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# On price risk and the inverse farm size–productivity relationship

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**Abstract**

The oft-observed inverse relationship between farm size and productivity has elicited several explanations having important policy and theoretical implications. Using advances in the analysis of price risk effects on producer behavior, and a simple two-period model of an agricultural household that both produces and consumes under price uncertainty at the time labor allocation decisions are made, this paper demonstrates analytically that a non-degenerate land distribution and price risk can together produce an inverse relationship, even absent any of the more common explanations. Empirical evidence from Madagascar confirms the plausibility of this intuitive explanation for the phenomenon.

*JEL classification:* D13; O13; Q12

*Keywords:* Agricultural household models; Inverse relationship; Price risk; Madagascar

## 1. Introduction

The existence of an inverse relationship between farm size and output per unit of cultivated land (physical yield) has been observed by agricultural and development economists for some time [1]. This common empirical finding has drawn several different, often competing explanations, with important implications for both policy and theory. For example, if we interpret observed physical yield differences as indicative of the inherently greater efficiency of small farms, this

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[1] Amartya Sen (1962) launched contemporary inquiry into the issue with his analysis of Indian data. Chayanov (1966) had identified the issue with respect to Russian agriculture much earlier this century.


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makes a strong case for redistributive land reform. However, if it is not farms' size but rather market failures that engender smallholders' apparent superior yields, this suggests a particular vision of the institutional structure and functioning of the agrarian economy, one in which the fundamental welfare theorems might not hold and textbook analytical methods might seem inappropriate. Finally, if the observed inverse relationship is the result entirely of mismeasurement or omitted variables, there is no reason to doubt either the efficiency of the prevailing land distribution nor the efficacy of conventional analytical methods. These different implications have long made the inverse relationship a topic ripe for examination and debate.

This paper re-examines the inverse relationship in light of advances in the analysis of price risk effects on producer behavior [2]. The basic intuition of the model developed here is that where land or credit market failures constrain small farmers' capacity to outbid larger farmers for land, food price risk creates food security stress that induces small, net buyer farms to utilize extraordinary amounts of labor, beyond even their shadow valuation of labor. Large, net seller farms, by contrast, behave as Sandmo (1971) posits, reducing the use of costly inputs when revenues are stochastic. Objectively identical risk exposure thus stimulates different behaviors conditional on endowments, generating the oft-observed inverse relationship.

A brief digression is warranted to address an issue of language. The literature on the inverse relationship habitually, perhaps cavalierly, equates physical yields with "productivity," although yields are only a partial productivity measure that fails to account for the differential use of other inputs (Binswanger et al., 1995). Cognizant of this shortcoming, I nonetheless retain the term "productivity" in this paper for three reasons. First, and perhaps least importantly, this maintains consistency with the substantial, preceding literature on the inverse relationship. Second, for the two-factor analytical model developed here, with one factor (land) fixed because of market failures, physical yield maps one-to-one to total factor productivity. Third, the weight of empirical evidence suggests that small farms are more productive than large farms even when we account for differences in other inputs' use, often because power relations preclude efficient resource transfer among farms (Binswanger et al., 1995). Thus the logical flaw in equating physical yields with productivity is bridged empirically, but should be kept in mind nonetheless.

## 2. Existing explanations of the inverse relationship

The most popular explanations of the inverse relationship rely on labor market dualism. Small, peasant households are believed to face a lower opportunity cost

[2] The seminal papers by Baron (1970), Sandmo (1971), Just (1974), and Finkelshtain and Chalfant (1991), especially the latter, are the analytical ancestors of this paper.

of labor than large, commercial farms. Consequently, small farms apply their own labor in such quantities that the expected marginal value product of household labor applied to homestead cultivation is less than a market-wage-based measure of the opportunity cost of labor (Chayanov (1966), Sen (1966), Vergopoulos (1978), Hunt (1979), Carter and Wiebe (1990)). Assuming agricultural production technology does not exhibit increasing returns to scale, peasants' presumed proclivity to labor, sometimes referred to as a 'peasant mode of production', directly produces an inverse relationship between farm size and productivity. Carter (1984) finds empirical support for this explanation in Indian data.

A second alternative turns not on differences in household behavior but on the nature of technology. Decreasing returns to scale technology would suffice to produce the observed phenomenon. However, decreasing returns should also favor a natural subdivision of land into smaller holdings to achieve a scale efficient equilibrium, a phenomenon far less commonly observed. Moreover, a substantial majority of the literature containing empirical production function estimates for developing country agriculture find that observed returns are nearly constant (Bardhan (1973), Berry and Cline (1979), Carter (1984), Cornia (1985)). Technology seems an unlikely source for the inverse relationship under most circumstances.

Unlike the two preceding explanations, which suggest that small farms possess intrinsic productivity advantages, a quite different theoretical approach relies on market failures to generate binding farm-level constraints and shadow prices of factors that vary with farm size. Small farms' more intensive application of inputs then follows from such differences in shadow prices. Perhaps the most common formulation assumes that principal–agent problems and the associated labor supervision costs render effective labor costs higher on farms large enough to hire labor than on peasant plots (Bardhan (1973), Sen (1981a, b), Feder (1985), Eswaran and Kotwal (1986), Carter and Kalfayan (1989), Taslim (1989)). Frisvold (1994) finds empirical support for this explanation in Indian data. A less formally developed labor market failure explanation has been posited by Binswanger and Rosenzweig (1986). They note that imperfect information in a labor search results in a positive probability of the misallocation of labor. Labor-selling households that fail to find casual labor reallocate the time they had planned for wage labor to work on their own farm instead, up to the point where the marginal utility of home production equals the marginal utility of leisure. Since the household wanted work, the marginal utility of the wage (and thus of production) necessarily exceeded that of leisure, so at least some windfall labor time goes to home farming. Exactly the reverse happens for labor-hiring households who fail to hire casual labor; they fall short of planned labor applications. In each of these two models, the inverse relationship depends on imperfect labor markets.

Feder (1985) demonstrates generally that multiple market failures are necessary for this sort of model to hold, since if labor cannot go to the land, land can come to the labor. If a second market fails, shadow prices necessarily diverge from market prices. This can generate an inverse relation between size and input allocation levels, and thus an inverse relationship between farm size and output per unit cultivated area. These market failure models suggest that the inefficiencies in agrarian developing economies stem from institutional and infrastructural bottlenecks that impede the free and full participation of all households in factor and product markets. In such an environment, market prices are imperfect and unpredictable signals (DeJanvry et al. (1991)).

