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Income risk and farm consumption behavior

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Abstract

Using panel data from Illinois grain farmers, a direct test of the relationship between income risk and farm consumption behavior is conducted. The estimation results indicate that income risk significantly affects farm consumption and the results are robust using alternative risk measures. This finding casts doubt on the relevance of the conventional life-cycle permanent income hypothesis, which implies that risk has no effect on consumption. © 1999 Elsevier Science B.V. All rights reserved.

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1. Income risk and farm consumption behavior

There are only a few empirical studies of the relationship between farm consumption and income in developed economies, even though farm consumption behavior has an important role in determining the ability of farms to invest and grow. Two major findings can be extracted from the limited literature on the topic (Friedman, 1957; MacMillan and Loyns, 1969; Girao et al., 1974; Mullen et al., 1980; Langemeier and Patrick, 1990, 1993; Carriker et al., 1993; Phimister, 1995). First, the conventional life-cycle permanent income hypothesis seems to describe farm consumption well.¹ Second, the marginal propensity to consume (MPC) for farm households is low com-

pared to non-farm families and the MPC from farm income is low compared to the MPC from non-farm income. Friedman (1957, p. 62) argues that individuals with volatile incomes have lower MPCs than those with stable incomes and farm income risk is often suggested as an explanation for the low MPC of farm families.

The two findings seem inconsistent with each other. The permanent income hypothesis implies that future income risk has no role in determining the consumption behavior of farm households, while the argument that low MPC are a result of income volatility implies that farm consumption is related to income risk. This apparent inconsistency has been overlooked in empirical studies and measures of income risk have been excluded from the theoretical and empirical models of farm consumption behavior. If income risk is an important determinant of the consumption behavior of farm households, then empirical consumption models that ignore income risk are biased and any policy implications drawn from these models are potentially

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¹One notable exception is Phimister (1995). The life-cycle model without borrowing restrictions is rejected using Dutch farm household data.

misleading. Blanchard and Mankiw (1987) show that a risk-augmented life-cycle model is consistent with behavior that is strikingly different from the predictions of the conventional life-cycle permanent income models.² Given the fact that the certainty equivalent life-cycle permanent income hypothesis has been widely used in modeling farm consumption, it is important to validate this assumption.

The objective of this paper is to test econometrically whether income risk affects farm consumption. In other words, whether farm households have a precautionary motive for savings. A particular effort is made to determine the proper measure of income risk. Traditional measures of risk are the normalized standard deviation or the variance of an underlying risky variable. However, these measures are reliable only if the underlying time series is stationary. Since economic data are rarely stationary, it is necessary to transform the underlying data to a stationary time series before a reliable measure of risk can be derived. To derive a stationary series requires knowledge about the time-series properties of the stochastic income series. Although farm income risk is widely discussed in the literature, few microeconomic studies of the structure of changes in farm income exist. Following Hall and Mishkin (1982), the individual income series are treated as time- and life-cycle trend-stationary processes. Since the econometric testing of risk behavior may be sensitive to the measure of risk used, several alternative measures of income risk are considered.

2. Characterizing a stochastic farm consumption function

The theoretical conditions under which income risk effects consumption have been studied by LeLand

²The conventional life-cycle permanent income model is generally rejected in the consumption literature (Hall, 1987). Searching for reasons to explain the failure of the life-cycle permanent income hypothesis has been an active research area. An explanation is that some consumers are not able to borrow the amount needed to maintain their optimal consumption level (Hayashi, 1985; Campbell and Mankiw, 1987; Zeldes, 1989; Jappelli and Pagano, 1989). For the farm household, Langemeier and Patrick (1993) rejected the importance of the borrowing constraint and accepted the life-cycle permanent income model.

(1968), Sandmo (1970), Dreze and Modigliani (1972), Kimball (1990) using two-period consumption models. Sibley (1975), Miller (1975) and Weil (1993) have extended the two-period framework to a multi-period model and established the same theoretical results. The main conclusion is that when income risk is uninsurable and utility is time-separable, earnings uncertainty increases saving and wealth accumulation if the third derivative of the utility function with respect to consumption is positive. A sufficient condition for risk to increase savings is that absolute risk aversion be non-increasing with wealth.

