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Introducing socioeconomic characteristics into production analysis under risk

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Abstract

A theoretical framework is developed to study the effects of socioeconomic factors on farmers' risk attitudes and production decisions. No maintained assumptions about the individual's utility are required. A key element in this framework is the categorization of socioeconomic factors by their effect on the farmer's risk attitudes. A simple methodology for this categorization, based on the equivalence between the Arrow–Pratt measure of risk aversion and the probability of winning demanded, is proposed. The latter is illustrated with data collected in a survey of 180 Israeli farmers from 20 villages.

1. Introduction

Worldwide, the operators of many agricultural enterprises are self-employed. Among other factors, the owner-operators of such farms differ from each other in their socioeconomic characteristics (SECs). Variations in SECs may cause differences in risk preferences, which in a risky environment will lead to heterogeneous production decisions. Successful empirical studies of risk attitudes among such a population and the design of risk-reducing policies require an understanding of the rules governing this heterogeneity.

The major objective of this paper is to lay out

a framework that enables the analysis of the effects of SECs on the heterogeneity of production decisions under risk. The main result is an intuitively appealing condition for determining the sign of the effect of any SEC that influences risk preferences on any choice variable that affects the farm's profit. Key elements in this condition are categorizations of SECs as risk-aversion increasing/decreasing and of inputs and other decision variables as marginally risk-increasing/decreasing. In the past decade, a fair amount of information related to the latter categorization has been accumulated in the agricultural economics literature (e.g. Just and Pope, 1979). On the other hand, information regarding the effects of SECs on risk attitudes is limited. Therefore, an approach for the empirical catego-

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rization of various SECs is developed based on the concept of probability of winning demanded (PWD). This method is then applied to a sample of 180 Israeli farmers, residing in 20 villages. The empirical results show that SECs play an important role in explaining variations in risk attitudes among Israeli farmers.

More than two decades of fruitful application of the expected utility theory have proven the importance of the Arrow–Pratt (Arrow, 1970; Pratt, 1964) measures of risk aversion in analysing decision-making under risk. As asserted by Yaari (1969), these measures have one major role in such analyses: they make a convenient analytical tool for formulating and testing certain hypotheses regarding the behavior of decision-makers under risk. Examples are (*ceteris paribus*): (1) the wealthier of two farmers can never be more risk averse; (2) of two farmers, the one with the larger family can never be less risk averse; (3) of two farmers, the more educated one can never be more risk averse. In combination with studies about the effects of risk aversion on production decisions (e.g. Sandmo, 1971; Pope and Kramer, 1979; Feder et al., 1980; Ramaswami, 1992), such axioms, coupled with empirical testing, can be used for policy design and general predictions.

To date, most of the work on the factors determining the sign and magnitude of an individual's measures of risk aversion have focused on the effect of wealth level (Binswanger, 1980; Bardsley and Harris, 1987; Antle, 1987; Love and Buccola, 1992 and Bar-Shira, 1992). The analysis presented here may be viewed as a generalization of a few studies (Binswanger, 1980; Dillon and Scandizzo, 1978; and Moscardi and de Janvry, 1977) that investigate not only the impact of wealth but also the role of various SECs in determining the farmer's level of risk aversion. Examples of such SECs are education level, wealth, family size, origin of the farm operator, operator's age, farm size and water availability. A special feature of the proposed framework is the ability to isolate the effects of SECs via risk aversion from their effects via the individual's ordinal preferences for commodities. Moreover, no maintained assumptions regarding the form of the farmer's utility function are required. The only ad

hoc practice used in the analysis is a linear regression model aimed at estimating the effects of various SECs on the PWD.

Finally, while several papers (e.g. Pope and Kramer, 1979) have studied the effect of increasing risk aversion on specific production decisions, a general theory is lacking. The current paper attempts to fill this gap. While such a theory exists for general decision problems (Diamond and Stiglitz, 1974), the current extension, which explicitly considers the agricultural production context, provides more intuition and is empirically applicable.

2. SECs, risk attitudes and production decisions

A convenient and general approach for studying various factors that influence decision-making under risk was introduced by Diamond and Stiglitz (1974), (hereafter, DS). Their framework allows an investigation of the impact of any factor affecting the decision-maker's risk attitudes, on choices under risk. They proposed several alternatives to measure the intensity of an individual's risk aversion, for example the absolute measure of risk-aversion and the Pratt risk premium. They then proved that information about the effect of a SEC on any measure of the individual's risk aversion suffices for finding the direction of the analogous effects on each of the alternative measures. Perhaps more importantly, such information also suffices for determining the direction of the effect of this parameter on choices under risk.

In this section we derive a sufficient condition for determining the effect of a SEC on a choice variable that affects producer's profit (hereafter a profit factor). This condition is an alternative to the 'single crossing' condition, proposed by DS. However, the newly proposed condition is based on the categorization of production factors (and other decision variables affecting profit) as marginally risk-increasing or risk-decreasing, which is intuitive and familiar to most agricultural economists and has been empirically investigated in a number of studies. Therefore, the newly proven condition may be applicable and useful for studying production decisions under risk.

3. The effect of SECs on production decisions via risk attitudes

The following notations and assumptions are used. Let the farmer's profit, π , be a function of a random variable ϵ , such as random output price risk, yield risk, input price risk, and a profit factor, α , that may represent the level of output, input choice, position in the futures market, etc.¹ The farmer's objective function is the expected value of the von Neumann–Morgenstern utility function, U . The general form of the utility function is $U(\epsilon, \alpha, \rho)$, where ρ is a SEC, possibly a vector, that is systematically associated with the farmer's risk attitudes. This formulation allows the control variable to affect utility directly and not only via profit or income. For example, α may represent the level of the operator's on-farm labor, which affects utility through both income and leisure.

