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FARM PRODUCTIVITY AND RURAL POVERTY IN INDIA

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ABSTRACT

To what extent do India's rural poor share in agricultural growth? Combining data from 24 household sample surveys spanning 35 years with other sources, we estimate a model of the joint determination of consumption-poverty measures, agricultural wages, and food prices. We find that higher farm productivity brought both absolute and relative gains to poor rural households. A large share of the gains were via wages and prices, though these effects took time. The benefits to the poor were not confined to those near the poverty line.

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1. INTRODUCTION

The following questions have been asked repeatedly in development research and policymaking over many decades:

- **How much do poor rural people share in the gains from higher average productivity in agriculture?** This has been the subject of (often vociferous) debates.[1] Contrast, for example, Ahluwalia's (1978, 320) conclusion that "there is evidence of some trickle down associated with agricultural growth" with Saith's (1981, 205) claim that "there can be little doubt that current growth processes have served as generators of poverty." One might imagine that these two authors were talking about different places or times; but both were using data for the same country—India—over roughly the same period (1957–73). The debate continues; in more recent literature on India, one finds claims that "rapid agricultural growth has benefitted all classes of the poor" (Singh 1990) and "acceleration in agricultural growth by itself is unlikely to make a dent in rural poverty" (Gaiha 1995, 285).
- **Does the relative position of poor people improve or worsen with agricultural growth?** This question has received less attention in the literature, but is

[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 (1995).

increasingly being asked. It is possible for the poorest to lose relatively, while gaining absolutely; this is one interpretation of the idea of "trickle down." And it is possible for a growth process to have sufficiently adverse effects on inequality that absolute poverty increases. By contrast, some growth processes can entail favorable distributional shifts, and, consequently larger gains to the poor than suggested by a mere "trickle down."

- **What role does the labor market play?** Early development theories assumed a rural economy in which extra employment would have no effect on the real wage (Lewis 1954; Ranis and Fei 1961). By this view, the rural sector has a large labor surplus and so there is little scope for the poor to gain 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.[2] Marked differences in the emphasis given to the role of the labor market can also be found in policy-oriented discussions.[3]
- **What role do food markets play?** The food economy of India was largely closed to external trade over the post-Independence period. Higher farm output is then very likely to reduce the relative price of food, though this effect will no doubt be

[2]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), Mukherjee and Ray (1992), and Datt (1996).

[3]Contrast, for example, the IFAD (1992) report with the World Bank (1990). The former emphasizes the scope for reducing rural poverty by developing smallholder agriculture, and pays rather little attention to the role of the unskilled labor market; by contrast, the World Bank's report emphasizes the importance of positive employment and wage effects in achieving pro-poor growth.

buffered to some extent by governmental procurement and storage policies. At the same time, the poorest in rural areas tend to be net consumers of food, in that they have insufficient land for their own consumption needs. So, to the extent that higher farm yields put downward pressure on food prices, the poorest will gain. Against this, there may be poor net producers of food who lose. Heterogeneous impacts among the poor of food price changes have been found in similar settings,[4] though in most of South Asia, the general presumption is that the poorest in rural areas will tend to gain from higher food output and, hence, lower food prices.

- **Are there significant lags in the process through which poor people gain, or lose, from agricultural growth?** The dynamics of the distributional impacts of growth are of obvious interest, though the topic has received surprisingly little attention. Past models have analyzed consumption-based poverty measures within a purely static framework. Yet labor markets are often found to exhibit short-run stickiness in wages, and there is some evidence that this is also true in agricultural labor markets in similar settings.[5] Food markets, too, in this setting may generate price stickiness; for example, governmental intervention in food markets, through producer price setting and storage, can no doubt buffer the effects of food prices somewhat. Stickiness in wage and price adjustment to higher farm yields suggests that long-run responses of poverty measures can exceed the short-run responses.

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[4]See Ravallion (1991) (for Bangladesh) and Ravallion and van de Walle (1991) (for Java, Indonesia).

[5]See the results of Boyce and Ravallion (1991) for Bangladesh.

Addressing these questions empirically calls for a time series of representative household-level surveys; yet such surveys are sporadic, at best, for most countries and their comparability over time is often questionable. India is an exception. There one can trace impacts on poverty over a long period, using reasonably comparable (at least by the standards of international comparisons) and nationally representative surveys of consumption.

