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Growth and Poverty in Rural India

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Higher agricultural yields reduced absolute poverty in rural India, both by raising smallholder productivity and by increasing real agricultural wages. But gains to the poor were far smaller in the short run than in the long run.

Background paper for *World Development Report 1995*

Summary findings

Unlike most developing countries, consistent poverty measures for India can be tracked over a long time. Ravallion and Datt used 20 household surveys for rural India for the years 1958–90 to measure the effects of agricultural growth on rural poverty and on the rural labor market and to find out how long it takes for the effects to be felt.

They found that measures of *absolute* rural poverty responded elastically to changes in mean consumption. But agricultural growth had no discernible impact — either positive or negative — on the share of total consumption going to the poor.

For the rural poor, Ravallion and Datt attribute the long-run gains from growth to higher average farm yields, which benefited poor people both directly and through higher real agricultural wages. And the benefits from higher yields were not confined to those near the poverty line — the poorest also benefited.

The process through which India's rural poor participate in the gains from agricultural growth takes time, although about half of the long-run impact comes within three years.

The long-run elasticity of the head-count index to farm yield was over 2 — of which 40 percent came through wages. Short-run elasticities were far smaller.

Inflation adversely affected the rural poor by eroding their real wages in the short run.

Growth and Poverty in Rural India

Martin Ravallion and Gaurav Datt[*]

1 Introduction

The scope for economic growth to make a real difference in the lives of the developing world's poor has been the subject of (often vociferous) debates in both academic and policy circles.[1] Measuring the long-run benefits to poor people from economy-wide changes calls for a time series of representative household-level surveys; yet such surveys are sporadic at best for most countries. India is an exception. There one can trace distributional impacts over a long period using reasonably comparable and nationally representative surveys of consumption.[2]

Here we use these surveys to examine how much India's rural poor have benefited from agricultural growth, what role the labor market has played, and whether the impacts were distributionally biased one way or another. We depart from past analyses for India and elsewhere in five main ways:

i) _The identification of distributional impacts_. It is theoretically possible for a growth process to have sufficiently adverse effects on inequality that poverty increases, and some have argued that this is also the reality of India's rural development. Another view—often termed "trickle down"—denies this, but still allows that distribution may worsen even though on balance the poor gain somewhat. By contrast, some growth processes can entail favorable distributional

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[1]   Recent surveys spanning the wide range of views concerning the impacts of growth in farm productivity on rural poverty include Saith (1990), Singh (1990) and Lipton and Ravallion (1994). Differences between countries or time periods may account for some of these differences, but certainly not all. Contrast, for example, Ahluwalia's (1978, p.320) conclusion that "..there is evidence of some trickle down associated with agricultural growth" with Saith's (1981, p.205) claim that "there can be little doubt that current growth processes have served as generators of poverty"; both were using data for the same country (India) over roughly the same period (1957-73).

[2]   It seems that not even for the U.S.A. can one track poverty measures as well. The were 15 Consumer Expenditure Surveys for the U.S.A. over the same period for which we have 20 National Sample Surveys for India.

shifts, and consequently larger gains to the poor than a "trickle down". Here we propose a method for identifying and testing the distributional impacts of growth on poverty.

ii) The role played by the labor market. Past work has often ignored or down-played the rural labor market's capacity to transmit the benefits of technical progress to the poor. Marked differences in emphasis can be found in policy-oriented discussions on this point.[3] Here we aim to quantify the role played by real wages in distributing the gains from aggregate growth.

iii) Allowing for dynamic effects. The dynamics of the distributional impacts of growth are of obvious interest, though the topic has received surprisingly little attention. Past models have analyzed the consumption-based poverty measures within a static framework, despite theory and evidence to the contrary. Stickiness in the adjustment of poverty measures also has an important implication: long-run responses can far exceed those in the short-run.

iv) The time period of analysis. Much of the scholarly debate for India has focused (often—though not always—for lack of data) on periods of rather little growth; these data may have low power in testing the effects of growth on poverty (Srinivasan, 1985). We use a new data set embracing a period (since the mid-1970s) of higher agricultural growth.

v) The treatment of survey spacing. Unlike past work, we deal consistently with the uneven survey spacing in the estimation. This can matter to estimating dynamic effects.

The following section outlines the model we will be using. Section 3 then presents our results, while conclusions can be found in section 4. Appendix 1 describes our data and its sources, while Appendix 2 discusses related work in the literature on rural poverty in India.

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[3] Contrast, for example, the IFAD (1992) report with World Bank (1990). The former emphasizes the scope for reducing rural poverty by developing smallholder agriculture, and pays little attention to the role of the unskilled labor market; by contrast the World Bank report emphasizes the importance of positive employment and wage effects in achieving pro-poor growth (though not to the exclusion of other channels).

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2   Modelling impacts of growth on poverty

2.1   Characterizing alternative growth processes

A poverty measure (P) can be written as a non-increasing function of the mean (μ), and a vector of parameters $\pi = (\pi_1,...,\pi_n)$ for the Lorenz curve:[4]

$$P = P(\mu, \pi) \tag{1}$$

Let the Lorenz parameters also vary with the mean $\pi(\mu)$ (other variables are ignored for now) and consider the effect of an increase in the mean.[5] Assume that all these functions are differentiable, and let subscripts denote partial derivatives. We can distinguish three cases:

i) <u>Immiserizing growth</u>; $dP/d\mu = P_\mu + \sum_{i=1}^{n} P_{\pi_i}\pi_{i\mu} > 0$. This requires that $\sum P_{\pi_i}\pi_{i\mu}$ is sufficiently positive to outweigh the direct impact of growth (since $P_\mu \leq 0$).

ii) "<u>Trickle down</u>"[6]; $dP/d\mu < 0$ but $\sum P_{\pi_i}\pi_{i\mu} \geq 0$. While the distributional shifts do not favor the poor, the growth effect is still strong enough that poverty falls. We call the special case in which $\sum P_{\pi_i}\pi_{i\mu} = 0$ "distribution neutral growth".

iii) <u>Redistribution with growth</u>; $\sum P_{\pi_i}\pi_{i\mu} < 0$. Here the redistribution is also pro-poor.

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[4]   The mean is taken to be normalized by the poverty line. For explicit formulae for these relationships in the case of the poverty measures used here see Datt and Ravallion (1992).

[5]   On the effects of a change in the mean on a poverty measure holding the Lorenz curve constant see Kakwani (1993). Here we allow the Lorenz curve to also vary.

[6]   This term is sometimes used in any situation in which the poor share in economic growth. But if <u>all</u> the growth was from gains to poor people, this is surely not "trickle down", which suggests a limited gain to the poor from a growth process substantially involving nonpoor people.

How can one distinguish these cases empirically? While the redistribution effects can be difficult to disentangle, we can readily determine the sign of $\Sigma P_{\pi_i} \pi_{i\mu}$ by constructing the simulated poverty measures:

$$P^* = P(\bar{\mu}, \pi(\mu)) \qquad (2)$$

for fixed $\bar{\mu}$ but using the actual Lorenz curve; the poverty measures are thus purged of the direct effect of growth, leaving only the effect via changes in the Lorenz curve. We then examine the relationship between P^* and μ, so as to distinguish the three cases above (noting that $dP^*/d\mu = \Sigma P_{\pi_i} \pi_{i\mu}$). To estimate P^* one multiplies all consumptions by $\bar{\mu}/\mu$ (thus preserving the actual Lorenz curve) and then calculates the poverty measure on the scaled distribution.

What is P^*? It is not an inequality measure as such,[7] but a measure of how inequality matters to the poor. One can interpret P^* as a measure of "relative poverty" in which the poverty line is set as a fixed proportion of the mean, as distinct from the "absolute poverty measure" $P(\mu_t, \pi_t)$.[8] But P^* does not have much appeal as a poverty measure in its own right (since it is unaffected by distribution-neutral changes, even when they entail substantial gains or losses to poor people); rather it is an analytic construct to help understand the distributional effects of growth. That is the way we will use it in the following analysis.

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[7]   For some possible poverty measures, it will not satisfy the Pigou-Dalton transfer axiom.

[8]   For discussion and references on this distinction, and the properties of both types of measures, see Ravallion (1994).

2.2 An econometric model

A time series of individual consumption is likely to be sticky, due to smoothing behavior. The poor are widely thought to be less well insured than others, yet it is also clear that they often self-insure (Alderman and Paxson, 1992; Deaton, 1992); the cost of not doing so can be prohibitively high. While there is evidence against the Permanent Income Hypothesis in rural India (Bhargava and Ravallion, 1993), consumption is clearly smoother than income (Walker and Ryan, 1990).

However, the way serial correlation in individual consumption is translated into the dynamic behavior of a time series of poverty measures (based on cross-sections of consumption) is likely to be complex.[9] This will also depend on the covariances across individuals. Suppose that there is no smoothing behavior at the individual level, in that current consumption is simply current income, but that incomes are negatively correlated across individuals; those that escape poverty are replaced by others. Then one could find serial dependence in the poverty measures even though individual consumptions fluctuate wildly. While we are skeptical about the prospects of disentangling these sources of dynamics in aggregate poverty measures, one cannot rule out the possibility that their adjustment to changes in current variables is far from instantaneous. Thus the econometric specification must allow for dynamic effects.

The choice of poverty measures is also an issue. It has been argued that the bulk of gains amongst the poor go to those near the poverty line (Lipton, 1983). To test this, one needs measures of poverty which better reflect its depth and severity than the popular "headcount index" given by the proportion of people deemed poor.

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[9] We shall not go into the issue of whether poverty should be measured in terms of consumption or income, for we have no choice in this context since India's NSS does not include income. For further discussion of this issue from various perspectives see Atkinson (1991), Slesnick (1993), and Ravallion (1994).

