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Rice Prices and Farmers' Welfare in Madagascar: A Nonparametric Analysis

Christopher B. Barrett
Paul A. Dorosh

CORNELL FOOD AND NUTRITION POLICY PROGRAM

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CONTENTS

	LIST OF FIGURES	iv
	ABSTRACT	v
1.	INTRODUCTION	1
2.	MODELING THE WELFARE EFFECTS OF FOOD PRICE INCREASES	3
3.	DATA AND ESTIMATION METHODS	6
4.	ESTIMATION RESULTS	15
5.	CONCLUSIONS	26
	REFERENCES	28

c

LIST OF FIGURES

1	—	Rosenblatt-Parzen Density for per Capita Income of Malagasy Rice Farmers, 1990	10
2	—	Rosenblatt-Parzen Density for per Capita Land Holdings of Malagasy Rice Farmers, 1990	11
3	—	Relationship Between per Capita Income and per Capita Land holdings of Malagasy Rice Farmers, 1990	13
4	—	Distribution of Rice purchases and Sales by Size of Land Holdings of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of C.5)	16
5	—	Distribution of Rice Purchases and Sales by Size of Land Holdings of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of 0.75)	17
6	—	Distribution of Gross Rice Transactions of Malagasy Rice Farmers, 1990	19
7	—	Rosenblatt-Parzen Density for Net Rice Sales and Marketable Surplus of Malagasy Rice Farmers, 1990	20
8	—	Distribution of Marketable Rice Surplus and Net Rice Sales by Income per Capita of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of 0.5)	22
9	—	Distribution of Marketable Rice Surplus and Net Rice Sales by Income per Capita of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of 1.0)	23

ABSTRACT

This paper uses nonparametric density estimation and kernel smoothing techniques to examine the distributional implications of increased and more variable rice prices in Madagascar. While a great many farmers do not participate in product markets as either buyers or sellers, and net sales or marketable surplus are fairly small for many others, the roughly one-third of rice farmers who fall below the poverty line face significant negative first-order welfare effects from increases in the mean or variance of rice prices. Conversely, the gains from rice price increases are highly concentrated among the largest rice farmers in Madagascar.

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1. INTRODUCTION

During the seventies and into the eighties, agricultural pricing policies, parastatal marketing boards, and overvalued exchange rates combined to keep agricultural producer prices low in much of sub-Saharan Africa. Inadequate producer incentives were seen as a major factor responsible for the stagnation of agricultural production and rural incomes.[1] Although other factors such as weak marketing infrastructure and the absence of improved agricultural technologies also were in part to blame for the poor performance of agriculture, raising agricultural producer prices became a major component of economic reform programs in sub-Saharan Africa.

The record of actual implementation of reforms and subsequent results, where reforms have actually been carried out, is mixed. Decreased taxation of export crops have been offset to some extent by contemporaneous adverse movements in world prices, particularly for cocoa, coffee, and tobacco. Liberalization of food crop marketing likewise often has had only minimal effects, either because of ineffective government intervention prior to reforms, as in Niger, or due to policy reversals, as in Kenya. In principle, however, liberalization of food crop marketing combined with exchange rate and trade policy reforms are designed to raise food prices for producers and stimulate agricultural growth and rural incomes.

Changes in agricultural producer prices have major welfare consequences as well, particularly among small farmers, who make up a large share of the poor in sub-Saharan Africa (Dorosh and Sahn 1993). While an increase in cash crops has an unambiguous positive impact on real incomes of households selling these crops, the effects of increases in food prices on the real incomes of small farmers is less clear. Because a substantial proportion of farmers are net food buyers, they may experience real income declines as food prices increase. Consequently, gains in real income from a food price increase may accrue mainly to large farmers producing a marketable surplus (Weber, Staatz, and Holtzman 1988). This is an important point often overlooked in the heated debate over economic adjustment, as both critics and supporters of market-oriented reforms tend to claim that increasing real food prices improves the welfare of poor smallholders.[2]

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[1] See Krueger, Schiff, and Valdés (1988), Schiff and Valdés (1992), or World Bank (1981) for strong articulation and empirical support of this position.

[2] Critics, such as Lipton (1986), Longhurst (1987), and Streeten (1987), lament that increasing food prices do not go far enough in redressing "urban bias," and call for more, not less, government intervention in support of smallholder farming. Supporters, notably the World Bank (1990; 1994), assert that the rural poor "still rely on selling their produce for a significant part

(continued...)

Several recent studies have explored the connections between economic reforms, food price responses, and agricultural growth (Jaeger 1993; Platteau 1993; Barrett and Carter 1994; World Bank 1994), so those issues are not revisited here. Instead, this paper concerns the links between real food prices and farmers' welfare. In particular, we examine the distributional implications of increased and more variable rice prices in Madagascar, a country which liberalized rice marketing, devalued its currency, and lowered rice imports as part of broader stabilization and structural adjustment efforts in the mid-eighties.

Section 2 outlines the theoretical framework for assessing the welfare effects of changes to the mean and variance of food prices. Section 3 describes the data from a survey of rice farmers and the estimation methods. Statistical results are presented in Section 4. The paper concludes with policy implications.

2(...continued)
of their income. So, changes in market prices matter" (World Bank 1994, 167) in efforts to reduce rural poverty

2. MODELING THE WELFARE EFFECTS OF FOOD PRICE INCREASES

Economic theory offers some clear direction on how to model the welfare effects of food price increases. Consider a household that consumes and possibly produces a staple food commodity and participates in other activities within an economy. The indirect utility function, V(p,y) offers a familiar proxy for household welfare or real income, where p is a price vector and y is household income. The welfare effect of changes in the price of the staple commodity, p_s, can be identified by Roy's Identity:[3]

$$\frac{\partial V}{\partial y} m_s = \frac{\partial V}{\partial p_s} \tag{1}$$

where $m_s = F_s - C_s$ is marketable surplus, i.e., output less consumption. So long as the marginal utility of income is positive, the welfare effect of a change in the staple price is of the same sign as the household's marketable surplus of that commodity: positive if it is a net seller, negative if it is a net buyer.

Welfare is also affected by price variability. Economists are familiar with the Arrow-Pratt income risk aversion. This is the assumption that utility is concave in income, implying that income variability reduces household welfare. In parallel fashion one can assess the welfare impact of price variability by examining the curvature of the indirect utility function in prices. The quasi-convexity of indirect utility in prices renders these effects ambiguous. But, following Newbery and Stiglitz (1981) and Turnovsky, Shalit, and Schmitz (1980), the local closed form expression for this curvature is:

$$\frac{\partial^2 V}{\partial p_s^2} = \frac{m_s \frac{\partial V}{\partial y}}{p_s} [\epsilon + \beta (\eta - R)] \tag{2}$$

where ϵ is the price elasticity of marketable surplus, β is the budget share of marketable surplus ($p_s m/y$), η is the income elasticity of marketable surplus, and R is the Arrow-Pratt coefficient of relative risk aversion.

$$R \equiv -y \frac{\partial^2 V}{\partial y^2} / \frac{\partial V}{\partial y} \tag{3}$$

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[3] See, for example, Varian (1989).

Simple algebra indicates that:

$$R > \eta + \frac{\epsilon}{\beta} \tag{4}$$

is a necessary and sufficient condition for:

$$\frac{\partial^2 V}{\partial p_B^2} < 0 \tag{5}$$

From that condition it should be apparent that β is the key to price risk aversion. R will almost always exceed η for a staple,[4] so the welfare effects depend on the term ϵ/β. Assuming the staple is not a Giffen good,[5] these two parameters are of the same sign. Although price elasticities are known to vary with household income (Timmer 1981; Waterfield 1985), the real impact in this quotient comes from variation in β. In particular, as β tends to zero, the necessary and sufficient condition almost surely does not hold, but as β grows in absolute value, price risk aversion becomes increasingly likely. This is an intuitive result. If a commodity is 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 greatly. However, if the crop is a household's primary source of income or nourishment, variable prices may impinge seriously on household well-being.

The first-order welfare effects of changing food prices can thus be summed up in the parameter β, which is a sufficient statistic for the household welfare effects of changing mean price and a close correlate of the effects of changing price variability. These first-order effects plainly ignore, however, the partial equilibrium consequences of food prices on quantities demanded and supplied, as well as the general equilibrium consequences on employment patterns, wages, the prices of other factors and products, technological innovation, and other determinants of welfare. Accounting for these effects requires modeling of markets of major food commodities, as in the multi-market model (Braverman and

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[4] The exception is for households in the neighborhood of pure subsistence, M=0. Since

$$\lim_{M \to 0} \eta = \infty$$

and R is a finite positive number, they will almost surely not be price risk averse.