This paper tests how much India's rural poor have benefitted from gains in average farm productivity, what role labor and food markets have played, and whether the impacts were distributionally biased one way or another. In the process, we try to resolve some long-standing debates in the literature. Some new methods are brought to bear on these questions. But, probably more important, some new data are used. 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. A new data set embracing a period of higher agricultural growth is used here.

The following section describes our structural model. Section 3 presents our results. Some implications are discussed in Section 4. In Section 5, we compare our results with those of related studies in the literature. Our conclusions can be found in Section 6.

2. MODEL

The bulk of the poor in rural India live on small farms with inadequate land for their own food needs, or are landless. They depend heavily on earnings from supplying unskilled wage labor to other farm or nonfarm enterprises. There are then two main channels through which the poor might benefit from higher farm productivity generally. One is by directly participating in the productivity gains, by producing more on their own land, or finding more employment, either on someone else's land or in some nonfarm enterprise made possible by higher farm yields. These gains can be expected at given wages and prices. The second potential channel is through higher wage rates, or lower prices for consumed agricultural goods, notably food.

Our model incorporates both channels. We assume that individual consumption is a function of the real wage rate for agricultural labor, the relative price of food, average farm productivity, other exogenous aggregate variables, and the individual endowment of land. (Standard micro-models can be used to justify such an assumption, so an elaboration is not necessary here.) Individual consumption normalized by the poverty line is thus given by

$$C = C(W,R,Y,X,\eta)\,, \qquad \textbf{(1)}$$

where W is the real wage rate, R is the relative price of food, Y is a measure of agricultural productivity (such as yield per acre), X is a vector of other relevant variables (including, for example, inflationary shocks), and η is the amount of land owned (measured in units of homogeneous quality), with interpersonal probability density, $f(\eta)$. The function C is assumed to be strictly increasing in η.

The poverty measure is defined on the distribution of consumption in a conventional manner. The poverty level of landholding (η^*) is obtained by inverting

$C(W, R, Y, X, \eta^*) = 1$ to obtain

$$\eta^* = \eta^*(W,R,Y,X) \,. \qquad \textbf{(2)}$$

Using the Foster-Greer-Thorbecke (1984) class of poverty measures, we have[6]

$$P_\alpha(W, R, Y, X) = \int_0^{\eta^*(W, R, Y, X)} (1 - C(W, R, Y, X, \eta))^\alpha f(\eta) d\eta \,, \qquad \textbf{(3)}$$

where α is a nonnegative parameter for the measure's aversion to inequality among the poor. In the empirical work, we allow this parameter to take three possible values: α=0, giving the *head-count index*; α=1, the *poverty gap index*; and α=2, the *squared poverty gap index*. Only in the latter case does the measure penalize inequality among the poor.

We will also be interested in identifying any pro-poor effects on the distribution of relative consumption, as distinct from absolute levels. For this purpose, we construct a set of simulated poverty measures, in which the current poverty line is set at a constant

[6] One can also allow the distribution of land to be influenced by W and X. This would not change our estimated model.

proportion of the current survey mean. The poverty measures are thus purged of the direct effect of growth in average consumption, leaving only the effect via changes in relative consumption levels (Datt and Ravallion 1992). Such measures can also be thought of as examples of “relative poverty” measures, which is what we will call them, though recognizing that there are potentially many examples of such measures.[7]

Market clearing conditions can readily deliver a model in which the wage rate and price of food are functions of Y and X. Consider the labor market. Analogously to equation (1), the net labor supply function (labor hired out minus labor hired in) is

$$L = L(W, R, Y, X, \eta)\,, \qquad \textbf{(4)}$$

and we assume that the market clearing condition,

$$\int_0^\infty L(W, R, Y, X, \eta) f(\eta)\, d\eta = 0\,, \qquad \textbf{(5)}$$

can be solved for the wage rate as a function of R, Y, and X. However, as noted above, this is only one possible economic model that would deliver a wage rate that depends on these variables. That could also be justified by (for example) a model of imperfect competition.

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[7]For example, one might also include measures in which the poverty line varies positively with the mean, but with an elasticity less than unity. For further discussion and references, see Ravallion (1994).

A similar representation of the food market will deliver a relative price of food that is a function of *W, Y*, and *X*. Using the food market clearing condition to eliminate the food price from the wage equation, and labor market clearing to eliminate wages from the food price equation, we end up with reduced form models giving both *W* and *R* as functions of *Y* and *X*.