Merton (1969) (see also Merton, 1973) has shown that it is not possible to find an analytical expression for consumption under uncertainty except in a few limited cases. Consequently, in order to obtain an explicit farm consumption function under income risk, it is assumed that the i th farm household has a constant absolute risk aversion contemporary utility function $U_{it} = -(1/R)\exp(-R \times C_{it})$ with a coefficient of constant absolute risk aversion, R . Assume that the farm household maximizes its expected lifetime utility as of time zero

$$E \left[\sum_{t=0}^T (1+\delta)^{-t} U(C_{it}|0) \right] \quad (1)$$

where $(1+\delta)$ is a discount factor, C_{it} the farm consumption, T the planning horizon, and $E(\cdot|0)$ denotes an expectation conditional on information available at time 0.

Suppose, at the beginning of the period t , the i th farm household possesses financial wealth (assets) with total real value W_{it} . In the same period, it also receives net farm income π_{it} , which is random. The sum of assets and net farm income is allocated between consumption C_{it} and a menu of assets $A_{it} = \sum_{j=1}^M A_{ijt}$. Before the beginning of the period $t+1$, the assets appreciate by a factor $(1+r_{it})$, where r_{it} is the interest rate. The evolution of the farm household's wealth is

$$W_{it+1} = (1+r_{it})(W_{it} + \pi_{it} - C_{it}) \quad (2)$$

For simplicity it is also assumed that farm net income follows a random walk with normally distributed innovations

$$\pi_{it} = \pi_{it-1} + \varepsilon_t, \quad \varepsilon_t \sim N(0, \sigma^2) \quad (3)$$

and that the discount rate and the risk-free rate of return are both zero.

Following Caballero (1987), the level of consumption is given by

$$C_{it} = \frac{1}{T-t} W_{it} + \pi_{it} - \frac{R(T-t-1)}{4} \sigma^2 \quad (4)$$

Eq. (4) shows that the level of consumption is a function of wealth, current income and income risk. The first two terms follow from the life-cycle permanent income hypothesis, while the third term is the contribution of the precautionary motive induced by the presence of income risk. Prudence is also reflected in the third term: the higher the risk, the lower the level of current consumption, given income and wealth. Eq. (4) illustrates a risk-augmented consumption model.

Decreasing absolute prudence is a desirable property of the underlying utility function (Kimball, 1990). To guarantee positive consumption along the optimal path and to permit decreasing absolute prudence, a flexible utility function should be used. An obvious candidate is the isoelasticity utility function. Unfortunately, its use requires numerical techniques to solve for optimal consumption. Zeldes (1989), using a stochastic dynamic simulation model, found that consumption under a CRRA utility function deviates dramatically from consumption under certainty. Two interesting results are:

- (i) the MPC from changes in current income or wealth is considerably larger than that predicted by a conventional life-cycle permanent income model; and
- (ii) the expected growth of consumption under CARA utility is considerably higher at lower levels of wealth than at higher wealth levels, while the expected growth of consumption under certainty is the same regardless of the level of wealth.

These results suggest that the impact of risk on consumption depends on wealth, and decreasing prudence is important. At higher levels of wealth, a larger portion of lifetime income is certain, and the variance of the percentage change in consumption decreases. This, in turn, implies a flatter consumption path for wealthy individuals as compared to those with lower levels of wealth. The marginal propensity to consume depends on the amount of risk and an increase in

income decreases the need for precautionary savings, leading to a larger response in consumption than would be predicted under certainty.