Just as production factors are categorized as either risk-increasing or risk-decreasing (Pope and Kramer, 1979), SECs can be categorized into those that increase the farmer's risk aversion and those that decrease it. Formally, the distinction is given by the following definition:

Definition 1. Consider a family of utility functions $U(\epsilon, \alpha, \rho)$; $\rho_i \in \rho$ is said to be a risk-aversion increasing (decreasing) (neutral) SEC if and only if

$$\frac{\partial A}{\partial \rho_i} = - \frac{\partial^2 \log U_\epsilon}{\partial \epsilon \partial \rho_i} > (<) (=) 0$$

where U_ϵ is the partial derivative of U with respect to ϵ , and A is the measure of absolute risk aversion.

That is, the index of absolute risk aversion, A , is increasing (decreasing) (does not vary) with ρ_i .

Linking the effects of SECs on risk attitudes with their effects on profit factors (namely, production-associated control variables), facilitates a

systematic introduction of socioeconomic factors into the theory of production under risk. This linkage is formally established below.

The typical family farm analyzed in the empirical section of this study is a legal consumer-producer economic unit, with a single owner-operator. Farm production decision and consumption behaviors are determined simultaneously. Socioeconomic factors enter directly into the farmer's utility and affect his welfare in combination with wealth and leisure. A change in a particular SEC may affect the production choices of such a household in two ways. First, it may affect the farmer's indifference map and, by doing this, change his choices. For example, a larger family may affect the farmer's marginal rate of substitution between income and leisure and change his time allocation between leisure and on-farm labor. These types of effects are not related to risk and have been extensively investigated elsewhere (e.g. Dawson, 1984).

Secondly, a change in a SEC may affect the farmer's risk attitudes and hence modify his production decisions. For example, a larger family may increase the farmer's aversion to risk and as a result he may show less readiness to adopt risky technologies on his farm. In some cases, the two effects may coincide in sign, but this is not necessarily the case.

Our interest is in the second type of effects, namely, the effects of SECs on choices via risk attitudes, rather than through their effects on the individual's indifference map. We therefore formulate a tractable model that allows separability between ordinal preferences and risk attitudes. This separability facilitates an isolation of the effects of various exogenous variables (SECs) on the individual's choices via risk attitudes.²

The approach taken in the literature of risk aversion with a multi-argument utility function for coping with such situations is best stated by Kihlstrom and Mirman, 1974, "In face of this

¹ The analysis can be generalized to cases in which α and/or ϵ are vectors, rather than scalars. However, these generalizations require significant complication of the mathematical model and are beyond the scope of this paper.

² The interaction between preferences for goods and risk attitudes in their effect on decisions under risk has been investigated in several studies (e.g. Finkelshtain and Chalfant, 1991).

difficulty the most natural approach, and the one taken here, is to limit comparisons of risk aversion to utility functions which represent the same ordinal preferences.” The analysis below adopts this solution. Formally, the requirement of identical indifference curves implies that the family of utility functions indexed by ρ , $U(\alpha, \epsilon, \rho)$, can be written in a separable form as $V(U(\epsilon, \alpha), \rho)$, where V is an increasing-monotone transformation of the familiar von Neumann–Morgenstern utility function, U and its concavity increases with any risk-aversion increasing SEC. In other words, comparisons are conducted between individuals who differ in their degrees of risk aversion and preferences for most commodities, but possess identical indifference curves between the random variable ϵ and the decision variable (profit factor) α .³

We categorize profit factors as marginally risk-increasing or risk-decreasing. Following the literature about production under risk (e.g. Just and Pope, 1979; Pope and Kramer, 1979), inputs are categorized as marginally risk-decreasing/increasing when their price exceeds/falls short of the expected value of the respective marginal products at the optimum for a risk-averse producer. The expected marginal profit of a marginally risk-decreasing/increasing input is negative/positive at the optimum. Assuming the Just and Pope (1979) production function, the above definition is intuitive, since marginally risk-decreasing/increasing inputs accordingly decrease/increase the variance of output.

The above categorization of inputs can be extended to apply to other profit factors (namely, decision variables affecting profits). For example, the output level of the firm can be thought of as a marginally risk-increasing profit factor, since it increases the variance of profits and the expected marginal profit indeed exceeds zero for risk-

averse producers (Sandmo, 1971). Similarly, the proportion of the area planted for a crop which has been chosen to be insured is a risk-reducing factor. Thus, we introduce the following definition:

Definition II. A profit factor, α , is said to be marginally risk-decreasing (increasing) (neutral) if and only if its expected marginal profit falls short of (exceeds) (equals) zero at the optimum for any risk and risk-averse producer⁴.

Lemma I in Appendix A shows that under the model assumptions, a profit factor α is marginally risk-increasing (decreasing) if $\text{sgn}(\pi_{\alpha\epsilon}) = (\neq) \text{sgn}(\pi_{\epsilon})$. The latter condition can be used as an alternative definition of a marginally risk-increasing/decreasing factor. The intuition is immediate, if the signs of the above derivatives are equal (not equal), α increases (reduces) the profit fluctuations caused by the random variable ϵ . In the following, it will be assumed that $U_{\epsilon} \geq 0$. This assumption is only for the sake of readability, and the same results follow with the alternative assumption, $U_{\epsilon} \leq 0$. We can now state and prove the main result of the paper.

Theorem I. Consider the above setting and let α^* be the level of the control variable that maximizes expected utility ($\alpha^* = \arg\max E(V[U(\pi(\alpha, \epsilon)), \rho])$) where π represents profit. Suppose that ρ_i is a risk-aversion increasing SEC⁵ ($\partial A / \partial \rho_i > 0$), then any marginally risk-increasing (decreasing) (neutral) profit factor decreases (increases) (does not change) with ρ_i ($\partial \alpha^* / \partial \rho_i < (>)(=) 0$).

Proof of Theorem I is provided in Appendix A. The next subsection demonstrates, by several

³ It should be noted, however, that the vast majority of existing studies of production under risk presume that the utility function is defined over wealth. In such models the assumption of identical indifference curves for ϵ and α is satisfied trivially, since all producers share identical ordinal preferences, i.e. they all prefer more to less.