We do not assume that markets adjust instantaneously. Short-run stickiness is common in 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 tacit collusion or other forms of resistance on the supply side could also yield downward stickiness. Long-run wage responses to agricultural growth can thus exceed short-run responses.[8] So we take it that our economic assumptions describe the steady-state equilibrium, but that observed current wages and food prices can deviate from that equilibrium. We assume a first-order autoregressive process of adjustment to a new equilibrium. Finally we assume that all these relationships can be adequately approximated by an econometric model that is linear in logarithms.

Combining these assumptions, we estimate the following triangular system for the observed data:

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[8]This also has implications for the welfare effects of external trade liberalization when domestic food prices are below world prices; see Ravallion (1991) on Bangladesh.

$$
\begin{aligned}
\ln P_{\alpha t} &= {}_{\alpha 0} + {}_{\alpha 2}\ln W_t + {}_{\alpha 3}\ln R_t + {}_{\alpha 4}\ln Y_t + {}^{/}_{\alpha 5}X_t + {}_{\alpha t} \\
\ln W_t &= {}_0 + {}_1\ln W_{t-1} + {}_2\ln R_{t-1} + {}_3\ln Y_t + {}^{/}_{4}X_t + {}^{W}_{t} \qquad \textbf{(6)} \\
\ln R_t &= {}_0 + {}_1\ln R_{t-1} + {}_2\ln W_{t-1} + {}_3\ln Y_t + {}^{/}_{4}X_t + {}^{R}_{t},
\end{aligned}
$$

where the ${}_{\alpha i}$'s, ${}_i$'s and ${}_i$'s are parameters to estimate, and ${}_{\alpha t}$, ${}^{W}_{t}$, and ${}^{R}_{t}$ are assumed to be normally distributed white-noise error processes. The parameter ${}_{\alpha}$ can be interpreted as the "direct effect" of higher farm yields on poverty, "direct" in the sense that it controls for wages and food prices (and *X*); here, we expect that the main direct effects would involve higher (farm and/or nonfarm) employment for poor underemployed workers, or higher yields on the farms of the poor, though our data do not permit us to disentangle these effects. By contrast, the "indirect" (or "general equilibrium") effects are identified as those that entail changes in wages and prices, which will also depend on and ${}_3$.

3. ESTIMATES

DATA

We use 24 rounds of India's National Sample Survey (NSS), spanning 1958–94. The gaps between surveys range from 11 months to 5.5 years. Nor do all the surveys cover a full year. Thus, while we can estimate the wage and food price equations using annual data, giving 35 observations, we can only estimate the poverty regressions on 24 observations.

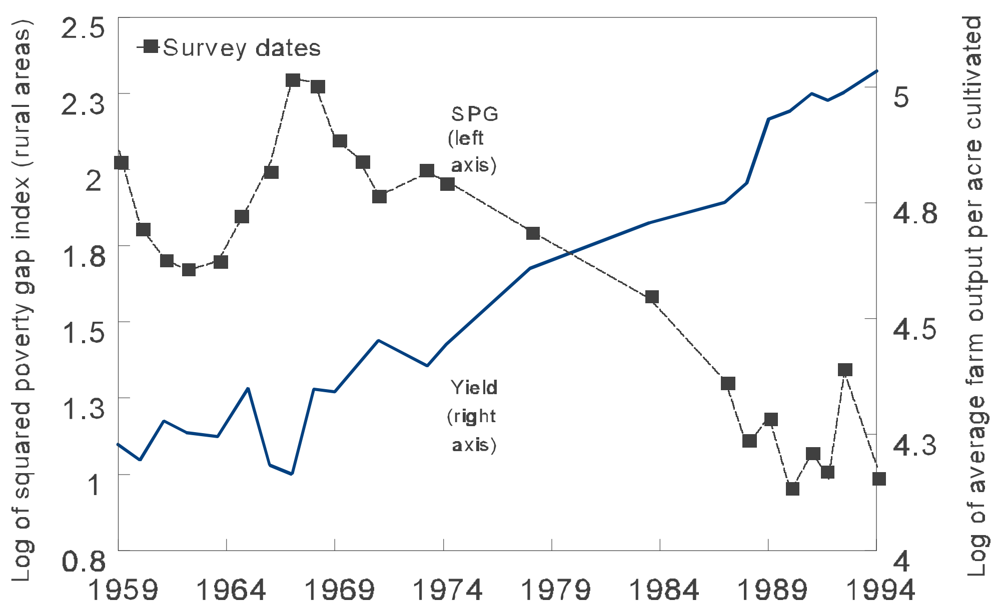
We calculate poverty measures based on distributions of household consumption expenditure (including from own production) per person. The poverty line is the one recommended by the Planning Commission's Expert Group (India 1993), and is set at a per capita monthly expenditure of Rs 49 at October 1973–June 1974 all-India rural prices. For the relative poverty measures, we fix mean consumption for each date at the overall mean for the whole period; this is equivalent to setting the poverty line for each date at a constant proportion of the mean for that date (specifically 83 percent of the mean). This purges the poverty measures of all effects of changes in mean consumption, leaving only distributional effects. However, since we are using a fixed-weight (Laspeyres) price index, we will not pick up substitution effects on the demand side, but only income effects.