To allow for non-constant prudence, the specification suggested by Eq. (4) must be modified. Although exactly how wealth interacts with risk is unknown, because a closed form solution under decreasing prudence cannot be obtained, it is likely to be non-linear. Consequently, an ad hoc interaction term between wealth and income risk is added to Eq. (4). This results in the following specification:

$$C_{it} = \frac{1}{T-t-1} W_{it} + \pi_{it} - \frac{R^a(T-t)}{4} (\sigma_{it}^\pi)^2 + \psi_{lev} W_t (\sigma_{it}^\pi)^2 \quad (5)$$

The parameter ψ measures how the effect of risk on consumption is influenced by the level of farm household wealth. This permits an empirical test of whether the farm household exhibits constant, decreasing, or increasing prudence. If $\psi=0$, Eq. (5) reduces to the case of the constant prudence. If $\psi<0$, the effect of risk on expected consumption increases with the household's assets, suggesting increasing prudence. If $\psi>0$, the effect of risk on expected consumption declines with the household's assets, suggesting decreasing prudence.

3. Data and measures of income risk

The data used in the analysis were obtained from Langemeier and Patrick.³ The sample consists of a panel of 144 observations on 18 Illinois grain farm households between 1979 and 1986. The most important household variables used in estimation are total consumption, disposable income, and the value of assets. Consumption is computed as the sum of all personal living expenses, which includes expenditures on durable goods.⁴ Because durable goods provide consumption beyond a single period, this flaw should be corrected and used to adjust the consumption

³The authors are grateful to Mike Langemeier for making data available for the study.

⁴Consumption is theoretically distinct from total expenditures on goods and services. Since it is impossible to measure the theoretical concept, the usual practice of treating observed expenditures and consumption as synonymous is followed.

figures. Unfortunately, the detailed information necessary to exclude durable goods from current consumption is unavailable. Disposable household income is computed as reported net farm income, plus net non-farm income and depreciation, minus taxes paid by the farm household.⁵ The value of assets is measured by net wealth, computed from the household's balance sheet. These series were inflated to 1986 dollars, using the implicit price deflator for personal consumption expenditures.

Comparing the coefficients of variation for real consumption, income, and wealth for the eighteen farms, for the period 1979–1986, two observations are important. First, the consumption series is much smoother than the income and net wealth series. The coefficient of variation of consumption is 15% of the coefficient of variation for income and 10% of the coefficient of variation for net wealth.⁶ Second, coefficients of variation for real consumption across farm households are quite similar, while those for household income and net wealth vary dramatically across the households. The implication of the second observation is that farm households appear to have ample opportunities to smooth their consumption. This is confirmed by the low correlation between income and consumption. The correlation between the farm household's income changes and consumption changes is only 0.10, while the correlation between the variance of income changes and the variance of consumption changes is only 0.15. Such low correlation coefficients between income and consumption imply that farmers engage in enough buffer savings to largely offset the effects of a risky income stream on consumption.

While it is easy to discuss the variation in income and consumption, determining the proper measure of risk to use in empirical work is difficult. There are two important choices when it comes to choosing a measure of risk. The first is which variable to use. Most studies, in agriculture, focus on measures of aggregate price, yield, or revenue risk which may be poor representations of the risk faced by farm households.

⁵As noted in Langemeier and Patrick (1990), this measure of income is more closely related to the actual funds available for consumption.

⁶These are simple averages of the coefficients of variation for the 18 households.

Theory suggests that individual income variations should be used to define risk. The second problem is how to measure risk, given the appropriate variable. The statistical definition of risk refers to the average deviation of a variable from its mean. Other measures of risk are the normalized standard deviation or the variance of the risky variable. However, these measures are reliable only if the underlying time series is stationary. If the data exhibits any form of non-stationarity, these measures are biased. Since economic data are rarely stationary, the underlying time series should be made stationary before calculating the risk measure.

Two highly regarded studies on the structure of income, which use panel data, are MaCurdy (1982) and Hall and Mishkin (1982). MaCurdy (1982) treats the income series as a first-difference stationary process, while Hall and Mishkin (1982) treat the income series as a time- and life-cycle trend-stationary process. There are a number of tests for the presence of unit roots that are capable, in principle, of separating between trend-stationary and difference-stationary processes. However, with a short data series the predictions of the two processes are normally so close that the tests cannot separate them (Deaton, 1992).⁷ Not surprisingly, both of the income processes are widely used in the consumption literature.

