Insurance Market Efficiency and Crop Choices in Pakistan

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Abstract

This paper tests the efficiency of insurance markets in the Pakistan Punjab by examining how crop choices are affected by the presence of price and yield risk. We estimate reduced-form and structural models of crop choices. Although we cannot reject the hypothesis that village members efficiently share risk among themselves, production choices are shown to depend on risk. Existing risk sharing and self-insurance mechanisms thus imperfectly protect Punjab farmers against village-level shocks. Results also indicate that households respond to consumption price risk, thereby suggesting that empirical and theoretical work on risk should avoid putting an exclusive emphasis on yield and output price risk.

Keywords: risk; agricultural development; household models; insurance markets; structural estimation.

JEL codes: O12; Q12.

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This paper tests the efficiency of insurance markets in the Pakistan Punjab, an area characterized by well developed markets for agricultural inputs, outputs and factors of production. Other tests of full insurance have examined whether households manage to smooth consumption *ex post* (e.g., Mace (1991), Cochrane (1991), Townsend (1994), Morduch (1991), and Jacoby and Skoufias (1995)). We investigate whether *ex ante* production choices depend on risk and, if yes, how. Our test is based on the recognition that, in the presence of perfect insurance markets, production choices should not depend on the consumption and risk preferences of the producer (e.g., Sandmo (1971), Fafchamps (1992a)). Although more demanding, our approach presents the advantage of throwing light on the form and magnitude of production inefficiencies induced by insurance failure.

Risk plays an important role in human livelihood, particularly for Third World farmers exposed to the vagaries of weather and price shocks. It has long been argued that poor farmers in developing countries attempt to minimize their exposure to risk by growing their own food (e.g., Roumasset (1976), Fafchamps (1992a)), avoiding new technologies (e.g., Feder (1980), Feder, Just and Zilberman (1985), Antle and Crissman (1990)), and diversifying their activities (e.g., Robinson and Brake (1979), Walker and Ryan (1990)). Risk avoidance inhibits gains from specialization and prevents Third World agriculture from achieving its full potential.

These concerns have spurred a large body of research on *ex post* risk management practices in developing countries. The literature has brought to light a variety of mechanisms that households use to absorb shocks. Third World households have been shown to accumulate grain, livestock, and financial assets as a form of precautionary saving (e.g., Paxson (1992), Deaton (1990, 1992b), Park (1995), Lim and Townsend (1994), Fafchamps, Udry and Czukas (1998), Rosenzweig and Wolpin (1993), Fafchamps and Pender (1997)).[1]

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[1] The empirical literature has followed renewed theoretical interest in the subject (e.g., Kimball (1990), Deaton (1991), Zeldes (1989b)) and has proceeded *pari pasu* with a similar empirical literature in the U.S. (e.g., Zeldes (1989a), Hubbard, Skinner and Zeldes (1995), Attanasio and Weber (1995)).

Gifts and mutual credit have also been identified as major conduits for the sharing of risk among members of the same community (e.g., Platteau and Abraham (1987), Rosenzweig (1988), Fafchamps (1992b), Foster and Rosenzweig (1995), Ligon, Thomas and Worrall (1996), Fafchamps and Lund (2001)), or with distant relatives (e.g., Lucas and Stark (1985), Rosenzweig and Stark (1989)).

Formal tests of the efficiency of these institutions indicate that, although much consumption smoothing occurs, household consumption remains dependent on income shocks and that full insurance is not achieved (e.g., Townsend (1994), Paxson (1992), Chaudhuri and Paxson (1994), Morduch (1991), Fafchamps and Lund (2001)).[2] Existing risk coping mechanisms thus appear unable to provide complete protection against exogenous shocks. Possible explanations for imperfect insurance are that credit constraints limit the effectiveness of precautionary saving (e.g., Deaton (1990, 1992a, 1992b), Morduch (1990)), commitment failure limits the enforcement of informal reciprocal arrangements (e.g., Fafchamps (1999), Ligon, Thomas and Worrall (1996)), and information does not flow perfectly among community members, thereby reducing the scope for formal or informal contingent contracts (e.g., Fafchamps (1992b), Ligon (1993)).

To the extent that full insurance is not achieved, exposure to risk is likely to affect the *ex ante* production choices (e.g., Sandmo (1971), Fafchamps (1992a), Chavas and Holt (1996)). There has been ample empirical work on risk avoidance (e.g., Roumasset (1976), Antle and Hatchett (1986), Antle (1987, 1989), Antle and Crissman (1990)). With the important exceptions of Rosenzweig and Binswanger (1993), Rosenzweig and Wolpin (1993), and Morduch (1990), empirical work on *ex ante* risk management that incorporates *ex post* consumption choices remains rare, however. In particular, it is unclear whether existing insurance institutions, in spite of their imperfections, are sufficient to ensure that the effect of risk on *ex ante* production choices is negligible and can be ignored. This paper aims at filling this gap by investigating a wider range of sources of variation than

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[2] This literature mirrors similar work in developed countries (e.g., Mace (1991), Cochrane (1991), Zeldes (1989a)).

previously studied and by tying our results together with explicit evidence on *ex post* consumption smoothing.

A structural model of household production and consumption choices is estimated using data from Pakistan. The Pakistan Punjab provides a good testing ground for completeness of insurance markets. Unlike the poor semi-arid areas on which much of the work on consumption smoothing and risk sharing is based, Punjab enjoys well developed factor and product markets. One would therefore expect that households either have accumulated sufficient precautionary savings or have sufficient access to insurance substitutes that they can ignore risk in their production choices. If, however, full insurance is not achieved in spite of the area's favorable situation, we can be doubtful that it can be achieved in less endowed areas of the Third World.

The model is used to ascertain what types of risk affect production behavior and, hence, are not insured. In addition to yield and output price risk, the model incorporates price risk for items that are either consumed by household members or used in household production activities. Test results show that risk matters, thereby providing evidence that complete insurance is not achieved. We cannot, however, reject the hypothesis that village members are able to efficiently share risk among themselves. The risk averse behavior of Punjab farmers thus appears due to insufficient channels for sharing collective price and yield risk with the rest of the world. Results also reveal that households respond to consumption price risk as well, thereby suggesting that empirical and theoretical work on risk should avoid putting an exclusive emphasis on yield and output price risk. These results are robust to alternative econometric methods and estimates of risk.

The paper is organized as follows. We give in Section 1 a brief overview of the conceptual framework. In Section 2, the data set is introduced and the estimating equations are derived. Actual estimation proceeds in two steps. We first derive in Section 3 a consistent estimate of the variance-covariance matrix of yields, input prices, and output prices. Estimation of the consumption and production choice equations is done in Section 4. We then formally test in Section 5 whether risk affects production decisions and examine whether households' desire to avoid consumption price risk has a significant effect on crop choices.

Section 1. The Conceptual Framework

The starting point of our analysis is an intertemporal allocation model that captures, in a stylized manner, the crop decisions made by small farmers in Pakistan and elsewhere. The model is used to illustrate the role that risk sharing and self-insurance play in production decisions, and to examine the interplay between price risk and consumption preferences.

An infinitely lived household is assumed to choose its consumption, production, and savings to maximize an expected discounted utility $u(c_t)$ with discount factor δ. Households can buy and sell $K+1$ securities S_t^a, with $a \in \{0, ..., K\}$. These securities represent, in a stylized manner, various savings instruments to which households have access, such as cash, bank deposits, financial assets, livestock, grain storage, and jewelry (e.g., Lim and Townsend (1994), Fafchamps, Udry and Czukas (1998)). Social obligations to share risk with others through transfers and quasi-credit can similarly be represented (e.g., Townsend (1995), Udry (1994), Fafchamps and Lund (2001)). Asset S_t^0 is thought of as a bank savings account and is assumed to have a constant return $\sigma_{t+1}^0 = r$ for all t. Other securities have stochastic returns σ_{t+1}^a, which may vary with household-specific shocks due to risk sharing arrangements.

Time is divided into discrete intervals during which decisions are made and exogenous price and output shocks are realized. The timing of shocks and decisions is as follows. At the beginning of period t the household observes its realized crop income z_t, the returns σ_t^a to its $K+1$ assets S_{t-1}^a, and consumption prices q_t.[3] Crop income and asset returns together determine the household's cash in hand x_t, to be spent on savings S_t^a and

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[3] In the theoretical argument presented in this section, consumption prices are assumed non-stochastic for simplicity. If they were stochastic, the argument can be generalized by assuming that insurance against consumption price fluctuation is available as well. This complication is omitted here for the sake of brevity. In the empirical sections, consumption prices are allowed to be stochastic and consumption price uncertainty plays an important role in planting decisions, as in Fafchamps (1992a).

consumption c_t. Households then allocate a fixed factor $\bar{L}$ to N competing production activities.[4] The share of the fixed factor devoted to activity j is written l_t^j. By definition, $\sum_{j=1}^N l_t^j = 1$. Agronomic constraints such as crop calendar complementarities and water availability requirements are represented by a technical relationship between the $N-1$ free production choices:

$$g(l_t^1, ..., l_t^{N-1}) = 0 \tag{1}$$

After having chosen how much to consume, what to save, and what to produce, households observe the realized prices of variable inputs w_{t+1}^n and choose the level of variable inputs v_{t+1}^n. A Leontieff production technology is assumed so that it is always optimal for households to set $v_{t+1}^n = \sum_{j=1}^N \kappa_{jn} l_t^j$ where κ_{jn} is a fixed input-output parameter.[5] To keep things simple, we assume that the distribution of input prices w_{t+1}^n is such that shutting down an activity is never optimal. Having chosen variable inputs, households then observe yields θ_{t+1}^j, output prices p_{t+1}^j, and returns to securities σ_{t+1}^a, which together determine their cash in hand at the beginning of the next period $t+1$. The cycle then begins anew.

We factor out consumption decisions as follows. Let y_t stand for total consumption expenditures in period t. Since utility is additively separable over time and all consumption prices are assumed non-stochastic, we can solve for c_t given total expenditures and prices q_t. This yields an indirect utility function $v(y_t, q_t)$ with the usual properties. Replacing $u(c_t)$ with $v(y_t, q_t)$, the household decision problem can be expressed as the following Belman equation:

$$\begin{aligned} V(x_t, q_t) &= \max_{l_t^j, S_t^a s.t.eq.(1)} v(x_t - \sum_{a=0}^K S_t^a, q_t) \qquad (2) \\ &+ \delta EV \left[\sum_{j=1}^N p_{t+1}^j l_t^j \bar{L} \theta_{t+1}^j - \sum_{j=1}^N \sum_{n=1}^M w_{t+1}^n \kappa_{jn} l_t^j \bar{L} + \sum_{a=0}^K (1+\sigma_{t+1}^a) S_t^a, q_{t+1} \right] \end{aligned}$$

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[4] The model can be expanded to include other production decisions, but we focus here on a model that closely resembles the one used in the estimation.