Some have suggested, however, that the inverse relationship might represent only poor measurement by analysts, not actual farm size-dependent differences in efficiency. In addition to smaller farms' more intensive application of labor inputs, Bharadwaj (1974) cites the sensitivity of cropping patterns to farm size as an explanation of the inverse relationship, with smaller farms allocating a higher proportion of area to more lucrative and labor-intensive crops. In particular, Bharadwaj notes that differences in cropping patterns resolve an apparent aggregation bias evident in empirical studies: inverse relationships that seem to hold at farm level often do not hold for individual crops. If aggregation bias accounts for observations of the inverse relationship, the implications for policy and theory are far less clear than under the peasant mode of production or market failures explanations.

Moreover, Sen (1975) reports that controlling for village-specific effects substantially reduces the observed effect of farm size on productivity, suggesting that prices, soils or wage rates cause the phenomenon. Regions of small farms might exhibit net food deficits, yielding a somewhat higher local price [3] than in net surplus regions populated by larger farms, and consequently inducing greater application of inputs. Or good soils might attract a greater population density, generating pressures to subdivide farms but still inducing a higher rate of input application than on farms with poorer soils. Finally, if households with small farms supply more labor to market than do households with large holdings, a region dominated by small farms may have more abundant, and thus cheaper labor, inducing greater application of labor in such a region. These hypotheses conflict somewhat, however, with estimation results reported by Bharadwaj (1974) and Carter (1984), each of whom considered village fixed effects and found that these could account for only a small portion of the observed variation in productivity across farm sizes. Bhalla (1988), Bhalla and Roy (1988), and Benjamin (1995) take these general points to farm-level, positing a size sensitivity in the quality of factor endowments, especially soil fertility. These authors claim that more intensive labor use on small farms reflects primarily the superior fertility of soils on smaller plots, thereby warranting more intensive labor per unit area. If correct, the village specific factors or farm-level soil quality variability arguments

[3] With the difference equal to a marketing margin between the prices prevailing in net surplus and net deficit areas.

undermine both efficiency arguments for land reform and claims that parsimonious neoclassical theory are inappropriate for analyzing low-income agriculture.

The stark differences in implications for policy and theory of the peasant mode of production, imperfect labor markets, and variable soil quality explanations are especially troubling because each is exceedingly difficult to prove empirically. The alleged peasant mode of production resists direct measurement and is routinely estimated as that portion of the inverse relation inexplicable by other, measurable variables. Reliable data on labor allocation and soil quality are rare and, to the best of my knowledge, no existing data set encompasses both, thereby precluding simultaneous direct testing of the competing hypotheses. An ideological stalemate results.

This paper offers an alternative explanation of the inverse relationship. It does not supersede the existing explanations, several of which are likely relevant in different places at different times, but offers an intuitive option that, unlike the existing explanations, is based on three empirically sound stylized facts readily verifiable in the data. First, farmers in low-income countries can neither fully hedge uncertain staple crop prices through futures or insurance contracts, nor through forward sales at the time that labor allocation decisions are made. Households may be averse to price risk, much as agents are often posited to exhibit Arrow–Pratt income-risk aversion [4]. Second, land is unevenly distributed across the agricultural population. Even if technologies and preferences are uniform among agents, differences in land endowments create a heterogeneous society. Finally, households' marketable agricultural surpluses are inversely related to landholdings. Small farms tend to be net purchasers while large farms tend to be net sellers. These three common features of low-income agriculture combine to induce an endogenous agrarian class structure, characterized by heterogeneous labor allocation behavior across classes of farmers, as in Roemer (1982). Section 3 proves analytically that price risk and distinct agrarian classes suffice in explaining the oft-observed inverse relationship between farm size and productivity. Even if technology does not exhibit decreasing returns to scale, all farms are able to hire or supply unlimited quantities of labor at a uniform, parametric wage, and there are no differences across farms in prices, cropping patterns or soil quality. The basic intuition is that small, net buyer farming households experience food price

[4] Stochastic prices form the base for the present model, although we can reproduce the general results by modelling stochastic output instead (Srinivasan (1972)). This paper by no means denies the importance of yield risk to cropping patterns and farm household behavior in developing country agriculture. Introducing both types of risk jointly, however, makes the model substantially more unwieldy without adding much in the way of insights, so the more tractable form of one source of risk is used here. With the dramatic changes taking place in recent years in agricultural pricing and markets, prices may be a more topical source of farm-level risk because there is little connection between policy reforms and exogenous shifts in yield risk while there can be strong links between policy reforms and commodity price risk (Barrett (forthcoming)).

risk as food security stress that induces hyper-exploitation of household labor, deviating even from their own shadow value of labor. Section 4 presents empirical findings from Madagascar that confirm the plausibility of this price risk-based model as an explanation for the inverse relationship. The concluding section, Section 5, draws out implications for both policy and theory.

## 3. An agricultural household model

Assume that a representative agricultural household exhibits Von Neumann–Morgenstern utility defined over consumption of leisure($L^L$) [5] and two goods: a staple food ($S$), and a non-staple ($N$) [6]. $U(\cdot)$ is quasi-concave, but concave in each argument individually, with $U_X|_{X=0} = \infty$ with respect to each argument $X$. The staple can either be produced or purchased, the non-staple is available only through market purchase.

The household has an endowment of land ($T$) and of labor time ($L^0$). Deterministic staple commodity production is strictly increasing in land and labor, and (weakly) concave in each. Agricultural labor is a function of household labor ($L^H$) and hired labor ($L^D$), but these may be imperfect substitutes. Just as the household can hire labor in, so can it hire out its time ($L^S$) at a parametric wage rate, $w$. Endowed land is a fixed factor of production, and neither land nor credit markets exist [7]. The household faces a time constraint, $L^S + L^L + L^H \leq L^0$. Exogenous transfers ($I$) supplement net wage earnings and agricultural revenues.

This is a two-period model. All product prices are unknown when production (i.e. labor allocation) decisions are made, but post-harvest prices are revealed before consumption decisions are made. The household's expected utility maximization problem can thus be expressed as

$$\begin{aligned}
&\underset{L^D, L^H, L^L}{Max}\ \mathrm{E}\underset{N,S}{Max}U(L^L, N, S)\\
&s.t.\ P^S S + P^N N \leq Y^*\\
&Y^* \equiv w[L^S - L^D] + P^S F(L,T) + I\\
&L \equiv h(L^D) + L^H\\
&L^0 \geq L^H + L^L + L^S\\
&h(L^D) \in [0, L^D] \qquad (1)
\end{aligned}$$

[5] Superscripts distinguish among goods across subcategories. Subscripts denote derivatives.