Combining these considerations, a simple test of whether the poor shared in growth is to regress the poverty measure against the mean of the distribution. On allowing for dynamic effects we estimate:

$$\ln P_{\alpha t} = \pi_{\alpha 0} + \pi_{\alpha 1}\ln P_{\alpha t-1} + \pi_{\alpha 2}\ln \mu_t + v_{\alpha t} \qquad (3)$$

where $P_{\alpha t}$ is the Foster-Greer-Thorbecke (1984) measure for date t (described below), and μ_t is mean consumption at date t. A significant (negative) estimate of $\pi_{\alpha 2}$ could reflect Lorenz curve shifts either for or against the poor. So, in addition to (3), we estimate the regression:

$$\ln P^{\bullet}_{\alpha t} = \pi^{\bullet}_{\alpha 0} + \pi^{\bullet}_{\alpha 1}\ln P^{\bullet}_{\alpha t-1} + \pi^{\bullet}_{\alpha 2}\ln \mu_t + v^{\bullet}_{\alpha t} \qquad (4)$$

where $P^{\bullet}_{\alpha t}$ is the relative poverty measure (equation 2) obtained by re-estimating $P_{\alpha t}$ applying a fixed reference mean to the Lorenz curve for date t.

The poverty measures are formed by taking a (household-size weighted) mean of household-level poverty measures based on consumption-poverty gaps; specifically:

$$P_{\alpha} = \sum_{i=1}^{n} \max[(1 - x_i/z)^{\alpha}, 0]s_i \qquad \alpha \geq 0 \qquad (5)$$

in which x_i is consumption expenditure per person in the i'th household of size s_i in a population of n households, z is the poverty line, and α is a non-negative parameter which we allow to take three possible values: i) $\alpha=0$, giving the headcount index; ii) $\alpha=1$, the poverty gap index; and iii) $\alpha=2$, the squared poverty gap index.[10] The headcount index is the most popular measure, but—unlike the other two measures—it is unresponsive to changes below the poverty line.

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[10] On the reasons for preferring additive measures to non-additive ones, such as the Sen (1976) index, see Foster and Shorrocks (1991).

Our principal concerns in estimating these various models for the poverty measures arise from two intrinsic limitations of our data: the first is that, after allowing for missing data, we have only 20 observations for the poverty regression, and the second is that those observations are unevenly spaced. Even if we achieve consistency, small sample biases in the types of models we are estimating will probably entail some underestimation of the true values of the $\pi_{\alpha 1}$'s.[11]

We shall try two approaches for dealing with the irregularity of the household surveys. Let τ_t denote the elapsed time since the last survey. On eliminating terms in the unknown poverty measures between surveys, we can then re-write equation (3) as:

$$\ln P_{\alpha t} = \pi_{\alpha 1}^{\tau_t} \ln P_{\alpha t - \tau_t} + \pi_{\alpha 0} \frac{(1 - \pi_{\alpha 1}^{\tau_t})}{(1 - \pi_{\alpha 1})}$$

$$+ \pi_{\alpha 2}(\ln \mu_t + \pi_{\alpha 1} \ln \mu_{t-1} + \pi_{\alpha 1}^2 \ln \mu_{t-2} + \ldots + \pi_{\alpha 1}^{\tau_t - 1} \ln \mu_{t - \tau_t + 1}) + \tilde{v}_{\alpha t} \qquad (6)$$

However, this is still not a readily estimable model; it requires the survey means for years between surveys and (in effect) it has a different number of variables in different observations (according to the variation in τ_t). The model becomes tractable if we re-write it as:

$$\ln P_{\alpha t} = \pi_{\alpha 1}^{\tau_t} \ln P_{\alpha t - \tau_t} + (\pi_{\alpha 0} + \pi_{\alpha 2} \ln \mu_t) \frac{(1 - \pi_{\alpha 1}^{\tau_t})}{(1 - \pi_{\alpha 1})} + \tilde{v}_{\alpha t}^* \qquad (7)$$

However, this does modify the properties of the error term, as it now includes deviations of $\ln \mu_{t-i}$ from its value at time t for $i \le \tau_{t-1}$. If $\ln \mu_{t-i}$ behaves as an autoregressive process during the period since the last survey, the new error term will still be heteroscedastic (since the uneven

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[11] See Hendry (1984) for Monte-Carlo evidence on this.

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spacing will mean that the error variance tends to be larger after longer gaps).[12] We shall later examine the properties of the model's residuals to see if this is a serious concern. Nonetheless, this specification is probably better than ignoring the gaps, or interpolating for the missing values.

A second approach can be suggested. If we start from a slightly different model by adding a term in $\pi_{\alpha 3} \ln\mu_{t-1}$ to the right hand side of (3) and apply common factor restrictions (Sargan, 1980) we can write the model in an estimable nonlinear form:

$$\ln P_{\alpha t} = \pi_{\alpha 1}^{\tau_t} \ln P_{\alpha t - \tau_t} + \pi_{\alpha 0} \frac{(1 - \pi_{\alpha 1}^{\tau_t})}{(1 - \pi_{\alpha 1})} + \pi_{\alpha 2}(\ln\mu_t - \pi_{\alpha 1}^{\tau_t} \ln\mu_{t - \tau_t}) + \tilde{v}_{\alpha t}^* \qquad (8)$$

This does not require data between surveys. Under the (here untestable) common factor restrictions the new error term ($\tilde{v}_{\alpha t}^*$) is $v_{\alpha t} + \pi_{\alpha 1} v_{\alpha t-1} + ... + \pi_{\alpha 1}^{\tau_t - 1} v_{\alpha t - \tau_t + 1}$. The implicit coefficient on $\ln\mu_{t-1}$ is $\pi_{\alpha 3} = -\pi_{\alpha 1}\pi_{\alpha 2}$. We shall also try this approach.

2.3 Direct and indirect effects of agricultural growth on poverty

The above discussion has outlined a method of assessing how pro-poor the distributional effects of growth are. But how does agricultural growth in particular affect the mean and Lorenz parameters of the distribution of consumption?

As in most of Asia, the vast bulk of the rural poor in India are either landless or live on small farms with inadequate land for even their food needs. They depend heavily on earnings from supplying unskilled wage labor to other farm or non-farm enterprises. In principle there are

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[12] The new error term contains terms in $\ln\mu_t - \ln\mu_{t-1}$. With $\ln\mu_t$ following an AR process, the error term will still have a zero mean.

two channels through which they might benefit from economic growth. One is by participating directly in that growth by producing more on their own land; the other is through the labor market.

The new farming technologies that started to be adopted in India from around 1970 allowed higher output by both raising yield per acre sown and by permitting multiple cropping of a given land area within one year. Of the two ways in which the landless rural poor might benefit from this growth—extra employment or higher wages—the latter channel is more contentious; indeed, early development theories assumed a rural economy in labor surplus, such that extra employment would have no effect on the real wage (Lewis, 1954; Ranis and Fei, 1961). By this view, there would be little scope for trickle down via real wages.

However, other models of the rural labor market allow a labor surplus to coexist with a process of wage determination in which labor-augmenting technical progress can lead to higher real wages.[13] The dynamics of such effects are also of concern here. Short-run stickiness is a widely observed property of formal sector wages, but it has also been observed in rural settings in which there are no trade unions or binding minimum wage rates (Boyce and Ravallion, 1991). In poor rural economies employers will resist wage increases at least initially, and the existence of tacit collusion or other forms of resistance on the supply side in village labor markets could also yield short run stickiness downwards. Long-run responses of real wages to agricultural growth can thus exceed short-run responses.

Motivated by the above arguments, we would also like to test whether or not: i) rural poverty depends on both average real wages in agriculture and average farm productivity; ii) real

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[13] For a survey of various models of rural labor markets in this setting see Drèze and Mukherjee (1989). Examples of the models we refer to include Osmani (1991) and Mukherjee and Ray (1992).

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wages are directly responsive to productivity; iii) both poverty and wages are positively serially dependent, in that (ceteris paribus) they will tend to be closer to their latest values than more distant ones; and iv) any pro-poor impact of higher farm productivity and wages persists after controlling for changes in mean consumption.

The economic model we have in mind for the long-run determinants of poverty gives individual consumption as a function of the real wage rate, a measure of technical progress in agriculture, other exogenous aggregate variables, and an unobserved idiosyncratic variable reflecting individual endowments. Integrating out the latter variable, the long run poverty measure defined on the distribution of consumption can be written as a function of the wage and the farm productivity variable. The long run wage we assume to also be a function of farm productivity, via (for example) a long run labor market clearing condition. The adjustment process toward these long run values is taken to be a simple autoregressive process, and we assume that the relationships of interest are linear in logarithms.

Thus we estimate the following triangular system:

$$\ln P_{\alpha t} = \beta_{\alpha 0} + \beta_{\alpha 1}\ln P_{\alpha t-1} + \beta_{\alpha 2}\ln W_t + \beta_{\alpha 3}X_t + \epsilon_{\alpha t}$$
$$\ln W_t = \gamma_0 + \gamma_1 \ln W_{t-1} + \gamma_2 X_t + \epsilon_t$$

(9)

where W_t is the real wage rate for agricultural labor at t, X_t is a vector of other relevant variables including average farm productivity, the $\beta_{\alpha i}$'s and γ_i's are parameters to estimate, and ϵ_t, v_t are assumed to be normally distributed white noise error processes. Similarly to equation (4) we shall then also re-estimate the model after replacing the poverty measures by $P_{\alpha t}^*$ to test for distributional impacts on poverty. To deal with uneven spacing we will estimate nonlinear models analogous to equations (7) and (8).