[5] Production flexibility considerations are thus ignored here (see Fafchamps (1993) for a discussion).

where M is the number of variable inputs. We are interested in assessing the conditions under which production choices l_t^j are independent from household preferences. Assuming that the value function is continuously differentiable,[6] we get the following first order conditions:

$$\delta E[V_x p_{t+1}^j \theta_{t+1}^j]\bar{L} - \delta \sum_{n=1}^{M} \kappa_{jn} E[V_x w_{t+1}^n]\bar{L} \quad - \quad \lambda_t \frac{\partial g}{\partial l_t^j} = 0 \qquad \forall j \tag{3}$$

$$-v_y + \delta E[V_x(1+\sigma_{t+1}^a)] = 0 \qquad \forall a \tag{4}$$

where V_x is the derivative of the value function with respect to cash in hand, v_y is the marginal utility of consumption, and λ_t is the Lagrange multiplier associated with equation (1). It is easy to see that if there exist securities a and a' (or linear combinations thereof) such that:

$$1+\sigma_{t+1}^a = \frac{p_{t+1}^j \theta_{t+1}^j}{\mu_t^j} \tag{5}$$

$$1+\sigma_{t+1}^{a'} = \frac{w_{t+1}^n}{\iota_t^n} \tag{6}$$

for all states of nature and all j, n, t, and $t+1$, then production choices are independent of consumption preferences, that is, l_t^j does not depend on δ, $v(.)$ and $V(.)$ for all $j \in N$ and all t.[7] For production choices not to be affected by risk attitudes, securities must therefore exist that span all the sources of output and price risk faced by households. Separability between production choices and household preferences follows from a simple

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[6] Precise conditions are given in Stokey and Lucas (1989), theorem 9.10. They hold essentially for concave problems with interior solutions. We assume these conditions are satisfied here.

[7] *Proof:* Plug equations (5) and (6) into first order condition (4). We get:

$$\delta E[V_x p_{t+1}^j \theta_{t+1}^j] = v_y \mu_t^j$$

$$\delta E[V_x w_{t+1}^n] = v_y \iota_t^n$$

Using the above results in first order condition (3) yields, after some simple manipulations:

$$\frac{\mu_t^j - \sum_n \kappa_{jn}\iota_t^n}{\partial g/\partial l_t^j} = \frac{\mu_t^h - \sum_n \kappa_{hn}\iota_t^n}{\partial g/\partial l_t^h} = \frac{\lambda_t}{v_y}$$

for all j and $h \in N$. Together with equation (1), the above system fully determines production choices.

arbitrage argument, μ_t^j and ι_t^n being thought of as securities' prices. These prices may vary over time without affecting separability. They do, however, influence production choices. It is not necessary that households fully insure for separability to hold. Under separability, households may choose to bear some risk — or even to insure other households — depending on their attitude toward risk.

Separability breaks down whenever a complete set of contingent securities does not exist. Production choices then must satisfy:

$$\begin{aligned} & \frac{1}{\partial g/\partial l_t^j} E[V_x(p_{t+1}^j \theta_{t+1}^j - \sum_{n=1}^{M} \kappa_{jn} w_{t+1}^n)] \\ = \; & \frac{1}{\partial g/\partial l_t^h} E[V_x(p_{t+1}^h \theta_{t+1}^h - \sum_{n=1}^{M} \kappa_{hn} w_{t+1}^n)] \qquad (7) \end{aligned}$$

for all production activities j and h. Production decisions depend on the covariance between net crop revenues and V_x, and thus on the concavity of the value function V (e.g., Sandmo (1971)). The concavity of V with respect to cash in hand depends not only on the curvature of indirect utility $v(.)$ with respect to consumption expenditures, but also on the presence of assets to self-insure or share risk with others. If, for instance, there are no market for securities and no assets, and households cannot borrow against future labor income, V boils down to v: the curvature of V simply reflects the risk preferences of the household (e.g., Zeldes (1989b), Deaton (1991)). At subsistence income levels, instantaneous utility v is likely to be steep and highly concave with respect to consumption because the household is concerned about starvation. Poor households without any option for insurance are thus expected to behave in a highly risk averse manner.[8]

Alternatively, suppose that there exists a fixed-return asset S_t^0, but no risk sharing and no borrowing against future income. The curvature of V then is a function of the household's accumulated wealth: the more precautionary saving the household has been able to accumulate, the more closely its value function approaches certainty equivalence

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[8] This assumes, of course, that utility is everywhere concave. If households are indifferent between various levels of starvation but fear falling below a minimum income level, risk loving behavior might actually increase the chances of survival (e.g., Roumasset (1976)).

(e.g., Zeldes (1989b), Deaton (1991)). For households without accumulated assets, V_x tracks v_y closely because variations in cash in hand are translated nearly one for one into variations of current consumption. In contrast, households with sufficient wealth have a less concave value function V because their wealth serves to buffer income shocks. By extension, the presence of insurance opportunities, even if imperfect, further dissociates V_x from v_y, thereby reducing the curvature of V even more. When complete insurance markets exist, V becomes linear and the household maximizes expected profit.

Having detailed the relationship between insurance, precautionary savings, and the curvature of V_x, let us now turn to the relationship between production decision and risk. As is well known, households with a less concave V_x behave in a less risk averse fashion (e.g., Dreze and Modigliani (1972), Diamond and Stiglitz (1974)). Define E_t^j as $E[V_x(p_{t+1}^j \theta_{t+1}^j - \sum_{n=1}^M \kappa_{jn} w_{t+1}^n)]$. Output prices and yields are, in general, positively correlated with total income and thus negatively correlated with V_x. Similarly, input prices are, in general, negatively correlated with total income and thus positively correlated with V_x. From equation (7) we see that, other things being equal, a high variance of output revenues and input prices for activity j relative to activity h lowers E_t^j relative to E_t^h, thereby inducing the household to reduce l_t^j. It follows that crops whose revenue and input prices are highly variable are less likely to be grown by risk averse households.

If, however, one output, say j, serves as input for another household activity, say h, output and input price risk partially cancel each other, inducing the household to engage in activities j and h more than it would on the basis of output and input price risk alone. Producing the input is a way for the household to self-insure against input price risk. Similarly, if the household consumes one of its output — or if the price of what it consumes is highly correlated with output revenue — producing it enables the household to self-insure against consumption price risk (e.g., Fafchamps (1992a), Finkelshtain and Chalfant (1991)). The absence of perfect insurance markets thus leads to predictions regarding the relationship between household wealth, production choices, and the risks the household face. Testing these predictions is the purpose of this paper.

Section 2. The Data and Empirical Model

The data used in this study comes from a series of surveys conducted by the Punjab Economic Research Institute (PERI) based in Lahore, Pakistan, between 1988/89 and 1990/91 (see Kurosaki (1998) for details). The purpose of the PERI study was to collect information on the production costs of major crops. Stratified sampling was used to select sample villages in thirteen study centers. A complete census of households was then carried out in each of the sample villages and used to draw a random sample of households from whom data on crop choices, output, wealth, and consumption were collected. To minimize any potential bias due to heterogeneity, we focus our analysis on a single agroclimatic zone — the rice-wheat zone — and limit our sample to five villages of the Sheikhupura district, in the Pakistan Punjab. This sample covers three years; there are 291 observations. Characteristics of sample households are summarized in Table 1.

In Sheikhupura as in the rest of Punjab, the cropping year is divided into two agricultural seasons called *Kharif* and *Rabi.* In each season surveyed farmers grow two major crops: Basmati and green fodder during Kharif; wheat and green fodder during Rabi. Basmati is a high quality rice sold for export; wheat is partly self-consumed, partly sold for local urban consumption. Most of the land not planted to one of the main crops is left fallow. Other crops are negligible in terms of cultivated acreage; they are ignored in the subsequent analysis. In addition to crops, households keep cattle and buffaloes for milk production, for which they require large quantities of green fodder.

Although the land of surveyed farmers is mostly irrigated and the number of their livestock animals is relatively high by Third World standards, their annual income and consumption remain fairly low (Table 1). Their agricultural equipment is moderately sophisticated: two thirds of households own a tubewell for irrigation purposes; one household out of six owns a tractor. Education levels are low, however, and the dependency ratio — that is, the share of kids in total household size — is high. Households spend close to half of their income on foods they produce, mostly milk, milk products, and wheat.

Active markets exist for all major outputs and inputs and surveyed households participate actively in all of them (Table 2). Proximity to local urban centers enables farmers to sell part of their milk production, thereby generating a steady stream of regular income (e.g., Zafar (1985), Lockwood (1982)). Surveyed farmers nevertheless face substantial price risk, particularly for green fodder whose bulky nature increases transport costs over long distances and isolates markets.[9] In contrast, variations in rice and wheat prices are dampened by government market intervention and by the spatial integration of these markets with the rest of Pakistan and the world (e.g., Mohammad (1985), Cornelisse and Naqvi (1987), Dorosh and Valdes (1990), Kurosaki (1996)).

The model presented in Section 1 is adapted to Sheikhupura realities as follows. Each household i in each year t is taken to allocate its Kharif land $L^k_{i,t}$ between grain g (Basmati), fodder f, and fallow, and its Rabi land $L^r_{i,t}$ similarly between grain g (wheat), fodder f, and fallow. The shares of Kharif land devoted to Basmati and fodder are denoted $l^{kg}_{i,t}$ and $l^{kf}_{i,t}$, respectively; the shares of Rabi land planted to wheat and fodder are written $l^{rg}_{i,t}$ and $l^{rf}_{i,t}$. Corresponding net revenues per acre are denoted $\pi^{sj}_{i,t} \equiv p^{sj}_t \theta^{sj}_{i,t} - \sum_{n=1}^{N} \kappa_{jn} w^n_t$ for $s = k, r$ and $j = g, f$. Agronomic constraints corresponding to equation (1) are imposed for each season separately and may vary across individuals; $g^{sj}_{i,t}$ denotes the derivative of the $g^s_{i,t}(l^{sg}_{i,t}, l^{sf}_{i,t})$ function with respect to its jth argument, with $s = k, r$. The fodder needs and milk output of household i in year t are assumed proportional to its number of milk animals $A_{i,t}$. Milk revenue is denoted $\pi^{sm}_{i,t}$ for $s = k, r$. Since land and livestock are predetermined long-term investments, they can be regarded as given as far as crop choices are concerned. Put differently, we examine crop choices conditional on land and

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[9] Within each surveyed village the differential between producers' selling and buying prices of green fodder is very small. This is because villages as a whole are net exporters of fodder to neighboring urban areas. Consequently, farmers who wish to buy green fodder can do so from other villagers. The local price of fodder, however, varies widely from day to day depending on availability.

livestock holdings.[10] Total income $z_{i,t}$ is:

$$z_{i,t} = \sum_{s=k,r} \sum_{j=g,f} \pi_{i,t}^{sj} l_{i,t}^{sj} L_{i,t}^{s} + \sum_{s=k,r} \pi_{i,t}^{sm} A_{i,t} + R_{i,t} \tag{8}$$

where $R_{i,t}$ represents non-farm income, taken as exogenous and non-stochastic. Allocative efficiency requires that:

$$FOC_{i,t}^{s} \equiv E\left[\frac{\partial V_{i,t}}{\partial x_t}(\pi_{i,t}^{sg} - \frac{g_{i,t}^{sg}}{g_{i,t}^{sf}}\pi_{i,t}^{sf})\right] = 0 \qquad \text{for} \quad s = k, r. \tag{9}$$

Equation (9) defines an implicit relationship between crop choices $l_{i,t}^{sj}$, for $s = k, r$ and $j = g, f$, and determinants of crop profitability, agronomic substitutability, risk attitudes, and consumption preferences. As shown in Section 1, crop choices should not depend on the risk attitudes and consumption preferences of the household if complete markets are present. A simple test of market completeness can thus be constructed by regressing $l_{i,t}^{sj}$, for $s = k, r$ and $j = g, f$, on variables capturing crop profitability and agronomic factors, and on determinants of risk attitudes and consumption preferences such as wealth and household composition. If markets are complete, the latter factors should be jointly non-significant (e.g, Benjamin (1992), Pitt and Rosenzweig (1986), Gavian and Fafchamps (1996), Udry (1996)).