[6] There is no consumption in the first period, only labor allocation (i.e. production) decisions.

[7] The existence of imperfect credit markets, in which access to capital increases with land holdings, reinforces the qualitative results of the analysis that follow (Carter and Kalfayan (1989)) but has been assumed away to simplify the presentation.

where E is the mathematical expectation operator, $P^S$ is the staple price, $P^N$ is the non-staple price, $Y^*$ is endogenous income, and the function $h(\cdot)$ is a hired labor efficiency index used to convert hired labor units into household labor units. It takes the value zero if hired labor is completely inefficient, and $L^D$ if hired labor is as efficient as household labor. By the strict monotonicity of $U(\cdot)$, the budget and time constraints will bind at any optimum, and technical efficiency (i.e., output on the production possibility surface) has been assumed.

The household allocates labor conditional on anticipated ex-post optimal choice of consumption quantities. Thus, by duality, we can work with the variable indirect utility function (Epstein (1975)). $V(\cdot)$ is homogeneous of degree zero in $(P^N, P^S, Y^*)$ and, therefore, invariant to a unit of measurement. So let $P^N$ be a numéraire, $P = P^S/P^N$ and $Y = Y^*/P^N$. Assume the household exhibits Arrow-Pratt income risk aversion ($V_{YY} < 0$).

The labor allocation decision can be represented as

$$\begin{aligned}
&\underset{L^D, L^H, L^L}{Max}\ \mathrm{E}V(L^L, P, Y)\\
&\text{s.t. } Y = w[L^0 - L^L - L^H - L^D] + PF(L,T) + I\\
&\text{where } V(L^L, P, Y) \equiv \underset{N,S}{Max}\, U(L^L, N, S)\\
&\text{s.t.}: P[S - F(L,T)] + N = w[L^0 - L^L - L^H - L^D] + I
\end{aligned} \tag{2}$$

The first-order necessary conditions for an optimum are

$$\text{w.r.t. hired labor: } \mathrm{E}\{V_Y[PF_{L^D} - w]\} \le 0\ (= 0 \text{ if } L^D > 0) \tag{3}$$

$$\text{w.r.t. household labor: } \mathrm{E}\{V_Y[PF_{L^H} - w]\} \le 0\ (= 0 \text{ if } L^H > 0) \tag{4}$$

$$\text{w.r.t. leisure: } \mathrm{E}\{V_{L^L} - V_Y w\} \le 0\ (= 0 \text{ if } L^L > 0) \tag{5}$$

Assume an interior solution to Eq. (4) since all households in this model have access to some land and will therefore dedicate some labor to their own production [8]. Note that the labor specification in Eq. (1) means that the marginal product of household labor is always at least as great as that of hired labor, an assumption supported in the empirical literature (Taslim (1989), Frisvold (1994)). Therefore, an interior solution to Eq. (4) does not imply an interior solution to Eq. (3). An interior solution to Eq. (5) follows from the Inada condition assumption.

There are two ways to proceed from here. The more direct, conventional method is to derive the comparative statics of the household labor/land ratio with respect to changes in land holdings. Unfortunately, as Chavas and Larson (1994), p.472) establish, 'a number of basic tools of economic analysis are not robust

[8] Formally, $F(\cdot)$, $P$, and $w$ must be such that $PF_{L_H}|_{L_H=0} > w$. Even under rudimentary technology and fixed, suboptimal prices in low-income agriculture, wages are unlikely to be so high as to violate that condition.

under temporal uncertainty and risk aversion,' and this is true in the present case. Rearranging the first-order condition, Eq. (4), then differentiating with respect to land holdings, $T$, yields the following expression.

$$\{V_Y[PF_{L^{\mathrm{H}}} - w]\} = 0$$

$$\frac{w}{F_{L^{\mathrm{H}}}} = \frac{\mathrm{E}\{V_Y P\}}{\mathrm{E}\{V_Y\}}$$

$$\frac{\mathrm{d}(w/F_{L^{\mathrm{H}}})}{\mathrm{d}T} = \frac{\partial \mathrm{E}\{V_Y P\}}{\partial \mathrm{E}\{V_Y\}} \frac{\partial \mathrm{E}\{V_Y\}}{\partial T} \frac{1}{E\{V_Y\}} - \frac{\mathrm{E}\{V_Y P\}}{(\mathrm{E}\{V_Y\})^2} \frac{\partial \mathrm{E}\{V_Y\}}{\partial T}$$

$$= \frac{\partial \mathrm{E}\{V_Y\}}{\partial T} \left[ \frac{\mathrm{E}\{V_Y\}(\partial \mathrm{Cov}(V_Y, P)/\partial \mathrm{E}\{V_Y\}) - \mathrm{Cov}(V_Y, P)}{(\mathrm{E}\{V_Y\})^2} \right] \tag{6}$$

Although the derivative of expected marginal utility of income with respect to land and the covariance term in the numerator of the righthand side expression can be signed, as will be done shortly, the derivative of the covariance of marginal utility of income and price with respect to the expected marginal utility of income is ambiguous. Although it might be tempting to throw up one's hands at this point, the literature on endogenously differentiated class behavior (Roemer (1982)) provides a way forward. An endowment continuum generates an endogenous partitioning of agents into distinct behavioral regimes. Although the allocation of agents across classes is endogenous, we will not attempt to measure those spaces but will instead move directly to characterizing the distinctive behaviors of the different classes [9]. This second path may be less direct or familiar than deriving comparative statics, but it provides superior intuition for the observed behavioral differences across farms of different sizes; moreover, it yields clear results.

Block and Heineke (1973) established that preferences and the nature of risk impact on labor allocation decisions. In that spirit, solving for Eq. (4), one finds that price uncertainty generally yields a gap between the marginal value product of household labor and the wage rate, a gap [10] proportional to the covariance of marginal income and output price.

$$\mathrm{E}\{V_Y PF_{L^{\mathrm{H}}}\} = \mathrm{E}\{V_Y w\}$$

$$\mathrm{E}\{V_Y[PF_{L^{\mathrm{H}}} - \mu F_{L^{\mathrm{H}}}]\} = \mathrm{E}\{V_Y[w - \mu F_{L^{\mathrm{H}}}]\}$$

$$F_{L^{\mathrm{H}}}\mathrm{Cov}(V_Y, P) = \mathrm{E}\{V_Y[w - \mu F_{L^{\mathrm{H}}}]\}$$

[9] The classes that concern us here are clearly non-empty sets, as will be established later.