This study adopts the approach used in Hall and Mishkin (1982). It is assumed that the change in income has a deterministic component that is a function of time- and life-cycle factors. In this case, the income process can be represented as

$$\pi_{it} = Z_{it}\beta_t + \tilde{\pi}_{it} \quad (6)$$

where β_t is a vector of parameters, $\tilde{\pi}_{it}$ the random component of farm income for the household i in the year t , and Z a vector of socio-demographic variables given by

$$Z_{it} = (A_{it}, A_{it}^2, Y_t, Y_t^2, S_{it}, S_{it}^2, \text{Household Dummies}_{it}) \quad (7)$$

where A_{it} indicates the age of the head of household i in the year t , S_{it} indicates the size of household i in the year t , Y_t the time trend common to all farm households, and $\text{Household Dummies}_{it}$ is a set of dummy

⁷Over a sufficiently long run, the two processes would look very different.

variables allowing both the intercept and slope to be different across farm households. This structure allows a different age/size/income profile, in each year, for each farm household. In estimation, the household slope dummies are excluded in order to insure a reasonable number of degrees of freedom. Under the usual assumptions, Eq. (6) can be estimated using ordinary least-squares yielding estimates of β_t . The estimated equation is then used to predict income in each year for each household. This procedure removes changes in the expected future income that were due to aggregate productivity growth and predictable life-cycle changes. The residuals of income can then be calculated, representing the stationary random component of income $\tilde{\pi}_{it}$, which can be used to measure risk.⁸ The standard deviation and variance of this ‘detrended’ income are taken as measures of income risk.

Since the econometric testing of prudence may be sensitive to the measure of risk used, several measures of income risk are considered. First, Friedman (1957) suggested that the expected future income can be estimated by a weighted average of current and past incomes. Langemeier and Patrick (1993) use a weighting scheme of $1/2\pi_t + 1/3\pi_{t-1} + 1/6\pi_{t-2}$. This formula is employed to derive expected future income in this study. The difference between the actual current income and the predicted future income gives the transitory income. The measure of risk can then be taken as either the standard deviation or variance of transitory income.

Second, Kimball (1990) derived an equivalent precautionary saving premium as a measure of risk. This measure should be a better measure of risk than the variance or the standard deviation of income, given that it is based on the theory of precautionary saving. Carroll (1994) showed that, under certain assumptions, it is possible to construct the equivalent precautionary saving premium. Modifying the Carroll (1994) formula for use with this data yields

$$\hat{\omega}_i = \bar{\pi}_i - \left[\frac{1}{8} \sum_{t=1979}^{1986} (\pi_{it})^{-R^r} \right]^{-1/R^r} \quad (8)$$

where $\hat{\omega}_i$ is an unbiased estimator of the true equivalent precautionary saving premium ω_i for farm household i , $\bar{\pi}_i$ the average of farm income from 1979 to 1986 for household i , and R^r is the coefficient of relative risk aversion, which is assumed to be 2.0.⁹ Using Eq. (8), the equivalent precautionary premium can be calculated for each household.

Third, Mao (1970) and Menezes et al. (1980) showed that utility maximizers are more likely to be down-side risk averse. As shown in Whitmore (1970), down-side risk aversion is implied by utility functions with a positive third derivative. This condition is the same as the condition for the existence of precautionary savings and it implies that downside risk should be a better measure of risk than a variable’s variance or standard deviation. Semi-variance is a widely used measure of downside risk, it is defined as

$$v_{it}^s = E(\pi_{it} - \pi_i^t)^2 \forall \pi_{it} < \pi_i^t \quad (9)$$

where v_{it}^s is the semi-variance and π_i^t the target value of farm income. Intuitively, semi-variance is the expected value of the squared deviation below a target value. Semi-variance equals zero when actual farm income is above its target value. Target income is set equal to the mean of farm income, $\pi_i^t = 1/8 \sum_{1979}^{1986} \pi_{it}$. From Eq. (9), the semi-variance can be calculated in each year for each household.