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3   Results

3.1   Data

We shall be using 20 rounds of India's National Sample Survey, spanning the period 1958-90. The gaps between surveys range from 11 months to 5.5 years. It is clearly less than ideal to be estimating a dynamic model from only 20 time series observations, and unevenly spaced as well. While more surveys are available (for the 1950s) other data (notably wages) are not. The distributions are of household consumption expenditure (including from own production) per person. The poverty line is the official one for India, produced by the Planning Commission. For the relative poverty measures we use the overall mean as the reference ($\bar{\mu}$ in equation (2)), equivalent to setting poverty lines at 84% of each date's mean. The wage data are for male agricultural wages.[14] The price deflator for both the poverty line and wages is the Consumer Price Index for Agricultural Laborers. We measure farm productivity by agricultural output per acre, output being measured by a value-weighted quantity index of production of all crops.[15] Appendix 1 describes the data more fully, while Appendix 2 describes the relationship with other specifications found in the literature.

The basic time pattern of the evolution of rural poverty has been one of fluctuation without a trend up to about the mid-1970s and a significant decline thereafter (see Table A2 and Figures 2, 3 in Appendix 1). The contrast between the two periods, 1958-75 and 1975-90, is also borne out by the annual growth rates in Table 1. The latter 15-year period saw significantly

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[14]   A complete time series is not available for women.

[15]   Alternatively one can use value added in agriculture, though this makes little difference since it is highly correlated with output (r=0.97 in logs).

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Table 1: Annual growth rates of selected variables

	Average annual rate of growth (%)	
	1958-75	1975-90
Head-count index	0.96	-3.87
Poverty gap index	1.09	-6.38
Squared poverty gap index	1.22	-8.15
Real mean per capita consumption	-0.93	1.76
Real agricultural wage rate	0.06	3.36
Agricultural output per acre of net sown area (YNS)	1.58	2.89
Net sown area per person in rural areas (NSAPC)	-1.57	-1.91
Agricultural output per person in rural areas (YPC)	0.00	0.97
Consumer Price Index for Agricultural Labourers (CPI)	7.41	6.84

higher rates of decline in all poverty measures alongside higher growth rates for real mean per capita consumption, real agricultural wages and agricultural yield.

3.2   Regressions for the poverty measures as functions of the mean

Figure 1 plots both the headcount index and squared poverty gap against mean consumption. We give both the absolute poverty measures implied by the actual mean and Lorenz curve for each date, as well as the relative poverty measures based on the actual Lorenz curve and a fixed reference mean so as to isolate the distributional effects (section 2.1). For the absolute poverty measures there is strong negative correlation with the mean. This vanishes for the relative poverty measures; there is no sign of distributional effects for or against the poor.

To estimate the poverty regressions against the mean in the form of equations (7) or (8) we used Amemiya's (1974) nonlinear least squares (NLS) and nonlinear two-stage least squares (N2SLS) methods, as programmed by Pesaran and Pesaran (1991). Our estimates of equations (3) and (4) are given in Table 2. For the N2SLS estimator the instrumental variables were the lagged poverty measures, the lagged mean, lagged real wages, lagged yield per acre and a time trend. We subjected all these regressions to a battery of standard residual diagnostic tests, including serial correlation (up to two years), normality, heteroscedasticity (squared residuals regressed against squared fitted values, though we also tried the length of time between surveys on its own), autoregressive heteroscedasticity (up to two years), and a miss-specification test for generalized IV estimation which can be interpreted as a test of exogeneity of instruments. All tests were passed comfortably.

The short-run elasticities to the mean vary from -1.1 to -2.2, being higher the higher the value of α. (We comment further on that property later.) The positive autocorrelation in the three poverty measures entails higher long-run elasticities which range from -1.8 to -3.6 again

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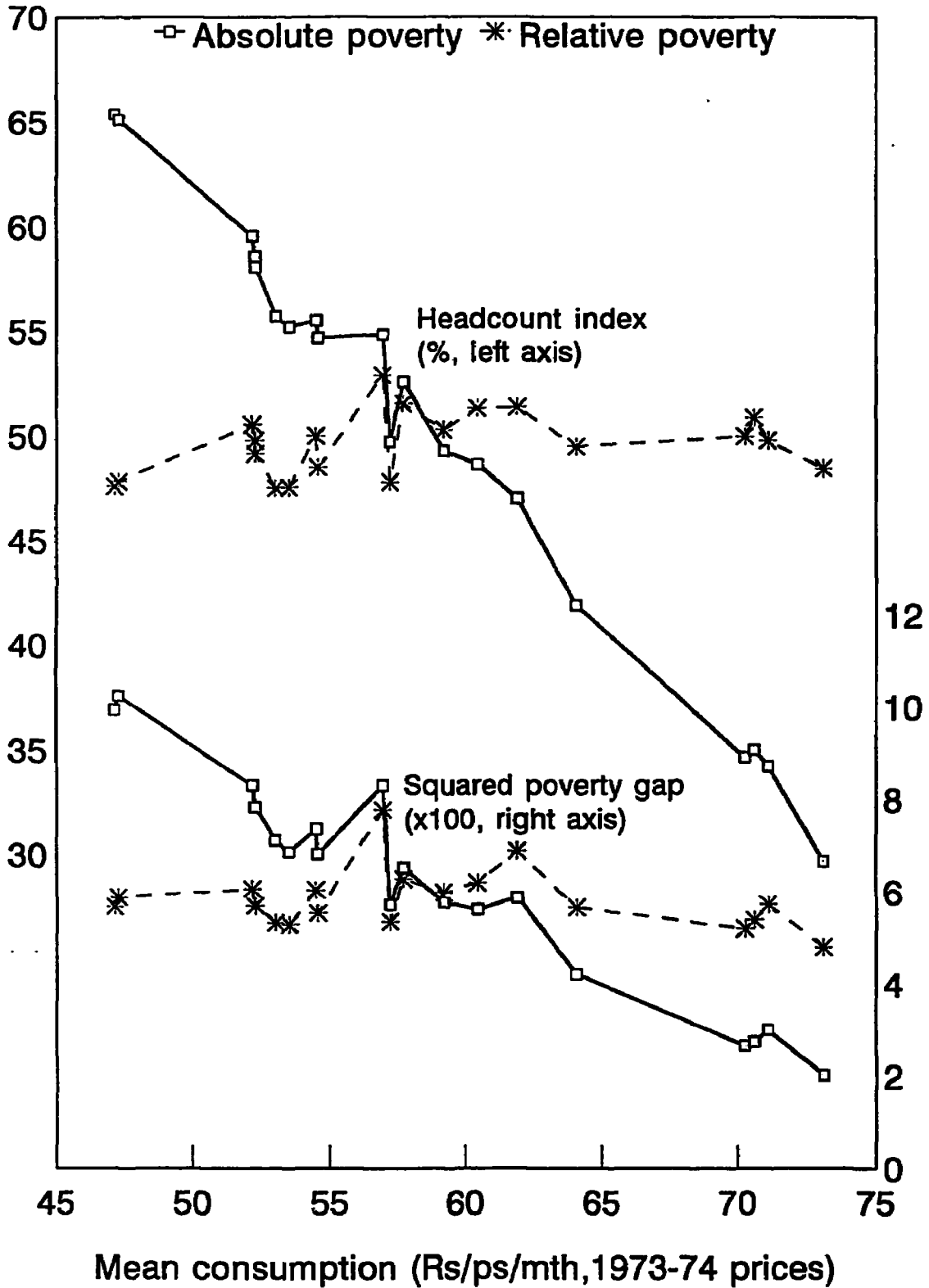# Figure 1: Poverty Measures Against Mean Consumption

Mean consumption (Rs/ps/mth,1973-74 prices)

Table 2: Poverty measures as functions of the mean

	Headcount index ($\alpha=0$)			Poverty gap index ($\alpha=1$)			Squared poverty gap ($\alpha=2$)		
	Absolute measure		Relative measure	Absolute measure		Relative measure	Absolute measure		Relative measure
	NLS	N2SLS	NLS	NLS	N2SLS	NLS	NLS	N2SLS	NLS
1	7.041 (5.342)	7.920 (4.904)	2.637 (2.969)	7.319 (4.727)	8.428 (4.239)	1.539 (2.758)	8.455 (4.575)	9.920 (4.228)	1.403 (2.313)
Lagged poverty measure	0.368 (2.856)	0.284 (1.837)	0.259 (1.063)	0.485 (3.925)	0.401 (2.607)	0.451 (2.622)	0.480 (3.751)	0.389 (2.471)	0.414 (2.371)
Real value of mean consumption	-1.128 (5.507)	-1.265 (5.002)	0.062 (1.318)	-1.466 (4.825)	-1.685 (4.298)	-0.020 (0.265)	-1.861 (4.622)	-2.183 (4.258)	-0.095 (0.834)
Adj.R^2	0.971	0.973	0.133	0.965	0.965	0.225	0.960	0.959	0.233
SEE	0.040	0.040	0.025	0.068	0.069	0.048	0.092	0.094	0.072
Mean dep.var.	3.889	3.889	3.900	2.649	2.642	2.664	1.726	1.717	1.742
Norm. Chi-Sq(2)	9.33	6.35	1.24	5.54	4.08	0.25	2.21	1.19	0.92
Hetero. Chi-Sq(1)	1.99	1.81	0.09	1.57	1.90	0.02	1.68	2.14	0.05
IV mis-spec. Chi-Sq(6)	n.a.	8.50	n.a.	n.a.	10.72	n.a.	n.a.	10.53	n.a.

Note: 20 irregularly spaced observations spanning 1958-90. Non-linear least squares estimates allowing for the uneven spacing. All variables in logs, except time. Absolute t-ratios in parentheses. NLS=nonlinear least squares estimates; N2SLS=nonlinear two-stage least squares estimates; the nonlinearity arises from the uneven spacing (see text). Norm.=Jarque-bera test of normality of residuals. Hetero.=test for heteroskedasticity base on squared fitted values. IV mis-spec.=mis-specification test for generalised IV estimation defined by the value of the IV minimand divided by the regression variance, with degress of freedom given by the number of overidentifying instruments.

depending on α. However, this strong effect of the mean vanishes when one isolates the distributional effect on the poverty measures using the simulated relative poverty measures. Growth in average living standards does reduce poverty, but the effect is roughly distribution neutral. The regressions thus confirm the impression from Figure 1.