[5] A Giffen good is a good with a negative income elasticity for which the (positive) income effect of a rise in the price of the good outweighs the negative substitution effect. Thus, an increase in the price of the good leads to an increase in consumption.

Hammer 1986), and ultimately factor markets, household incomes, real exchange rates and trade flows in a computable general equilibrium (CGE) framework. The advantages of these latter approaches in terms of analytical completeness, however, must be weighed against the loss of richness of detail in describing the characteristics of households (for reasons of tractability) and the uncertainties surrounding model parameters and the possible distortions caused by imposition of substantial modeling structure on the problem.

In analyzing the effects of changes in rice prices on farm households in Madagascar, a focus on the instantaneous impacts of changes in prices on household welfare is appropriate for three reasons. First, short-run elasticities of demand and supply for rice are small, so that first-order effects of rice price changes capture much of the story. Second, although rice production is labor intensive, wage employment in rice is not substantial. The next section documents these two claims. Third, rice price changes in Madagascar have been caused mainly by changes in domestic policies on marketing and imports of rice, and less so by changes in macroeconomic policy (e.g., real exchange rate devaluation). These latter two points suggest general equilibrium effects are likely to be small.

3. DATA AND ESTIMATION METHODS

Madagascar has undergone relatively wide-ranging, desperately needed economic reforms since the early 1980s. Among other measures, the government had by 1989/90 substantially devalued the currency, decontrolled crop prices, terminated consumer rice subsidies, and freed domestic and international trade in rice. Real rice prices increased slowly, by only five percent between 1982 and 1985, but have jumped further since the rice distribution network was fully liberalized and major real devaluations began in 1986 (Dorosh, Bernier, and Sarris 1990; Bernier and Dorosh 1993; Barrett 1994b).[6]

The data employed come from a 1990 survey of 825 rice farming households in Madagascar conducted by the Cornell University Food and Nutrition Policy Program (CFNPP) in collaboration with a Malagasy research firm.[7] In the wake of extensive policy reforms but disappointing rice supply response in Madagascar, the survey sought to shed light on farmers' perspectives on the constraints to expanding rice production. Among the data collected were detailed measures of rice consumption, production, purchases, and sales, along with disaggregated information on income and land holdings.

The sample households were randomly selected within ten purposively selected regional strata (*tivondronana*, or districts), selected 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, in comparison to the 1984/85 national agricultural census (MPARA 1988), larger farms are over-represented in the sample.[8] Yet, because the nonparametric results 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 (as reflected in wider confidence bands) at the lower tail of the conditioning domain, and to increase the number of observations and decrease dispersion in the upper tail. In the few instances where global statistics are presented to

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[6] The percentage change in price varies considerably across regions, ranging from -6.9 in the far north to 73.2 in the far west. Different aggregation methods and periods under comparison yield estimated national average rice price increases since 1986 of 8.2 to 19.2 percent (Bernier and Dorosh 1993; Barrett 1994a).

[7] Details on the survey, undertaken with the Office Statistique et Informatique pour le Programmation du Developpement (OSIPD), are available in Bernier and Dorosh (1993).

[8] A χ^2 goodness of fit test of the null hypothesis that the sample data came from the same land distribution as the census yields a test statistic of 557.65, which has a p-value of zero against the $\chi^2(7)$ (indeed, virtually any) distribution.

characterize Madagascar, these are corrected for selection bias by reweighting land strata-specific statistics using census data.

Our analysis focuses on rice, the major food commodity in Madagascar, which accounted for 54 and 52 percent of calorie and protein supply, respectively, in 1979-1981 (FAO 1984) and 52 percent of all cultivated land nationally in 1984/85 (MPARA 1988). In the survey data, the value of rice production accounted for 63 percent of average household income and represented more than three-quarters of total income in 43 percent of survey households.[9] Rice thoroughly dominates food consumption and production in Madagascar, so the distributional effects of changes in rice prices can be powerful, as reflected in the popular aphorism *vary ny politika* ("rice is politics").

Econometric evidence suggests that both rice supply and demand are highly price inelastic in Madagascar in the short run. The price elasticity of supply has been estimated at 0.18 to 0.24 (AIRD 1984), while estimated price elasticities of demand for rice range from -0.26 to -0.35 (Bernier and Dorosh 1993; AIRD 1984). Relatively little wage labor is used in food production, even for a labor-intensive crop like rice. Less than half of the farmers in the survey hired wage labor in the preceding year, and the average for the sample was just 59 worker days per year.

Because our concern is with issues of poverty and distribution, it becomes important to disaggregate analysis. Although the rice farming sector as a whole might benefit from higher rice prices, aggregation might disguise a highly concentrated intrasectoral distribution of the benefits — and the costs — of rice price changes. It is customary in poverty analysis to use per capita income as a conditioning variable. We do so here, with income denominated in Malagasy francs (FMG). We supplement the analysis conditional on per capita income with parallel estimation conditioned on per capita land holdings, measured in ares.[10] Although the poor are generally identified by their position in an income distribution, income is often, in practice, a cumbersome indicator for guiding targeted policy interventions. Because household income comes from multiple sources (both in terms of activities and individuals), collecting reliable data on household incomes is a notoriously difficult and time-consuming process. Moreover, income is inherently stochastic in agricultural communities. Where land is the principal factor constraining agricultural output, as in Madagascar, land endowments are crucial to output and sales outcomes. Land is often substantially easier to measure, particularly in the case of irrigated agriculture, subject to less error in respondent reporting, and nonstochastic,

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[9] Income is here defined as the sum of the net value of agricultural production, sales of non-agricultural goods, salaries, rents, gifts, pensions, transfers, and income of undeclared source. There may be some underreporting of off-farm incomes and the value of non-rice crop production, but these are unlikely to be substantial or, thus, to affect the results materially.

[10] One are = 100 square meters = 0.01 hectares = 0.025 acres.

so it may be a more useful indicator than income in many settings. Finally, we normalize households by an adult male equivalent measure,[11] rather than headcount, and transform both income and land per capita by taking natural logarithms, so as to improve the symmetry of the variables and reduce computational expense by compressing the conditioning domain.

Angus Deaton has recently introduced nonparametric regression techniques to development analysis, especially on matters of agricultural pricing policy (Deaton 1989a, 1989b, forthcoming). The advantage of nonparametric density estimation and kernel smoothing is that they "do justice to the richness of the household survey data, and . . . allow convincing demonstration and presentation of results with a minimum of unnecessary assumption" (Deaton 1989a, p.2). We direct readers elsewhere for detailed treatment of the simple nonparametric methods employed here,[12] but offer the following brief overview.

In nonparametric kernel smoothing, one estimates the conditional expectation of the dependent variable at k distinct points in the conditioning domain. This generates a sequence of k estimates that together present a readily comprehensible graphical representation of a potentially complex high-order polynomial or nondifferentiable function. The manner in which individual observations influence estimation results depends on the weighting scheme chosen. The weighting scheme is itself a function of both the kernel estimator, which assigns weights to observations based on their distance from the conditioning value, and the bandwidth, which establishes the window within which observations receive positive weight. Here we use an Epanechnikov kernel, a parabolic kernel optimal for this sort of problem (Härdle 1990), and a global bandwidth, h, varied as necessary with the application.[13] As a byproduct of this sort of nonparametric regression, the weights generated in computing the kernel provide grist for estimating the empirical density of variables, a descriptive statistic

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[11] 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. Thus, the term "per capita" used henceforth really means per adult male equivalent.

[12] Deaton (1989a, 1989b) and Budd (1993) include appendices with the basic formulae and methodological considerations. These appendices may be good places to start for those seeking a casual understanding. Deaton (forthcoming) offers fuller treatment, while Eubank (1988), Härdle (1990), and Silverman (1986) provide more comprehensive and technical coverage.