⁴ A recent paper (Ramaswami, 1992) advanced the definition of risk-increasing factors and provided an alternative useful definition, which is consistent with the notion of an increase in risk (e.g. Diamond and Stiglitz, 1974).

⁵ An analogous theorem, with the obvious modifications, exists for the cases of risk-aversion decreasing and neutral SECs.

examples, that Theorem I is applicable to a wide range of production decision problems under risk.

Examples

(a) Output level under price risk. Let ϵ and α represent the random output price and the output level, respectively. In this case the signs of π_ϵ (= output level) and $\pi_{\epsilon\alpha}$ (= 1.0) are both positive, implying (see Lemma I in Appendix A that output is a marginally risk-increasing profit factor (MRIPF). Therefore, if an increase in a specific SEC, ρ_i , represents an increase (decrease) in risk aversion, then the optimal output level decreases (increases) with ρ_i .

(b) Adoption of a new crop variety (or a new cultivation technology). Let ϵ and α denote the random profitability per unit of land area and the area planted for the new crop. The total farm level profits, π , are given by $\pi = \alpha\epsilon + (\bar{L} - \alpha)K$ where \bar{L} represents total land area and K is the deterministic profit per unit area planted for the traditional crop. As in the previous example, the signs of π_ϵ (= α) and $\pi_{\epsilon\alpha}$ (= 1.0) are both positive, implying that the area planted for the new crop is a MRIPF. Therefore, if an increase in a specific SEC, ρ_i , represents increases (decreases) in risk aversion, then the optimal level of α decreases (increases) with ρ_i .

(c) Part-time farming. Suppose ϵ represents the random profit associated with farming activities and α represents the share of on-farm labor. Assuming that the farmer's alternative wage rate is deterministic, it can be verified that the share of on-farm labor is a MRIPF and therefore, $\partial A / \partial \rho_i > (<) 0$, implying that the optimal share of on-farm labor decreases (increases) with ρ_i .

(d) Crop insurance. Denote by ϵ the random crop yield per acre and by α the area chosen to be insured out of the total farm land, L , planted for the crop under consideration. Consider an insurance contract which provides the farmer with a 'guaranteed' yield level, equal to the yield expected value $\bar{\epsilon}$. The profit associated with this crop can be written as

$$\pi = \alpha \{ P\epsilon + P \max[0; (\bar{\epsilon} - \epsilon)] - \xi \} \\ + (\bar{L} - \alpha) P\bar{\epsilon} - C\bar{L}$$

where P is the market price, ξ is the per unit area insurance-premium function and $C(L)$ is the cost of production. In this case

$$\pi_\epsilon = \begin{cases} P(\bar{L} - \alpha), & \text{if } \epsilon \leq \bar{\epsilon}; \\ P\bar{L}, & \epsilon > \bar{\epsilon} \end{cases}$$

In both cases $\pi_\epsilon \geq 0$. The sign of $\pi_{\epsilon\alpha}$ is either negative (= $-P$, for $\epsilon \leq \bar{\epsilon}$) or zero. Therefore, as expected, the area chosen to be insured is a marginally risk-decreasing profit factor (MRDPF). Therefore, $\partial A / \partial \rho_i > (<) 0$ implies that the optimal crop area chosen to be insured increases (decreases) with ρ_i .

(e) Participation in futures markets. Let ϵ represent an output price risk and let α represent the output sold in the futures market at a known price, denoted by ϵ^f . The relevant profit function is $\pi = \epsilon^f \alpha + \epsilon(y - \alpha) - C(y)$, where y is total output (assumed to be deterministic) and $C(\cdot)$ is a cost function. It can be shown (e.g. Holthausen, 1979 Feder et al., 1980) that for a risk-averse producer, the optimal output level is $y = C^{-1}(\epsilon^f)$ and is independent of both the producer's risk attitudes and the level of risk. Note that $\text{sgn}(y - \alpha) = -\text{sgn}(\epsilon^f - E(\epsilon))$. Since $\pi_\epsilon = y - \alpha$ and $\pi_{\epsilon\alpha} = -1$, the output sold in the future market is a MRDPF (MRIPF) if $\epsilon^f < (>) E(\epsilon)$. Thus, $\epsilon^f < E(\epsilon)$ implies that if an increase in a specific SEC, ρ_i , represents an increase (decrease) in risk aversion, then the optimal level of output chosen to be sold in the future market increases (decreases) with ρ_i . If $\epsilon^f > E(\epsilon)$, the conclusion is reversed.

4. A simple method for empirical categorization of SECs

Given information about the effect of various SECs on farmers' risk preferences, Theorem I provides a vehicle to predict the effects of such factors on production decisions. This section presents a simple method for eliciting this information. The next section presents an empirical study which applies this method.

Our interest is to elicit the direction of the effect of each SEC on the absolute measure of

risk aversion. One way to proceed with this problem is via two steps. First, to assume some specific form for the farmer's utility function (e.g. the constant absolute risk aversion (CARA) utility function), estimate its parameters, and based on the estimation, evaluate the absolute measure of risk aversion. In the second stage, one would regress the estimate of absolute measure of risk aversion against the levels of SECs to determine the sign of their effects. The drawback of this method is the need to presume a functional form for the farmer's utility function. The latter could result in biased estimates. Another source of potential bias in the estimates is related to the specification of the regression model. As our empirical application is based on a regression analysis, our result are not immune to this problem.

Since the only information required to categorize SECs is the direction of their effects on the measure of risk aversion, an alternative indirect method of eliciting this information is proposed. The methodology is based on the equivalence between several measures of the intensity of an individual's aversion to risk. It has been shown previously (DS) that if a SEC increases the absolute measure of risk aversion it also increases the Pratt risk premium and vice versa. We now extend their theory by proving an analogous result regarding the PWD, defined below.

Denote by $[p, \epsilon^1, \epsilon^2]$ a lottery with two possible prizes ϵ^1, ϵ^2 ($\epsilon^1 > \epsilon^2$), and p is the probability of winning the larger prize. Suppose an individual is facing the choice between such a lottery and a fixed amount of money $\hat{\epsilon}$, such that $\epsilon^1 > \hat{\epsilon} > \epsilon^2$. The PWD, p , is defined by

$$pU(\epsilon^1, \rho) + (1-p)U(\epsilon^2, \rho) = U(\hat{\epsilon}, \rho)$$

We can now state the following theorem.