The alternative approach based on equation (8) gave similar results. The coefficient on lagged poverty was slightly higher, as was the short-run elasticity to the mean, though the long-run elasticities (including the implicit coefficient on the one year lagged mean, eliminated from (8) using the common factor restriction) were very similar. On replacing the absolute poverty measures by the relative measures purged of the direct growth effect, the mean again became insignificant, with results similar to Table 2.

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3.3   *Regressions for the poverty measures as functions of the real wage and agricultural yield*

Turning to the model in equation (9), the X_t vector comprises the log of agricultural output per unit net sown area (YNS), the log of net sown area per head of rural population (NSAPC), a time trend, and two variables to pick up any short-term effects of fluctuations, namely the change in log YNS and the change in the log of the consumer price index for agricultural laborers (CPI). To deal with the uneven spacing we first estimated the poverty regression in a form analogous to (7). While the real agricultural wage rate, the lagged poverty measure and log YNS were significant at the 5% level or better for all three poverty measures, the other variables were not. There was a deterioration in overall fit (as measured by adjusted R^2) when the other variables were dropped entirely. However, there was negligible drop in fit when one dropped log NSAPC and the inflation rate, leaving the time trend in. We also estimated the restricted model treating both the current year's wage and yield as endogenous, and

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using a N2SLS estimator with lagged values up to two years of the X_t vector, the gap between

surveys, the time trend, and the lagged poverty measure as instrumental variables.[16]

We subjected all these regressions to the same residual diagnostic tests as in section 3.2.

All tests were passed comfortably except serial correlation, with indications of significant

negative serial correlation in the residuals for all three poverty regressions (for both NLS and

N2SLS estimators). On testing for the possibility of omitted lagged effects of some or all of the

X_t variables from the last survey date, we found that lagged output per acre had a strong effect,

similar in size to (and not statistically different from) the effect of current output. On replacing

current output per acre by the sum of current and lagged values we obtained the results in Table

3. There was no sign of serial correlation in this model's residuals, for either NLS or N2SLS

estimators. All other diagnostic tests passed comfortably. We also deleted the first n

observations or the last n, with n=1,2,3,4, 5. The results were quite robust to these changes.

Our results indicate strong evidence of serial dependence in all three poverty measures,

with autoregression coefficients around 0.5-0.6. The real wage rate has a strong short run

impact, with a higher elasticity for the poverty gap index than the headcount index, and highest

for the squared poverty gap, indicating strong effects well below the poverty line. The same

comment applies to the agricultural yield variable. Direct effects of yield growth on poverty are

evident; these may involve either wage-labor employment effects or own-farm productivity gains

(as noted, the data do not allow us to distinguish the latter two effects empirically). Controlling

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[16] The lagged poverty measure was multiplied by the NLS estimate of its coefficient raised to the power of the gap between surveys; this provided a good instrument for the first term on the RHS of equation (6).

Table 3: Poverty measures as functions of the real wage and agricultural yield

	Headcount index ($\alpha=0$)			Poverty gap index ($\alpha=1$)			Squared poverty gap ($\alpha=2$)		
	Absolute measure		Relative measure	Absolute measure		Relative measure	Absolute measure		Relative measure
	NLS	N2SLS	NLS	NLS	N2SLS	NLS	NLS	N2SLS	NLS
1	-19.809 (3.312)	-18.286 (2.554)	7.925 (1.334)	-36.688 (4.895)	-35.014 (3.900)	-0.802 (0.088)	-51.360 (5.194)	-49.781 (4.321)	-2.865 (0.229)
Lagged poverty measure	0.602 (6.499)	0.645 (5.685)	0.092 (0.264)	0.600 (8.079)	0.610 (6.164)	0.338 (1.175)	0.581 (7.571)	0.550 (5.335)	0.159 (0.577)
Real wage rate	-0.374 (3.503)	-0.338 (2.539)	0.090 (0.865)	-0.578 (4.556)	-0.519 (2.805)	0.063 (0.381)	-0.715 (4.331)	-0.771 (2.750)	0.187 (0.767)
Current plus lagged yield	-0.302 (3.246)	-0.281 (2.705)	0.011 (0.140)	-0.524 (4.464)	-0.531 (4.110)	-0.092 (0.732)	-0.738 (4.728)	-0.726 (4.289)	-0.233 (1.227)
Time	0.012 (3.612)	0.011 (2.763)	-0.002 (0.803)	0.022 (5.108)	0.021 (4.088)	0.002 (0.345)	0.030 (5.357)	0.029 (4.501)	0.003 (0.454)
Adj.R^2	0.971	0.970	0.012	0.982	0.982	0.190	0.981	0.981	0.297
SEE	0.039	0.041	0.028	0.049	0.050	0.049	0.064	0.065	0.069
Mean dep.var.	3.889	3.884	3.900	2.649	2.642	2.664	1.726	1.717	1.742
Norm. Chi-Sq(2)	2.18	2.20	1.67	0.67	0.62	0.23	0.98	0.83	0.33
Hetero. Chi-Sq(1)	0.37	0.29	0.45	0.29	0.23	0.64	0.26	0.38	0.88
IV mis-spec. Chi-Sq(7)	n.a.	10.98	n.a.		11.15			7.40	

Note: See Table 2.

for these variables, there is a significant underlying positive trend, entailing annual rates of increase of 1.2%, 2.2% and 3.0% for headcount index, poverty gap and squared poverty gap.

We also tried the alternative approach analogously to (8). Using the same starting point for the right hand side variables, but adding the one-year lags, and imposing the common factor restrictions we again obtained significant effects of the real wage rate and output per acre, with strong serial dependence in the poverty measures. There were differences, however. The coefficient on the lagged poverty measures was higher (0.82-0.84 for all three measures), and not significantly less than unity (standard errors of 0.10-0.11), implying explosive long run properties. Also, the time trend was insignificant. The coefficients on wages and output per acre were also somewhat higher; around 0.62-1.06 for the wage rate and 0.52-1.11 for output per acre, all with t-ratios around four. We rejected these results as implausible.

3.4   Regressions for the real wage rate

Turning to the wage equation, we began with the same variables for X_t as used in the

poverty regressions. This time the inflation rate was highly significant, though none of the other variables were individually. There was, however, a noticeable loss of fit when all were dropped. Yet log yield was significant when the others were dropped, and none of the others were significant (except inflation) if they were put back in the equation individually.[17] The parameter restriction that the coefficients on current and lagged log YNS were equal again passed easily. Table 4 gives both OLS and IV estimates of the final wage equation; the instrumental variables

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[17]   A joint F-test convincingly accepted the parameter restrictions (F(2,25)=0.81).

Table 4: Regressions for real agricultural wage rate

	OLS	IV
1	-0.717 (3.017)	-0.732 (2.964)
Lagged real wage rate	0.766 (8.600)	0.761 (8.223)
Current plus lagged yield	0.126 (3.262)	0.128 (3.191)
Inflation rate	-0.636 (7.040)	-0.642 (5.423)
Adj.R^2	0.957	0.957
F	223.125	223.017
SEE	0.042	0.042
Mean dep.var.	1.529	1.529
AR(1): $Chi^2(1)$	1.16	1.12
RESET: $Chi^2(1)$	6.99	0.56
Normality: $Chi^2(2)$	0.13	0.09
Hetero.: $Chi^2(1)$	5.34	5.21
IV misspecification test: $Chi^2(4)$	n.a.	6.78

Note: 32 annual observations, 1958-89. All variables in logs. Absolute t-ratios in parentheses.

were the two year lagged values of real wages, the price index, output and the time trend. The equations passed the same set of residual diagnostic tests described above.

The dynamic effect is clearly strong, and there is also a sizable short-run adverse effect of inflation. The short-run elasticity of the real wage to output per acre is about 0.13, rising to over eight times that figure in the long-run. Thus, while we do find strong support for real wage sluggishness, there is a detectable though small short-run impact of current yield.

One can also write this model in terms of the first differences of these variables, augmented with an error correction mechanism. The variables of interest—the logs of W, CPI, and YNS—were all stationary in first differences (using both Dickey-Fuller and Augmented Dickey-Fuller tests). Regressing the log wage on the other two variables in levels, the residuals were found to be stationary, implying cointegration and hence the existence of an error-correction model (Engel and Granger (1987):

$$\Delta \ln W_t = \pi_1 \Delta X_t + \pi_2 (\ln W_{t-1} - \pi_3 X_{t-1}) + v_t \tag{10}$$

On estimating the model in this form one obtains (absolute t-ratios in parentheses):

$$\Delta \ln W = \underset{(6.91)}{-0.637 \Delta \ln CPI} + \underset{(1.37)}{0.137 \Delta \ln YNS} - \underset{(2.55)}{0.233} (\underset{(5.50)}{\ln W_{-1}} - \underset{(3.74)}{1.074 \ln YNS_{-1}} - 3.057) \tag{11}$$

(where $R^2 = 0.749$; $SEE = 0.043$).[18] However, the parameter restriction needed to yield the same model as in Table 4 (namely that $2\pi_1^{YNS} + \pi_2 . \pi_3^{YNS} = 0$ in obvious notation) performs very well.[19]

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[18] The long-run price level effect in the error correction term was highly insignificant—indicating that only the real variables matter in the long-run—and was dropped.

[19] A Wald test gives a Chi-square of 0.014.