The validity of such a test rests critically on the existence of other markets. Table 2 indicates that markets for agricultural outputs exist in the study area; but what of factor markets? Studies suggest that rural factor markets are quite active in Pakistani countryside (e.g., Strosser and Meinzen-Dick (1993), Alderman et al., (1996), Adams (1997), Fafchamps and Quisumbing (1999)). Given that the area studied here is more commercialized than most parts of Pakistan, chances are that factor markets are even more active than elsewhere in the country. To investigate this issue further, we examine whether reduced-form results are indicative of missing factor markets. If, for instance, labor markets were incomplete, one would expect households with a higher person per

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[10] Given the lumpiness of livestock and the short duration of the Kharif and Rabi growing seasons, this is a reasonable assumption. Treating these semi-fixed variables as endogenous to the input decision process, as in Antle and Hatchett (1986), is left for future research.

land ratio AL to specialize in labor intensive crops such as Basmati rice and Rabi fodder. Similarly, if finance for agricultural production was limited, one would expect households with extra cash to put more emphasis on input intensive crops, namely, Basmati.

Regression results are presented in Table 3, together with a Wald test of the exclusion restrictions. Cultivated acreage L is included to control for scale economies.[11] Other sources of wealth, proxied by non-farm income R and total number of livestock heads $\bar{A}$, are regarded as pre-determined and used to capture ability to bear risk; the dependency ratio DEP, i.e., the share of children below age ten in the total household size, is used as consumption preference shifter. The livestock variable also controls for households' possible desire to be self-sufficient in green fodder, a desire that would be absent if perfect insurance against fodder price risk is available and the market-to-farm transaction costs are not prohibitively high for green fodder.[12]

The null hypothesis of complete markets is rejected in three of the four crop choices (Table 3). Results further suggest that concerns about fodder availability play an important role in farmers' planting decisions: households with more livestock devote a larger share of their cultivated acreage to fodder than grain. The effect is particularly strong in the dry Rabi season when alternative sources of animal feed, such as common pasture, are limited. Reduced form regressions cannot, however, distinguish between the risk and transactions costs motives: self-sufficiency could be a result of uninsurable price risk or the presence of high transactions costs that farmers must incur to buy or sell fodder. The latter possibility is indeed suggested by the fact that some households neither buy nor

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[11] Given the high multicollinearity between owned land and cultivated acreage, we did not include the two variables together. Cultivated acreage should thus be interpreted as controling for both scale economies and possible wealth effects due to land ownership. Similar results are obtained if land owned is used instead of cultivated acreage.

[12] Although transactions costs and price risk may both induce farmers to be self-sufficient, the two explanations are nevetheless distinct: uninsurable price risk refers to price variations in the market itself, while the transactions costs hypothesis refers to unobservable variations in the shadow cost of fodder to individual farmers.

sell fodder (Table 2).

Upon closer inspection, however, the transaction costs interpretation loses much of its appeal. First of all, surveyed farmers customarily harvest fodder crops by hand in the fields and bring them home to feed their livestock, instead of letting livestock harvest the fodder themselves. Harvesting is done daily so that fodder is fresh and milk output maximized. But within-village transport costs are essentially the same whether the household sources fodder from the market or from its own fields. If transport costs are not so prohibitive as to prevent daily transportation from fields to farm, there is little reason to believe that they would prevent transportation from fields to market.[13] This is not to say that there are no transactions costs in fodder. Inter-village transport costs are high relative to the value of fodder. This probably explains why fodder markets are not fully integrated spatially and why fodder prices vary wildly. Surplus green fodder from rural areas is nevertheless transported to town and cities where it is used in the production of milk for urban consumption (e.g., Zafar (1985), Lockwood (1982)). Transport of fodder over long distances does, therefore, take place, at least on well travelled routes.[14] Third, when sample farms purchase green fodder, the transaction usually takes place in the village and the purchase price is equal to the neighboring urban price *less* transportation cost (i.e., the effective selling price in the village). Since sample households are geographically concentrated in the village residential area, this implies that the shadow cost of green fodder does not vary much among sample households as far as direct transportation costs are concerned. There are thus good reasons to suspect that it is concerns about the timely availability and price of green fodder that motivates farmers to be self-sufficient, rather

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[13] Although we do not have hard data on the distances from field to farm and from field to market in the survey area, we observed that fields are concentrated but often distant from homesteads where livestock are kept.

[14] Transport remains problematic over long distances, especially on seldom travelled routes such as village to village haulage. Over such routes, high transport costs of green fodder primarily result from the fact that the freshness of green fodder deteriorates with time spent on a tractor. This leads to a rapid drop in value since the milk yield of green fodder depends on its freshness.

than market-to-farm ransactions costs.

To test which of the two motives is most compatible with the data, we re-estimate the reduced from regressions, allowing the effects of livestock on crop choices to differ depending on the household's self-sufficiency status in green fodder. Key, Sadoulet, and de Janvry (2000) show that if a household is subject to high transaction costs, the marginal effect of a change in input requirement on crop choices depends on the household's market participation status: if the household is self-sufficient, crop choices adjust to exogenous changes; in contrast, if the household is a net buyer or seller, crop choices are unaffected because separability applies locally. The situation is different if market-to-farm transactions costs are small but households worry about fodder price variability: in this case, fodder production choices depend on livestock ownership irrespective of whether the household is self-sufficient or not.

These observations enable us to test the presence of (large) market-to-farm transaction costs in fodder. We create a dummy variable D that takes value one for households that never participated in green fodder market during the study period. The livestock variable in Table 3 is replaced by its cross terms with D and $(1 - D)$. If high transactions costs are present, the coefficients on these cross terms should be different; if transactions costs are low but farmers worry about fodder price variations, the two coefficients should be significant and equal.

The results of this test are presented in Table 4. Let b_0 denote the coefficient on livestock times D and b_1 the coefficient on livestock times $(1 - D)$. We see that both b_0 and b_1 for Kharif and Rabi fodder crops are positive and the difference between the two is not statistically significant. The hypothesis that $b_1 = 0$ is rejected in both cases. Therefore, the market-to-farm transaction cost hypothesis is not supported by the data. As before, however, livestock variables are significant in three of the four regressions. Results are thus consistent with the price risk hypothesis.

Turning to factor markets, estimated coefficients are not consistent with a missing labor market: the person to land ratio AL has a negative coefficient in first and fourth regressions (significantly so for Basmati), indicating that households with more manpower per land cultivate less Basmati and Rabi fodder, not more as would be the case if households could not hire themselves out. Similarly, we find that non-farm income has no effect on crop choices. This is contrary to what would have obtained if production finance constraints were binding and households with more cash were better able to purchase inputs for rice production.

From these results we conclude that the assumption of complete markets can be rejected in the Sheikhupura district.[15] Although it is unclear which market is missing, estimated coefficients are not consistent with missing labor markets or binding credit constraints. A strong relationship is observed between livestock ownership and fodder cultivation, in spite of evidence that an active fodder market exists in the study area. This relationship is equally strong irrespective of whether the household is self-sufficient in fodder. This is consistent with an uninsurable risk in the timely availibility and price of green fodder. By themselves, however, reduced-form results have little if anything to say on the effect of risk on production choices. This is why we turn to structural estimation. The contribution of this paper is not to test transactions costs versus incomplete insurance, but rather to characterize how incomplete insurance markets affect crop choices when market-to-farm transactions costs are of the same order of magnitude as field-to-farm transport costs and can thus reasonably be ignored.

To develop an estimable structural model, we begin by taking a first order approximation of $\frac{\partial V_{i,t}}{\partial x_t}$ as in Fafchamps (1992a). After some straightforward manipulations and

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[15] To control for the possible presence of household specific unobservables, e.g., determinants of income that are correlated with asset holdings, we rerun the above regressions with fixed effects and random effects, using the subset of households surveyed in at least two years (247 observations) or using only the 59 households for whom we have three years of data. These results (available on request from the authors) show that coefficients of significant variables are quite similar with or without fixed effects. It therefore does not appear that asset holdings (which change very little over time) are correlated with unobserved determinants of income. Consequently, we are probably safe in using asset holdings to indirectly identify the structural estimation below.

after taking expectations, equation (9) can be rewritten as:

$$FOC^s \approx \quad \bar{V}_x \Big[E[\pi^{sg}] - \frac{g^{sg}}{g^{sf}} E[\pi^{sf}] + \sum_{j=1}^{C} \frac{\bar{V}_{xq_j}}{\bar{V}_x} E[(q^j - E[q^j])(\pi^{sg} - \frac{g^{sg}}{g^{sf}} \pi^{sf})] + \frac{\bar{V}_{xx}}{\bar{V}_x} E[(z - E[z])(\pi^{sg} - \frac{g^{sg}}{g^{sf}} \pi^{sf})] \Big] = 0 \tag{10}$$

where i, t subscripts and superscripts have been dropped to improve readability. Variable q^j stands for the price of consumption good j. The set of consumption goods C comprises wheat w, milk products m, rice b, and other consumption goods o. Expressions of the form $\bar{V}_k$ denote the derivative of V with respect to k, evaluated at $E[z]$ and $E[q^j]$. To derive a consistent set of expressions for $\bar{V}_x$, $\bar{V}_{xx}$, and $\bar{V}_{xq_j}$, we assume that V takes the form:

$$V(z_{i,t}, q_t^1, ..., q_t^C) = \frac{1}{1 - \Psi_{i,t}} \left[\frac{z_{i,t} - \sum_{j=1}^{C} q_t^j \gamma^j}{\prod_{j=1}^{C} (q_t^j)^{\beta_{i,t}^j}} \right]^{1 - \Psi_{i,t}} \tag{11}$$

where $\sum_{j=1}^{C} \beta_{i,t}^j = 1$ for all i, t. The reader will have recognized the indirect utility function associated with the linear expenditure system:

$$q_t^j c_{i,t}^j = q_t^j \gamma^j + \beta_{i,t}^j (z_{i,t} - \sum_{h=1}^{C} q_t^h \gamma^h) \tag{12}$$

for $j \in C$. This parameterization is chosen because it is sparse in parameters and yet allows income elasticities to differ from zero.[16] To control for household size, $z_{i,t}$ and γ^j are defined on a per-capita basis. $\Psi_{i,t}$ represents relative risk aversion with respect to income *after* necessary consumption has been covered. The standard coefficient of relative risk aversion r is obtained after correcting for necessary consumption using the following transformation:

$$r_{i,t} = \Psi_{i,t} \frac{E[z_{i,t}]}{E[z_{i,t} - \sum_{i=1}^{C} q_t^j \gamma^j]} \tag{13}$$

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[16] Newbery and Stiglitz (1981), Finkelshtain and Chalfant (1991) and Fafchamps (1992a) have indeed shown that income elasticity plays an important role in how production decisions are made in the presence of consumption price risk.