Theorem II. Any risk-aversion increasing (decreasing) (neutral) SEC, according to Definition I, increases (decreases) (does not affect) the PWD, i.e. $(\partial p / \partial \rho_i > (<)(=) 0)$.

The proof of the theorem is shown in Appendix A. This theorem proves that the sign of the derivatives of the PWD and the absolute

measure of risk aversion, A , with respect to a specific SEC are equal. The concept of PWD is only defined for two outcome risks. However, it should be emphasized that the derivative of the PWD with respect to a specific SEC indicates the sign of the derivative of the absolute risk aversion with respect to this SEC, in the presence of any risk whatsoever.

Thus, it facilitates the categorization of SECs based on information regarding the effects of these SECs on the PWD. Accordingly, our indirect method is based on elicitation of the PWD, a measure of risk-aversion, which is both intuitive and easily explained to farmers, and a regression of this measure against the level of SECs. The next sections illustrate the method with Israeli family farm data.

5. The data set

This section describes the data used for the empirical investigation. The data set is based on a survey which was designed to collect information about socioeconomic factors and risk attitudes of a random sample of farmers from the northern part of Israel. The survey included 180 family farms, located in 20 cooperative villages (moshavim) from four geographical and climatic regions: the Hefer Valley, the Jezre'el Valley, Galilee and the Golan Heights. The survey was performed via personal interviews conducted by a graduate student⁶.

⁶ There is extensive literature regarding survey methodologies and possible biases, especially in the context of natural resource valuations (e.g. Kealy and Turner, 1993 and references therein). The questions that consumers are asked in such surveys do not resemble daily decisions that consumers must take. The current survey is based on decision problems which are similar to those that farmers face daily and hence we expect its results to be less susceptible to the above criticisms. Moreover, our interest focuses mainly on the direction of the effects and not their size, thereby further reinforcing the argument. This is not to say, however, that the empirical analysis conducted here is free from error. A different survey design would probably yield some variation in the size of the effects, but hopefully not in their signs.

Each farm operator was asked to choose the minimum probability for success (or PWD) at which he would adopt a new agricultural technology in his area of specialization. Adoption of the new technology was presented as a two-prize lottery in which the farmer may win or lose a fixed amount of dollars. Two levels of prizes were considered: (1) $h = \$1000$ and (2) $h = \$10000$. The PWD was determined as the minimum probability of winning which would trigger adoption. The areas of specialization included livestock, orchards, greenhouses and field crops. In addition to the above, the questionnaires elicited information about various characteristics and socioeconomic factors of the farm and its operator.

The resultant data set contains panel data regarding farmers' choices of PWD corresponding to the two lotteries and an array of characteristics of the farm and its operator. The following characteristics are included:

Farm size (dunams, where 1 dunam = 1000 m²) and annual water quota (m³); herd inventory and annual milk quota (l) for farms specializing in livestock; broiler and/or layer inventory, and annual egg quota for farms specializing in poultry; qualitative dummy variables for the above-mentioned major areas of specialization; age, education level (years of schooling) and family size of the farm operator. In addition, we utilize several qualitative dummy variables. First, we have dummy variables for part-time farmers and full-time farmers, where the 'left-out' category consists of farmers who work on their farm for only a few hours a week. Second, we have dummies that indicate the period during which the cooperative was established. Specifically, there are dummies for cooperatives that were established after the Six Day War in 1967, and cooperatives under the supervision of the Jewish Agency that were established after 1948. The 'left-out' category are co-

Table 1
Descriptive statistics of recorded variables

Variable	Minimum	Maximum	Mean	Standard Deviation
PWD (%)	30	99	62.3	19.9
Farm size (1000m ²)	3	600	49.9	57.9
Water quota (1000m ³)	6	100	22.5	12.8
Milk quota (L)	60 000	550 000	301 075	85 478
Eggs quota	28 000	500 000	388 095	76 591
Family members	1	10	5.2	1.5
Schooling years	4	18	12.0	2.1
Age	23	73	43.8	9.6

Farm type	Number
Green house	14
Plantation	57
Cattle	22
Dairy	53
Poultry	21
Others (field crops)	13
Few hours farming	5
Part time farming	52
Full time farming	123
Agency settlement ^a	91
New settlement ^b	36
Established settlement	53

^a Agency settlement, villages that receive financial and managerial support by the Jewish Agency.

^b New settlement, villages that were established after the Six Days War (1967). These villages are characterized by relatively experienced and educated farmers as well as capital intensive production technology.

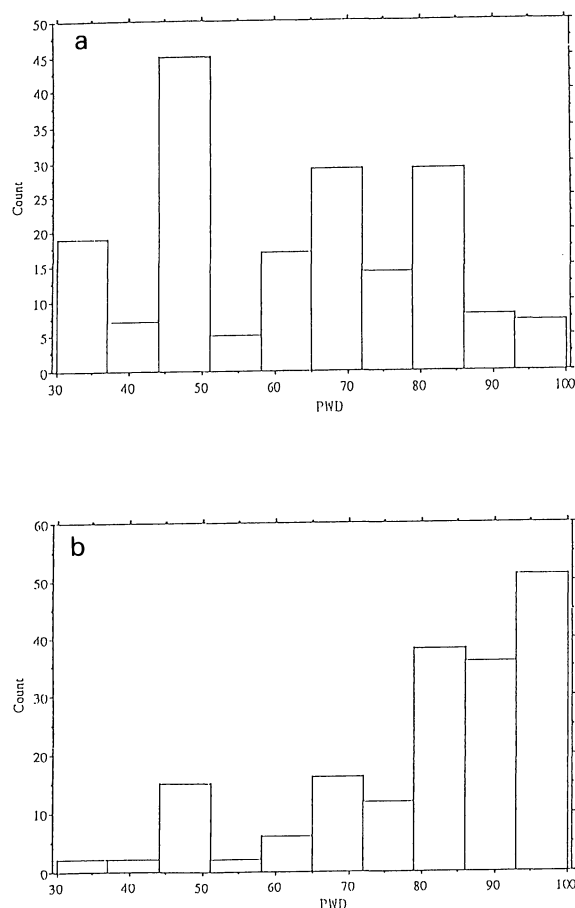


Fig. 1. The distribution of the PWD with prize levels of $h = 1000$ (1a) and $h = 10000$ (1b).

operatives that were established prior to 1948. Table 1 presents descriptive statistics of the principal variables from the data set.

