I.8 reports GMM results where current and lagged OV^* are added to model (19) (deleting insignificant coefficients for P^{DIV} and VP^{DIV}). In 8A with lagged investment, estimated coefficients and t-ratios are +.136 (1.96), -0.007 (2.80) for P^{DIV} and VP^{DIV} , +0.405 (1.74), +0.579 (1.56) for current and lagged W^0 , and -0.003 (0.50) and -0.003 (1.78) for current and lagged OV^* . R^2 is .662, and overidentifying restrictions are accepted. In 8B excluding the insignificant lagged investment, estimated coefficients are +0.122 (1.84), -0.007 (2.82) for P^{DIV} and VP^{DIV} , +0.497 (1.96) and +0.404 ((1.19) for W^0 , and -0.004 (0.46) and -0.003 (2.06) for current and lagged OV^* . The Durbin-Watson statistic is 1.93, indicating zero autocorrelation. R^2 is .658, and overidentifying restrictions are accepted. Note that lagged OV^* is statistically significant at the .05 level with the anticipated negative sign. Somewhat similar results were obtained using alternative measures of option value mentioned above.

These exploratory results suggest that it may be possible to incorporate option value of waiting into empirical Euler equation models. However, much more research is required on this subject.

An attempt was made to accommodate asymmetric responses. At the Manitoba level capital investment is always positive, so irreversibility in investment could not be modelled. Instead a (0,1) dummy variable distinguishes between positive and negative first differences of investment I , and interactions between this dummy and other variables were added to the econometric model. However, these interaction variables were statistically insignificant, so such asymmetric responses remained out of the model.

*Annex IV.***Parameter Assumptions for Simulations**

Parameter estimates of the ADL investment model are considered more reliable than estimates of the Euler equation, as discussed above. 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, lags in expectations are assumed to follow the simple naive model, so that short run lags in response are identified.

The Lucas critique raises the question of stability of expectations to policy. However, the author is convinced from his previous studies (and related studies) that the simple naive model of expectations formation postulated here (as opposed to a rational expectations process) is a valid model of expectations for Manitoba crop agriculture and is likely to remain stable across foreseeable policy changes. Note that the crop acreage allocation model in this study can reasonably be specified as a static model (with no lags in response), so the somewhat reasonable estimates of this model can be interpreted as qualified support for this naive model of expectations. A report accompanying this study (Coyle 2005) also provides support for this naive model over a rational expectations (GARCH) model.

The relevant long run elasticities (sum of lag coefficients) from Table I.3.ABC (*i.e.* columns ABC in Table I.3) in Annex I for ADL investment in machinery and equipment for Manitoba crops are as follows:

Table IV.1. Long Run Elasticity of Investment

	Table I.3.A	I.3.B	I.3.C
Expected Output Price (P^{Div})	+1.0903	+1.1920	+1.1141
Variance of Output Price (VP^{Div})	-0.0520	-0.0497	-0.0473
Initial Wealth (W_0)	+0.7593	+0.8404	+0.9966
Insurance Effect (CovGp)	-0.0738	-0.0728	-0.0908

The estimates from Table I.3.B (OLS, excluding the insignificant lagged dependent variable) will be used later.

Short run elasticities are coefficients of the current period (un-lagged) variables. These are estimated as follows:

Table IV.2. Short Run Elasticity of Investment

	Table I.3.A	I.3.B	I.3.C
Expected Output Price (P^{Div})	+0.4053	+0.5162	+0.4173
Variance of Output Price (VP^{Div})	-0.0234	-0.0276	-0.0230
Initial Wealth (W_0)	+0.5645	+0.6152	+0.6973
Insurance Effect (CovGp)	0	0	0

These short run effects are less statistically significant than the long run elasticities (sum of lag coefficients). Based on Table I.3.B, the ratio of long run to short run elasticities are 2.31, 1.80 and 1.37 for expected price, price variance and initial wealth, respectively.

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