[13] More sophisticated, locally-adaptive bandwidth selection methods were not employed because (i) they are computationally more expensive, (ii) they produced trial results essentially indistinguishable from the ones presented here, and (iii) they are not necessarily superior to global smoothing in constructing confidence intervals (Härdle and Bowman 1988).

of considerable value in social welfare analysis (Deaton 1989b). For example, Figures 1 and 2 present the Rosenblatt-Parzen densities for per capita income and land holdings, respectively, computed from an Epanechnikov kernel with a global bandwidth of 0.5.[14] As an indication of the extent of the poverty in Madagascar's farming communities, the modal values of the two distributions are approximately US$70 and 0.6 hectares per capita.[15]

The variability of the estimator, which conveys important information in assessing the robustness of the results for drawing policy implications, is indicated by confidence intervals around the regression estimates. One can estimate a confidence interval using the asymptotic properties of the estimator. Unfortunately, the asymptotic approach does not reflect potential asymmetries in the conditional distributions. Moreover, the choice to employ nonparametric methods is one in favor of letting the data speak for themselves as much as possible, so it seems somewhat peculiar to rely on asymptotic theory rather than the data themselves to establish the variability of the estimates. In light of these concerns, we opted to bootstrap confidence bands. Again, the reader can find detailed explanation of this simple technique elsewhere,[16] but we provide a basic summary.

The bootstrap generates an empirical distribution of a parameter estimate through resampling. A bootstrap replication is a parameter estimate computed from a random sample of size n drawn with replacement from:[17]

$$\{Y_i, X_i\}_{i=1}^{n} \tag{6}$$

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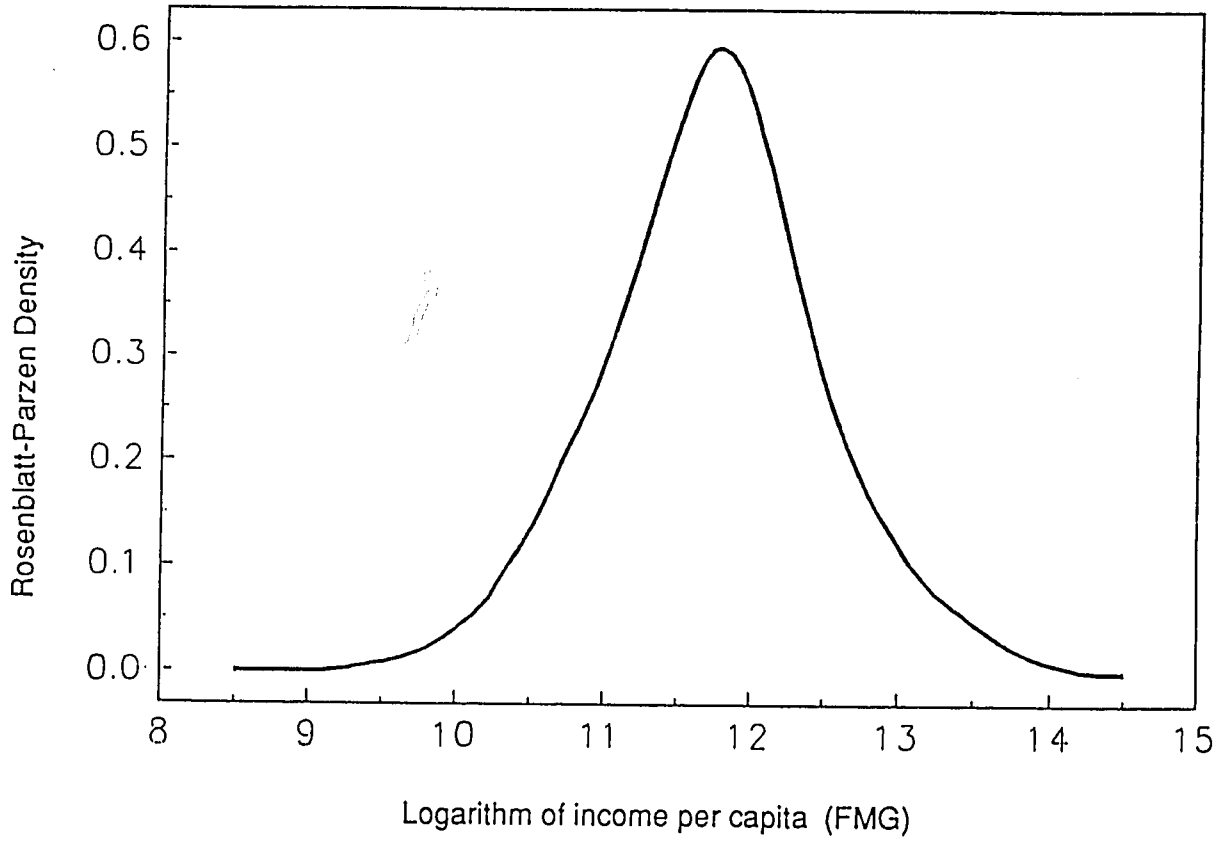
[14] The apparent normality of the log income distribution indicates of course substantial positive skewness in the underlying, untransformed income distribution.

[15] Remember that due to sample selection these figures are upwardly biased estimates of the broader farming sector in Madagascar. The 1984/85 national agricultural census shows modal farm size (per household, not per capita) in the 1.00 to 1.49 hectare class with modal farm household size of 6 to 9 members (MPARA 1988). Mean farm size was 1.20 hectares. Although these are not directly comparable figures, the survey farms are clearly larger than the census standard.

[16] Efron and Tibshirani (1986, 1993), Efron (1987), and Härdle and Bowman (1988) provide reasonably accessible, detailed treatments of the bootstrap and related methods in a nonparametric setting.

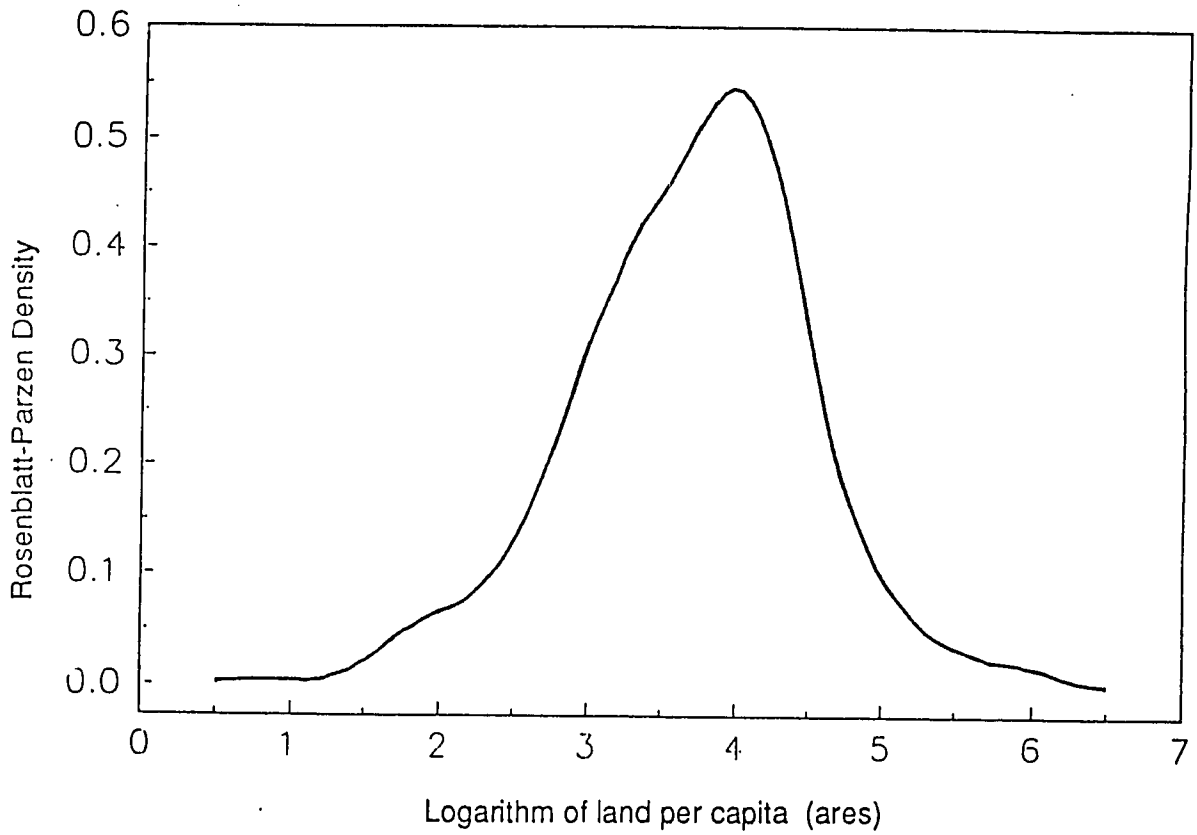
[17] In the present case, we resampled from the joint distribution of the dependent and independent variables $\{Y_i, X_i\}$ to get a heteroskedasticity-consistent bootstrap. If homoskedasticity can be assumed, one can simply resample from the regression residuals.