[10] The wedge between the wage rate and the marginal value product of household labor equals the marginal income-risk premium ($\partial R/\partial Y$) obtained from evaluating the certainty equivalent of Eq. (7), wherein a risk premium, $R$, is subtracted from $Y$ in exchange for the removal of uncertainty.

$$\mathrm{Cov}(V_Y, P) > (<) 0 \Leftrightarrow w > (<) \mu F_{L^H}$$
$$\mu = \mathrm{E}\{P\} \tag{7}$$

Finkelshtain and Chalfant (1991) pointed out that in pure producer theory, Arrow-Pratt income risk aversion ($V_{YY} < 0$) implies Cov($V_Y$, $P$) $< 0$; thus the familiar result that production under price uncertainty is less than under certainty (Sandmo (1971)). In a household model, however, the relation of $V_Y$ to $P$ is ambiguous, varying inversely through the household's production activities and directly through its consumption activities. Thus, if a commodity of uncertain price is more important to a household as a consumption good than as a source of income, it is intuitive that the household's marginal utility of income varies positively with prices and it will rationally overemploy [11] labor. Conversely, if the staple is mainly a source of income, Sandmo's result will obtain; that is, the farm will underemploy labor.

In signing the covariance in Eq. (7), the household's marketable surplus of the staple ($M \equiv F\text{-}S$) and its preferences regarding price risk become crucial. The marginal utility of income can be identified by Roy's Identity. Differentiation then produces the following expression:

$$\mathrm{sign}(\mathrm{Cov}(V_Y, P)) = \mathrm{sign}(V_{YP})$$
$$\text{Roy's Identity yields } V_Y = \frac{V_P}{M}$$
$$\text{thus } V_{YP} = \frac{V_{PP}}{M} - \frac{V_P}{M^2}\frac{\partial M}{\partial P} = \frac{1}{M}\left\{V_{PP} - \frac{\partial M}{\partial P}V_Y\right\} \tag{8}$$

Eq. (8) indicates that we must know something about household preferences regarding price risk, as represented locally by the curvature of indirect utility in prices, $V_{PP}$, in order to sign the covariance of the staples price and the marginal utility of income. If and only if $V_{PP} < 0$ do producers favor stable to variable prices. Quasi-convexity of the variable indirect utility function in prices renders these preferences ambiguous. Turnovsky et al. (1980) and Newbery and Stiglitz (1981) demonstrate in the case of pure consumer theory that the curvature of indirect utility in prices depends on the income and price elasticities of gross demand for staples, the budget share of gross expenditures on the product, and on the household's coefficient of relative risk aversion. Modifying their results to account for the agricultural household model context, one can derive the following expression [12]:

$$V_{PP} = \frac{MV_Y}{P}[\epsilon + \beta(\eta - R)] \tag{9}$$

[11] The term 'overemploy' is used to contrast the present result with that which would obtain under the certainty equivalent counterfactual. The term 'underemploy' will be used in a similar spirit in the remainder of the paper.

[12] A full derivation is presented in Appendix A.

where: $\epsilon$ = the price elasticity of marketed surplus = $((\partial M/\partial P))/((P/M))$; $\beta$ = the budget share of marketed surplus = $(PM)/(Y)$; $R$ = Arrow–Pratt coefficient of relative risk aversion = $-(YV_{YY})/(V_Y)$; $\eta$ = income elasticity of marketed surplus = $((\partial M/\partial Y))/((Y/M))$. The necessary and sufficient condition for $V_{PP} < 0$ is $R > \eta + \epsilon/\beta$ [13].

Reduce the expression on the righthand side of Eq. (9) to readily estimable or observable variables, and one obtains a price risk analog to Pratt's (1964) coefficient of absolute income risk aversion. Let the coefficient of absolute price risk aversion, $A$, be defined as follows:

$$A \equiv -\frac{V_{PP}}{V_Y} = \frac{M}{P}[\beta(R-\eta) - \epsilon] \tag{10}$$

A positive coefficient of absolute price risk aversion, $A > 0$, implies $V_{PP} < 0$, and thus is subject to the same necessary and sufficient condition.

It should be apparent that $\beta$ is the key to price risk aversion. R will almost always exceed $\eta$ for staples, so the sign of A turns on the term $\epsilon/\beta$. If a 'staple' is, in fact, not especially important in either expenditure or revenue terms (i.e. $\beta \approx 0$), then uncertainty surrounding its price is unlikely to concern a household significantly. This is the case for virtually all commodities in the developed world; only a small coalition of specialized producers have much at stake in a particular commodity price and they demonstrate significant price risk aversion. This is the prevailing belief in the extant literature which thus generally finds commodity price stabilization to be welfare reducing (Waugh (1944), Turnovsky et al. (1980), Newbery and Stiglitz (1981), Behrman (1987)). That belief follows directly from the developed world context in which most of the analysis has taken place. No commodity is more than 5% or 10% of American and European consumers' budgets, while most farmers derive less than half their income from any single crop. Newbery and Stiglitz have implicitly recognized the importance of $\beta$ when they make the point that price stabilization is more likely to be beneficial in monoculture than in diversified production systems. Indeed, if the crop is the key to the household's earnings ($\beta \to 1$) or is heavily dominant in its diet ($\beta \to -1$), variable prices may impinge seriously on household well-being. Household budget shares for staple commodities have been shown to be quite high in Sub-Saharan Africa, often reaching 60–70% Hazell and Roell (1983), Weber et al. (1988), Budd (1993), Barrett and Dorosh (1996). It should come as little surprise that agents' preferences with respect to commodity price stability can vary across radically different economic environments, i.e., that price risk aversion might exist among poor agrarian populations even though it is generally thought unlikely in wealthier, industrial countries.

[13] A proof of the necessary and sufficient condition is found in Appendix B.

Continuing with the analytics, substituting Eq. (9) into Eq. (8) yields:

$$V_{YP} = \frac{V_Y}{P}[\epsilon + \beta(\eta - R)] - \frac{\partial M}{\partial P}\frac{V_Y}{M} = \frac{V_Y}{P}[\beta(\eta - R)]$$

$$\mathrm{sign}(\mathrm{Cov}(V_Y, P)) = \mathrm{sign}(\beta[\eta - R]) \tag{11}$$

Apparently, whether a staple food matters more to a household as an income source or as a consumption good depends on the household's coefficient of relative income risk aversion (i.e., its elasticity of marginal utility with respect to income), and the budget share and income elasticity of its marketable surplus. This leads directly to Proposition 1 which depends solely on a household's observable marketable surplus.

*Proposition 1*: If S is a normal good, and a household is a net seller, then it will underemploy labor.