Using equation (11), we can compute closed-form expressions for all the $\bar{V}_k$ terms in equation (10). The resulting, fully parameterized version of equation (10) is given in Appendix.

Next, we assume that the concavity of V depends on household's ability to bear risk and we parameterize $\Psi_{i,t}$ as follows:

$$\Psi_{i,t} = \Psi_0 + \sum_h \Psi_h W_{i,t}^h \tag{14}$$

where variables $W_{i,t}^h$ are total land and livestock owned, plus human capital measured in years of formal education. We also assume that consumption preferences depend on household composition and posit that:

$$\beta_{i,t}^j = \beta_0^j + \beta_1^j DEP_{i,t} \tag{15}$$

for $j \in C$, where $DEP_{i,t}$ is the dependency ratio, i.e., the share of children below age ten in total household size. Finally, we specify an explicit functional form for $g(.)$. Agronomic constraints are assumed to take the following normalized quadratic form (e.g., Chavas and Holt (1996)):[17]

$$g_{i,t}^s(l_{i,t}^{sg}, l_{i,t}^{sf}) \equiv \alpha_0^s + l_{i,t}^{sf} + \alpha_1^s l_{i,t}^{sg} + \alpha_2^s (l_{i,t}^{sg})^2 + \alpha_3^s D_{i,t}^T = 0 \tag{16}$$

for $s = k, r$, where the α's are parameters to be estimated. To account for differences in households' capacity to control irrigation water, we have added a dummy $D_{i,t}^T$ for ownership of a tubewell. Indeed, although 100% of the study area has access to canal irrigation, the timing of water supply is somewhat unreliable due to technological and bureaucratic problems. Tubewell ownership guarantees more timely irrigation (e.g., Strosser and Meinzen-Dick (1993), Strosser and Kuper (1994)).

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[17] Agronomic constraints may include intertemporal requirement such as crop rotations and the need to fallow land. These requirements are ignored here because they would require incorporating lagged values of endogenous variables in the estimation, something that would singularly complicate structural estimation. Since intertemporal requirements are not the focus of the paper, we only consider agronomic constraints that are independent of past crop choices, such as crop complementarities within a single season and water availability requirements.

After replacing for equations (14) and (15) into (10) and (12), the basic model to be estimated boils down to three groups of equations: two first order conditions for optimal land allocation (10), one for Kharif, one for Rabi; four consumption demand equations for wheat, rice, milk products and other consumption goods (12); and two agronomic constraints (16), one for Kharif, one for Rabi. Because of adding-up restrictions, the last demand equation can be dropped.[18] Note that land allocation decisions are made *ex ante*, i.e., on the basis of anticipated prices (see equation (10)); consumption decisions are made *ex post*, i.e., on the basis of realized prices and income q_t^j and $z_{i,t}$.

We estimate the model by finding the parameter vector that minimizes individual deviations from these seven equations. To do so, we postulate the following:

$$\begin{aligned} FOC_{i,t}^{k}(l_{i,t}^{kg}, l_{i,t}^{kf}, l_{i,t}^{rg}, l_{i,t}^{rf}) &= \varepsilon_{i,t}^{1} && (17) \\ FOC_{i,t}^{r}(l_{i,t}^{kg}, l_{i,t}^{kf}, l_{i,t}^{rg}, l_{i,t}^{rf}) &= \varepsilon_{i,t}^{2} && (18) \\ g_{i,t}^{k}(l_{i,t}^{kg}, l_{i,t}^{kf}) &= \varepsilon_{i,t}^{3} && (19) \\ g_{i,t}^{r}(l_{i,t}^{rg}, l_{i,t}^{rf}) &= \varepsilon_{i,t}^{4} && (20) \\ -q_t^w c_{i,t}^w + q_t^w \gamma^w + \beta_{i,t}^w (z_{i,t}(l_{i,t}) - \sum_d q_t^d \gamma^d) &= \varepsilon_{i,t}^{5} && (21) \\ -q_t^m c_{i,t}^m + q_t^m \gamma^m + \beta_{i,t}^m (z_{i,t}(l_{i,t}) - \sum_d q_t^d \gamma^d) &= \varepsilon_{i,t}^{6} && (22) \\ -q_t^b c_{i,t}^b + q_t^b \gamma^b + \beta_{i,t}^b (z_{i,t}(l_{i,t}) - \sum_d q_t^d \gamma^d) &= \varepsilon_{i,t}^{7} && (23) \end{aligned}$$

where superscripts w, m and b stand for wheat, milk products, and rice, respectively. The last three equations are nothing but a rearranged LES demand system. Per capital consumption expenditure $z_{i,t}(l_{i,t})$ is given by:

$$z_{i,t} = \sum_{s=k,r} \sum_{j=g,f} E[\pi_{i,t}^{sj}] l_{i,t}^{sj} L_{i,t} + \sum_{s=k,r} E[\pi_{i,t}^{sm}] A_{i,t} + R_{i,t} + \hat{\varepsilon}_z$$

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[18] To test whether the adding-up restriction can be safely imposed, we separately estimated the demand system with and without it. The right hand side variable $z_{i,t}$, which is the ex post level of per capital expenditure, is instrumentalized using all the exogenous variables included in the structural model below. Using a likelihood ratio test, we cannot reject the adding-up restriction at the 5% level. Coefficient estimates with and without the restriction imposed are very similar.

where $\hat{\varepsilon}_z$ is the realized level of income shock, which is exogenous to *ex post* consumption decisions. Expected values of profits and realized values of $\hat{\varepsilon}_z$ are estimated in Section 3. Let $\varepsilon_{i,t} \equiv \{\varepsilon^1_{i,t}, ..., \varepsilon^7_{i,t}\}$. The disturbance vector $\varepsilon_{i,t}$ is assumed jointly normal with variance-covariance matrix Σ. The joint likelihood function of the system of seven equations is:

$$\begin{aligned} & \ln L(\alpha, \beta, \gamma, \Psi, \Sigma | l, c, \text{exogenous variables}) \\ = \quad & -\frac{21N}{2} \ln 2\pi - \frac{3N}{2} \ln |\Sigma| + \sum_{i=1}^{N} \sum_{t=1}^{3} \left[\ln |J_{i,t}| - \frac{1}{2}\, \varepsilon_{i,t}{}' \Sigma^{-1} \varepsilon_{i,t} \right] \end{aligned} \tag{24}$$

where N is the number of sample households. Since equations (17) to (20) do not yield an explicit solution for $l^{kj}_{i,t}$ and $l^{rj}_{i,t}$ with $j = g, f$, the determinant of the Jacobian transform matrix $J_{i,t}$ must be included (e.g., Mood, Graybill and Boes (1974), Fafchamps (1993)). The Jacobian matrix is block triangular between the sets of equations for land allocation and agronomic constraints, and the set of consumption demand equations. Parameters can be estimated by maximizing likelihood function (24), thereby yielding Full Information Maximum Likelihood estimates of the structural parameters of the model.

Before we can proceed with the estimation proper in Section 4, we must first construct a consistent estimate of the expectation vector and variance-covariance matrix of prices, revenues, and incomes. This is achieved in Section 3.

Section 3. Prices, Revenues and Incomes

Production choices depend on the joint distribution of prices, revenues, and incomes. In order to estimate the structural model, we need, therefore, to construct consistent estimates of their distribution. We assume that households form rational expectations about yields and prices, and we proceed as follows. We first estimate price and revenue risk common to all households. We then estimate idiosyncratic yield risk and combine it with estimates of collective risk to compute year-specific expectation vectors and variance-covariance matrices for prices and revenues. The details of this estimation process, which is briefly summarized in the following pages, are given in Kurosaki (1997).

Let p_t^j be the producer price of commodity j in year t and let Π_t^j be the regional average gross revenue from commodity j in year t. We assume that prices and gross revenues follow a stationary process with drift:[19]

$$\ln p_t^j = a_0^j + a_1^j t + u_t^j \tag{25}$$

$$\ln \Pi_t^j = b_0^j + b_1^j t + v_t^j \tag{26}$$

where disturbance terms follow a first-order auto-regressive process:

$$u_t^j = \rho_a^j u_{t-1}^j + \eta_t^j \tag{27}$$

$$v_t^j = \rho_b^j v_{t-1}^j + \nu_t^j \tag{28}$$

with η_t^j and ν_t^j white noise. We estimate equations (25) and (26) for the period 1970/71 to 1990/91 using data on annual market prices and district average yields that are published for the Sheikhupura district by the Government of Pakistan.[20] Variable Π_t^j is constructed as the product of annual price p_t^j and average per-acre yield θ_t^j. The Basmati support price, which is typically announced at the beginning of the cropping year and may influence farmers' price expectations, is added to the rice equations (25) and (26) to improve efficiency. The resulting estimates give one-year ahead predictions of prices $\hat{p}_t^j$ and average

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[19] Given the small number of observations and the limited purpose of the price and revenue equations, namely, to generate one-year ahead predictions, it is numerically irrelevant whether prices are assumed to follow a stationary process with drift, or a non-stationary process.

[20] For Basmati, wheat, and milk prices, we use monthly wholesale prices in the town of Sheikhupura (Government of Pakistan, Federal Bureau of Statistics, *Monthly Statistical Bulletin*, various issues). Fodder prices are taken from provincial government data available only for the nearby city of Faisalabad (Government of Punjab, Directorate of Agriculture, *Economic Situation and Price Trends of Agricultural Commodities in the Punjab*, various issues). Average prices at harvest time (May and June for Rabi, December and January for Kharif) are used. Average crop yields for the Sheikhupura district are taken from Government of Pakistan, Ministry of Food, Agriculture and Cooperatives, Economic and Policy Analysis Project's computerized databases, *AGDAT/PC*, 1992). Average fodder yield data is not available at the district level. Fodder yields for the Punjab province are used instead. Reported *jowar* yields are taken for Kharif fodder, *berseem* for Rabi fodder. See Kurosaki (1997) for details.

gross revenues $\hat{\Pi}_t^j$ in principle shared by all farmers. Residuals are used to construct an estimate of the variance-covariance matrix of prices and average gross revenues.