4. Econometric considerations

The null hypothesis is that the current consumption of farm households is negatively related to income risk. To test this hypotheses Eq. (10),

$$C_{it} = \alpha_0 + \alpha_1 W_{it} + \alpha_2 \pi_{it} + \alpha_3 \sigma_{it}^2 + \alpha_4 W_{it} \sigma_{it}^2 + \varepsilon_{it} \quad (10)$$

is estimated, where α ’s are parameters and ε is the error term. Eq. (10) represents a generalized panel data regression model. Before estimation, the proper estimator needs to be chosen; ordinary least squares can be used if the disturbances are homoskedastic and independent. But, often this is not a realistic assumption.

⁸Strictly speaking, this variable may include some systematic variation, since farm households have different information sets to predict the future income.

⁹A range between one and five is considered to be reasonable for the magnitude of the coefficient of relative risk aversion. However, most economists believe a smaller number, hence, two is chosen here.

tion. The variance of ε_{it} might vary with t or i , or both. Moreover, the error terms ε_{it} and ε_{jt} might be correlated for some $i \neq j$, if random shocks affect several farm households at the same point of time. Similarly, the error terms, ε_{it} and ε_{is} , might be correlated for some $t \neq s$, if shocks affect the same farm household at more than one point in time. Consequently, an estimated covariance matrix based on OLS, if these failures occur, is inefficient and may lead to serious errors of inference. In some circumstances the parameter estimates are inconsistent.

Are any of the econometric problems likely to arise in this data? Descriptive analysis of the data suggests that

- (i) the variance is different for the consumption, income, and wealth time series; and
- (ii) the variance is different for consumption, income, and wealth across farm households.

Further, autocorrelation may be a problem because of the time dimension of the panel data. These considerations suggest that both, autocorrelation and heteroskedasticity are likely to occur. It is thus important to test the hypothesis of no autocorrelation and heteroskedasticity. On account of the absence of a general joint test, autocorrelation and heteroskedasticity are tested independently.

In view of the nature of panel data conventional serial correlation and heteroskedasticity tests are not meaningful. There are two potential forms of heteroskedasticity in the panel data, groupwise heteroskedasticity and cross-sectional correlation. In the first case, only the variance is allowed to vary across farm households. The second case allows for both, non-constant variance and correlation of the disturbances across farm households. Following Greene (1993), these two forms of heteroskedasticity can be tested independently. First, a null hypothesis of no groupwise heteroskedasticity can be tested using White's test. To carry out this test, Eq. (10) is estimated using OLS and then the squared residuals are regressed on a constant, independent variables, and squared independent variables. The chi-squared statistic is $N^*T^*\bar{R}^2$, where N is the number of cross sections, T the number of observations in each cross section, and \bar{R}^2 the adjusted R -square (referred to as R -square). The degrees of freedom for this chi-squared distribution equals the number of independent variables in the OLS residual

regression. Second, a null hypothesis of no cross-sectional correlation is tested using the Lagrange multiplier test developed by Breusch and Pagan (1980)¹⁰

$$LM = T \sum_{i=1}^N \sum_{j=1}^{i-1} r_{ij}^2 \quad (11)$$

where r_{ij}^2 is the ij th residual correlation coefficient. To conduct this test, a groupwise heteroskedastic model is estimated based on Eq. (10). LM is then calculated based on the residual correlation matrix.¹¹ LM is a chi-square distribution with $N(N-1)/2$ degrees of freedom.