*Proof*:

$$\text{normal good, net seller household} \Rightarrow \beta, R > 0, \eta < 0 \Rightarrow \beta(\eta - R) < 0$$
$$\Leftrightarrow \mathrm{Cov}(V_Y, P) < 0 \Leftrightarrow w < \mu F_{L^H} \ \square$$

This is simply a restatement of the well-known Sandmo (1971) result: price uncertainty reduces income risk averse firms' hiring of factors of production, thereby reducing output. By contrast, the rational labor allocation regime obtaining in the case of a net buyer households is one of labor overemployment if the household is price risk averse.

*Proposition 2*: If S is a normal good and a household is a price risk averse net buyer, then it will overemploy labor.

*Proof*:

(i) *normal good, net buyer* $\Rightarrow R, \eta > 0, \epsilon, \beta < 0$

(ii) $A > 0 \Leftrightarrow R > \eta + \epsilon/\beta$

(i) with (ii) $\Rightarrow R > \eta \Rightarrow \beta[\eta\text{-}R] > 0 \Leftrightarrow \mathrm{Cov}(V_Y, P) > 0 \Leftrightarrow w > \mu F_{L^H}$ $\square$

Price-risk aversion is thus a sufficient condition—but not a necessary condition—for the overemployment of labor by net buyer households [14].

Since production is strictly increasing in labor, these propositions together are sufficient for an inverse relationship between farm size and productivity per unit of land cultivated if, (1) marketable surplus is positively related to farm size and (2) some significant subset of net buyers are price risk averse. The next section shows that these conditions are met in at least one contemporary setting in which we find an inverse farm size productivity relationship: rice farming in Madagascar.

[14] The necessary and sufficient condition is that $R > \eta$. I pursue the price risk aversion sufficient condition for the additional intuition it provides with respect to class-dependent preferences over uniform objective risk.

## 4. Empirical evidence of an inverse relationship and price risk aversion in Madagascar

The data employed come from a 1990 survey of 825 rice farming households in 51 villages in Madagascar. Among the data collected were detailed measures of rice consumption and production, along with disaggregated information on income and land holdings. Details on the survey are available in Bernier and Dorosh (1993).

The first point to establish is that there is an inverse relationship between farm size and output per unit cultivated area in Madagascar. Fig. 1 presents a non-parametric regression of output value per are [15] on the logarithm of household land holdings per adult male equivalent [16]. These regressions were computed using a Nadaraya-Watson estimator with an Epanechnikov kernel of bandwidth 0.75 [17]. The variability of the estimator is conveyed by the lighter confidence bands, which have been estimated by bootstrapping [18].

The solid line shows that expected total output per are falls tenfold, from FMG6660/are to FMG600/are, as farm size increases over the range of the land distribution in Madagascar. The dashed line in Fig. 1 similarly presents a pronounced inverse relationship for paddy. The paddy regression curve appears slightly flatter, in general, than that for total agricultural output, indicating that in Madagascar, as in India (Bharadwaj (1974)), some size sensitivity in cropping patterns contributes to the observed aggregate inverse relationship. But cropping pattern differences fail to explain why paddy output per are falls by a factor of nine across the land-holding continuum, so cropping patterns clearly do not suffice in explaining the inverse relationship.

To put the conditioning domain of Fig. 1 in better context, Fig. 2 presents the estimated density of household land holdings per capita, also computed using an Epanechnikov kernel, here with a bandwidth of 0.5. The small size of these farms

[15] 1 are = 100 m$^2$ = 0.01 hectare = 0.025 acres.

[16] This reflects the different nutrient intake requirements and physical labor capacities of household members of different ages and gender. As computed here, children under age 15 received a weight of 0.5, adult women (15 and older) and senior men (65 and older) received a weight of 0.8, and adult men (15–64), a weight of 1.0. The term 'per capita' is henceforth used interchangeably with 'per adult male equivalent'.

[17] In essence, this method generates a smoothed sequence of weighted conditional means, enabling the data to speak for themselves as much as possible. Deaton (1989) first applied these methods to development analysis. Technical details can be found in Silverman (1986), Härdle (1990) or Deaton (1995).

[18] The purpose of bootstrapping is, again, to let the data speak for themselves as much as possible. Some 500 resampled replicates were drawn from the joint distribution of the dependent and independent variables, yielding a heteroskedasticity-consistent bootstrap. Efron (1987), Härdle (1990) and Efron and Tibshirani (1993) provide technical details on this method and its application to non-parametric regression techniques.

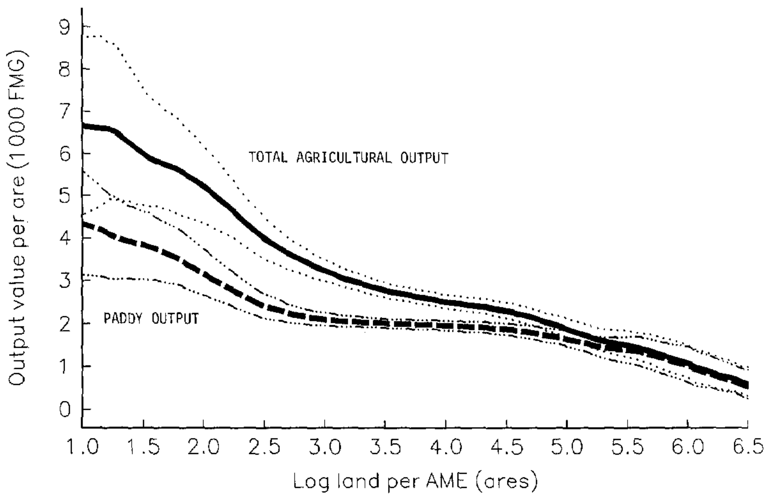


Fig. 1. Inverse relationship in Madagascar (non-parametric regressions with bootstrapped confidence bands).

is striking; the bulk of Malagasy rice farming households fall in the interval 20–90 ares per capita. Indeed, average farm size in the 1984/1985 national agricultural census was but 1.2 hectares (MPARA/FAO (1988)) [19]. The Gini coefficient for land was 0.411 in the sample (0.408 in the census). The small and relatively equal land distribution might tempt one to assume virtual uniformity in behavior and experience among Malagasy rice farming households. But the empirical evidence, such as the inverse relationship evident in Fig. 1, suggests that the agrarian structure, compressed though it is, spans some important class differences. Class-differentiated behavior of this sort is often overlooked in African agriculture.