Figure 1 — Rosenblatt-Parzen Density for per Capita Income of Malagasy Rice Farmers, 1990



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

Figure 2 — Rosenblatt-Parzen Density for per Capita Land Holdings of Malagasy Rice Farmers, 1990



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

A large number, b, of replications is drawn. Here b=500, so each nonparametric regression was reestimated 500 times. In using the percentile method to construct a 95 percent confidence interval, the replications are ordered at each of the k values in the conditioning domain, and the sequence of 2.5 and 97.5 percentiles of the empirical conditional distributions were designated as the bounds of the interval for the regression.[18]

Deaton (1989a) and Budd (1993) employ nonparametric kernel smoothing methods in compelling analyses of the effects of food price changes on real incomes in Thailand and Côte d'Ivoire, respectively. However, those analyses cover countries that the World Bank (1993) labels "lower middle-income economies." Their results are not necessarily transferable to "low-income economies," such as Madagascar's, where per capita GNP (US$210 in 1991) is one-third to one-seventh that of Côte d'Ivoire or Thailand, respectively. Since the welfare of the rural poor in low-income agrarian economies is of greatest concern, it is desirable to extend the existing nonparametric analysis to this setting. Likewise, this nonparametric irvestigation provides a useful complement to published parametric models already applied to low-income agrarian economies in sub-Saharan Africa (Sahn and Sarris 1991; Dorosh and Sahn 1993; Sahn 1994).

Methodologically, we extend the existing nonparametric literature on agricultural price policy in three ways. First, bootstrapping confidence bands to enhance the evaluation of estimation results is an innovation.[19] Second, Deaton (1989a) and Budd (1993) employed slightly different dependent variables; we examine both variables. This permits direct comparison with the results of both earlier studies, but also, by comparing the two different indicators, allows some inference regarding consumption smoothing behavior. Third, we estimate conditional expectations not only over household income per capita, but also over household land holdings per capita. If there is a close correspondence between income and land, the latter might, paradoxically, be a more useful marker for guiding policies targeted at the income poor, for reasons outlined earlier. Figure 3 presents the Nadaraya-Watson regression of log income per capita on log land holdings per capita, using a bandwidth of 1.25. The strong, positively monotonic relationship between these two variables suggests that the land may indeed serve as a reliable proxy for income in Madagascar. It should be noted, however, that the relation is less strong at the lower tail of the land holdings distribution. There are two reasons for this. One is statistical: the densities at these values are low, generating relatively imprecise estimates. The other is that an important subset of the smallest farmers have primary employment in government or trading. Salaries are the single largest source of

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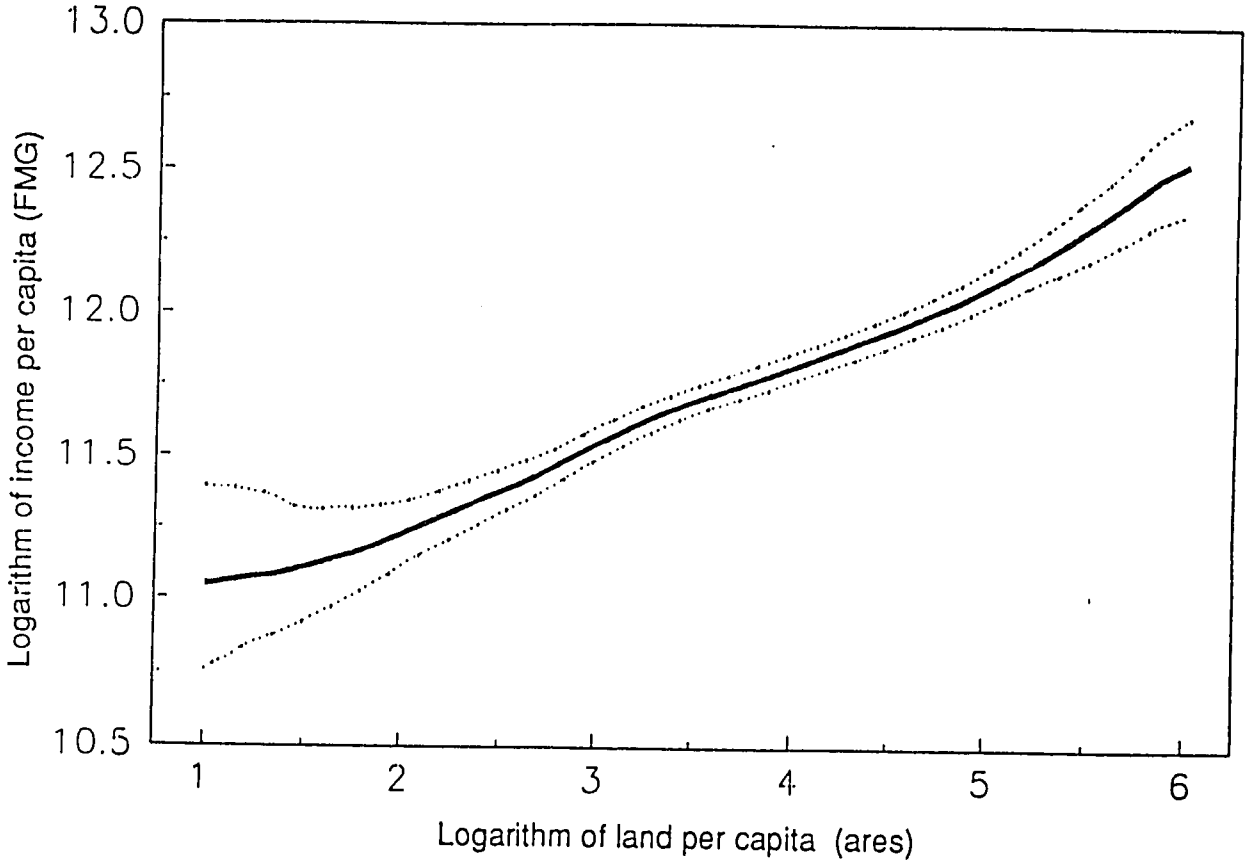
[18] Note that this method neither (i) minimizes the size of the confidence interval in the presence of an asymmetric parameter distribution, nor (ii) necessarily yields an interval centered about the regression.

[19] Deaton (1989a) did not incorporate any measures of dispersion. Budd (1993) estimated confidence bands, but using the asymptotic distribution of the estimator.

Figure 3 — Relationship Between per Capita Income and per Capita Land Holdings of Malagasy Rice Farmers, 1990



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

income only for farmers with small land holdings, and the average incomes of households in the smallest land holding stratum (0-24 ares) is greater than that for any other stratum except the largest (500 or more ares).[20] Thus at the very lower end of the land distribution tail, the income-land relationship breaks down somewhat as farm income-generating activities become more diverse. Nonetheless, for the bulk of the distribution, including most of the lower end where the welfare impacts of rice price changes are of greatest concern, land holdings may be a fair proxy for welfare status, as is commonly true in land-scarce agrarian economies.

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[20] In the 1984/85 agricultural census (MPARA 1988), the smallest stratum comprised 7 percent of agricultural households and the largest stratum comprised 1 percent.

4. ESTIMATION RESULTS

As shown in the analytical discussion of modeling technique in Section 2, the welfare effects of food price changes on a particular household depend very much on that household's behavior with respect to the production, consumption, sales, and purchase of food. It is possible to look within the agricultural sector at particular identifiable subpopulations and establish subgroup-specific behavioral tendencies. This sort of information is crucial to targeting poverty alleviation measures, including any efforts to mitigate inevitable effects of necessary reforms.

The simplest way to characterize behavior is using dichotomous variables taking unit value if the household undertakes an activity and zero if it does not. In Madagascar, the distribution of the dichotomous variables indicating whether a household purchased rice, sold rice, or did neither is quite revealing. Figures 4 and 5, computed with Epanechnikov kernels with a global bandwidth of 0.5 and 0.75, respectively, present the conditional expectation of the fraction of households making gross purchases, gross sales, or neither. We refer to the latter as autarky with respect to rice.