The fact that the panel data used here has only eight time-series observations for each farm household makes a test for autocorrelation difficult. Langemeier and Patrick (1990) use the same data and applied the Q -statistic proposed by Box and Pierce (1970) to test for first-order autocorrelation:¹²

$$Q = T \sum_i \hat{r}_i^2 \quad (12)$$

where \hat{r}_i is the sample autocorrelation coefficient for each cross section. The Q statistic is based on a chi-square distribution with $(N-1)$ degrees of freedom. Langemeier and Patrick used the estimated first-order autocorrelation statistics for each cross section as the sample autocorrelation coefficient. When T is eight, however, the estimated first-order autocorrelation statistic for each cross section is unreliable. Instead, \hat{r}_i can be estimated directly as

$$\hat{r}_i = \frac{\sum_{t=2}^T \varepsilon_{it} \varepsilon_{it-1}}{\sqrt{\sum_{t=2}^T \varepsilon_{it}^2} \sqrt{\sum_{t=2}^T \varepsilon_{it-1}^2}} \quad (13)$$

¹⁰This test is different from the usual Breusch and Pagan (1980) test of heteroskedasticity which can be readily computed in most applied econometric software.

¹¹Though most econometric software have the capacity to estimate the model, their ability to manipulate panel data is extremely limited. For this reason, a program was written using MATLAB to facilitate the calculation of the LM as well as the Q -statistics.

¹²Davidson and MacKinnon (1993) point out that Box and Pierce's test may be invalid when used with the residuals from linear or non-linear regressions. However, given the data in this study, Box and Pierce's test appears to be the only alternative.

Table 1
 Estimation results for the reduced-form consumption equations ^a (Dependent variable: Total consumption C_{it+1})

Independent variable	Risk measures					
	SD of detrended income (1)	variance of detrended income (2)	equivalent precautionary premium (3)	downside risk (4)	SD of transitory income (5)	variance of transitory income (6)
Current income	0.1107 (5.9467) ^c	0.0354 (2.5685) ^c	0.0139 (1.1728)	0.0404 (2.8626) ^c	0.0489 (2.8200) ^c	0.0300 (2.1440) ^c
Wealth	0.0105 (5.3546) ^c	0.0126 (7.5233) ^c	0.0118 (5.0519) ^c	0.0131 (6.7781) ^c	0.0125 (7.0400) ^c	0.0126 (7.7763) ^c
Risk measure	−0.1096 (4.9241) ^c	−0.7090E-07 (1.0142)	−0.1489 (3.8698) ^c	−0.1391 (3.1347) ^c	−0.0697 (2.2954) ^c	−0.3306E-06 (1.1550)
Interaction between risk and wealth				0.1643E-06 (2.0171) ^c		
Year	129.9200 (0.5067)	47.5340 (0.1616)	98.655 (0.3279)	−226.9300 (0.8363)	−375.1000 (1.5316)	−50.8490 (0.1694)
Age	221.7500 (1.8675) ^b	−0.0008 (0.3065)	−0.0008 (0.3058)	−0.0007 (0.2883)	−0.0005 (0.20001)	−0.0008 (0.3114)
Household size	1299.6000 (2.1431) ^c	115.2100 (0.1370)	−89.919 (0.1383)	748.5800 (1.4505)	823.5700 (1.5104)	298.77 (0.3380)
Constant	6599.4000 (1.0842)	14483.0000 (1.8964) ^b	18352.0000 (2.5334) ^c	10409.0000 (1.5786)	10092.000 (1.5659)	13306.0000 (1.6950) ^b
Buse R^2	0.44	0.38	0.28	0.41	0.50	0.38
White's test	25.92 ^c	26.68 ^c	17.28 ^c	43.20 ^c	17.28 ^b	23.04 ^c
LM	189.79	189.49	188.28	186.24	192.54	188.68
Q	13.45	13.27	18.92	12.51	13.38	13.42
RESET	11.63 ^c	0.15	/ ^d	8.00	8.40	3.12
F test for dummies	5.94 ^c	8.53 ^c	6.40 ^c	6.01 ^c	6.35 ^c	8.53 ^c
Implied saving rate (percent)	6.96	2.62	5.71	3.99	3.87	2.59

^a t -Statistics are given in parentheses.

^b Significantly different from zero at the 10 percent level of significance.