6. Empirical findings

The collected data pertaining to risk attitudes indicate wide variations in the PWD (Fig. 1(a) and Fig. 1(b)). Moreover, a comparison of these figures clearly demonstrates that the distribution of the PWD shifts to the right as the prize level increases from \$1000 to \$10000. The mean PWD at prize level \$1000 is 0.62, and at prize level \$10000 it increases to 0.818. Moreover, the skewness statistics -0.04 at \$1000 and 1.03 at \$10000

indicate that the distribution of the PWD is skewed much more to the left at the higher prize level.

One is tempted to conclude from this result that farmers do not exhibit CARA. However, this need not be the case. Consider a farmer with certain income W^0 and a random income z . Let $z = [h, -h, p, 1-p]$ be a bet to gain or lose a fixed amount $h \in (0, W^0)$ with probabilities p and $1-p$, respectively, where p is the farmer's PWD. It can be easily verified that the mean income for a farmer who requires a PWD p is given by $W^0 + h(2p - 1)$, and its variance is $4h^2p(1-p)$. Thus, it becomes apparent that when h changes, both the mean of the lottery and higher moments change. Therefore, without putting additional structure on the farmer's utility function, no conclusion about the CARA or the DARA (decreasing absolute risk aversion) hypotheses can be drawn.

Moreover, since both the mean and the variance are affected by a change in h , it is not clear a priori at which level of h the PWD should be larger. The assumption of maximization of the expected value of a monotonic and concave utility function is not sufficient to predict at which level the PWD should be larger.

We turn, now, to a regression analysis, in an attempt to explain the variation in PWD by socioeconomic variables. In general, the econometric analysis fits the framework of the two-limit probit regression model (Rosett and Nelson, 1975) since the dependent variable, the PWD, can only take values between zero and one. However, in the present case the dependent variable was observed for all cases, since no observation reached the limits. Accordingly, application of an OLS regression would yield the unbiased estimates. Nevertheless, since the data are from a panel, in which the choices of PWD by each farmer are observed twice, at two prize levels, a GLS procedure, which takes into account this structure, is preferable. Such a procedure yields more efficient estimators and facilitates hypothesis testing.

Alternative regression models are suggested in the literature for the analysis of panel data. An appropriate model for the current case is applied by the 'pool' procedure in the Shazam package,

which is based on Kmenta's model (Kmenta, 1986). This model assumes heteroscedasticity among the error terms of various farmers and a positive correlation between the error terms of the two choices of each farmer. Thus the model can be described as:

$$PWD = X\beta + \epsilon_{ij}, \quad i = 1, \dots, 180, j = 1, 2$$

where X is a 360×15 matrix of SECs, β is a 15×1 vector of coefficients, and ϵ_{ij} is a 360×1 random vector with the following properties:

$$E(\epsilon_{ij}) = 0 \quad \forall i, j,$$

$$E(\epsilon_{ij}\epsilon_{mn}) = \begin{cases} \sigma_i, & \text{if } i = m \text{ and } j = n; \\ \rho\sigma_i, & \text{if } i = m \text{ and } j \neq n; \\ 0; & \text{if } i \neq m \end{cases}$$

The somewhat restrictive assumption of this model is that the correlation coefficients between the error terms of the two observations for each farmer are identical among the farmers. Note, however, that this does not imply identical covariances⁷.

The regression results are reported in Table 2 and show that a small, but significant share ($R^2 = 0.38$) of the variations in the farmers' choices of PWD are explained by the socioeconomic variables⁸.

An examination of the regression coefficients

Table 2

The effects of socioeconomic variables on risk attitudes

Variable	Regression	T ratio
Farm size	$-0.13e^{-1}$	-0.6
Water quota	-0.25^*	-4.0
Milk quota	$-0.16e^{-4}^*$	-1.9
Eggs quota	$-0.29e^{-1}^*$	-3.7
Green house ^a	-0.58	-0.26
Plantation ^a	-10.07^*	-4.1
Cattle ^a	-10.3^*	-2.8
Family size	1.27^*	2.5
Schooling years	$-0.54e^{-1}$	-0.16
Operator's age	-0.11	-1.26
Part time farming ^a	9.47^*	2.3
Full time farming ^a	1.66	.4
Agency settlement ^a	0.65	0.0
New settlement ^a	3.47^{**}	1.5
Lottery prize	0.79^*	7.2
Constant	75.57^*	8.4
ρ	0.18	

* Significant at the 5% level. ** Significant at the 20% level. Buse $R^2 = 0.38$ with 360 observations and 342 degrees of freedom.

Dependent variable, PWD.

^a Dummy variable. Base classes, 'Field crop' and 'Few hours farming'.

and their statistical significance (Table 2) leads to several interesting findings. First, the coefficients of the farm's size and those of the water, milk, and egg quotas are all negative. The latter three are significant at the 5% level. These variables are measures of the farm scale (or potential scale) and are not related to any particular crop. Thus, they can be thought of as fairly reasonable proxies of the farmer's level of wealth. The fact that their coefficients are negative and statistically significant supports Arrow's hypothesis of DARA (Arrow, 1970). The same qualitative results were found by Binswanger (1980), for landlords, and Moscardi and de Janvry (1977). It is worth mentioning that the estimated coefficient of the lottery-prize level is positive and significant. However, as explained above, this need not be interpreted as additional support or lack of support for the DARA hypothesis.