The analytical literature on price risk (Waugh (1944), Turnovsky et al. (1980), Newbery and Stiglitz (1981), Behrman (1987)) has tended to dismiss concerns about price instability, claiming that $V_{PP} > 0$ almost always obtains. There are two potential pitfalls in those analyses. First, claims are based on parameter estimates

[19] Sample households were randomly selected within 10 purposively selected regional strata (*fivondronana*, or districts) so as to give coverage to regions with both small and large average farm sizes within each of Madagascar's four major agro-ecological zones. As a result, larger farms are overrepresented in the sample, in comparison with the census. Yet, because the non-parametric results and subsample parametric estimates presented below are conditioned on land endowments, the effect of this bias is simply to reduce the number of observations and increase the dispersion of estimates at the lower tail of the conditioning domain, and to increase the number of observations and decrease dispersion in the upper tail. There is no evidence of any sample selection bias within land holding strata.

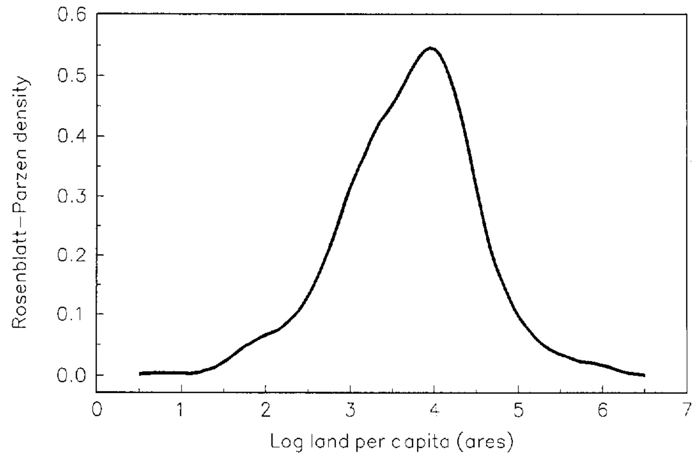


Fig. 2. Land distribution in Madagascar.

from high-income country data. Second, sample mean parameter estimates are used, although these parameters are known to vary with income. Disaggregated empirical investigation using data from Madagascar, consistent with the class disaggregated theoretical analysis of Section 3, suggests that price risk aversion may indeed obtain among important subsets of the agricultural population, thereby ratifying the plausibility of price risk as an explanation of the inverse relationship, as we will now demonstrate.

It is widely recognized that price and income elasticities, budget shares and degree of income risk aversion—the parameters that make up A, the coefficient of absolute price-risk aversion—may vary with income. Because the relationship between income and land holdings tends to be quite strong in low-income agrarian nations, the predictable variation in these parameters carries over to the land distribution. Fig. 3 demonstrates the statistically robust positive (and strongly monotonic) relationship between expected household land holdings per capita and income per capita among Malagasy rice farmers [20].

Not surprisingly, a strong positive relationship also exists between a household marketable rice surplus, $M$, and land holdings, $T$. The sample correlation coefficient between $M$ and $T$ is 0.372, which has a $t$-statistic of 13.45 and a $p$-value of zero against the $t(823)$ distribution. One can estimate a subsistence endowment of land, $T^*$, such that $F(\bar{L}_H, T^*) = \bar{S}$, where an overbar represents the population mean. In other words, a farm household endowed with $T_0 = T^*$ is expected to

[20] This Naradaya–Watson regression employs an Epanechnikov kernel with a bandwidth of 1.25.

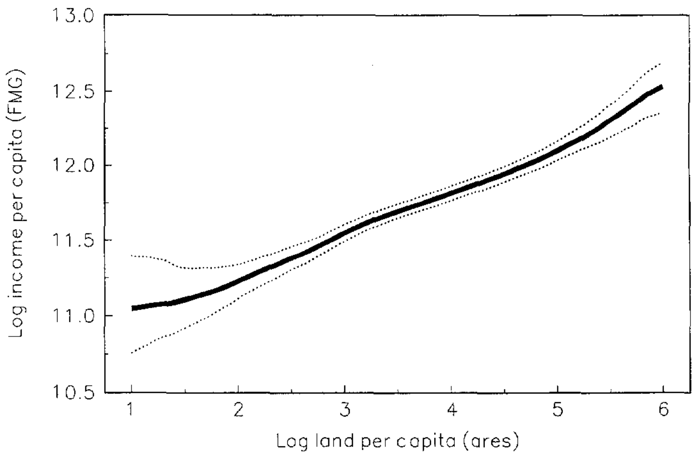


Fig. 3. Land–income relationship in Madagascar (non-parametric regression with bootstrapped confidence bands).

generate a marketable surplus of zero [21]. Differences in relations of exchange thus result systematically, if not deterministically, from different land endowments. Households with land endowments greater than the subsistence endowment ($T > T^*$) tend to be net sellers, while households with endowments $T < T^*$ tend to be net buyers.

Estimation of risk-preference parameters is notoriously difficult and has not yet been successfully performed in a marketable surplus context [22]. That is beyond the scope of the present work, so a simpler approach is taken: estimate a reduced form equation for marketable surplus

$$M_i \equiv F_i - S_i = g(L_i, T_i, C_i, P_i, Y_i) \quad i = 1, \ldots, n_k \tag{12}$$

to obtain parameter estimates ($\beta$, $\eta$, $\epsilon$) from each of k strata. The threshold level of $R$ that would yield $A = 0$ is then computed and $A$ is estimated under alternative assumptions about $R$.

The reduced form marketable surplus equation estimation expands upon the approach taken by Toquero et al. (1975), wherein marketable surplus was presumed not to vary with household income or community fixed effects. While Toquero et al. (1975) used gross agricultural output as an independent variable,

[21] In Madagascar, $T^*$ appears to be about 0.75–0.85 hectares. This range comes from solving for $M = 0$ in econometric estimation of the full sample using the model in Eq. (12), and alternatively, by computation using mean yields, fallow and post-harvest loss rates, household size, subsistence farmer calorie requirements and rice calorie content.

[22] Antle (1987), Chavas and Holt (1993) and Saha et al. (1994) have estimated income risk preferences jointly with farm technology, but consumption preferences have been omitted in these papers. One obstacle is that income-risk preferences are generally derived from utility defined over wealth but consumer preferences follow from utility defined over goods.

that implicitly admits stochastic yield, so production inputs are the regressors here. The dependent variable, $M$, is the difference between recorded rice consumption and production volumes [23]. The total cultivated rice area represented land inputs, $T$, and the sum of household size, measured in adult male equivalents, and hired worker days divided by 200 (to adjust for season length) served as the independent variable for labor inputs, $L$. Household income, $Y$, was a third regressor. Each of these variables ($M$, $T$, $L$, $Y$) was normalized by dividing through by household size, in adult male equivalents. Average household rice unit values [24] provided a household-specific price measure, $P$. Nine dummy variables, $C$, were included to represent the community level fixed effects of the distinct zones within the survey.