Rice purchases are more broadly distributed across the rice farming population than are rice sales. A household with modal income or land holdings is expected to exhibit gross purchases and gross sales with equal probability. Expressed differently, a greater share of households below the mode make gross purchases than gross sales, and vice versa for households above the mode. Adjusting for sample selection bias, the estimated share of Malagasy rice farming households making gross rice purchases was 60 percent, while the estimated share of households making gross rice sales was only 32 percent.[21] But there are enormous intrasectoral differences in propensities to purchase, sell, or be autarkic in rice.

The conditional expectation is equivalent to the estimated probability of a household endowed with the conditioning income or land holdings engaging in the activity. With this interpretation, the smallest farms purchase rice with probability in excess of 0.9. The probability of making gross purchases is monotonically decreasing in both income and land, while the probability of making sales is increasing in both income and land.

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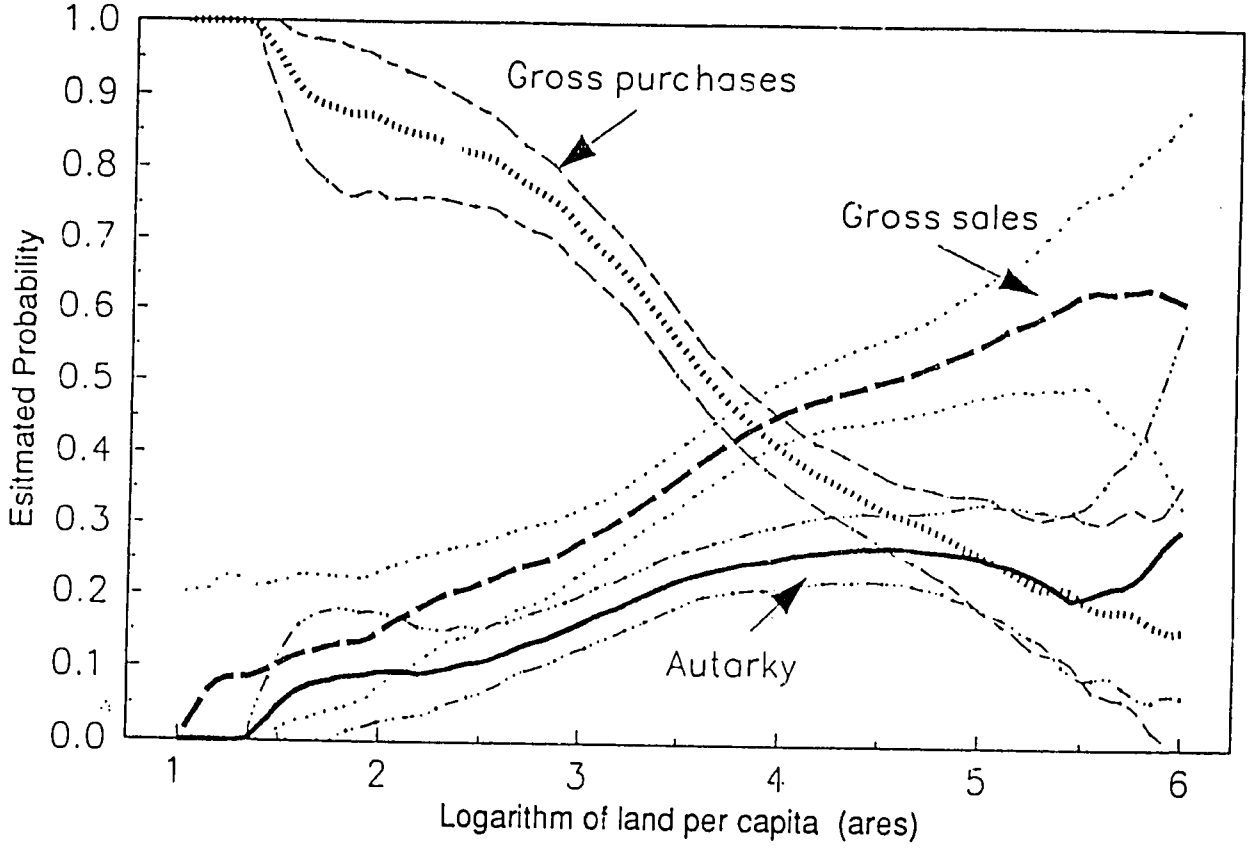
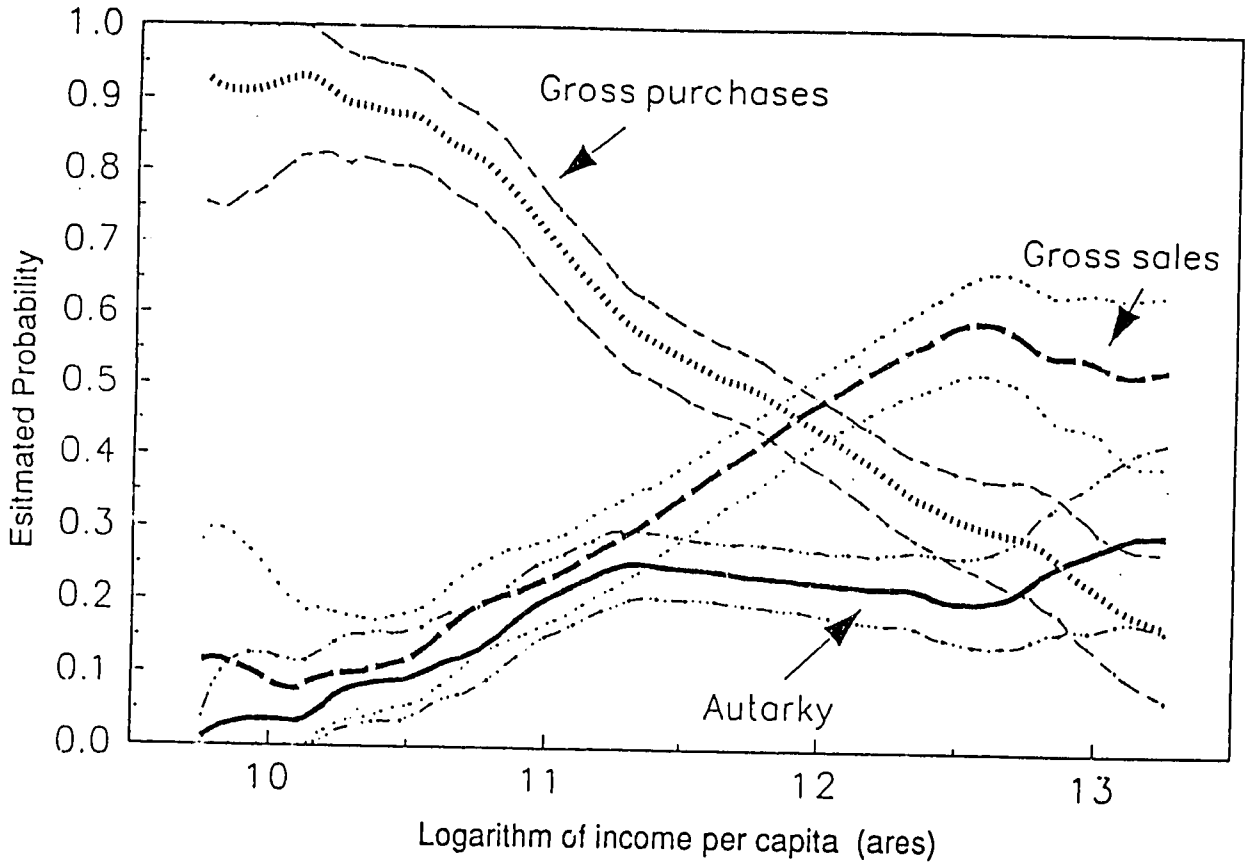
[21] If anything, this difference might be slightly understated because the 1989-90 crop year for which data were collected was marked by relatively good rains and high rice yields. Much of the household data probably include a positive shock to stochastic yields, which likely increases the estimated probability of gross sales and decreases the estimated probability of gross purchases.

Figure 4 — Distribution of Rice Purchases and Sales by Size of Land Holdings of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of 0.5)



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

Figure 5 — Distribution of Rice Purchases and Sales by Size of Land Holdings of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of 0.75)



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

The probability of attaining autarky with respect to rice increases with income and land in the submodal regions and is essentially flat thereafter. This result is interesting in that social scientists usually conceive of small farmers as self-provisioning households liable to retreat from the market. The empirical evidence from Madagascar, by contrast, suggests that medium and larger farmers are more likely to be disengaged from the market for the staple commodity than are small farmers.