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3.5   The elasticities of poverty to agricultural yield

The <u>short-run elasticity</u> of the poverty measure to a change in output per acre is given by

$$\frac{\partial \ln P_{\alpha t}}{\partial \ln YNS_t} = \beta_{\alpha 2} \gamma_2^{YNS} + \beta_{\alpha 3}^{YNS} \tag{12}$$

(where the superscripts "YNS" refer to the parameter on log YNS in each equation). The steady-state reduced form of equation (9), as estimated in Tables 3 and 4, gives the poverty measures and real wage rate as a function of agricultural yield per acre and time. Using " ~ " to denote steady-state values, the <u>long-run elasticity</u> of the poverty measure to yield is[20]

$$\frac{\partial \ln \tilde{P}_\alpha}{\partial \ln Y\tilde{N}S} = \frac{2\beta_{\alpha 2}\gamma_2^{YNS}}{(1-\beta_{\alpha 1})(1-\gamma_1)} + \frac{2\beta_{\alpha 3}^{YNS}}{(1-\beta_{\alpha 1})} \tag{13}$$

In each case, the first term on the right-hand-side is the effect operating through the wage rate, while the second reflects other channels, as discussed in Section 2.2.

Table 5 gives the elasticities implied by both our estimates from Tables 3 and 4. In the short-run, the growth effect via the wage is small, and dominated by other channels. But in the long-run the wage effects account for about 30-40% of the total elasticity. Total elasticities are at least six times higher in the long-run. Half or more of the long-run impact is reached within three years; 80% is reached in six years. The elasticities are also higher for higher values of α.

The long-run poverty measures are also an increasing function of time. This presumably reflects an omitted variable, such as population growth, which is difficult to distinguish from time. (Appendix 2 discusses this point further in the context of related work in the literature.)

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[20]   Notice that it is the sum of the current and lagged log YNS on the right hand side of both the poverty and wage regressions. Thus one collects two terms in log YNS in the steady-state.

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Table 5: Elasticities of absolute poverty to growth in output per acre

Poverty measure	Estimation method	Short-run elasticities			Long-run elasticities			Elasticity at the end of	
		Through the wage rate	Other channels	Total	Through the wage rate	Other channels	Total	3 years	6 years
Headcount index	NLS/OLS	-0.047	-0.302	-0.349	-1.010	-1.514	-2.524	-1.365	-2.047
	N2SLS/IV	-0.043	-0.281	-0.325	-1.024	-1.584	-2.606	-1.316	-2.051
Poverty gap	NLS/OLS	-0.073	-0.524	-0.597	-1.554	-2.620	-4.174	-2.313	-3.426
	N2SLS/IV	-0.067	-0.531	-0.597	-1.430	-2.720	-4.152	-2.317	-3.429
Squared poverty gap	NLS/OLS	-0.090	-0.738	-0.828	-1.832	-3.520	-5.352	-3.128	-4.500
	N2SLS/IV	-0.099	-0.726	-0.825	-1.840	-3.226	-5.064	-3.057	-4.313

Note: Refers to the estimators for the poverty and wage equations respectively.

The existence of a positive trend, controlling for yield, implies that without the yield gains over this period, poverty would have increased. Thus the gains to India's rural poor since about 1970—with the incidence of poverty roughly halving between 1970 and 1990 (Appendix 1)—are attributed to yield increases outweighing the underlying adverse trend. From the results in Table 3, one can calculate the minimum rate of growth in output per acre needed assure that poverty does not increase in the long-run; for the headcount index the minimum rate of yield growth needed is 1.19% per year, while it is 1.30% and 1.27% for the poverty gap and squared poverty gap respectively. The historical rate of growth in yield over this period was 2.32% per year.[21]

3.6 Why are all elasticities higher for higher values of α?

It can be seen in Tables 2, 3 and 5 that the (absolute) elasticities of the poverty measures to both mean consumption and yield are higher for higher values of the parameter α. From (5):

$$P_1 = P_0(1 - \mu^P/z); \quad P_2 = P_1\left[1 + \mu^P/z + \frac{(\sigma^P/z)^2}{1 - \mu^P/z}\right] \tag{14}$$

where μ^P and σ^P are the mean and standard deviation of consumption by the poor. So the higher elasticity for P_1 than P_0 indicates that growth also increases the average consumption of the poor (i.e., that the income gap ratio $1 - \mu^P/z$ falls). Furthermore, inequality amongst the poor—as measured by the CV (σ^P/μ^P)—must be decreasing as average productivity increases (since μ^P is increasing with yield, a higher elasticity for P_2 than P_1 must imply that σ^P is falling). Thus, with growth there is also an improvement in distribution amongst the (changing

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[21] This least squares estimate using the annual data, with a correction for the (mild) serial correlation in the OLS error terms.

numbers of) poor. None of this is inconsistent with distribution neutrality overall; for example, all consumption levels could increase at the same rate while the mean consumption of those below the poverty line stays the same or even falls, depending on the density of consumptions in a neighborhood of the poverty line.[22]

4   Conclusions

India appears to be the only developing country for which consistent poverty measures can be tracked over a long time. We find that measures of absolute rural poverty responded elastically to changes in mean consumption over the period 1958-90. This response vanishes when one focuses on measures of relative poverty; the impact of growth on poverty was roughly distribution-neutral in the long run. Our results strongly reject the "immiserizing growth" hypothesis. But nor was there "redistribution with growth".

We have also collated the household survey data with data on agricultural wages and outputs, and estimated a dynamic model determining rural poverty measures and real wages. We find that a range of absolute poverty measures responded in the short run to changes in agricultural wages as well as to average farm yields. And wages responded to farm yields, presumably through effects on labor demand, such as due to multiple cropping. Higher yields thus helped reduce absolute poverty through induced wage effects, as well as the more direct channels, including effects on both employment and own-farm productivity. The bulk of the consumption gains to poor people since about 1970 are attributed to the direct and indirect impacts of agricultural growth. Nor were the gains confined solely to those near the poverty

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[22] When all consumptions grow at the same rate, it can be shown that the necessary and sufficient condition for the absolute elasticity of P_1 to the mean to be greater than that of P_0 is that $f(z)/P_0 < \mu/z$ where $f(z)$ is the density of consumption at the poverty line z.

line; higher yields also benefited those well below it. However, agricultural growth had no discernable impact on relative poverty. In the short-run, there was also a strongly adverse impact of inflation on real agricultural wages and (hence) absolute poverty.

Neither the consumption-based poverty measures nor real wage rates adjusted instantaneously. The combined effect of this stickiness in both variables is that the short-run gains to poor people of labor-demanding productivity growth are far lower than the long-run impacts. Also, the short-run effects operating via the wage rate are minor compared to those operating through other channels. But in the long run, the wage effects do matter, accounting for about 30-40% of the steady-state elasticity of absolute poverty to a yield increase. Overall elasticities in the long run are at least six times higher than the short-run values, and long-run elasticities exceed two for the incidence of poverty, and are over four for the poverty gap indices. Small sample biases entail that (if anything) we have probably underestimated the true long-run elasticities. The process through which India's rural poor participate in the gains from agricultural growth does take time, though about half of the long-run impact occurs within three years.

Appendix 1: Data sources and estimation methods

Poverty measures

The poverty measures are based on the published National Sample Survey (NSS) data on distributions of per capita monthly expenditure. The distributions are available for 33 NSS rounds, starting with the 3rd round for August 1951-November 1951 and going up to the 47th round for July 1991-December 1991. In keeping with data availability on other variables (see below), the poverty estimates used here are for a shorter period from NSS round 14 (July 1958-June 1959) to round 45 (July 1989-June 1990). The time interval between surveys varies from 11 months to 5.5 years. The poverty line is defined by a per capita monthly expenditure of Rs. 49 in rural areas at October 1973-June 1974 prices. This poverty line was originally proposed by the Task Force on Projections of Minimum Needs and Effective Consumption Demand (Planning Commission 1979), and has recently been endorsed by the Expert Group on Estimation of Number and Proportion of Poor (Planning Commission 1993). The deflator we have used to adjust for temporal changes in the cost of living in the rural sector is the Consumer Price Index for Agricultural Laborers (CPI). Point estimates of the poverty measures are constructed using either the beta Lorenz curve of Kakwani (1980) or the general quadratic model of Villasenor and Arnold (1989), depending on which fits the data best (both satisfied the theoretical conditions needed for a valid Lorenz curve in all survey rounds). Using the formulae in Datt and Ravallion (1992), the poverty measures are calculated from the estimated parameters of the Lorenz curve and the mean per capita consumption expenditure.[23] Figure 2 gives the headcount index. A complete series of the poverty measures with a detailed discussion of our methodology and data sources can be found in Datt (1994).

The price index

The CPI for agricultural laborers for the period we cover is compiled from the Labour Bureau's monthly series on consumer price indices (published in the *Indian Labour Journal* and the *Indian Labour Yearbook.*) Beginning September 1964, the index is directly available at the

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[23] A number of checks are made on the results, including both the theoretical conditions for a valid Lorenz curve, and consistency checks, such as that the estimated value of the head-count index must lie within the relevant class interval of the published distribution. The estimation technique has been set-up in a user-friendly computer program which is available on request, so interested readers can readily check our calculations and their sensitivity to our assumptions.

all-India level. For the earlier period, September 1956 through August 1964, however, only state-level indices are available. We have aggregated these into an all-India index using the same weighting diagram as used in the Labour Bureau series for the later period.

A problem with this price index is that the Labour Bureau ignored increases in firewood prices after 1960-61; firewood is typically a common property resource for agricultural laborers, but it is also a market good, and so the Labour Bureau's practice is questionable.[24] To test for bias due to this problem we estimated an alternative deflator replacing the firewood sub-series by one based on mean rural firewood prices (only available from 1970) and a series derived by assuming that firewood prices increased at the same rate as all other items in the Fuel and Light category (prior to 1970); Datt (1994) discusses this index further. Figure 2 also gives the headcount indices and real wages implied by the alternative deflator. We re-estimated the poverty and wage regressions; the change made little difference to our results (either coefficients or standard errors); elasticities to yield tended to be slightly lower (at most 10% in the long run).