^c Significantly different from zero at the 5 percent level of significance.

^d Estimation failed due to non-positive matrix.

The calculated White, LM, and Q statistics are presented at the bottom of Table 1. The White statistics are significant at either the 5 or 10% level in all cases, while the LM and Q statistics are significant at either the 5 or 10% level in only a few cases. It is concluded that the data exhibits strong evidence of groupwise heteroskedasticity and weak evidence of autocorrelation and cross-sectional correlation.¹³

¹³In contrast, Langemeier and Patrick found strong evidence of autocorrelation. This may be due to the difference in the calculation of the sample autocorrelation coefficient.

A number of different econometric models can be used with panel data (Kmenta, 1986; Hsiao, 1988; Matyas and Sevestre, 1992), including Kmenta's model, a covariance model, an error-components model and a random-coefficient model. The Kmenta and error-components models are popular in applied research (Baltagi, 1986) and differ in the assumptions imposed on the disturbances. The error-components model has homoskedastic disturbances, whereas the Kmenta model has heteroskedastic disturbances. Both models allow for serial correlation, but in the usual error-components model this correlation is

constant across time, whereas it decays over time with the Kmenta model. Moreover, the Kmenta model allows for correlation among the different cross sections, whereas the error-components model assumes independence among disturbances. Monte-Carlo results by Baltagi (1986) indicate that while the usual error-components model is generally expected to perform better with large cross-sectional units and a small time series, the Kmenta model is expected to perform better otherwise. Recent studies using panel data have attempted to account for a more sophisticated error structure within the error-components model (Baltagi and Griffin, 1988, 1991), but these models have not been extended to empirical use. Given the fact that both the autocorrelation and cross-sectional heteroskedasticity are present in the data used, it is more appropriate to use Kmenta's model.

There are various forms of the Kmenta model, depending on the specific assumptions with regard to the disturbances. Two alternatives are:

- (i) the groupwise heteroskedastic and timewise autoregressive model; and
- (ii) the cross-sectionally correlated and timewise autoregressive model.

The characterization of models (i) and (ii) are given in Kmenta (1986). The question of which model to use is determined using the statistical tests discussed earlier.

To take into account the household-specific effects, household dummy variables are added to the regression equation. To have a reasonable number of degrees of freedom, only intercept dummies are added. The significance of the household specific dummy variables is determined using an *F*-test. The set of household dummy variables, which have a significant impact at the >10% significance level, are included in the final regression. No attempt is made to consider each dummy variable independently. Similarly, a time variable (year) is added to account for common shifts such as taste, weather and price. Age of the household's head is added to account for life-cycle effects. The size of the farm household is added to account for the size effect in consumption. In estimation, the risk-free interest rate and discount rate are assumed to be constant across the farm households. The interaction term between wealth and risk is included in the final

regression only when it is statistically significant at the >10% level.

After the test for autocorrelation and heteroskedasticity, the Ramsey (1969) RESET test is conducted to detect possible misspecifications or omitted variables. RESET is constructed by means of a regression of the dependent variable on the original variables and the powers of the predicted values of the dependent variable. The number of added variables is the square, cube, and fourth power of the predictions obtained from the Kmenta model.¹⁴ The RESET statistic follows an *F*-distribution with degrees of freedom equal to the number of independent variables in the second-stage regression.

5. Empirical results

The RESET test results are presented at the bottom of Table 1. For most of the equations, the null hypothesis of no misspecification is not rejected. The estimation results for Eq. (10), corresponding to each of the six measures of income risk, are presented in Table 1. Because the standard *R*-square measure may not fall between 0 and 1 under generalized least-squares estimation, the Buse *R*-square, which does fall within these bounds, is reported (Buse, 1973). The *R*-square values range from 0.28 to 0.50 across equations indicating that the estimated models fit the data reasonably well.