Given the behavioral patterns evident in Figures 4 and 5, there is little reason to expect an egalitarian distribution of purchases or sales across farming households. Indeed, these dichotomous patterns are reinforced by the differential magnitudes of sales and purchases volumes across the income and land spectra. Gross rice purchases and sales distributions appear highly unequal among Madagascar's rice farming households, as reflected in the Lorenz curves mapped in Figure 6. The Gini coefficients for both the gross purchases (0.707) and gross sales (0.829), in value terms, are quite high.[22] Just 5 percent of the rice farming households accounted for 50 percent of rice sales by value (and 16 percent of households for 80 percent of sales), a figure consistent with the findings reported for Rwanda, Senegal, Mali, and Zimbabwe in Weber, Staatz, and Holtzman (1988).

The behavioral patterns indicated by the dichotomous variables and the plotted Lorenz curves for gross sales and purchases all suggest that within the rice farming sector there are distinct groups of winners and losers from rice price changes. As Section 2 demonstrated, the sign and magnitude of the proportion of income comprised by marketable rice surplus, ß, provide substantial information about household welfare effects from changes to the rice price distribution. As Deaton (1989b) points out, this ratio is equivalent to the elasticity of real income with respect to rice price. A related, but slightly different measure is net rice sales as a share of income. Examination of how marketable surplus and net sales vary across the rice farming sector, ordered by income or land holdings per capita, conveys substantial information about how income distribution and poverty are affected by exogenous rice price changes.

The difference between marketable surplus and net sales as a share of income is that the former excludes storage and gifts. The Rosenblatt-Parzen densities of these two variables, computed with an Epanechnikov kernel of 0.2 bandwidth, are presented in Figure 7. Note that both variables have a mode at net subsistence, i.e., zero percent of household income. The net rice sales distribution, however, has a far higher modal peak than the marketable rice surplus distribution. This indicates that Malagasy rice farmers manage stocks and gifts so as to neutralize prospective cash gains or losses from stochastic production experience, i.e., they smooth income. An interesting finding that emerges consistently in the upcoming results is that households blessed with

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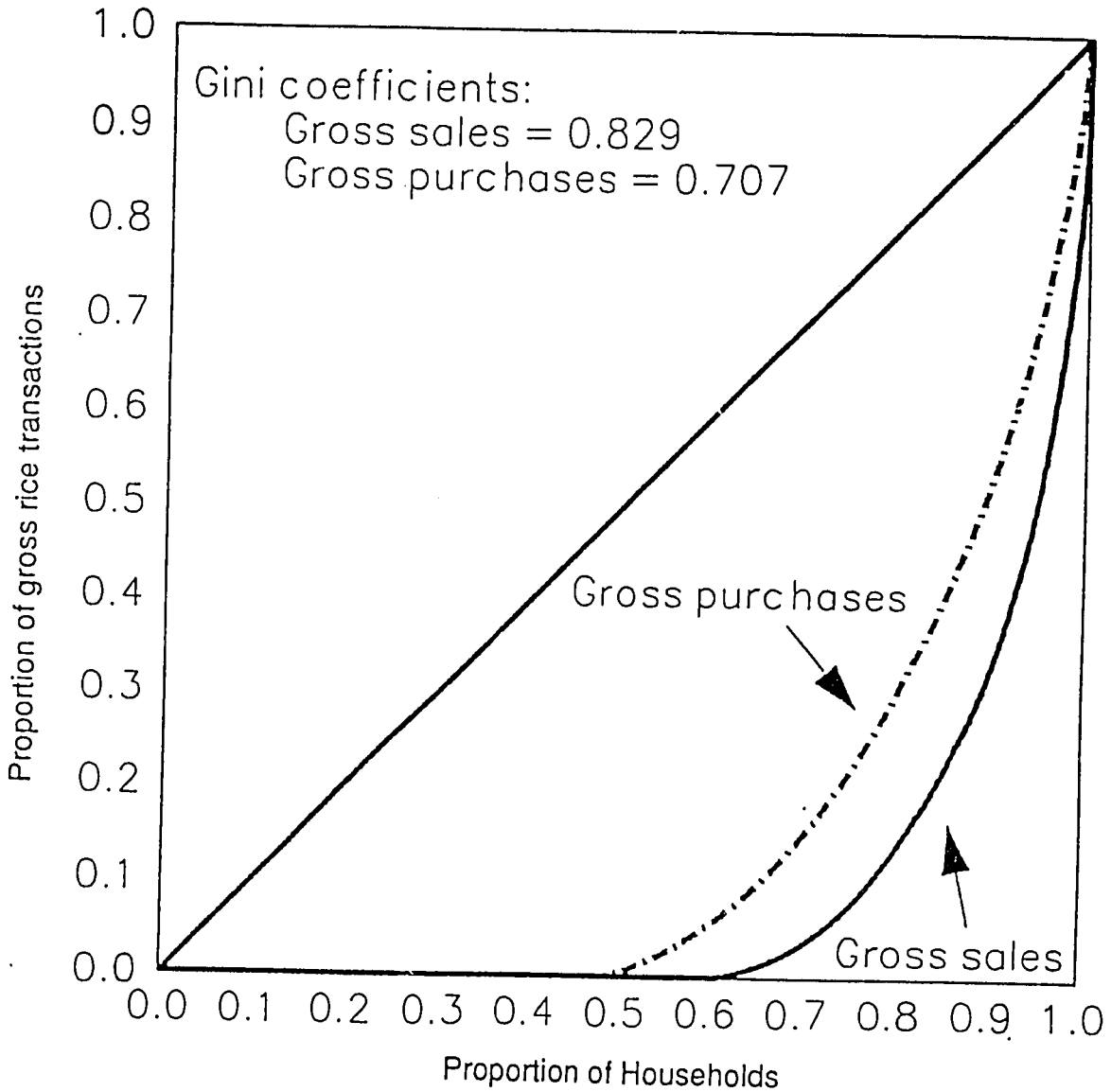
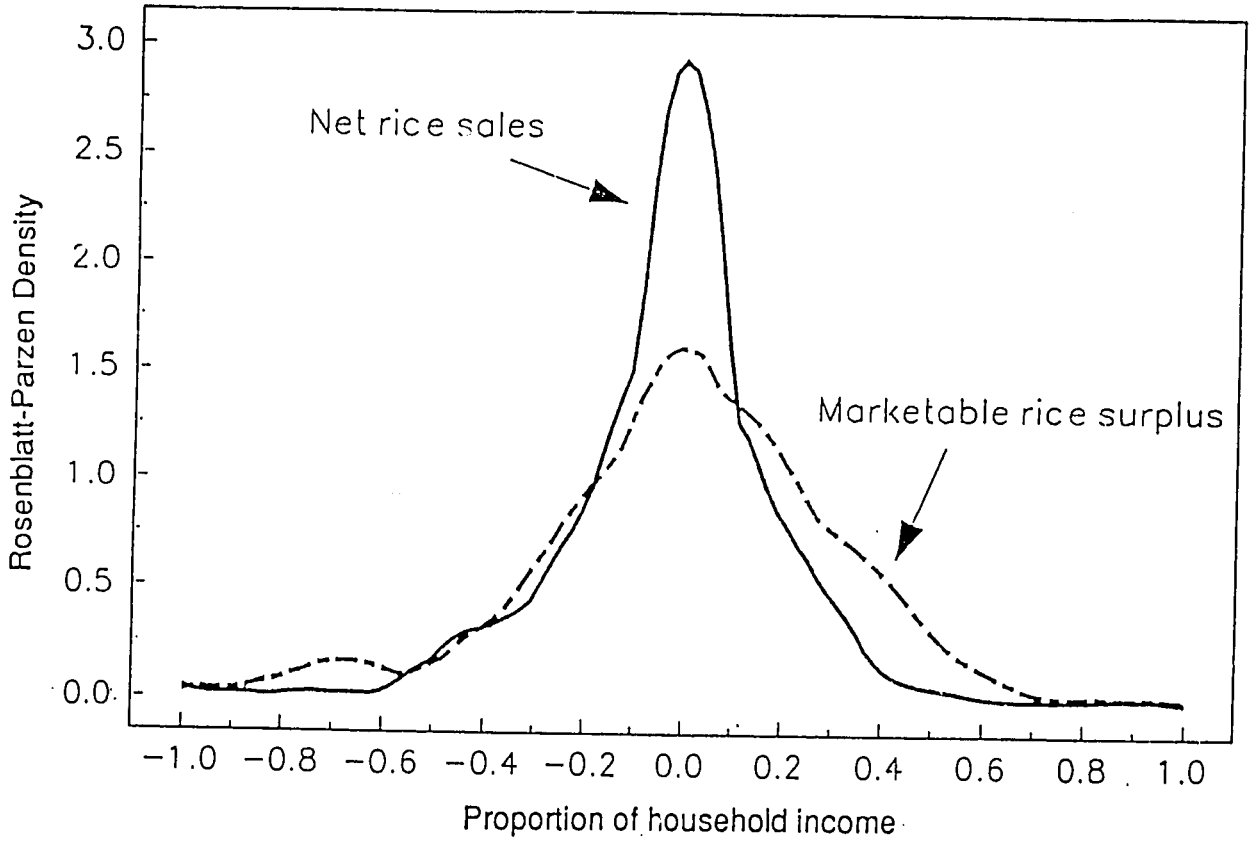
[22] The Gini coefficients for the same variables are virtually identical in volume terms: 0.700 for gross purchases and 0.826 for gross sales.