Next, we construct estimates of farm specific yield risk. Individual yields $\theta_{i,t}^j$ are assumed to be related to average yields θ_t^j as follows:[21]

$$\frac{\theta_{i,t}^j}{\theta_t^j} = \chi_0^j + \sum_{h=1}^{K} \chi_h^j X_{i,t}^h + \nu_{i,t}^j \tag{29}$$

where variables $X_{i,t}^h$ are household characteristics thought to influence individual yields, such as ownership of a tubewell D^W, ownership of a tractor D^T, number of family members per acre AL, education of the household head EDU, cultivated area L, livestock per land $\bar{A}/L$, dependency ratio DEP, and non-farm income per land R/L. Cultivated area controls for possible scale economies. Other variables capture the effect that missing or incomplete markets might have on yields.[22] Equation (29) is estimated from survey data, using only households for whom three years of consecutive data are available. The results, presented in Table 5, offer additional evidence regarding missing factor markets. If labor markets were incomplete, for instance, one would expect households with more land per person to achieve lower yields, especially in labor intensive crops such as Basmati. Estimated coefficients provide only limited support to this hypothesis: the AL coefficient is significant in only one regression; it is not significant for Basmati. We similarly find little evidence of binding credit constraints: the coefficient on non-farm income per acre

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[21] Other functional forms were tried but do not affect final results (see Kurosaki (1997)).

[22] Note that regressors do not include factors that may directly influence the magnitude of transactions costs, such as distance between fields and markets and the like. Consequently, in the structural estimation below, constructed measures of fodder revenue risk are unlikely to proxy for unobserved variation in transactions costs across households. This makes it possible to identify the effect of fodder price risk on production choices independently from transactions costs, without ruling out the presence of transactions costs effects as well.

Transactions costs nevertheless remain a potential source of bias if, in its effort to fit the empirical correlation between livestock holdings and fodder cultivation, the structural model attributes to risk aversion what is in fact due to transactions costs. The likelihood and magnitude of such bias are, unfortunately, difficult to ascertain in the absence of data on factors influencing transactions costs.

is small and non-significant in all regressions. Results also indicate that yields are higher among better educated farmers. This finding is consistent with the fact that farming in the survey area underwent a rapid technological change since the 1960's; education has indeed been shown to facilitate technology adoption (e.g., Jamison and Lau (1981); see, however, Fafchamps and Quisumbing (1999) for recent conflicting evidence in the Pakistan).

Predicted values from equation (29) are taken to represent differences in production technology across households that are pre-determined when crop choices are made. Multiplying $\hat{\Pi}^j_t$ by $\hat{\chi}^j_0 + \sum_{h=1}^K \hat{\chi}^j_h X^h_{i,t}$ yields an estimate of gross revenue at the household level; it is different for each household and each year. The standard deviation of $\nu^j_{i,t}$ is used as estimate of the variance of idiosyncratic yield shocks.

To go from gross to net revenues, we assume that total production costs per acre are proportional to expected farm-level revenues. The share of gross revenues that covers the production costs of crop h for household i in season j is denoted ϕ^{hj}_i and is estimated from the household survey data. Since, by definition, $\Pi^j_t = p^j_t \theta^j_t$ and:

$$\pi^{hj}_{i,t} = p^j_t \theta^{hj}_t (\chi^{hj}_0 + \sum_{n=1}^K \chi^{hj}_n X^n_{i,t} + \nu^{hj}_{i,t}) - \sum_{n=1}^K \kappa_{jn} w^n_t \tag{30}$$

an estimate of the expected net revenues per acre can be computed as:

$$E[\pi^{hj}_{i,t}] \approx \hat{\Pi}^j_t (1 - \hat{\phi}^{hj}_i)(\hat{\chi}^{hj}_0 + \sum_{n=1}^K \hat{\chi}^{hj}_n X^n_{i,t}) \tag{31}$$

Combining equation (31) with previous results and assumptions, an estimate of the coefficient of variation of net revenues $CV(\pi^{hj}_{i,t})$ can also be derived as:

$$CV(\pi^{hj}_{i,t}) \approx \frac{\hat{CV}(\Pi^h_t)}{1 - \hat{\phi}^{hj}_i} \left[1 + \frac{\hat{Var}(\nu^{hj}_{i,t})}{(\hat{\chi}^{hj}_0 + \sum_{n=1}^K \hat{\chi}^{hj}_n X^n_{i,t})^2} \left(1 + \frac{1}{\hat{CV}(\Pi^h_t)^2} \right) \right]^{1/2} \tag{32}$$

Covariance terms are similarly constructed (see Kurosaki (1997)).

Average coefficients of variation are reported in Table 6 for the prices and net returns per acre of Basmati, wheat, and fodder. They show that fodder price risk is much higher than Basmati or wheat price risk. Except for Rabi fodder which is shown to have extremely volatile returns, the variation of net returns is generally comparable between grain and fodder due to substantial yield risk. Examination of the variance-covariance matrix, shown in Table VI in Kurosaki (1997), reveals that returns to milk are very sensitive to fodder prices, thereby introducing a strong negative correlation between net profits from fodder and milk production. This negative correlation confers a risk diversification advantage to households who combine both activities. This is because fodder production reduces the variance of total net income by insulating milk profits from fodder price movements. This is true even though fodder production *per se* is more risky than grain production (Table 6).

Section 4. Estimation of the Structural Parameters

Armed with the above estimates, we can now proceed with the estimation of the structural parameters of equations (17) to (23).[23] The absence of consistent evidence of missing factor markets in this highly commercialized agricultural region of Pakistan enables us to focus on risk issues. FIML estimates of the complete model are given in Table 7. Judging by the asymptotic standard errors reported in the Table, parameters are estimated with remarkable precision: standard errors are less than half of the corresponding point estimate in 19 out of 22 cases. The reader should nevertheless keep in mind that the reported accuracy is only as good as the assumptions underlying the structural model and the restrictions that it imposes on the data.

Three of the four parameters characterizing risk attitudes are significant at the 5% level; livestock is non-significant. Predicted values of $\Psi_{i,t}$ are positive for all observations and range between 1.34 and 4.11. The mean $\Psi_{i,t}$ is 1.83 (Table 8), twice as large as the similarly constructed coefficient reported by Rosenzweig and Wolpin (1993) for Indian farmers — 0.96. The (estimated) coefficient of relative risk aversion $\hat{r}_{i,t}$ is obtained using

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[23] Because of computational burden, estimates of the means and variances of prices and revenues are treated as fixed points in FIML estimation.

equation (13). Results, presented in Table 8, yield estimates of relative risk aversion that fluctuate between 1.8 and 20. Large values occur for households which are very close to minimal necessary consumption, i.e., for whom $z_{i,t} \approx \sum q_t^j \gamma^j$. Our average estimate of relative risk aversion — 3.6 — is somewhat larger than estimates reported for India by Morduch (1990) — 1.39 — and Fafchamps and Pender (1997) — 1.8 to 3.1. It nevertheless remains of the same order of magnitude as those elicited in experimental games with South Asian farmers by Binswanger (1980) and Binswanger and Sillers (1983). The similarity of magnitude of our risk preferences and experimental results is consistent with Antle's (1987) finding that econometric estimates of risk aversion based on the production choices of Indian farmers are similar to Binswanger's results. This is reasonable if risk preferences elicited from experiments better approximate the curvature of V_x — "after-insurance" preferences — rather than the curvature of v_y — intrinsic preferences — a likely outcome if farmers expect to share gains and losses from experimental games with other villagers. This issue deserves further investigation.

Table 8 also reports the elasticity of $\hat{r}_{i,t}$ with respect to education, land, and livestock.[24] Results indicate that more land and livestock reduce risk aversion, a finding in line with the theoretical prediction that households with more liquid assets and more access to credit are better able to bear risk. The effect of education on risk aversion is zero on average, with some variation across households. Consumption parameters are also in line with expectations. Implied income elasticities at the sample mean are .95 for wheat, .84 for milk, .61 for rice and 1.12 for other consumption goods. There is, however, substantial variation among surveyed households (Table 8).

Coefficients capturing technological constraints are all individually significant and imply a mildly curved relationship between $l_{i,t}^{sg}$ and $l_{i,t}^{sf}$ for $s = k, r$, and hence moderately binding agronomic constraints. Crop choices appear to be influenced more by risk con-

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[24] These elasticities are obtained by differentiating equation (13), controling for the indirect effect of education, land, and livestock on income and, thus, on total expenditures $z_{i,t}$. This explains why reported elasticities differ in sign from estimated coefficients reported in Table 7. Similar results are obtained if the elasticity of absolute risk aversion is used instead.

cerns than by technical considerations of joint production and competition between crops. The coefficient of tubewell ownership is negative in both seasons, consistent with the hypothesis that farms with better access to water can allocate land more freely. The tubewell coefficient is larger in Kharif than Rabi, in line with the fact that timely water application is more important for Basmati paddy than wheat.

Section 5. Testing Alternative Specifications

One advantage of structural estimation is that it makes it possible to test alternative specifications of the same decision problem. To further our understanding of decision making under risk, we now compare the performance of the structural model presented in Section 4 (Model 1) with that of alternative structural specifications. We first compare Model 1 with results obtained by assuming that surveyed households are perfectly insured and behave as expected profit maximizers (Model 2). Coefficient estimates are reported in the first column of Table 9. To test whether expected profit maximization fits the data better than Model 1, we apply Vuong (1989)'s likelihood-based test.

Vuong's test statistic T_{AB} compares two non-nested representations of the same data, say, Model A and Model B and is computed as $T_{AB} = \frac{LR}{\hat{\omega}\sqrt{3N}}$ where LR is the difference in log-likelihood between the two models, $\hat{\omega}$ is given by:

$$\hat{\omega}^2 \equiv \frac{1}{3N}\sum_{i=1}^{N}\sum_{t=1}^{3}(d_{i,t}^A - d_{i,t}^B)^2 - \left[\frac{1}{3N}\sum_{i=1}^{N}\sum_{t=1}^{3}(d_{i,t}^A - d_{i,t}^B)\right]^2 \qquad (33)$$

and $d_{i,t}^h$ is the contribution of observation i, t to the log-likelihood of Model h, for $h = A, B$. Under the null hypothesis that the two models equally fit the data, T_{AB} is distributed as a standard normal variable. If Model A provides a better fit than Model B, $T_{AB} \rightarrow \infty$. If, in contrast, B is better than A, then $T_{AB} \rightarrow -\infty$. Applied to Models 1 and 2, $T_{12} = 6.98$. The hypothesis that the two models fit the data equally well can be rejected in favor of Model 1: the hypothesis of complete insurance is again inconsistent with production choice data.

We then test whether full insurance exists within each of the five surveyed villages but not with the rest of the world (Model 3). In this case, households remain exposed to aggregate yield and price shocks but their production decisions should be insensitive to idiosyncratic yield risk. To test this hypothesis, we recompute the variance-covariance matrix of revenues and incomes, assuming that surveyed households are only subject to aggregate risk. We then reestimate equations (17) to (23) using the alternative variance-covariance matrices.