Table 1 contains the estimation results for the reduced-form consumption (Eq. (10)). All risk measures, except the variance, are statistically significant at the >5% level. The negative signs for the risk measures show that risk affects consumption negatively. Households that face more risk defer their consumption, as predicted by the theory of precautionary saving. The interaction term between risk and wealth is not statistically significant, except in one equation when risk is measured by downside risk. The positive sign implies that farm households exhibit decreasing prudence. Ramsey's RESET statistics of 8.0 and 8.6 are obtained for the model in columns 4 and 5, which are statistically insignificant, thus not rejecting the null hypothesis of no misspecification.

¹⁴Household dummies are excluded in all regressions.

The attempt to calculate RESET in Column 3 fails due to a non-positive matrix. A Ramsey's RESET statistic of 11.63 is obtained for the model in Column 1, which is statistically significant at the 5% level, thus rejecting the null hypothesis of no misspecification. Alternative specifications were tried but none provided a better result.

While the estimation results suggest that the coefficients for the risk variables are statistically significant, a more interesting question is the magnitude of these effects. If they are so small as to imply negligible effects on savings, then the results are less interesting. A crude way to judge the magnitude of these effects is by examining what would happen to consumption if each measure of income risk increased by a given amount. Suppose that each risk measure is increased by an amount equal to its own cross-sectional standard deviation. The standard deviation of the equivalent precautionary saving premium is \$24 591. The estimated coefficient on the premium is -0.1489 . This implies that if the equivalent precautionary saving premium were increased by \$24 591, consumption would fall by $0.1489 \times 24\,591 = \3688 . This represents 5.71% of the average income, which implies that the effect of the savings rate on income is 5.71%. These changes are quite large relative to an aggregate personal saving rate of approximately 5.6% in the US. The increases in the savings rate caused by a one standard-deviation increase in the other measures of income risk are reported at the bottom of Table 1. The implied savings rate ranges from 3.99 to 6.96. These results cast substantial doubt on the relevance of the certainty model which implies that income risk has no effect on farm consumption. The precautionary saving motive appears to be an important part of the consumption behavior of farm households.

Given that the data used in this study applies to a panel of 18 farmers over only eight years, it is reasonable to be concerned about the generality of the findings. One method of collaborating the results is to compare them with the findings from previous research. Dardanoni derived a closed form solution for present consumption as a function of future income risk (in addition to conventional variables), under the assumption of constant absolute risk-aversion and a normal distribution of future income. He found, using cross-sectional data from the 1984 UK Family Expen-

diture Survey, that income risk had a significant effect on current consumption. However, due to the cross-sectional nature of the data, his measure of income risk is rather ad hoc. He grouped data into cohorts by occupation and used the variance of labor income within each cohort as a proxy for the variance of future income. Conversely, Carroll (1994) measured income risk by the 'equivalent precautionary saving premium', defined by Kimball (1990), and found that income risk has a significant effect on individual current consumption. Unhappy with the use of proxies for income risk, Guiso et al. (1992) used data from an Italian household survey to see if consumption was related to a self-reported expected variance of the following year's income. They found a statistically significant, though small, relationship between self-reported income risk and consumption. Therefore, the empirical findings for farm households are consistent with the limited empirical evidence of the impact of risk on consumption in non-farm families, although the magnitude of the estimated risk effect is larger in this study. Considering that farmers are widely believed to face more income risk than other groups in society, a larger risk effect on consumption is not a surprising result.

6. Conclusions

Using panel data from Illinois grain farmers a direct test of the relationship between income risk and farm consumption behavior is conducted. The estimation results indicate that income risk significantly affects farm consumption and precautionary savings are a significant proportion of total savings. The results are robust to alternative risk measures. These findings cast substantial doubt on the relevance of the conventional life-cycle permanent income hypothesis, which implies that risk has no effect on consumption. The consumption/income divergence which is inconsistent with the Keynesian model can be explained using Friedman's logic; consumption does not respond strongly to transitory shocks to income because assets are used to buffer consumption against such shocks. A careful reading of Friedman (1957) suggests that the risk-augmented life-cycle model may be a closer approximation to his original thinking than the modern interpretations.

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