Figure 6 — Distribution of Gross Rice Transactions of Malagasy Rice Farmers, 1990



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

Figure 7 — Rosenblatt-Parzen Density for Net Rice Sales and Marketable Surplus of Malagasy Rice Farmers, 1990



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

positive marketable rice surplus regularly dedicate some of this surplus to consumption smoothing, as evidenced by the positive and statistically significant margin between marketable rice surplus and net rice sales. Symmetrically, expected net sales are consistently greater (i.e., less negative) than expected marketable surplus for rice deficit households, although that margin is not statistically significant. Furthermore, it is worth noting that the crossover point (i.e., the point at which expected marketable surplus and expected net sales are exactly equal) occurs in the negative share range, not at the (higher) income neutral point. This indicates that rice farming households routinely share or store even from a negative marketable surplus position. Net dissaving or gift receiving is clearly reserved for the most destitute households or adverse shocks.

Figures 8 and 9[23] present the nonparametric regressions of marketable rice surplus and net rice sales on income and land holdings per capita, respectively. Both are clearly upward sloping curves, suggesting that larger, wealthier farmers benefit relatively more than do smaller farmers from rice price increases. This basic feature of the data contrasts somewhat with Deaton's (1989a) findings from Thailand, where net sales are greatest among middle-income households. This is in part due to the different nature of the data sets used; the Thai data are not confined to farmers, so Deaton's (1989a) results mix intersectoral and intrasectoral distributional issues. Rice price increases in Thailand seem to redistribute real income from wealthy non-farmers to middle-income farmers.

These regressions were estimated over a smaller domain than the densities depicted in Figures 1 and 2 because in bootstrapping the confidence intervals it becomes likely that random resampling will generate a zero-valued kernel at some point(s) on the tails of the distribution, and that kernel enters the denominator of the regressor. So Figures 8 and 9 depict only income and land values with some weight to them. This is an important point because the income shares reflected on the graphs are substantial, above all on the negative side. An estimated 63 percent of rice farming households had marketable rice deficits in 1990, a relatively good year in terms of rains and yields.[24] Moreover, while the modal households had deficits or surpluses quite close to zero, suggesting relative insulation from welfare shocks induced by rice price changes, the lower one-third of each distribution was exposed to rice price changes of 20 percent or more of estimated income. Although it is likely that incomes deriving from the value of auto-consumption of production of non-rice crops are slightly underreported, the data strongly suggest that the poorest rice farmers are quite vulnerable to an increase in the price of their principal crop. An estimated 37

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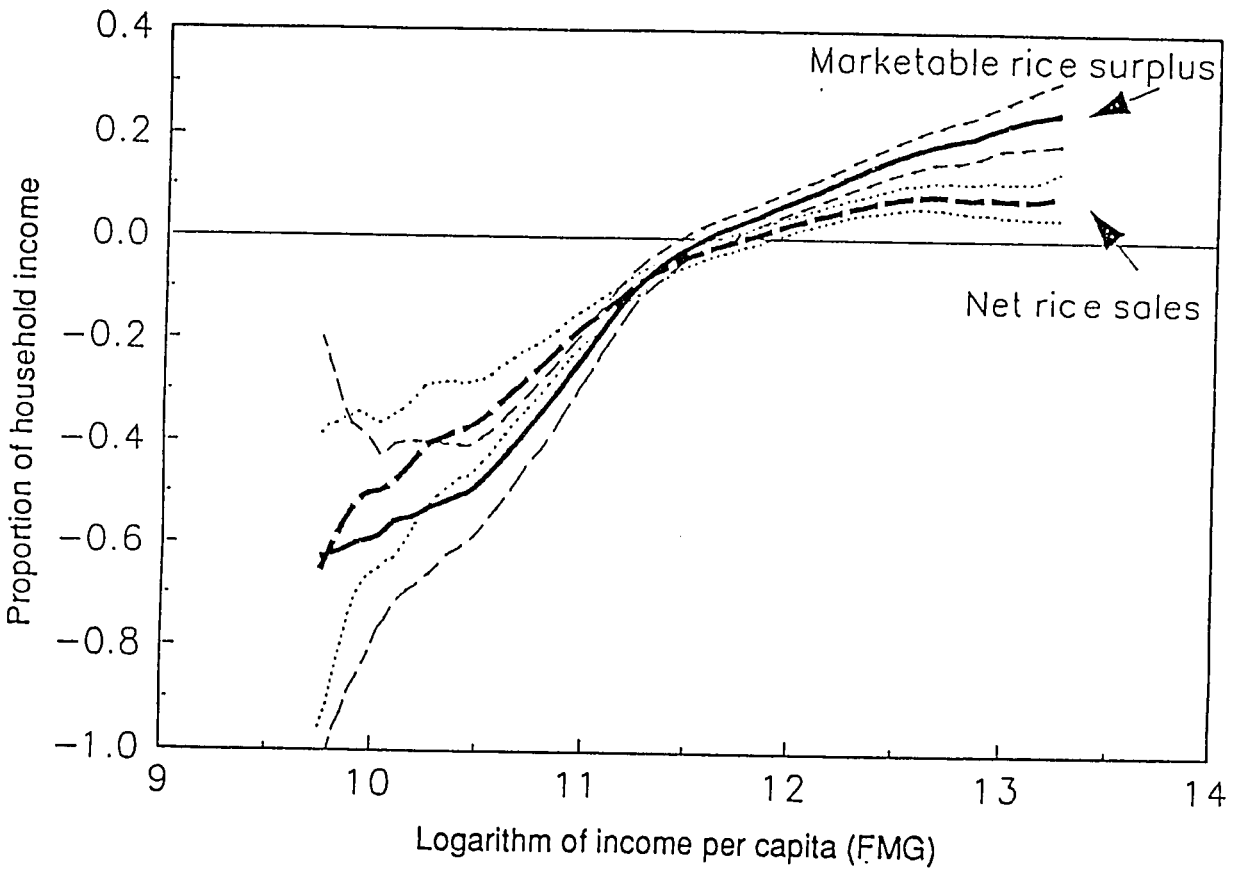
[23] These Nadaraya-Watson nonparametric regressions were computed with an Epanechnikov kernel with global bandwidth equal to 0.5 and 1.0 for Figures 8 and 9, respectively.