We measure average farm yield (Y) by agricultural output per acre of net sown area, output being measured by a value-weighted quantity index of production of all crops. (We comment further on this choice in Section 5.) The wage data are for male agricultural wages. The price deflator for both the poverty line and wages is the Consumer Price Index for Agricultural Laborers (CPI), corrected for inflation and firewood prices (see Appendix for details). The index of the relative price of food was obtained by dividing the food component of the CPI by the value of the general index.

The appendix describes how we mapped the annual data on other variables into the NSS survey periods, and further details on data sources.

As can be seen from Figure 1, there is a high correlation between the squared poverty gap and average farm yield; the correlation coefficient is –0.88.[9] (On the reasons for the sharp increase in 1992, see Datt and Ravallion 1997a.) The contrast between the first and second halves of the period is borne out by the annual growth rates in Table 1. The second half saw significantly higher rates of decline in all poverty measures alongside higher growth rates for real mean per capita consumption, real agricultural

Figure 1 Squared poverty gap and average farm yield, rural India 1959-94



[9]The figure looks similar for the other poverty measures. The correlation is even stronger if one uses a moving average of farm yield; the correlation coefficient is then –0.90.

Table 1 Annual growth rates

	Average annual rate of growth (%) 1958–75	1976–94
Head-count index of poverty	1.18	–1.91
Poverty gap index	1.42	–3.65
Squared poverty gap index	1.65	–4.97
Mean real consumption per person	–0.93	1.76
Real agricultural wage rate	0.33	2.84
Agricultural output per acre of net sown area	1.51	2.91
Net sown area per person in rural areas	–1.53	–1.76
Agricultural output per head of rural population	–0.01	1.15
Consumer price index for agricultural laborers	7.17	8.11
Consumer price index for food	7.79	7.96
Head-count index (relative)	–0.49	–0.05
Poverty gap index (relative)	–1.07	–0.86
Squared poverty gap index (relative)	–1.45	–1.56

wages, agricultural yield, and nonagricultural output per person. In what follows, we try to sort out the relative importance of these factors.

RESULTS

In estimating equation (6), the *X* vector comprised the net sown area per head of rural population, the consumer price index for agricultural laborers for all commodities (CPI), and a time trend. All variables except the time trend were measured in natural logarithms. Both the current and lagged values of *W*, *R*, *Y*, and *X* were included, with up to two year lags in *Y*.[10]

Starting with this unrestricted specification, we were led to a restricted form for each poverty regression with just three explanatory variables, agricultural output per unit net sown area, the real wage rate, and the relative price of food. Both current and lagged values of yield and wages were retained. And the coefficients on both real wages and the agricultural yield variables were not significantly different from each other for any of the poverty measures, and so we imposed the restriction that they had the same coefficient. Thus, in addition to the relative price of food, the data suggest that a key explanatory variable is the (log of the) product of the real wage rate in agriculture and agricultural output per acre. One can interpret this variable as the product of average labor productivity in agriculture (output per worker) and agricultural labor earnings per acre, a measure of labor intensive in agriculture. We shall call this variable "wage-weighted farm yield."

Thus we are drawn to a model in which the poverty measures are determined by the productivity of labor in agriculture weighted by a measure of wage labor intensity, and the relative price of food. The estimates of this model are given in Table 2. We subjected all the regressions to a battery of standard residual diagnostic tests (serial correlation,

[10]The lagged values refer to values a year before the mid-point of the current survey period, and are estimated by interpolation using $X_{it-1} = (1 - (1/\ _t))X_{it} + (1/\ _t)X_{it-\ _t}$ (and similarly for other lagged variables), where is the time between the midpoints of survey periods. We resort to interpolation because the NSS survey periods do not coincide with the annual periodicity of the time-dependent variables, which are thus not centered at the mid-point of the survey periods.

normality, heteroscedasticity) and the model passed all tests at the 5 percent level or better. All parameter restrictions imposed on the initial model were tested and passed. We also estimated the model, treating both the current year's wage and yield as endogenous, and using lagged values up to two years and a time trend as instrumental variables. The results were very similar.