Limited observations in the smallest (< 25 ares) and largest (≥ 500 ares) census land holding strata necessitated creating an amalgam of the two smallest (0–49 ares) strata and of the two largest strata (≥ 200 ares). Parameters were then estimated for six different size-differentiated subsets of the survey households. The righthand side of Eq. (12) was specified as translog in the continuous variables to provide a second-order flexible approximation to marketable surplus

$$\begin{aligned}\tilde{M}_i = {} & \alpha_0 + \alpha_1 \ln\tilde{T}_i + \alpha_2 \ln\tilde{L}_i + \alpha_3 \ln\tilde{Y}_i + \alpha_4 \ln P_i + \alpha_5\left(\ln\tilde{T}_i\right)^2 + \alpha_6\left(\ln\tilde{L}_i\right)^2 \\ & + \alpha_7\left(\ln\tilde{Y}_i\right)^2 + \alpha_8\left(\ln P_i\right)^2 + \alpha_9 \ln\tilde{T}_i \ln\tilde{L}_i + \alpha_{10} \ln\tilde{T}_i \ln\tilde{Y}_i + \alpha_{11} \ln\tilde{T}_i \ln P_i \\ & + \alpha_{12} \ln\tilde{L}_i \ln\tilde{Y}_i + \alpha_{13} \ln\tilde{L}_i \ln P_i + \alpha_{14} \ln\tilde{Y}_i \ln P_i + \sum_{j=1}^{9} \gamma_j C_{ij} + e_i \qquad (13)\end{aligned}$$

where a tilde (~) indicates a normalized variable and $e_i$ is a stochastic error with mean zero.

The model was estimated by ordinary least squares, which a battery of diagnostic statistics confirmed as satisfactory [25]. The strata-specific estimates for $\eta$

[23] Paddy output was converted to rice using an average yield coefficient of 0.67.

[24] Equal to the sum of gross rice sales receipts and expenditures divided by the sum of gross sales and purchases volumes.

[25] The regressions fit the data reasonably well, especially considering the limited degrees of freedom available. The strata sample sizes, adjusted $R^2$, $F$-statistics of the global hypothesis of all slopes equal zero, and Breusch–Pagan tests for heteroskedasticity were:

| Stratum | $n$ | $\bar{R}^2$ | $F$-statistic | Breusch–Pagan |
|---|---|---|---|---|
| 0–49 ares | 51 | 0.223 | 4.26 | 23.62 |
| 50–74 ares | 85 | 0.406 | 4.47 | 13.27 |
| 75–99 ares | 78 | 0.307 | 2.46 | 30.62 |
| 100–149 ares | 185 | 0.509 | 9.30 | 18.53 |
| 150–199 ares | 134 | 0.666 | 14.93 | 25.03 |
| ≥ 200 ares | 292 | 0.696 | 36.43 | 12.61 |

All the $F$-statistics are significant at the one% level. None of the Breusch–Pagan statistics are significantly different from zero at the 10% level (the null hypothesis is homoskedasticity). Regressions were run in Shazam, version 7.0; full results are available from the author by request.

Table 1
Land strata-specific marketable surplus parameter estimates

| Stratum | Census Proportion | $\eta$ | $\epsilon$ | $R$ for $A=0$ | $\beta$ | $A$ under $R_1$ | $A$ under $R_2$ |
|---|---|---|---|---|---|---|---|
| 0–49 ares | 20.5% | .256 | -.284 | 1.02 | -.369 | 0.08 | 0.17 |
| 50–74 ares | 17.4% | .322 | -.250 | 1.55 | -.203 | 0.01 | 0.03 |
| 75–99 ares | 17.1% | 1.743 | -1.897 | 25.47 | -.080 | -0.33 | -0.33 |
| 100–149 ares | 21.6% | -1.761 | 1.926 | 50.98 | .037 | -0.26 | -0.26 |
| 150–199 ares | 11.5% | -.698 | .240 | 1.90 | .092 | 0.02 | 0.01 |
| ≥ 200 ares | 11.9% | -.603 | .233 | 0.17 | .300 | 0.74 | 0.58 |

and $\epsilon$, calculated at strata means, are reported in Table 1. The rice price elasticity of marketable surplus estimates are similar to those found by Strauss (1984) and Toquero et al. (1975), who used household data from Sierra Leone and the Philippines, respectively. To the best of my knowledge, no other studies have published estimates for the income elasticity of marketable surplus. The variation in $\eta$ across strata are somewhat counterintuitive on the basis of pure consumer theory, since they suggest that the income elasticity of demand for the staple commodity is higher among larger (i.e. wealthier) farms than smaller ones. Most likely this captures the effect of increased income in alleviating working capital constraints to increased output faced by small farmers.

The three rightmost data columns in Table 1 suggest the relevance of price risk as an explanation for the inverse farm size–productivity relationship apparent in the data from Madagascar. As indicated earlier, marketable surplus, ß, varies directly with land holdings. Estimates of $A$ are plainly sensitive to the choice of values for $R$, the coefficient of relative income risk aversion. The threshold values of $R$, that which sets $A=0$ for non-zero marketable surplus, are listed for each stratum in Table 1. Choosing values for $R$, $R_1$, corresponding to existing empirical evidence (Binswanger (1980), Newbery and Stiglitz (1981), Antle (1987), Chavas and Holt (1993), Saha et al. (1994)) and consistent with the hypothesis of increasing relative risk aversion [26], gives the $A$ estimates shown in the next-to-last column. Note that patterns of risk preferences corresponding to constant relative risk aversion, $R_2$, make the smaller farms change only the magnitudes of the price-risk coefficients, not the qualitative results. The smallest (net buyer) farms are quite likely price-risk averse, which Proposition 2 identified as a sufficient condition for overemployment of labor and overproduction of rice, relative to a certainty equivalent [27]. The econometric evidence thus supports the analytical link between price risk and the existence of an inverse farm size–productivity relationship.

[26] $R_1=1.5$ for the 0–49 ares stratum and increases by 0.2 for each stratum, to 2.5 for the largest stratum. $R_2=2.0$, the mean of $R_1$.

[27] Note that $R>\eta$ holds everywhere, so the necessary and sufficient condition identified in footnote 14 to Proposition 2 is also met. The difference in behaviors between net buyers and net sellers is evident.