Agricultural wages

Our data on nominal daily agricultural wages are compiled from Jose (1974, 1988), supplemented with the data reported in the *Report of the National Commission on Rural Labour (Volume I)* [GOI, 1991], and *Agricultural Wages in India* reports since 1984-85. The primary source of all these data is the Ministry of Agriculture's annual publication *Agricultural Wages in India* (AWI). The series we use is for male agricultural wages only; data on female agricultural wages are not available on a consistent basis for the whole period. The AWI data from the aforementioned sources are available as aggregated up to the state level. We have aggregated these data further to derive an all-India male agricultural wage using state and year specific weights. These weights are constructed as follows.

Ideally, we would have liked to weight the state-level wage rates by the total (or male) state-level agricultural wage employment for that year. However, the data to do this are not available. As a second best, the weights could be constructed as proportional to the number of total (or male) agricultural laborers in the state. The primary source of data on the latter is from the decennial censuses. However, even this turns out to be problematic because of the non-

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[24] The NSS values firewood consumption from own-production at local market prices. Also see Minhas et al., (1987) for further discussion.

comparability of the definition of "work" in the 1971 census (Krishnamurty, 1984). The more stringent definition in this census actually shows up as a decline in the <u>absolute</u> number of rural workers in many states, when compared with the 1961 census. Our procedure is as follows.

We have comparable state-level data from the NSS on the proportion of agricultural labor households for the years 1956-57, 1964-65, 1974-75, 1977-78, 1983, 1987-88.[25] We use this as a proxy for the proportion of agricultural laborers in the rural population. Next, we combine this with the state-level rural population for these years, estimated by interpolating across the census years, to obtain an estimate of the number of agricultural laborers in different states for these six years. Assuming a constant rate of growth between any two years, we derive annual estimates of the number of agricultural laborers in each state over the period 1956-57 to 1991-92. The latter expressed as a proportion of the total number of agricultural laborers in all states provides us with the state-year specific weights used in constructing an all-India agricultural wage.

Figure 2 gives the estimated wage rates corresponding to each NSS round.

Agricultural output and area

These data are collated from the Ministry of Agriculture (1993) publication *Area and Production of Principal Crops in India 1991-92*. The data are in the form of three annual indices: (i) the index of agricultural production which is a Laspeyres quantity index of production of all crops, where the weight for a particular crop is given by the average value of that crop's output over the triennium ending 1981-82; (ii) the index of gross cropped area under all crops (including area sown more than once during the year) with the same base period, i.e. the triennium ending 1981-82; and (iii) the index of net sown area also with the same base period. All three indices refer to the agricultural year from July to June.

Figure 3 gives the index of output per unit net sown area corresponding to each NSS round.

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[25] These are data from the *Second Agricultural Labour Enquiry* for 1956-57, the *First Rural Labour Enquiry* for 1964-65, the *Second Rural Labour Enquiry* for 1974-75, and the second, third and fourth *Quinquennial Surveys on Employment and Unemployment* for 1977-78, 1983 and 1987-88. After 1974-75, *Rural Labour Enquiries* were integrated with the *Quinquennial Surveys*. These surveys have adopted a consistent definition of agricultural labor households, *viz.* rural labor households who derived more than 50% of their income over the preceding year from wage-paid manual labor in agricultural occupations.

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Figure 2: Rural Poverty and Agricultural Wages
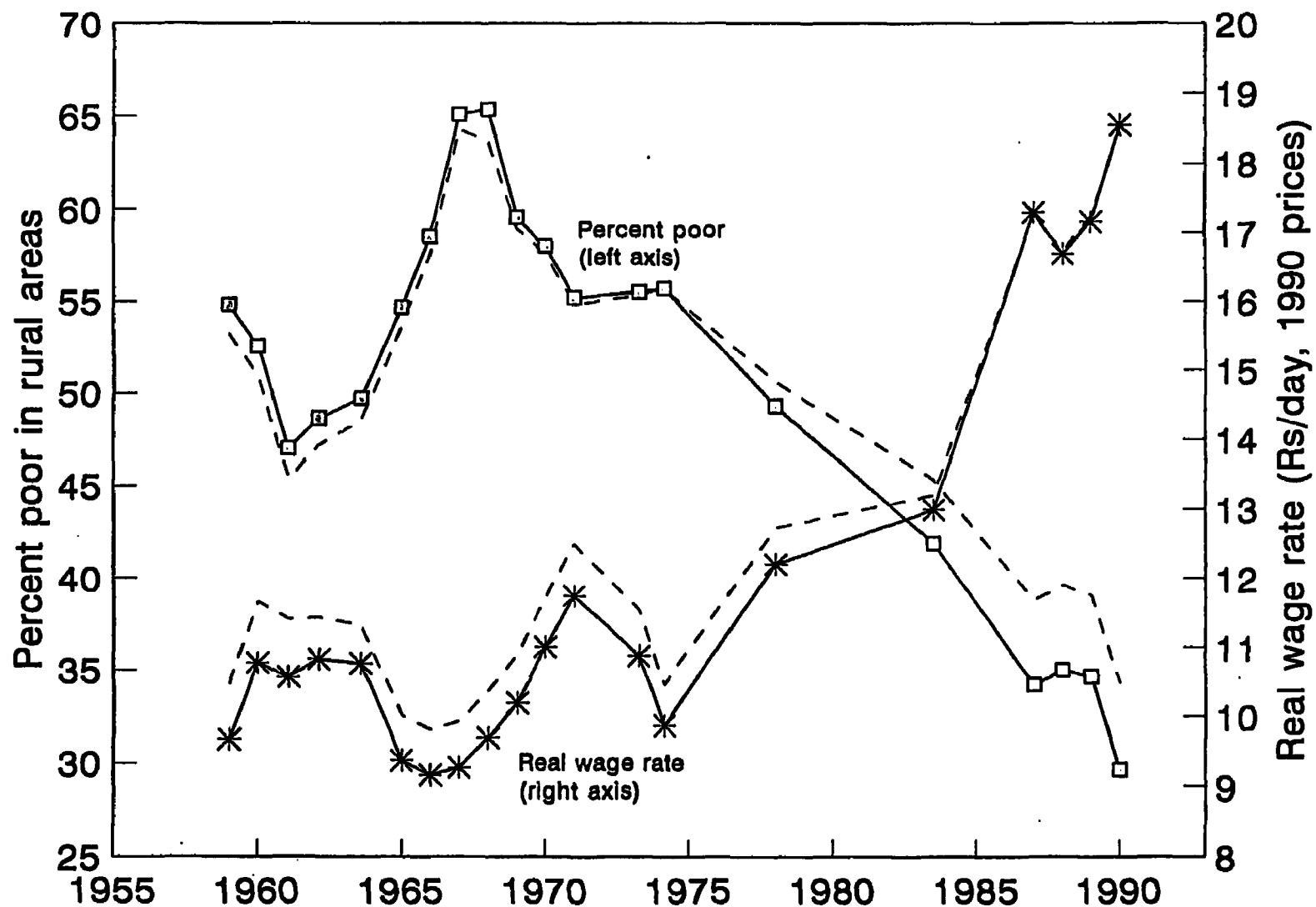
Dashed lines use an alternative deflator (Appendix 1)

Figure 3: Rural Poverty and Agricultural Yields
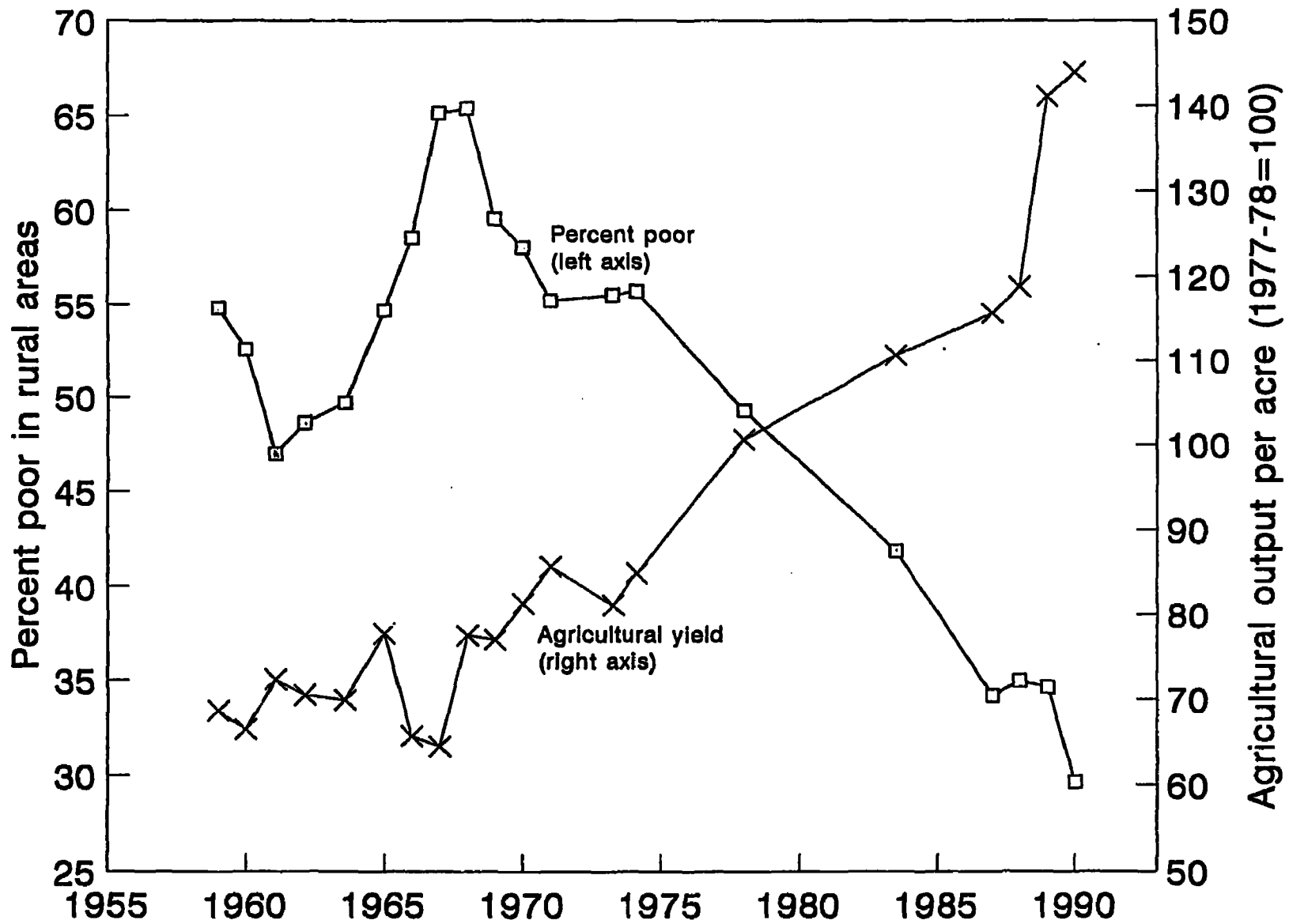
Population

Annual estimates of the rural population are constructed using census data from all five censuses conducted in the post-independence period. Sectoral populations are assumed to grow at a constant rate between censuses. The population estimates are centered at the beginning of each calendar year which coincides with the mid-point of the corresponding agricultural year.