Our results in Table 2 indicate that both higher agricultural wages and higher yields reduce rural poverty, and that they do so with about the same elasticity. It is higher yields combined with higher wages that matter.

There is also an independent adverse effect of higher food prices. The elasticities to food price are high, though the relative price of food varied little over the period (in Table 1, note that the growth rates for the food price index are very close to those of the general index.) The elasticities tend to be higher as the value of α is higher, i.e., the more “distribution-sensitive” is the poverty measure.

Table 2 Regressions for rural poverty measures

	Head-count index		Poverty gap index		Squared poverty gap	
	Absolute measure	Relative measure	Absolute measure	Relative measure	Absolute measure	Relative measure
Constant	5.454 (49.1)	3.886 (51.6)	5.459 (33.3)	3.115 (25.2)	5.508 (24.9)	2.629 (15.2)
Real wage rate times farm yield (current plus lagged values)	–0.119 (12.5)	–0.001 (0.19)	–0.219 (15.5)	–0.042 (3.95)	–0.296 (15.6)	–0.080 (5.34)
Relative price of food	2.701 (8.86)	–0.309 (1.50)	3.991 (8.88)	–0.483 (1.42)	5.054 (8.33)	–0.498 (1.05)
R^2	0.954	0.112	0.965	0.441	0.964	0.608
Mean of dependent variable (Standard error of regression)	3.864 (0.043)	3.886 (0.023)	2.593 (0.064)	2.626 (0.048)	1.642 (0.086)	1.682 (0.067)
Parameter restrictions (Wald) Chi-Sq(11)	5.30	18.84	5.34	12.17	7.21	7.62

Note: Twenty-four observations spanning 1958–94. All variables in logs. Absolute t-ratios in parentheses. All regressions passed (at 5 percent level) tests of the residuals for normality (Jarque-Bera), heteroscedasticity (White), and autocorrelation (Breusch-Godfrey, up to two lags). The Wald test presents a test of parameter restrictions where the unrestricted model included current and lagged values of *W, Y,* net sown area per person, *CPI, R*, nonagricultural output per capita, and time.

To test whether or not there are distributional effects, we use the same specification, but switch to the simulated "relative poverty" measures described above. These behave quite differently. The food price effect vanishes (Table 2). (Note that it is not inconsistent to find that the food price elasticity is higher for higher values of α, yet it is zero for the relative poverty measures. We will return to this point in the next section.) For the head-count index, the wage-weighted productivity effect also vanishes when we switch to the relative poverty measure. However, this changes when we move to the higher order measures of relative poverty. Then a significant effect of the wage-weighted yield variable emerges (Table 2).

Is there also evidence of an effect of agricultural yields via the real wage rate? Turning to the wage equation in (6), we began with the same variables for *X* as in the poverty regressions, though now we can add the extra observations possible by using annual data. The restriction that the coefficients on the current and lagged (log) price level add up to zero was very comfortably accepted, so the inflation rate was used instead. However, aside from the lagged real wage and the inflation rate, no other variables in *X* were individually significant. Nor was the lagged food price significant. There was, however, a significant loss of fit when all were dropped. Yet log farm yield per acre was significant when the others were dropped. The restriction that the coefficients on current and lagged ln*Y* were equal passed easily. Table 3 gives the ordinary least squares (OLS) estimate; an IV estimate using only lagged values instrumental variables gave very similar coefficients and only slightly higher standard errors. The regression passed the same set

Table 3 Regressions for the real wage rate and relative price of food

	Wage rate	Food price
Constant	–0.641	0.060
	(4.04)	(3.096)
Lagged dependent variable	0.763	0.890
	(10.85)	(21.118)
Farm yield variable[a]	0.117	–0.016
	(4.24)	(3.791)
Inflation rate	–0.641	0.135
	(9.46)	(9.981)
R^2	0.980	0.955
Mean of dependent variable	1.554	–0.048
(Standard error of regression)	(0.032)	(0.006)

Note: Thirty-five annual observations, 1958–93. Absolute t-ratios in parentheses. All variables in logs. All regressions passed (at 5 percent level) tests of the residuals for normality (Jarque-Bera), heteroscedasticity (White), and autocorrelation (Breusch-Godfrey, up to two lags).