Estimated coefficients are presented in Table 9, column 2. Risk aversion estimates are higher in Model 3 than Model 1. Vuong's test is $T_{13} = 0.17$: the hypothesis that Models 1 and 3 equally fit the data cannot be rejected. We cannot, therefore, reject the hypothesis that village members are able to efficiently share risk among themselves. It is possible that the risk averse behavior of Sheikhupura farmers is due to insufficient channels for sharing aggregate risk with the rest of the world. If villagers are insured against idiosyncratic risk, individual consumption should only vary with village-level consumption, not with individual income (e.g., Mace (1991), Cochrane (1991)). To test whether rural households efficiently share risk at the village level, we regress changes in individual consumption on changes in village-level consumption and in individual income, i.e., we estimate:

$$z_{i,t} - z_{i,t-1} = a + b(z^*_{i,t} - z^*_{i,t-1}) + c(Y_{i,t} - Y_{i,t-1}) + u_{i,t} \tag{34}$$

where $z_{i,t}$ denotes total consumption expenditures per capita of household i in year t, $z^*_{i,t}$ is the village-level average of total consumption per capita (minus that of household i, to avoid spurious correlation), and $Y_{i,t}$ is the total income per capita of household i in year t. The intuition behind equation (34) is that, if risk is shared efficiently at the village level, idiosyncratic income shocks should not affect changes in individual consumption once changes in average village consumption are controlled for. Equation (34) can be derived from utility maximization using a constant absolute risk aversion (CARA) utility function. In case of constant relative risk aversion (CRRA), consumption and income in levels need to be replaced by logs.[25] See Mace (1991), Cochrane (1991), and Townsend

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[25] Neither CRRA nor CARA formulations are fully consistent with our structural model since they

(1994) for derivation and details.

Equation (34) is estimated for the entire sample and for the sub-sample of households with three years of data, in levels (CARA) and in logs (CRRA). The results, summarized in Table 10, show that village-level risk pooling cannot be rejected: the F-statistics are all below the critical value with two and N degrees of freedom.

Cochrane (1991) and Ravallion and Chaudhuri (1997) have criticized the estimation of equations such as (34) as too sensitive to the choice of specific functional form for utility, and as vulnerable to measurement error. In particular, Ravallion and Chaudhuri (1997) show that the bias is large in the results reported by Townsend (1994). They suggest running the following regression instead:

$$z_{i,t} - z_{i,t-1} = \sum_v \sum_t a_{v,t} D_{v,t} + c(Y_{i,t} - Y_{i,t-1}) + u_{i,t} \tag{35}$$

where $D_{v,t}$ denotes village-time dummies and $a_{v,t}$ are parameters to be estimated. Efficient risk sharing cannot be rejected if c is not significantly different from 0. Results from the estimation of equation (35) are reported in the second half of Table 10, in both level and log form. Again, we cannot reject full insurance at the village level. These results are by and large consistent with other studies that have demonstrated that household consumption in South Asia and other parts of the Third World is largely, though not perfectly, insulated from idiosyncratic shocks (e.g., Townsend (1994), Paxson (1992), Morduch (1991), Jacoby and Skoufias (1995)), and that rural dwellers actively trade risk through transfers and consumption credit but face information and enforcement constraints (e.g., Rosenzweig (1988), Udry (1994), Fafchamps (1992a), Foster and Rosenzweig (1995), Ligon, Thomas and Worrall (1996), Fafchamps and Lund (2001)). The risk averse behavior of sample farmers appears due to insufficient channels to share collective risk with the rest of the world.

Next, we test whether consumption preferences influence decisions under risk (Model 4). We reestimate the model assuming that the value function does not depend on con-

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ignore the effect of minimal subsistence consumption on risk preferences. They nevertheless provide useful benchmarks given that they have been used extensively in the literature.

sumption prices, i.e., that $V(z, q^j)$ takes the familiar form $\frac{z^{1-\Psi}}{1-\Psi}$. The purpose of this test is to investigate whether ignoring consumption price risk, as is done in most of the work on decision under risk, is a valid approximation in practice. Results from Model 4 are presented in the third column of Table 9. Vuong's test statistic is 5.13, sufficient to reject Model 4 in favor of Model 1. Consumption preferences are thus shown to affect consumption choices. Coefficients estimated from Model 4 are similar to those obtained in Model 1, except that risk aversion is estimated to be a little higher — $\hat{\Psi}$ is 1.93 on average compared to 1.83 in Model 1. Ignoring the effect of consumption price risk on production decisions can thus lead to a bias in estimating risk aversion. The bias is fairly small in this case because the coefficient of variation of food prices is low in the study area (see Table 6). If, however, a similar test was conducted in an area with more volatile food prices, the difference in risk parameter estimates would probably be larger. We also reestimate Model 3 (without idiosyncratic risk) assuming that consumption price risk does not affect crop choice (Model 5). Vuong's test statistic T_{35} is 4.44, again allowing us to safely reject the hypothesis of no consumption price effect.

To test the robustness of our results, we reestimate Models 1 to 5 using only the 59 households for whom we have all three years of data. To control for omitted variable bias, household fixed effects are introduced in estimating the matrix of means and variance-covariances of prices and revenues (see Section 3). Crop revenues may indeed depend on the unobserved heterogeneity of farmers' human capital — e.g., experience and talent for farming — which itself may be correlated with livestock and land holdings. Vuong's test statistics, reported in the second panel of Table 11, lead to the same conclusions as before.

Finally, we investigate the sensitivity of our results to measurement error in the expectations and variance-covariance matrix of prices and revenues and reestimate the models with alternative assumptions regarding their computation. Results are summarized in Table 11. First, we experiment with a multiplicative specification of equation (29) and run it in logs instead of levels (alternative assumption AA1). Equation (30) and (31) are modified accordingly. The resulting variance-covariance matrix is qualitatively similar to the one used so far, and so are all the test results. Next, to account for the fact that the available data on fodder prices may be less reliable and thus may overestimate their variation over time, we reduce the coefficient of variation of fodder prices by 25% and rerun models 1 to 5 (AA2). It is also possible that subjective crop production costs are lower than those imputed on the basis of market prices (e.g., Feder (1980)). We therefore rerun the models with lower imputed production costs for fodder, resulting in estimates of the coefficient of variation of Kharif and Rabi fodder of 0.44 and 0.64, respectively (AA3). Finally, we rerun the models with lower production cost estimates, and hence lower coefficients of variation for all crops (AA4).

Not surprisingly, results show that reducing the variance-covariance of prices and profits leads to higher estimates of risk aversion: higher values of $\Psi_{i,t}$ are required to reconcile observed crop choices with utility maximization. Other results are essentially unaffected by alternative assumptions. Model 2 with expected profit maximization is consistently rejected in favor of Model 1 or Model 3 with risk averse behavior. Models in which production choices depend on consumption preferences are consistently preferred to models without ordinal preference effects (e.g., Model 4 vs. Model 1 and Model 5 vs. Model 3). Models 1 and 3 are, in general, equivalent; only with alternative assumptions AA3 and AA4 does Vuong's test lead to the rejection of Model 1 in favor of Model 3. We interpret this result as additional but limited evidence of risk pooling at the village level. Finally, ignoring ordinal preferences always leads to overestimate risk aversion. Our main conclusions thus appear quite robust.

Before concluding, we investigate the magnitude of the effect that various sources of risk have on welfare and production choices. The model estimated in Table 7 is simulated under alternative risk environments; more detailed results can be found in Kurosaki (1998), Chapter 8.[26] A hypothetical household with median characteristics is used as

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[26] In the simulations reported here, market prices are assumed unaffected by farmers' production choices. Kurosaki (1998) reports simulation results in which the fodder market price adjusts in response to farmers' aggregate production decisions.

reference. Livestock poor and land poor households simulations are obtained by halving median livestock and median land, respectively. Results, presented in Table 12, show that surveyed farmers respond to risk by reducing their Basmati area: an elimination of all risk would raise rice cultivation by close to 30%, thereby increasing the expected income of the representative household by 2%. Risk also has a dramatic impact on welfare: an elimination of all price and output risk faced by the reference household would raise its welfare by 9.4%. An elimination of fodder price risk alone would raise welfare by 5.0% — more than half the cost of risk from all sources. In contrast, an elimination of yield risk would only raise welfare by 3.6%. Results for livestock poor and land poor households do not fundamentally alter the above conclusions. The most noticeable difference is that the welfare gain from risk elimination is larger for land poor households, especially for fodder price risk. Taken together, these simulation results illustrate vividly the magnitude of the welfare cost of risk.

Conclusions

We have investigated the efficiency of insurance markets in the Pakistan Punjab by examining how crop choices are affected by the presence of price and yield risk. To do so, we estimated reduced-form and structural models of crop choices using household survey data. Results are by and large consistent across the two approaches, but the structural model provides deeper insights into the relationship between risk and production decisions.

Although we cannot reject the hypothesis that village members efficiently share risk among themselves, production choices are shown to depend on risk. The production decisions of Punjabi farmers thus appear to be affected by the presence of risk in spite of a large body of evidence suggesting that South Asian farmers self-insure and share risk with others (e.g., Townsend (1994), Morduch (1991), Foster and Rosenzweig (1995), Fafchamps and Pender (1997), Walker and Ryan (1990), Jacoby and Skoufias (1995)). The reason appears to be that households find it difficult to protect themselves against collective shocks that affect yields as well as output and input prices. Evidence of risk sharing and precautionary saving by Third World households should thus not be interpreted as a sign that existing institutions are efficient. Barriers to the pooling of risk across villages and regions remain. Unless these barriers are removed, government intervention is needed to mitigate village-level shocks that Third World households face, such as famines and floods.

Results further reveal that households adapt production to respond to consumption price risk. This suggests that empirical and theoretical work on risk should avoid putting an exclusive emphasis on yield and output price risk. In this respect, our results nicely complement previous work by Roumasset (1976), Antle (1987, 1989), Antle and Crissman (1990), and Walker and Ryan (1990). Our work also builds a bridge between the literature on production decision under risk and that on food self-sufficiency with or without missing markets (e.g., de Janvry, Fafchamps and Sadoulet (1991)): even when food markets are present, households may produce their own food in order to self-insure against price fluctuations. The same reasoning explains why they may choose to produce inputs, e.g., green fodder, for other farm activities, e.g., milk production.