Matching annual data with NSS rounds

As already noted, the survey periods of the NSS rounds do not always coincide with the agricultural year, nor do all of them cover a full 12-month period. In adapting the data for the modelling of poverty by NSS rounds, we have used the following procedure. The data on the CPI are originally collated on a monthly basis and hence permit easy aggregation in the form of averages over months for the corresponding NSS survey periods. Population estimates were made for the mid-point of each NSS survey period. However, data on the other variables are only available on an annual basis for each agricultural year. For these variables, we have constructed values corresponding to a given NSS round as (i) the value of the variable for the agricultural year if the survey period coincides with (or falls entirely within) the agricultural year, or otherwise, as (ii) a weighted average of the values for agricultural years overlapping with the survey period of that round. Thus, for example, for the 18th round, which covers the period February 1963 to January 1964, we construct the corresponding agricultural wage by combining the wage for 1962-63 and 1963-64 in the proportion (5/12) and (7/12) respectively. The annual data are in Table A1, while the data by NSS round are given in Table A2.

Appendix 2: Antecedents in the literature
Poverty regressions

Much of the scholarly debate over agricultural growth and rural poverty in India centered on a seminal paper by Ahluwalia (1978), who regressed measures of rural poverty from 12 surveys between 1957 and 1974 against agricultural output per head of the rural population and a time trend. He found that higher output was associated with lower poverty, and that there was no trend independently of this. The debate which followed questioned sensitivity to changes in the period of analysis (Griffin and Ghose, 1979; Saith, 1981). But with only 12 observations, and relatively little sustained agricultural growth over the period, this may not have been a

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Table A1: Agricultural wages, area, production, and consumer price indices

Year	Daily male agricultural wage		Agricultural output per person in rural areas (index)	Agricultural output per hectare of net sown area (index)	Net sown area per person in rural areas (index)	Consumer Price Index for Agricultural Laborers
	Nominal	Real (Oct 73-Jun 74 prices)	1979/80-1981/82 = 100			Oct 73-Jun 74 = 100
1956-57	1.21	3.65	93.7	64.2	146.0	33.2
1957-58	n.a.	n.a.	85.8	60.7	141.4	34.3
1958-59	1.32	3.73	97.4	68.7	141.7	35.4
1959-60	1.45	4.17	93.2	66.5	140.2	34.9
1960-61	1.40	4.07	100.1	72.5	137.9	34.5
1961-62	1.46	4.17	97.1	70.6	137.5	35.1
1962-63	1.49	4.01	94.3	69.4	135.7	37.3
1963-64	1.66	4.18	93.7	70.3	133.3	39.8
1964-65	1.75	3.61	102.7	77.7	132.2	48.4
1965-66	1.92	3.54	84.0	65.7	127.9	54.3
1966-67	2.34	3.58	81.5	64.5	126.3	65.4
1967-68	2.65	3.74	97.9	77.5	126.3	70.9
1968-69	2.51	3.93	93.6	77.0	121.6	63.8
1969-70	2.82	4.25	98.0	81.3	120.5	66.4
1970-71	2.99	4.53	102.7	85.7	119.9	66.0
1971-72	3.12	4.53	100.0	85.5	117.0	68.8
1972-73	3.26	4.20	90.2	79.8	113.0	77.6
1973-74	3.81	3.91	98.0	84.9	115.4	97.3
1974-75	4.30	3.39	93.5	85.6	109.2	126.8
1975-76	4.72	4.32	105.2	95.4	110.2	109.1
1976-77	5.09	4.90	95.9	89.9	106.7	103.9
1977-78	5.24	4.71	107.3	100.6	106.6	111.2
1978-79	5.46	5.00	108.8	103.1	105.5	109.3
1979-80	5.94	4.80	90.3	89.6	100.8	123.8
1980-81	6.44	4.57	102.1	102.2	99.9	141.0
1981-82	7.40	4.79	107.3	108.0	99.3	154.4
1982-83	8.15	4.92	101.1	104.5	96.7	165.8
1983-84	9.49	5.27	112.4	116.7	96.3	179.9
1984-85	10.53	5.82	109.7	117.5	93.4	180.8
1985-86	11.83	6.19	109.3	119.1	91.7	191.0
1986-87	13.28	6.67	103.4	115.6	89.5	198.9
1987-88	14.43	6.44	101.6	118.8	85.5	224.0
1988-89	16.51	6.62	121.5	141.1	86.1	249.3
1989-90	18.54	7.16	121.6	143.9	84.5	259.1

Table A2: Agricultural wages, area, production, and consumer price indices: NSS rounds 14 through 45

NSS round	Survey period	Poverty measures (x100)			Daily male agricultural wage		Agricultural output per person in rural areas (index)	Agricultural output per hectare of net sown area (index)	Net sown area per person in rural areas (index)	CPI for agricult-ural laborers
		Headcount index (α=0)	Poverty gap index (α=1)	Squared poverty gap (α=2)	Nominal	Real (Oct 73-Jun 74 prices)	1979/80-1981/82 = 100			Oct 73-Jun 74 = 100
14	Jul 58-Jun 59	54.83	18.51	8.31	1.32	3.73	97.4	68.7	141.7	35.4
15	Jul 59-Jun 60	52.56	16.06	6.53	1.45	4.17	93.2	66.5	140.2	34.9
16	Jul 60-Aug 61	47.00	14.28	5.88	1.41	4.08	99.6	72.3	137.9	34.6
17	Sep 61-Jul 62	48.64	14.21	5.62	1.47	4.17	96.9	70.5	137.3	35.2
18	Feb 63-Jan 64	49.72	14.36	5.72	1.59	4.16	93.9	69.9	134.3	38.3
19	Jul 64-Jun 65	54.69	16.53	6.83	1.75	3.61	102.7	77.7	132.2	48.4
20	Jul 65-Jun 66	58.55	18.42	7.84	1.92	3.54	84.0	65.7	127.9	54.3
21	Jul 66-Jun 67	65.11	22.46	10.27	2.34	3.58	81.5	64.5	126.3	65.4
22	Jul 67-Jun 68	65.37	22.16	9.97	2.65	3.74	97.9	77.5	126.3	70.9
23	Jul 68-Jun 69	59.56	19.25	8.32	2.51	3.93	93.6	77.0	121.6	63.8
24	Jul 69-Jun 70	58.04	18.46	7.85	2.82	4.25	98.0	81.3	120.5	66.4
25	Jul 70-Jun 71	55.21	16.71	6.88	2.99	4.53	102.7	85.7	119.9	66.0
27	Oct 72-Sep 73	55.52	17.43	7.37	3.40	4.19	92.1	81.1	113.6	81.1
28	Oct 73-Jun 74	55.72	17.18	7.13	3.81	3.81	98.0	84.9	115.4	100.0
32	Jul 77-Jun 78	49.29	14.47	5.78	5.24	4.71	107.3	100.6	106.6	111.2
38	Jan 83-Dec 83	41.89	11.34	4.22	8.82	5.01	106.8	110.6	96.5	175.8
42	Jul 86-Jun 87	34.19	8.39	3.01	13.28	6.67	103.4	115.6	89.5	198.9
43	Jul 87-Jun 88	34.99	8.13	2.75	14.43	6.44	101.6	118.8	85.5	224.0
44	Jul 88-Jun 89	34.63	7.99	2.67	16.51	6.62	121.5	141.1	86.1	249.3
45	Jul 89-Jun 90	29.65	6.36	2.02	18.54	7.16	121.6	143.9	84.5	259.1

particularly good data set for this purpose (Srinivasan, 1985). The substantial growth that has occurred since then offers hope for a more powerful test (Table 1). Ahluwalia (1985) and Bell and Rich (1994) returned to the Ahluwalia regressions adding data for another year (1977/78) and broadly confirmed his conclusions.[26] By adding new data for the 1980s, we have spanned a period of considerably larger changes in all variables though, as noted in section 3.3, our results turn out to be quite robust to changes in the period of analysis—our model estimated on the Ahluwalia (1978) data set would have given similar results. But we have also adopted a rather different model to this literature. This Appendix elaborates on the reasons.