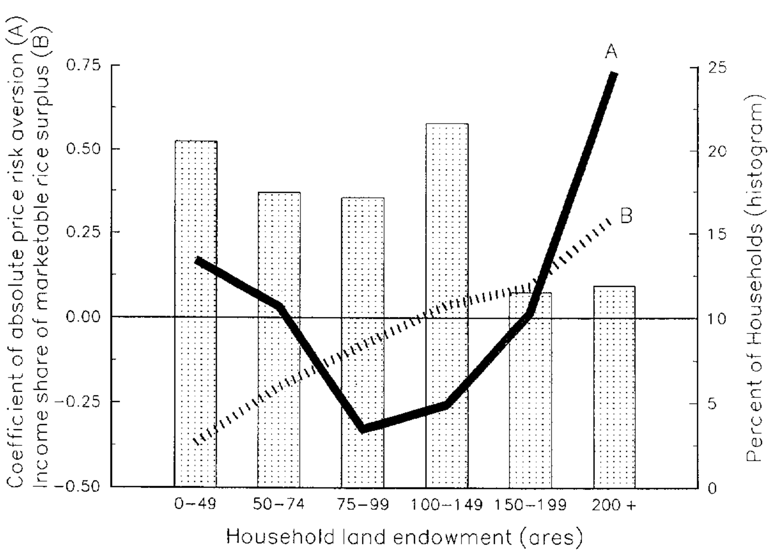


Fig. 4. Marketable surplus and price-risk aversion.

Fig. 4, based on Table 1 and $R_1$, suggests that a farm household's perspective on exogenous changes to the staple commodity price distribution depends on its land endowment. If a household is a net buyer (i.e., $\beta < 0$), an increase in the mean price is equivalent to a real income loss. In contrast, a net seller ($\beta > 0$) gains from an increased mean. Recognizing that the sign of the coefficient of absolute price risk aversion is the same as that of the premium the household would pay to stabilize prices at their mean, it is apparent that the households with land endowments near the subsistence endowment (i.e., the endowment generating $\beta = 0$) favor volatile prices ($A < 0$). Conversely, the dominant mass of households on either tail are price-risk averse ($A > 0$). This subpopulation encompasses those with substantial net purchases or sales, cumulatively 62% of farms by census proportions, as depicted on the righthand Y axis. However, production behavior in the face of temporal price risk is quite different depending on whether they are net buyers or net sellers, as Propositions 1 and 2 demonstrated. Overall, this suggests class-differentiated perspectives on price distributions, and therefore distinct production behaviors sufficient to engender an inverse relationship between farm size and productivity.

## 5. Conclusion and implications

The foregoing analysis demonstrates that differences in households' marketable surplus in an environment of uncertain prices suffices to explain an inverse relationship between farm size and productivity if some small farmers are price-risk averse. While the formulation of the model was general enough to allow for labor market imperfections, nowhere were such imperfections assumed. Neither were any differences admitted in the quality of land endowments, in cropping patterns or in village-level effects. Thus, the present model provides an alternative to existing explanations of the inverse relationship between farm size and productivity per unit cultivated area. This explanation suggests that food security stress placed on food-deficit farmers by staple price uncertainty elicits supranormal labor acitivity. Given ubiquitous price uncertainty and uneven land distribution in low-income agriculture, this is an intuitive explanation for an oft-observed phenomenon.

What are the implications of this explanation for theory and policy? First, for theory it most obviously highlights the importance of incorporating uncertainty into models. Second, and perhaps more profoundly, it emphasizes the importance of recognizing 'structure' less in the form of market rigidities, as does 'structuralist' macroeconomics, and more in the analytical Marxist (and empirically demonstrable) sense of heterogeneous endowments. That objectively identical conditions often beget diametrically opposed responses by equally rational agents is a recognized but underdeveloped dimension of neoclassical theory, despite the pioneering work of John Roemer.

In policy terms, the present results are somewhat ambiguous. Because the relative productivity of small farmers results from their relative food insecurity (i.e. their dependence on markets for food), the productivity effects of land reform would depend upon the ex post endowments of the ex ante class of net food buyers. Modest land transfers to small farmers could improve average yields, while land redistribution that left peasants sufficiently endowed to provide for their own needs could actually reduce agricultural yields. The key to agricultural efficiency is not so much the land distribution as the state of agricultural technologies available across the full range of the distribution.

The observation that greater apparent productivity, as proxied by yields, can result from threatening the food security of net buyer households suggests that improved yields cannot be taken blindly as an indicator of either technology or welfare improvements among smallholders. Intriguingly, domestic staple food price stabilization to reduce the underemployment of labor by larger farmers, combined with modest land redistribution, to take advantage of the stress-induced diligence of land poor peasants, might be the most effective extra-technological means by which to stimulate agricultural productivity. This seems to have been the path followed over the past few decades by several east Asian countries that have enjoyed above-average yield increases.

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## Appendix A. Deriving a measure of local price risk aversion

Take Roy's Identity, modified for marketable surplus, rearrange terms and differentiate with respect to both P and Y. Then substitute expressions and manipulate the resulting terms to obtain familiar parameters, albeit with an adjusted interpretation due to the marketable surplus term.

$$M = \frac{V_P}{V_Y} \Leftrightarrow V_P = V_Y M$$

$$V_{PP} = V_{YP} M + V_Y \frac{\partial M}{\partial P}$$

$$V_{PY} = V_{YY} M + V_Y \frac{\partial M}{\partial Y} = V_{YP} \ (\textit{by symmetry})$$

$$V_{PP} = M\left[ V_{YY} M + V_Y \frac{\partial M}{\partial Y} \right] + V_Y \frac{\partial M}{\partial P}$$

$$V_{PP} = \frac{MV_Y}{P}\left[ \epsilon + \beta(\eta - R) \right]$$

The last expression is (8) above.

## Appendix B. A proof of the necessary and sufficient condition

The necessary and sufficient condition for $V_{PP} < 0$ can be proved as follows.

$$V_{PP} = \frac{MV_Y}{P}\left[ \epsilon + \beta(\eta - R) \right] < 0$$

*Case (a): household is a net buyer* $\Leftrightarrow$ $M,\ \beta < 0$

$$\epsilon + \beta(\eta - R) > 0$$

$$\beta(\eta - R) > -\epsilon$$

$$\eta - R < -\epsilon/\hat{\beta}$$

$$\eta + \epsilon/\beta < R \ \square$$

*Case (b): household is a net seller* $\Leftrightarrow$ $M,\ \beta > 0$

$$\epsilon + \beta(\eta - R) < 0$$

$$\beta(\eta - R) < -\epsilon$$

$$\eta - R < -\epsilon/\beta$$

$$\eta + \epsilon/\beta < R \ \square$$

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