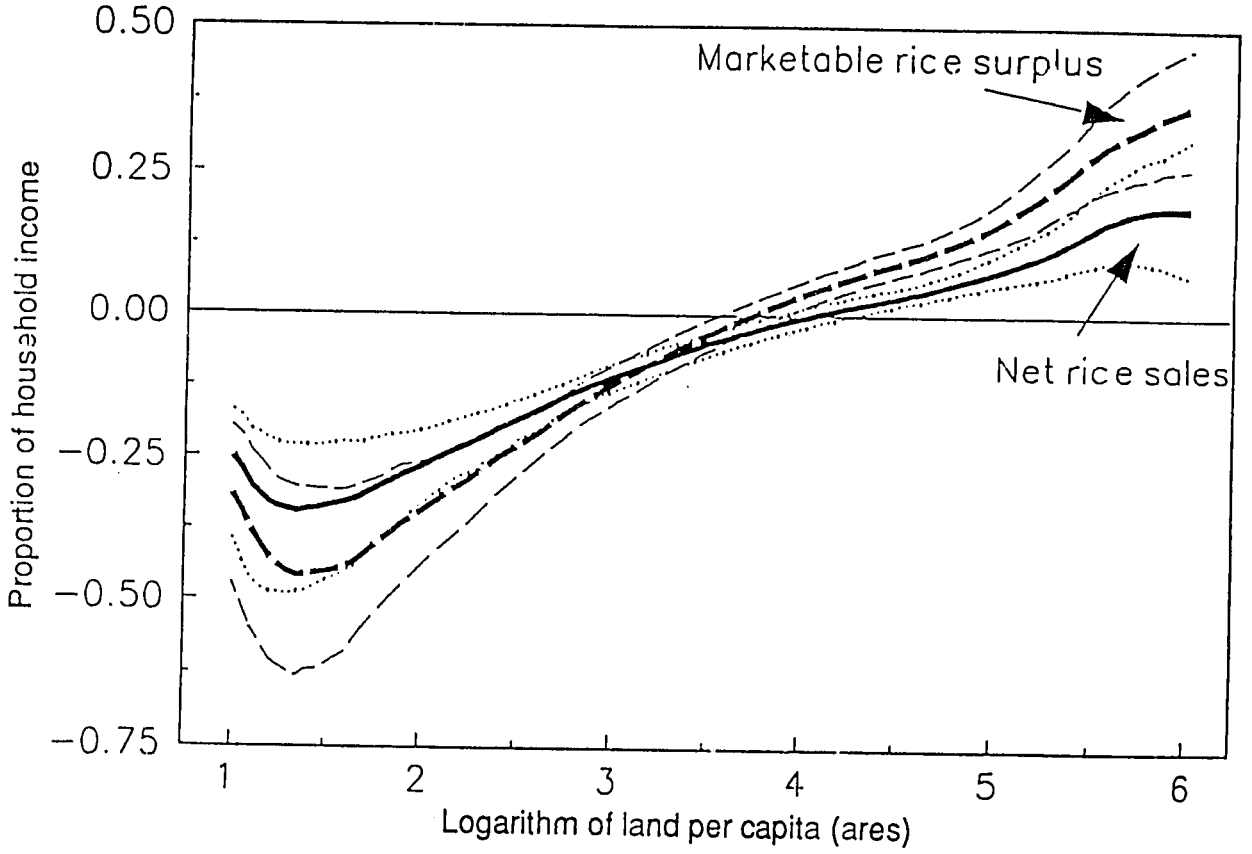
[24] The sample statistic is 49 percent.

Figure 8 — Distribution of Marketable Rice Surplus and Net Rice Sales by Income per Capita of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of 0.5)



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

Figure 9 — Distribution of Marketable Rice Surplus and Net Rice Sales by Income per Capita of Malagasy Rice Farmers, 1990 (Epanechnikov kernel with global bandwidth of 1.0)



Source: 1990 Cornell Survey of Rice Farmers in Madagascar and authors' calculations.

percent of rural households are already below the poverty line,[25] and virtually all rural Malagasy households produce rice, so the biggest losers from rice price rises are likely poor, small farmers. Conversely, the largest, wealthiest 10 or so percent of farmers stand to benefit significantly from rice price increases. The confidence intervals indicate that these inferences regarding the markedly different prospects for poor and rich rice farmers are statistically robust. These results differ considerably from Budd's (1993) findings from Côte d'Ivoire. There, no single food crop, nor an amalgam of food crops, displays estimated deficits greater than 5 percent of expenditures, and "the majority of the distribution is closer to 1% or 2%, with the confidence interval encompassing zero. This suggests not only that the differences across welfare levels are small but that the magnitude of any effect would also be quite small" (Budd 1993, 595). In Madagascar, neither the differences nor the magnitudes are small.

The magnitudes of the budget shares of marketable rice surplus among the poorest farmers suggest that rice price risk aversion is quite plausible. Moreover, because the poorest farmers are net rice buyers, price risk amplifies rather than mitigates yield risk.[26] If a net buyer household experiences an adverse production shock that is felt widely, dampening aggregate supply, its real income will be reduced not just by lower output but also by the higher price of its primary consumption item. This price effect might provide some insight on the political economy motivations behind the government's creation of a rice buffer stock facility in 1986, after rice prices skyrocketed shortly after the freeing of distribution channels and prices and a change in government rice imports management. The principal constituencies for which the socialist government competed with opposition forces were the urban poor and small farmers. The former, like poor farmers, spend a substantial (> 0.5) share of their income on rice, making them excellent candidates for price risk aversion. The latter, even if they are not price risk averse in the sense presented in Section 2, face heightened income and consumption instability from more variable rice prices due to their status as net rice buyers. The language used in announcing and

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[25] This estimate comes from Dorosh, Bernier and Sarris (1990), which also estimates that the rural poor comprise 88.2 percent of Madagascar's poor households.

[26] Newbery and Stiglitz (1981) make the case that it is not price risk per se, but income or consumption risk that matters. This point is well taken, although income risk is explicitly accounted for in the present context through the indirect utility function formulation. Newbery and Stiglitz (1981) then note that if a producer is subject to yield risk, then price risk is in fact desirable if the producer faces a downward sloping demand curve. In the case of smallholders this most likely occurs due to production shocks that are highly covariate among a subpopulation of farmers large enough to affect market supply nontrivially.

establishing the buffer stock suggests an appeal to these two groups, not a play to larger farmers.[27]

[27] See Barrett (1994a) for more detail on this episode.

5. CONCLUSIONS

Increased real food prices, an important component of policy reforms in many developing countries, have significant impacts on farmers' welfare. Though real food price increases raise gross incomes of most farmers, since most make gross commodity sales, small farm households that are net purchasers of food may suffer substantial declines in welfare. The food price dilemma (Timmer, Falcon, and Pearson 1983), reflected in the choice between high food prices to spur production and employment versus low food prices to stimulate consumption by the poor, thus has an acute intra-rural dimension, along with the more obvious urban-rural conflict.

In the case of Madagascar, a great many farmers do not participate in product markets as either sellers or buyers, and for many others net sales or marketable surplus are fairly small. Nonetheless, the roughly one-third of rice farmers who fall below the poverty line have substantial net purchases of rice, suggesting important negative first-order effects of increases in rice prices on household welfare. Greater variability in rice prices induced by economic reforms (Barrett 1994b) likely imposed additional welfare losses by threatening household food security and destabilizing incomes. Of course, a number of factors have influenced changes in rice prices in Madagascar in the period analyzed. Weather-induced production shocks, local market imperfections, and policy reforms (such as changes in the level of government-controlled rice imports and exchange rate devaluations): all have affected rice prices and household welfare. Moreover, the net effects of policy reforms on households include the impacts through non-rice activities as well, and for some households these may be positive, e.g., as for the effects of trade and exchange rate liberalization on rice-deficit clove and coffee farmers on Madagascar's east coast (Dorosh 1994). Nonetheless, the dominance of rice in Malagasy agricultural households' consumption and production patterns, as documented in Section 3, suggests a need for cautious and disaggregated analysis of the effects of exogenous shocks and rice policy on household welfare. With international donors pressing for further market-oriented reforms and devaluation of the FMG, and a new government anxious to revitalize a seemingly stagnant agricultural sector, the results here suggest that considerable attention may need to be given to complementary interventions to cushion small farmers against further changes in rice prices.

These results stand in contrast to those obtained in methodologically similar studies of Thailand and Côte d'Ivoire by Deaton (1989a) and Budd (1993), respectively, where changes in food prices had minimal impacts on the rural poor. Madagascar differs because the major food crop, rice, accounts for a much larger share of income and food expenditures than in these wealthier and more agriculturally diverse economies.

The disaggregated analysis possible with the nonparametric methods used here highlights the potentially important effects of policy on vulnerable rural households. Where demand and supply elasticities are low, general equilibrium effects are likely small, and sizeable household surveys of sufficient quality are available, these techniques may be a useful complement to more complex modeling exercises that rely on parametric methods and more broadly defined household groups.

Finally, the findings do not make a case against either economic reforms or, necessarily, food price increases. Distortions in price signals have important costs in terms of reduced allocative efficiency and lower economic growth. Complementary policy measures may be needed, however, especially in low-income countries where food consumption and production are heavily concentrated on one commodity, to ensure that food price increases designed to spur agricultural growth and rural incomes do not have significant adverse effects on food-deficit, small, poor farmers.

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