Past work has not tested the distributional effects. Our approach using simulated measures of relative poverty to isolate distributional shifts in the econometric model of poverty measures is new, though it has an antecedent in the method used by Datt and Ravallion (1992) to decompose changes in poverty into growth and redistribution components.[27]

Higher-order poverty measures—reflecting distribution below the poverty line—have often been used in the literature, though their role has been rather incidental. Ahluwalia (1978, 1985) and others following him did include the Sen (1976) poverty index, but did not draw out any implications concerning the depth of impacts on the poor. Our results indicate differences in the impact of productivity growth between these measures, arising from changes in distribution below the poverty line. We find no support for claims that productivity gains in Indian agriculture have by-passed the poorest.

Turning to the right-hand side variables, almost all past studies of the evolution of India's poverty measures have followed Ahluwalia in using output or income per head. The log of output per person is simply the sum of the logs of output per acre and acres per person, so our specification (initially using both variables) is more general. Our results suggest that the output effects identified in the literature are to do with yield not land per person. Our finding that lagged output matters, and it has a similar effect as current output, echoes results in Ahluwalia (1985). It is also clear that the output variable is picking up more than year-to-year fluctuations as Saith (1981) suggests (for then current and lagged output would have opposite signs).

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[26] Bell and Rich used an earlier version of our data set, rather than Ahluwalia's, and they made some changes to Ahluwalia's specification, as discussed below.

[27] There are other methods of doing such decompositions which do not lend themselves to this extension; see Kakwani (1993), and for further discussion of the differences see Ravallion (1994b).

Our long-run elasticity of the headcount index to output per acre is higher than Ahluwalia's estimate of the elasticity to agricultural income per person, though the difference is not large; our elasticity is -2.6 (including the wage response; see Table 5), while Ahluwalia (1985) obtained -1.9.[28] However, our short-run elasticity is far lower; while we estimate a short-run elasticity of -0.3, Ahluwalia (1985) obtained -1.0; the difference appears to reflect the fact that (in common with almost all this literature) Ahluwalia did not include the lagged poverty measure. (We comment further on dynamic specification later.) According to our estimates, the impact of growth on the poor is a slower process than Ahluwalia's results suggested, although in the end the impact is even larger than he had predicted.

We have also added real wages. It is odd that the real agricultural wage rate has not figured more prominently in this literature, given how much India's rural poor depend on agricultural labor markets. We know of only one other study which has looked at the impact of the real wage rate on the evolution of rural poverty in India (van de Walle, 1985, using the earlier Ahluwalia poverty measures for 1959-71), though numerous observers have conjectured that this is an important variable (including Acharya and Papanek, 1992). Since this is a very strong predictor in our results, it appears that this may have been an important omitted variable in other studies. If we drop the real wage rate from our model then the NLS estimates of the short-run yield elasticities of poverty to yield rise to -0.41 (t=3.2) for the headcount index (instead of -0.30), and -1.0 for the squared poverty gap (instead of -0.74).

Saith (1981), Narain (see Desai, 1985) and others (Mathur, 1985; Gaiha, 1989) added the nominal price level to the original Ahluwalia (1978) model.[29] It is difficult to believe that a monetary variable such as this could have long-run real effects in a correctly specified model (Bliss, 1985). The price-level effect may well be picking up other omitted income sources or financial assets which matter to the poor (agricultural labor and services, or cash holdings) and are not fully indexed for price changes (Desai, 1985; Bliss, 1985; Sen, 1985). We would conjecture that the significant price level effect identified in the literature largely reflects an

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[28] See equation 5 in Table 7.2 of Ahluwalia (1985). (Since Ahluwalia's model is static, the long-run elasticity for Ahluwalia's model is simply the sum of the coefficients on current and lagged output.) Ahluwalia's augmented model including the price level gives an elasticity of -1.8.

[29] In some cases this was the log CPI (Narain), while in others it was the deviation from trend (Saith, Gaiha). But with the trend already included in the regression this difference will only affect the interpretation of the time trend; we return to this issue.

omitted variable bias, the key omitted variable being the real wage rate. While we have allowed short-run effects of price levels, this proved to be insignificant once the real wage was added.[30]

What would a regression more consistent with past work look like when estimated on our extended series? The Ahluwalia specification for the headcount index gives:

$$\ln P_{0t} = 26.00 - 0.75 \ln YPC_t - 1.12 \ln YPC_{t-\tau_t} - 0.01 t$$
$$(4.12) \quad (2.03) \qquad (2.92) \qquad (1.74)$$

where YPC is agricultural output per capita. (The R^2 is 0.81, though there is clearly strong serial correlation in the error term, so that this R^2 and the t-ratios are probably a poor guide to fit and significance.[31]) The Narain-Saith specification gives:

$$\ln P_{0t} = 122.86 - 0.74 \ln YPC_t + 0.71 \ln CPI_t - 0.06 t$$
$$(2.82) \quad (1.89) \qquad (2.16) \qquad (2.61)$$

(The R^2 is 0.77, though there appears to be even stronger serial correlation in the residuals.[32]) An encompassing model can be used to jointly test our changes to both specifications. By adding both (current and lagged) net sown area per person and the price level (all in logs) to our preferred model in Table 2, we can do a nested test against the Ahluwalia-Narain-Saith models. On doing so we found that all three variables are (highly) insignificant, while the other coefficients in our poverty model were quite robust, and remained significant at the 5% level or better.[33] Our model is more consistent with the data.

Notice that, in keeping with the literature, we have included a time trend, and (as usual) this is not readily interpretable. The coefficient on time depends, of course, on what other

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[30] Bell and Rich (1994) find significant effects on the headcount index of their measure of the unanticipated component of inflation, though they do not include the real wage rate.

[31] The Durbin-Watson statistic is 0.80, and the LM tests gives a Chi-square of 8.08 (significant at the 0.005 level). Though these tests are not strictly valid given the uneven spacing, the rejection of serial independence is so strong that one must be concerned about inferences made on the basis of the OLS standard errors. The problem of serial correlation in the errors have been ignored in this literature, even when using data with fairly even spacing.

[32] The Durbin-Watson statistic is 0.56, and the LM tests gives a Chi-square of 9.69.

[33] Wald tests convincingly accept the restriction that the coefficients on all three variables are zero; the rejection is even stronger if one adds the restriction that the coefficient on the lagged poverty measure is also zero.

variables are included; the Ahluwalia-Narain-Saith models above all indicate a negative coefficient; in our specification the sign switches. Van de Walle (1985) attributed her positive trend to adverse effects of population growth, presumably operating through effects on the distribution of land. This is a conjecture, though a seemingly plausible one.

Aside from the specification of the right-hand side variables, studies in the literature have differed in the assumptions (often implicit) they have made about the dynamic structure of their models, and the properties of the error term. Virtually all of the regressions in this literature are static. While the shortage of surveys for estimating the poverty measures will entail some small-sample bias in dynamic models, it is plain from our investigation that the true coefficient on lagged poverty is well above the value of zero that has been widely assumed.

The only other study to allow for serial dependence is Bell and Rich (1994). To deal with the uneven spacing they filled the gaps by two alternative methods: linear interpolation and a forecasting model relating current poverty to its lagged value and current rainfall. They did not use observations after 1977/78, being concerned about the gap to the next round (namely 1983). Aside from other differences in specification already noted, our method for dealing with the dynamics has the advantage that we do not lose observations (five would be lost, spanning the 1980s, in which poverty fell considerably). Nor do we need to introduce a separate (and, under the maintained hypotheses, biased) interpolation model in order to estimate the main model.

Wage regressions

Our results are plainly inconsistent with the classic labor surplus theories—there is a detectable effect on real wages of productivity changes in the short-run. However, there is also considerable sluggishness in the adjustment to such changes. Compared to the poverty regressions, there has been less empirical work on the dynamics of agricultural wage formation in India with which to compare our results. Acharya and Papanek (1992) estimate models of agricultural wage determination for India using time series data, but do not allow for dynamic effects. Omitting the effect of lagged wages also lead to overestimation of the short-run effect of agricultural productivity on wages; this also holds if an autoregressive estimator is used.[34]

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[34] Dropping the lagged wage from the model in Table 4 but using such an AR1 estimator we obtained an elasticity of the wage to yield of 0.38, with a t-ratio of 6.9.

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Lal (1988) estimates a model for the real agricultural wage in India as a function of agricultural output and land under cultivation. Lal's short-run elasticity of wage to yield for 1958-78 is about twice the size of ours. This appears to be due to a difference in specification. Lal differences all variables in his model. The omission of an error-correction term (recalling that we have found that wages and yield are cointegrated) entails that his model does not have a long-run solution, and it may well impart an upward bias on the short-run yield effect on wages.

While we find no support for past claims that inflation hurts the poor independently of its impact on real wages (and yields), we do find strong adverse effects through real wages which are clearly not fully indexed in the short run (Table 4). Thus inflation may be an important omitted variable in "real" models of wage determination. For example, if one re-estimates our wage regression dropping the rate of inflation and adding the log of the relative price of food (food component of the CPI normalized by the CPI) then the latter has a significant negative coefficient (-1.48 with a t-ratio of 3.12); yet this is clearly spurious, for as soon as one returns the rate of inflation to the same regression the relative price effect vanishes (the coefficient drops to 0.002, with a t-ratio of 0.005).[35] This short-run non-homogeneity also has implications for the social impacts of adjustment policies, though this takes us beyond our present scope.[36]

Models of rural-urban migration imply a parity between the agricultural wage and the expected urban-sector wage, where that expectation allows for a (non-trivial) chance of being unemployed in the urban sector (the classic model of this sort is that of Harris and Todaro, 1970). While not claiming to properly test such models, we did try adding an indicator of expected urban sector earnings, namely wage earnings in the formal factory sector per head of the urban population. This was highly insignificant (t=0.32), and had negligible impact on other coefficients. This potential source of gains to the poor from growth is not evident.[37]

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[35] Boyce and Ravallion (1991) get similar results for Bangladesh.

[36] For further discussion see Ravallion and Subbarao (1992).

[37] Ravallion and Datt (1994) find no positive distributional impact of urban economic growth on rural poverty in India, but rural growth did improve urban distribution.

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