Second, it is worth noting that the farm-type coefficients are all negative and the plantation and cattle coefficients are significant at the 5% level. These dummy variables (coupled with the milk and egg quotas) allow comparison with the

⁷ It was suggested by a reviewer that since we have only two observations for each farmer, the variances' estimates may be not reliable. The reviewer proposed an alternative procedure based on White's consistent estimator of the covariance matrix. We applied the reviewer's suggestion and found a significant smaller value of R^2 with White's procedure, than with Kmenta's procedure (0.18 as compared to 0.38). It is not clear whether White's procedure takes care of the autocorrelation problem and it requires to add 179 binary variables to the regression model. We therefore, choose to report the results obtained via Kmenta's procedure.

⁸ This is consistent with Moscardi and de Janvry (1977), who found $R^2 = 0.37$, and Dillon and Scandizzo (1978), who reported $R^2 = 0.72$ in a regression of their measure of risk aversion as the dependent variable and socioeconomic variables as the explanatory ones. However, they differ from the results of Binswanger (1980) who found that socioeconomic variables in his sample of rather resource-poor farmers did not seem to explain much of the variations in risk attitudes ($R^2 = 0.088$).

‘left-out’ category of extensive field crop farms. Thus we find, as expected, that the operators who are less risk averse choose a riskier career-practicing intensive agriculture which requires a sizeable capital investment.

Third, several characteristics of the farmer and his family affect risk-taking behavior. As expected, increased family size of the operator leads to more cautious and conservative behavior. This finding agrees with that of Moscardi and de Janvry (1977), but contradicts the result of Dillon and Scandizzo (1978) that farmers with larger families are less risk averse. Note that a priori, family size can have two opposing effects on farmers’ risk attitudes. On the one hand, a larger family represents an increased labor force for the household and can therefore be expected to have a negative effect on risk aversion. On the other hand, a larger family means more mouths to feed, which might well increase aversion to risk. In Israeli farm households most of the children are students, devoting most of their time to educational or social activities. Moreover, at the peak of their physical ability (18 years of age) they are drafted into the military. Therefore, the contribution of children to the household’s earning assets is negligible, which explains our finding that family size increases risk aversion. This is not necessarily the case with peasant households, as found by Dillon and Scandizzo (1978).

Contrary to the effect of family size, more experienced (measured by age) and educated (measured by years of schooling) farmers show more readiness to try new treatments on their farms. Since education and experience are forms of human capital, these results are expected. It should be noted, however, that while in the expected directions, the estimates of these effects were found insignificant. The result concerning experience is supported by the positive and significant (at the 20% level) coefficient of the new settlement. Farmers from new settlements are the least experienced farmers in the sample and show less readiness to take risks.

Finally, it is interesting to examine the effect of the share of time devoted to on-farm labor. We find that this variable (represented by the full- and part-time farming dummies) increases

risk aversion. As compared to farmers who spend only a few hours a week doing on-farm work, these farmers are more risk averse. Since farmers who devote more time to on-farm labor have less time for off-farm labor, this effect is expected. Off-farm jobs usually represent a safer income source and, in any case, represent opportunities for diversification. Therefore, farmers who are less dependent on the farm’s profit for living are willing to take higher risks on that farm. This final result is consistent with the findings of both Binswanger (1980) and Moscardi and de Janvry (1977).

7. Concluding remarks

The theoretical framework developed in this paper allows a systematic analysis of the role of SECs in production decisions. A major advantage of the proposed method is its usefulness for the analysis of farm households with complex objective functions. It allows a deduction of the effects of socioeconomic variables on production decisions from an observation on their effects on any of a variety of measures of individual risk preferences. It also enables the categorization of the decision variable to marginally risk-increasing/decreasing. Information regarding the latter is widely available in the literature. It should be emphasized that our analysis does not establish explicit casual effect between SECs and risk preferences. However, it does show that they might be systematically associated with each other. The empirical results suggest that the degree of association might be quite significant.

The paper yields several hypotheses. The analysis in the paper predicts that the more wealthy, experienced and smaller family farmers are less susceptible to risk. Therefore, in the presence of output price risk for example, *ceteris paribus*, such farmers are expected to choose a larger scale of production. Testing these and other resultant hypotheses presents a promising avenue for further research.

Finally, the proposed theoretical framework combined with a complementary empirical study (such as the one conducted here) may provide a

useful tool for the analysis of policies like stabilization programs, direct and indirect regulations and agricultural insurance and support programs. For example, suppose that, as found here, farm size as measured by milk and eggs quotas decreases risk aversion. Then one would expect the supply response to a price stabilization policy to be smaller in sizeable farms.

Appendix A

Lemma I. Suppose that π_ϵ and $\pi_{\alpha\epsilon}$ exist and are uniformly signed. Then, a profit factor, α , is marginally risk-increasing/decreasing if and only if $\text{sgn}(\pi_{\alpha\epsilon}) = (\neq) \text{sgn}(\pi_\epsilon)$.

Proof. The first-order condition for the choice of α is given by

$$EV_\alpha = EV_u U_\pi \pi_\alpha = 0$$

or equivalently

$$E\pi_\alpha = \frac{-\text{COV}(V_u U_\pi; \pi_\alpha)}{EV_u U_\pi}$$

Since V_u and U_π are positive, the profit factor α is marginally risk-increasing (decreasing) if and only if $\text{COV}(V_u U_\pi; \pi_\alpha) \leq (\geq) 0$

But the covariance is positive (negative) if and only if

$$\text{sgn}\left(\frac{\partial(V_u U_\pi)}{\partial\epsilon}\right) = (\neq) \text{sgn}(\pi_{\alpha\epsilon})$$

It can be verified that the concavities of V and U (i.e. $V_{UU} \leq 0$ and $U_{\pi\pi} \leq 0$) imply that $\text{sgn}\left(\frac{\partial(V_u U_\pi)}{\partial\epsilon}\right) = -\text{sgn}(\pi_\epsilon)$. It follows that the covariance is negative (positive), i.e. α is a marginally risk-increasing (decreasing) profit factor, if and only if $\text{sgn}(\pi_{\alpha\epsilon}) = (\neq) \text{sgn}(\pi_\epsilon)$.