Appendix. First Order Conditions in Detail

A fully parameterized estimable approximation to $FOC^s_{i,t}$, for $s = k, r$, is derived as follows. First, use equation (11) to compute $\bar{V}_x$, $\bar{V}_{xx}$ and $\bar{V}_{xq_j}$ evaluated at the one-year ahead expected values for prices, revenues, and incomes. Then replace these expressions in each of the $FOC^s_{i,t}$. After some tedious but straightforward algebraic manipulations, a system of second-order polynomial equations in the four endogenous variables $l^{sj}_{i,t}$ is obtained:

$$\begin{aligned} 0 &= FOC^s_{i,t} \approx F^g_{i,t,s} - F^f_{i,t,s}(\alpha^s_1 + 2\alpha^s_2 l^{sg}_{i,t}) \\ &\quad + \sum_{h=\{k,r\}} \sum_{j=\{g,f\}} \left[G^{hjg}_{i,t,s} - G^{hjf}_{i,t,s}(\alpha^s_1 + 2\alpha^s_2 l^{sg}_{i,t}) \right] l^{hj}_{i,t} \end{aligned} \tag{36}$$

for $s = k, r$, where $F_{i,t,t}$ and $G_{i,t,s}$ terms are functions of constructed variables and of the parameters to be estimated, β, γ and Ψ, i.e.:

$$\begin{aligned} F^a_{i,t,s} &\equiv (1 - \Psi_{i,t}) \sum_j \gamma^j E[q^j] \sum_d \beta^d_{i,t} Z^{a,d,1}_{i,t,s} + \Psi_{i,t} \sum_d \gamma^d Z^{a,d,2}_{i,t,s} \\ &\quad -(1 - \Psi_{i,t}) \sum_d \beta^d_{i,t} Z^{a,d,3}_{i,t,s} + \sum_d \gamma^d Z^{a,d,4}_{i,t,s} + \Psi_{i,t} Z^{a,5}_{i,t,s} + Z^{a,6}_{i,t,s} \end{aligned}$$

for $a = g, f$, and:

$$G^{hja}_{i,t,s} \equiv (1 - \Psi_{i,t}) \sum_d \beta^d_{i,t} W^{h,j,a,d,1}_{i,t,s} + \Psi_{i,t} Z^{h,j,a,2}_{i,t,s} + Z^{h,j,a,3}_{i,t,s}$$

for $a = g, f$, $h = k, r$, and $j = g, f$, and where the $Z_{i,t,s}$ and $W_{i,t,s}$ are constructed variables defined as follows:

$$\begin{aligned}
Z^{a,d,1}_{i,t,s} &\equiv \frac{1}{L^{s}_{i,t}} \frac{Cov(q^{d}_{t}, \pi^{sa}_{i,t})}{E[q^{d}_{t}]E[\pi^{sf}_{i,t}]} \\
Z^{a,d,2}_{i,t,s} &\equiv \frac{E[q^{d}_{t}]}{E[\pi^{kf}_{i,t}]} Z^{a,d,1}_{i,t,s} \\
Z^{a,d,3}_{i,t,s} &\equiv \left[A_{i,t} \sum_{h=k,r} E[\pi^{hm}_{i,t}] + R_{i,t} \right] \frac{Z^{a,d,1}_{i,t,s}}{E[\pi^{kf}_{i,t}]} \\
Z^{a,d,4}_{i,t,s} &\equiv -\frac{1}{L^{s}_{i,t}} \frac{E[\pi^{sa}_{i,t}]}{E[\pi^{sf}_{i,t}]} \frac{E[q^{d}_{t}]}{E[\pi^{kf}_{i,t}]} \\
Z^{a,5}_{i,t,s} &\equiv \frac{1}{L^{s}_{i,t}} \frac{A_{i,t} \sum_{h} Cov(\pi^{sa}_{i,t}, \pi^{hm}_{i,t})}{E[\pi^{kf}_{i,t}]E[\pi^{sf}_{i,t}]} \\
Z^{a,6}_{i,t,s} &\equiv \frac{1}{L^{s}_{i,t}} \frac{E[\pi^{sa}_{i,t}]}{E[\pi^{sf}_{i,t}]} \frac{Z^{a,d,3}_{i,t,s}}{Z^{a,d,1}_{i,t,s}} \\
W^{h,j,a,d,1}_{i,t,s} &\equiv L^{h}_{i,t} \frac{E[\pi^{hj}_{i,t}]}{E[\pi^{kf}_{i,t}]} Z^{a,d,1}_{i,t,s} \\
W^{h,j,a,2}_{i,t,s} &\equiv -\frac{L^{h}_{i,t}}{L^{s}_{i,t}} \frac{Cov(\pi^{hj}_{i,t}, \pi^{sa}_{i,t})}{E[\pi^{kf}_{i,t}]E[\pi^{sf}_{i,t}]} \\
W^{h,j,a,3}_{i,t,s} &\equiv \frac{L^{h}_{i,t}}{L^{s}_{i,t}} \frac{E[\pi^{hj}_{i,t}]}{E[\pi^{kf}_{i,t}]} \frac{E[\pi^{sa}_{i,t}]}{E[\pi^{sf}_{i,t}]}
\end{aligned}$$

The parameterization of $\Psi_{i,t}$ is given in equation (14). Variables $A_{i,t}$ and $R_{i,t}$ stand for livestock assets and non-farm income, respectively. In the estimation, expectations, variances and covariances are replaced by the consistent estimates described in Section 3. Closed form solutions for $l^{sj}_{i,t}$ cannot be computed from the set of polynomial equations (36), which is why a Jacobian transform term appears in likelihood function (24).

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Table 1. Main Characteristics of the Survey Area

Variable	**Symbol**	**Unit**	**Mean**	**Std. Dev.**
Land owned	$\bar{L}$	acres	11.170	9.175
Cultivated land in Kharif	L^k	acres	9.071	7.266
Cultivated land in Rabi	L^r	acres	8.846	6.773
Livestock owned	$\bar{A}$	ad. equiv.	6.361	3.534
Milking animals	A	ad. equiv.	4.742	2.987
Non-farm income	R	’000 Rupees	6.688	3.609
Tubewell ownership	D^W	1=yes	0.694	
Tractor ownership	D^T	1=yes	0.137	
Years of education of head	EDU	years	1.794	3.495
No. persons per cultivated area	AL	pers./acre	1.153	0.892
Dependency ratio	DEP	child./hh size	0.268	0.184
Share of Basmati in Kharif cultivated land	l^{kg}	share	0.653	0.162
Share of fodder in Kharif cultivated land	l^{kf}	share	0.289	0.140
Share of wheat in Rabi cultivated land	l^{rg}	share	0.734	0.106
Share of fodder in Rabi cultivated land	l^{rf}	share	0.209	0.097
Total consumption expenditures per capita		’000 Rs./pers.	3.055	0.606
Share of wheat in total expenditures	s_w	share	0.135	0.021
Share of rice in total expenditures	s_r	share	0.041	0.008
Share of milk prod. in total expenditures	s_m	share	0.271	0.028

No. of observations = 291

Note: During the study period, the exchange rate of Pakistani Rupee was approximatively 21 Rupees per US dollar.

Table 2. Market Participation

	Sales		Purchases	
	No. of observations	%	No. of observations	%
Basmati paddy	290	99.7%	0	0.0%
Wheat	182	62.5%	28	9.6%
Milk	226	77.7%	16	7.7%
Kharif fodder	97	33.3%	15	5.2%
Rabi fodder	110	37.8%	9	3.1%

Table 3. OLS Estimation of the Reduced-Form Model

	Basmati			Kharif fodder			Wheat			Rabi fodder		
	coef.	t-stat		coef.	t-stat		coef.	t-stat		coef.	t-stat	
Intercept	0.559	15.135	***	0.371	11.523	***	0.755	29.141	***	0.199	8.580	***
Tubewell ownership	0.100	4.670	***	-0.057	-3.079	***	0.000	0.001		0.009	0.695	
Tractor ownership	0.005	0.169		-0.004	-0.156		0.022	0.981		-0.015	-0.755	
Person/acre	-0.029	-2.262	**	0.028	2.500	**	-0.004	-0.431		-0.004	-0.527	
Education of head	-0.002	-0.733		0.002	0.837		-0.001	-0.634		0.001	0.534	
Year 1989 dummy	0.066	3.210	***	-0.087	-4.855	***	-0.020	-1.396		0.001	0.098	
Year 1990 dummy	0.067	3.230	***	-0.088	-4.809	***	0.001	0.039		-0.010	-0.774	
Non-farm income	-0.031	-0.840		0.001	0.040		0.000	0.017		-0.021	-0.889	
Cultivated area	0.002	1.373		-0.003	-1.984	**	0.004	3.892	***	-0.005	-5.057	***
Livestock owned	-0.002	-0.678		0.005	1.744	*	-0.010	-4.643	***	0.012	6.223	***
Dependency ratio	0.075	1.527		-0.068	-1.573		0.011	0.314		0.011	0.343	
R-squared	0.271			0.253			0.155			0.202		
Wald test that the last four variables are jointly significant:												
F(4,280)	1.581			2.177	*		6.593	***		11.493	***	

Note: For coefficient estimates, *, **, and *** mean that the coefficient is significant at the 10%, 5%, and 1% levels, respectively. For Wald tests, *, **, and *** mean that the null hypothesis is rejected at the 10%, 5%, and 1% levels, respectively.

Table 4. Effect of Fodder Market Participation on Crop Choices

	Basmati			Kharif fodder			Wheat			Rabi fodder		
	coef.	t-stat		coef.	t-stat		coef.	t-stat		coef.	t-stat	
Intercept	0.561	14.993	***	0.372	11.386	***	0.751	28.67	***	0.205	8.753	***
Tubewell ownership	0.100	4.665	***	-0.057	-3.073	***	0.000	0.004		0.010	0.708	
Tractor ownership	0.007	0.210		-0.004	-0.146		0.019	0.816		-0.011	-0.533	
Person/acre	-0.029	-2.276	**	0.028	2.484	**	-0.003	-0.343		-0.005	-0.649	
Education of head	-0.002	-0.781		0.002	0.806		-0.001	-0.390		0.000	0.203	
Year 1989 dummy	0.066	3.198	***	-0.087	-4.847	***	-0.020	-1.374		0.001	0.067	
Year 1990 dummy	0.067	3.224	***	-0.088	-4.801	***	0.001	0.043		-0.010	-0.782	
Non-farm income	-0.000	-0.821		0.000	0.043		0.000	-0.046		-0.000	-0.804	
Cultivated area	0.002	1.211		-0.003	-1.909	*	0.005	4.070	***	-0.005	-5.331	***
Livestock*D	-0.003	-0.749		0.004	1.507		-0.008	-3.550	***	0.010	4.745	***
Livestock*$(1-D)$	-0.002	-0.527		0.005	1.659	*	-0.011	-4.771	***	0.013	6.427	***
Dependency ratio	0.075	1.523		-0.068	-1.571		0.011	0.321		0.010	0.334	
R-squared	0.271			0.253			0.159			0.209		
Wald test that the two livestock variables have the same coefficient:												
F(1,279)	0.103			0.004			1.398			2.691		

Note: D is a dummy variable taking the value of 1 for households that never participated in green fodder market during the study period. For coefficient estimates, *, **, and *** mean that the coefficient is significant at the 10%, 5%, and 1% levels, respectively. All the Wald test statistics are statistically insignificant at 10%.