Proof of Theorem I. The first-order condition for maximization of $EV(\cdot)$ is given by

$$EV_\alpha = EV_u U_\pi \pi_\alpha = 0$$

Totally differentiating with respect to α and ρ_i yields

$$\frac{d\alpha}{d\rho_i} = -\frac{EV_{u\rho} U_\pi \pi_\alpha}{\Delta}$$

where Δ is the second-order condition, which is negative by assumption. Thus, it remains to show that $EV_{u\rho} U_\pi \pi_\alpha$ is negative (positive) for a marginally risk-increasing (decreasing) profit factor, α .

We now multiply and divide the term $EV_{u\rho} U_\pi \pi_\alpha$ by V_u to get

$$EV_{u\rho} U_\pi \pi_\alpha = E\left[\frac{V_{u\rho}}{V_u} V_u U_\pi \pi_\alpha\right]$$

But since $E[V_u U_\pi \pi_\alpha] = 0$ (by the first-order condition), we can subtract from the above expression $E\left[\frac{V_{u\rho}}{V_u} \Big|_{\epsilon^*} V_u U_\pi \pi_\alpha\right]$, where ϵ^* is the level of ϵ for which $\pi_\alpha = 0$, at the optimal level of α . Following the subtraction we get

$$EV_{u\rho} U_\pi \pi_\alpha = E\left[\left(\frac{V_{u\rho}}{V_u} - \frac{V_{u\rho}}{V_u} \Big|_{\epsilon^*}\right) V_u U_\pi \pi_\alpha\right]$$

We now examine the signs of the two terms on the RHS, beginning with $\left(\frac{V_{u\rho}}{V_u} - \frac{V_{u\rho}}{V_u} \Big|_{\epsilon^*}\right)$.

First, note that

$$\begin{aligned} \frac{\partial \frac{V_{u\rho}}{V_u}}{\partial\epsilon} / \frac{\partial\epsilon}{\partial\epsilon} &= \frac{\partial \frac{V_{u\epsilon}}{V_u}}{\partial\epsilon} / \frac{\partial\epsilon}{\partial\epsilon} = \\ \frac{1}{U_\epsilon} \frac{\partial \frac{V_{\epsilon\epsilon}}{V_\epsilon}}{\partial\epsilon} / \frac{\partial\epsilon}{\partial\epsilon} &= \frac{-1}{U_\epsilon} \frac{\partial A}{\partial\epsilon} \end{aligned}$$

where A is the measure of absolute risk aversion. But, under the assumption that $U_\epsilon = U_\pi \pi_\epsilon > 0$, and that ρ is a risk aversion-increasing SEC (i.e. $\partial A / \partial\rho \geq 0$), it follows that $V_{u\rho} / V_u$ is decreasing with ϵ and, hence, the parentheses are nonnegative for $\epsilon \leq \epsilon^*$ and negative for $\epsilon > \epsilon^*$.

We turn now to the second term, $V_u U_\pi \pi_\alpha$. Since $V_u U_\pi > 0$, the sign of this term is determined by the sign of π_α , which, by the definition of ϵ^* , vanishes at ϵ^* . Thus, if the profit factor is risk-increasing ($\pi_{\alpha\epsilon} > 0$, $\pi_\epsilon > 0$), $V_u U_\pi \pi_\alpha$ is negative for $\epsilon \leq \epsilon^*$ and positive for $\epsilon > \epsilon^*$. Therefore, an increase in risk aversion-increasing SEC will always lead to a decrease in the utilization of a marginally risk-increasing profit factor. Similarly it can be shown that an increase in a risk aversion-increasing SEC will always lead to an increase in the utilization of a marginally risk-increasing profit factor.

Proof of Theorem II. We will prove the case for a risk-increasing SEC. The proofs for the other two cases follow immediately. Recall that the PWD, p , is defined by

$$pU(\epsilon^1, \rho) + (1-p)U(\epsilon^2, \rho) = U(\hat{\epsilon}, \rho)$$

Totally differentiating the above equation yields

$$\frac{dp}{d\rho} = \frac{pU_\rho(\epsilon^1, \rho) - (1-p)U_\rho(\epsilon^2, \rho) + U_\rho(\hat{\epsilon}, \rho)}{U(\epsilon^1, \rho) - U(\epsilon^2, \rho)}$$

The denominator is clearly positive. Hence, the sign of $\partial p / \partial \rho$ is identical to that of the numerator.

The cumulative distribution functions (CDFs) of ϵ and $\hat{\epsilon}$ are denoted $F(\epsilon)$ and $F(\hat{\epsilon})$, respectively and are given by

$$F(\epsilon) = \begin{cases} 0, & \text{for } \epsilon < \epsilon^2 \\ 1-p, & \text{for } \epsilon^2 \leq \epsilon < \epsilon^1 \\ 1, & \text{for } \epsilon \geq \epsilon^1 \end{cases}$$

and

$$\hat{F}(\epsilon) = \begin{cases} 0, & \text{for } \epsilon < \hat{\epsilon} \\ 1, & \text{for } \epsilon \geq \hat{\epsilon} \end{cases}$$

It should be noted that while ϵ is a discrete random variable and $\hat{\epsilon}$ is a degenerate one, their respective CDFs are everywhere continuous to the right (Hogg and Craig, 1978).