Table 5: OLS Estimation of the Reduced-Form Yield Model

	Basmati			Kharif fodder			Wheat			Rabi fodder		
	coef.	t-stat		coef.	t-stat		coef.	t-stat		coef.	t-stat	
Intercept	0.829	12.099	***	0.927	13.035	***	0.862	16.741	***	1.019	19.699	***
Tubewell ownership	0.016	0.404		-0.053	-1.282		0.038	1.305		-0.016	-0.531	
Tractor ownership	-0.173	-2.518	**	-0.094	-1.325		0.102	1.990	**	-0.122	-2.408	**
Person/acre	0.049	1.575		0.073	2.246	**	-0.023	-0.968		0.031	1.322	
Education of head	0.021	4.732	***	0.019	4.192	***	0.002	0.626		0.009	2.687	***
Cultivated area	0.009	2.686	***	0.010	2.832	***	-0.001	-0.430		0.003	1.369	
Non-farm inc./land	-0.003	-1.487		-0.002	-0.924		0.003	1.515		-0.002	-1.191	
Livestock/land	0.039	1.062		-0.011	-0.298		0.081	2.995	***	-0.052	-1.934	*
Dependency ratio	-0.060	-0.650		-0.303	-3.184	***	0.161	2.390	**	-0.086	-1.274	
R-squared	0.176			0.181			0.130			0.106		
Nber obs.	177			177			171			176		

Notes: The dependent variable is defined in equation (29). t-statistics is indicated in parentheses; *** indicates significance at 1%, ** at 5%, and * at 10%, respectively (two-sided t-test).

Table 6. Variability of Prices and Revenues

Coefficient of variation of	Price	Net Revenue
Basmati paddy	0.141	0.486
Wheat	0.086	0.559
Milk	0.146	0.862
Kharif fodder	0.353	0.481
Rabi fodder	0.415	1.262

Table 7. Estimation Results for Model 1

Willingness to Bear Risk			
Ψ_0	intercept of $\Psi_{i,t}$	1.452	(0.155)
Ψ_e	effect of education of $\Psi_{i,t}$	0.054	(0.018)
Ψ_l	effect of land ownership on $\Psi_{i,t}$	0.039	(0.013)
Ψ_a	effect of livestock ownership on $\Psi_{i,t}$	-0.024	(0.017)
Consumption Preferences			
β_0^w	intercept of $\beta_{i,t}^w$	0.035	(0.016)
β_1^w	effect of dependency ratio on $\beta_{i,t}^w$	0.284	(0.046)
β_0^m	intercept of $\beta_{i,t}^m$	0.192	(0.009)
β_1^m	effect of dependency ratio on $\beta_{i,t}^m$	0.214	(0.047)
β_0^b	intercept of $\beta_{i,t}^b$	0.014	(0.008)
β_1^b	effect of dependency ratio on $\beta_{i,t}^b$	0.062	(0.019)
γ^w	per-capita subsistence consumption for wheat	4.148	(0.120)
γ^m	per-capita subsistence consumption for milk	5.799	(0.159)
γ^b	per-capita subsistence consumption for rice	0.907	(0.028)
γ^o	per-capita subsistence cons. for other items	7.238	(0.221)
Technological Constraints on Crop Choices			
α_0^k	intercept in Kharif	-0.636	(0.010)
α_1^k	linear term in Kharif	0.130	(0.037)
α_2^k	quadratic term in Kharif	0.624	(0.032)
α_3^k	effect of tubewell ownership in Kharif	-0.028	(0.007)
α_0^r	intercept in Rabi	-0.642	(0.018)
α_1^r	linear term in Rabi	0.280	(0.037)
α_2^r	quadratic term in Rabi	0.433	(0.031)
α_3^r	effect of tubewell ownership in Rabi	-0.014	(0.009)
Log-likelihood		-3282.98	

Note: The number of observations is 291. Asymptotic standard errors are given in parentheses. They are computed by inverting the sum of squares matrix of the outer product of the gradient of the likelihood function with respect to relevant parameters (see Berndt et al. (1974)).

Table 8. Summary Statistics of Model Parameters

	Sample Mean	Standard Dev.	Minimum	Maximum
Income elasticity of consumption:				
wheat	0.955	0.228	0.406	1.645
milk products	0.844	0.073	0.635	0.976
rice	0.610	0.150	0.265	0.911
other items	1.117	0.087	0.907	1.495
Risk aversion parameter:				
$\hat{\Psi}_{i,t}$	1.830	0.381	1.341	4.116
Arrow-Pratt coefficient of relative risk aversion:				
Value	3.596	2.381	1.786	20.682
Its elasticity w.r.t				
Education of head	0.032	0.072	-0.153	0.251
Land ownership	-0.424	0.956	-9.558	0.411
Livestock ownership	-0.450	0.579	-4.543	0.000

Note: The number of observations is 291 for demand elasticity and distribution of $\hat{\Psi}_{i,t}$. It is 282 for Arrow-Pratt measures, after eliminating observations that have negative or zero expected income after meeting LES subsistence requirements.

Table 9. Estimation Results for Models 2-4

	Model 2 Expected profit maximization		Model 3 Complete village level insurance		Model 4 No consumption price effects	
Willingness to Bear Risk						
Ψ_0			2.055	(0.180)	1.573	(0.166)
Ψ_e			0.058	(0.095)	0.058	(0.018)
Ψ_l			0.078	(0.017)	0.040	(0.014)
Ψ_a			-0.071	(0.037)	-0.030	(0.019)
Consumption Preferences						
β_0^w	0.033	(0.028)	0.035	(0.017)	0.036	(0.016)
β_1^w	0.282	(0.054)	0.284	(0.046)	0.281	(0.046)
β_0^m	0.193	(0.010)	0.192	(0.012)	0.193	(0.009)
β_1^m	0.212	(0.062)	0.214	(0.049)	0.211	(0.047)
β_0^b	0.015	(0.009)	0.014	(0.008)	0.014	(0.012)
β_1^b	0.064	(0.023)	0.062	(0.024)	0.062	(0.019)
γ^w	4.162	(0.143)	4.149	(0.119)	4.133	(0.122)
γ^m	5.778	(0.129)	5.801	(0.155)	5.778	(0.156)
γ^r	0.904	(0.028)	0.907	(0.030)	0.904	(0.028)
γ^o	7.202	(0.213)	7.242	(0.221)	7.206	(0.218)
Technological Constraints on Crop Choices						
α_0^k	-0.471	(0.042)	-0.636	(0.010)	-0.636	(0.010)
α_1^k	-0.528	(0.221)	0.143	(0.035)	0.125	(0.035)
α_2^k	1.211	(0.215)	0.606	(0.030)	0.633	(0.031)
α_3^k	-0.030	(0.007)	-0.027	(0.008)	-0.029	(0.007)
α_0^r	0.038	(0.027)	-0.505	(0.042)	-0.642	(0.017)
α_1^r	-1.893	(0.070)	-0.051	(0.077)	0.274	(0.035)
α_2^r	2.089	(0.080)	0.630	(0.041)	0.440	(0.030)
α_3^r	-0.010	(0.008)	-0.017	(0.011)	-0.014	(0.009)
Log-likelihood	-3520.91		-3284.23		-3290.34	
Vuong's Test Statistics						
$\hat{\omega}$	1.999		0.424		0.084	
T	6.977		0.172		5.129	
p-value	0.000		0.864		0.000	

Note: Vuong's test is applied against Model 1.

Table 10. Risk Sharing Test Results

(t-value in parenthesis)

	In levels (CARA)		In logs (CRRA)	
	Balanced panel	Unbalanced panel	Balanced panel	Unbalanced panel
A. With Changes in village consumption				
Intercept	11.93	-0.84	0.003	0.001
	(0.311)	(-0.001)	(0.250)	(0.083)
Change in village consumption	0.907	0.922	0.947	0.939
	(5.704)	(6.492)	(7.015)	(6.854)
Change in household income	0.011	0.014	0.024	0.020
	(0.786)	(1.077)	(0.889)	(0.769)
F-statistic	0.584	0.942	0.611	0.510
B. With village-time dummies				
Change in household income	0.014	0.017	0.029	0.026
	(0.945)	(1.280)	(1.049)	(0.915)
Number of observations	118	152	118	152

Note: The F-statistic tests the joint hypothesis that $b = 1$ and $c = 0$.

Table 11. Summary of Model Specification Tests

	Log-likelihood	Sample mean of $\hat{\Psi}_{i,t}$	Vuong test statistic			
			vs. Model 1		vs. Model 3	
Default estimation:						
Model 1	-3282.98	1.830	n.a.		-0.172	
Model 2	-3520.91	n.a.	6.977	***	7.953	***
Model 3	-3284.23	2.581	0.172		n.a.	
Model 4	-3290.34	1.931	5.129	***	n.a.	
Model 5	-3294.47	2.877	n.a.		4.439	***
Balanced-panel estimation:						
Model 1	-1959.55	1.993	n.a.		-1.663	*
Model 2	-2102.38	n.a.	6.179	***	6.029	***
Model 3	-1965.97	2.526	1.663	*	n.a.	
Model 4	-1964.44	2.123	4.738	***	n.a.	
Model 5	-1966.31	2.728	n.a.		0.182	
Multiplicative yield equation (AA 1):						
Model 1	-3280.63	1.783	n.a.		-0.546	
Model 2	-3520.91	n.a.	7.151	***	7.953	***
Model 3	-3284.22	2.581	0.546		n.a.	
Model 4	-3288.10	1.875	5.109	***	n.a.	
Model 5	-3294.47	2.877	n.a.		4.438	***
Lower variance of fodder prices (AA 2):						
Model 1	-3291.01	2.404	n.a.		0.358	
Model 2	-3520.91	n.a.	7.188	***	7.992	***
Model 3	-3288.58	3.704	-0.358		n.a.	
Model 4	-3305.47	2.795	2.210	**	n.a.	
Model 5	-3304.00	4.176	n.a.		4.009	***
Lower variance of fodder profits (AA 3):						
Model 1	-3317.26	1.938	n.a.		2.036	**
Model 2	-3520.91	n.a.	6.786	***	7.952	***
Model 3	-3303.00	3.159	-2.036	**	n.a.	
Model 4	-3325.41	2.305	1.374		n.a.	
Model 5	-3316.84	3.592	n.a.		3.939	***
Lower variance of all profits (AA 4):						
Model 1	-3309.26	3.772	n.a.		2.108	**
Model 2	-3520.91	n.a.	8.647	***	8.492	***
Model 3	-3304.00	4.673	-2.108	**	n.a.	
Model 4	-3311.91	3.825	0.796		n.a.	
Model 5	-3313.35	5.013	n.a.		2.728	***

Note: Model 5 assumes perfect insurance within the village and no consumption price effects. For Vuong tests, *** means that the hypothesis that the model performs as well or better than Model 1 or Model 3 is rejected at the 1% level.

Table 12. Simulation Results for Welfare Cost of Risk

	From elimination of:			
	all risk	yield risk	output price risk	fodder price risk
Median household				
Change in Basmati acreage	29.5%	9.6 %	17.2%	11.7%
Change in expected income	2.0%	0.8%	1.7%	1.5%
Equivalent variation (1)	9.4%	3.6%	5.3%	5.0%
Livestock poor household				
Change in Basmati acreage	33.0%	13.4%	14.1%	6.4%
Change in expected income	1.9%	1.1%	1.1%	0.7%
Equivalent variation (1)	7.7%	4.4%	2.7%	2.1%
Land poor household				
Change in Basmati acreage	32.2%	6.9%	23.9%	20.0%
Change in expected income	2.2%	0.4%	2.1%	2.1%
Equivalent variation (1)	13.1%	3.5%	8.8%	8.8%

Notes: (1) Expressed as a percentage of initial expected income.