It is convenient to use these CDFs and to rewrite the numerator as

$$\begin{aligned} & -pU_\rho(\epsilon^1, \rho) - (1-p)U_\rho(\epsilon^2, \rho) + U_\rho(\hat{\epsilon}, \rho) \\ & = -\int_{-\infty}^{\infty} U_\rho(\epsilon, \rho) dF(\epsilon) + \int_{-\infty}^{\infty} U_\rho(\epsilon, \rho) d\hat{F}(\epsilon) \end{aligned}$$

where $dF(\epsilon) = (1-p)\delta(\epsilon - \epsilon_2) + p\delta(\epsilon - \epsilon_1)$, $d\hat{F}(\epsilon) = \delta(\epsilon - \hat{\epsilon})$ and δ is Dirac's function satisfying

$$\delta(x) = \begin{cases} x, & \text{for } x \neq 0; \\ \infty, & \text{for } x = 0, \end{cases}$$

and

$$\int_{-\infty}^{\infty} \delta(x) dx = 1$$

Integrating the above equation by parts (and noting that $F(\epsilon) = \hat{F}(\epsilon) = 0(1)$ for $\epsilon < \epsilon^2(\epsilon \geq \epsilon^1)$) yields:

$$\begin{aligned} & \int_{-\infty}^{\infty} U_\rho(\epsilon, \rho) [d\hat{F}(\epsilon) - dF(\epsilon)] \\ & = -\int_{-\infty}^{\infty} U_{\rho\epsilon} [\hat{F}(\epsilon) - F(\epsilon)] d\epsilon \end{aligned}$$

Multiplying and dividing through by U_ϵ , integrating by parts once more and rearranging terms yields

$$\begin{aligned} & -\frac{U_{\rho\epsilon}}{U_\epsilon} \int_{-\infty}^{\epsilon} U_\epsilon(\tau, \rho) [\hat{F}(\tau) - F(\tau)] d\tau \Big|_{-\infty}^{\infty} \frac{\partial^2 \ln U_\epsilon}{\partial \epsilon \partial \rho} \\ & \times \left[\int_{-\infty}^{\epsilon} U_\epsilon(\tau, \rho) [\hat{F}(\tau) - F(\tau)] d\tau \right] d\epsilon \end{aligned}$$

Evaluation of the first term at $-\infty$ and ∞ yields

$$\begin{aligned} & -\frac{U_{\rho\epsilon}}{U_\epsilon} \int_{-\infty}^{\epsilon} U_\epsilon(\tau, \rho) [\hat{F}(\tau) - F(\tau)] d\tau \Big|_{-\infty}^{\infty} \\ & = 0 + \int_{-\infty}^{\infty} U_\epsilon [\hat{F}(\epsilon) - F(\epsilon)] d\epsilon \end{aligned}$$

Integrating this latter term by parts yields

$$\begin{aligned} & \int_{-\infty}^{\infty} U_\epsilon [\hat{F}(\epsilon) - F(\epsilon)] d\epsilon \\ & = U [\hat{F}(\epsilon) - F(\epsilon)] \Big|_{-\infty}^{\infty} \\ & \quad - \int_{-\infty}^{\infty} U [d\hat{F}(\epsilon) - dF(\epsilon)] \end{aligned}$$

But, since $F(\infty) = \hat{F}(\infty) = 1$ and $F(-\infty) = \hat{F}(-\infty) = 0$, the first term on the RHS of the above equation vanishes. Moreover, since by the definition of the PWD, the expected utility is the same under both $F(\epsilon)$ and $\hat{F}(\epsilon)$, the second term vanishes as well.

With all this information in mind, it follows that the numerator of the expression $\frac{dp}{d\rho}$, which should be signed can be written as

$$\begin{aligned} & pU_\rho(\epsilon^1, \rho) - (1-p)U_\rho(\epsilon^2, \rho) + U_\rho(\hat{\epsilon}, \rho) \\ & = \int_{-\infty}^{\infty} \frac{\partial^2 \ln U_\epsilon}{\partial \epsilon \partial \rho} \\ & \quad \times \left[\int_{-\infty}^{\epsilon} U_\epsilon(\tau, \rho) [\hat{F}(\tau) - F(\tau)] d\tau \right] d\epsilon \end{aligned}$$

Note that the first term $\frac{\partial^2 \ln U_\epsilon}{\partial \epsilon \partial \rho} = -\frac{\partial A}{\partial \rho}$, where A is the Arrow–Pratt measure of absolute risk aversion, and thus by the theorem assumption, it is negative. Note, in addition, that the difference $F(\epsilon) - F(\hat{\epsilon})$ equals: (i) $1 - p$ for $\epsilon^2 \leq \epsilon < \hat{\epsilon}$; (ii) p for $\epsilon \leq \epsilon < \epsilon^1$; and (iii) zero for any ϵ outside the range $[\epsilon^2, \epsilon^1]$. Therefore, the second term in the above product $\left[\int_{-\infty}^{\epsilon} U_\epsilon(\tau, \rho) [F(\tau)] d\tau \right]$ is equal to

$$(i) \quad -(1-p) \int_{\epsilon^2}^{\epsilon^*} U_\epsilon(\tau) d\tau$$

$$= -(1-p) [U(\epsilon^*, \rho) - U(\epsilon^2, \rho)] < 0 \quad \text{for} \\ \epsilon^2 \leq \epsilon^* < \hat{\epsilon},$$

$$(ii) \quad -(1-p) \int_{\epsilon^2}^{\hat{\epsilon}} U_\epsilon(\tau) d\tau + p \int_{\hat{\epsilon}}^{\epsilon^{**}} U_\epsilon(\tau) d\tau \\ = (1-p) U(\epsilon^2, \rho) + p U(\epsilon^{**}, \rho) - U(\hat{\epsilon}, \rho), \\ \text{for } \hat{\epsilon} \leq \epsilon^{**} < \epsilon^1$$

and

$$(iii) \quad 0, \quad \text{for } \epsilon^1 \leq \epsilon < \epsilon^2$$

Since U increases with ϵ and $\epsilon^{**} < \epsilon^1$, it is easily verified via inspection of the definition of PWD that expression (ii) is negative. Thus both terms in the product

$$\int_{-\infty}^{\infty} \frac{\partial^2 \ln U_\epsilon}{\partial \epsilon \partial \rho} \left[\int_{-\infty}^{\epsilon} U_\epsilon(\tau, \rho) [\hat{F}(\tau) - F(\tau)] d\tau \right] d\epsilon$$

are negative and the theorem follows.

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