[a] For the wage equation, the farm yield variable is the sum of current and lagged output per acre. For the food price equation, it is current output per acre plus the lagged first difference of output per acre.

of residual diagnostic tests described above, as well as a Wald test for omitted variables (including the lagged food price and the full set of original variables in *X*).[11]

The dynamic effect in the wage equation is strong, and there is also a sizable short-run effect of inflation. The short-run farm-yield elasticity of the real wage is about 0.12,

[11]As an aside, it may be noted that one can also write our wage determination model in terms of first differences, augmented with an error correction term. The variables of interest—the logs of *W*, *CPI*, and *Y*—all had unit roots (using Augmented Dickey-Fuller tests). Regressing the log wage on the other two variables in levels, the residuals were found to be stationary, consistent with cointegration. However, the parameter restriction needed to yield the same model as in Table 3 performed very well.

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.

Are there also effects via food prices? Starting with the same set of *X* variables, and following the same model selection and testing procedure, we were drawn to the model in Table 2, with just two explanatory variables (in addition to the lagged food price), namely the sum of current agricultural yield and its lagged first difference, and the rate of inflation. The former variable is suggestive of an effect of yield on aggregate food supply (with lagged effects through storage behavior); the elasticity is, however, very small (probably reflecting governmental efforts to buffer food prices from shocks to farm output). There is also a significant positive effect of inflation on the relative price of food. This could well be due to an effect of the current agricultural wage on the relative price of food (recalling that we have solved out the current wage rate). There is a potential concern about endogeneity here, since the CPI appears on both sides of the equation. We also estimated the same equation treating the current level of the CPI as endogenous, using its lagged values over two years as an instrument. The results were very similar to Table 3; the coefficient of the inflation rate was slightly higher (0.147) and remained highly significant (a t-ratio of 4.97), while other coefficients and standard errors were almost unaffected.

We also checked for a common omitted variable determining both real wages and food prices, by testing for a correlation between the residuals of the two equations.[12] There was no sign of a common omitted variable; the correlation coefficient between the residuals was 0.12.

4. IMPLICATIONS

The above results suggest that India's poor gained)both absolutely and relatively)from both higher agricultural wages and higher average farm yields, and with the same elasticity. Lower food prices also helped. There is evidence of an effect of higher farm yields on both the wage rate and food price. This section will examine the implications of these results for understanding the total effect of agricultural growth on rural poverty.

ELASTICITIES OF RURAL POVERTY TO FARM YIELD

Since agricultural wages and food prices do not adjust instantaneously to yield gains, short-run responses of the poverty measures will differ from long-run responses. The short-run (contemporaneous) elasticity of the poverty measure to a change in yield is given by

[12] Given the difference in the data sets used, there is no straightforward test for common omitted variables between the poverty regressions and either of the wage and price equations.

$$\frac{\partial \ln P_{\alpha t}}{\partial \ln Y_t} = \quad _{\alpha 2} \quad _3 \; + \quad _{\alpha 3} \quad _3 \; + \quad _{\alpha 4} \, . \qquad \textbf{(7)}$$

We shall refer to the last term on the right-hand side (RHS) as the "direct effect," while the first two terms are the "indirect effects," via real wages and relative food prices, respectively. Using "*" to denote steady-state values, the long-run elasticity of the poverty measure to yield is[13]

$$\frac{\partial \ln P_{\alpha}^{*}}{\partial \ln Y^{*}} = \frac{2 \quad _{\alpha 2} \quad _3}{(1 - \quad _1)} + \frac{\quad _{\alpha 3} \quad _3}{(1 - \quad _1)} + 2 \quad _{\alpha 4} \, . \qquad \textbf{(8)}$$

Again, the first two terms on the RHS are the cumulative indirect effects operating through the wage rate and food prices, respectively, while the last term is the direct effect.

Table 4 gives the elasticities implied by our parameter estimates from Tables 2 and 3. In the short run, the indirect effect via the real wage is small, and dominated by other channels, particularly the direct effect. But in the long run, the wage effects account for about 30 percent of the total elasticity. The indirect effect via the relative price of food is even higher. So the direct effect dominates in the short run, while the indirect effects via wages and prices dominate in the long run. This reflects, in part, the high degree of serial

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[13]This formulae assumes that $\quad _2 = \quad _2 = 0$, which was accepted as a parameter restriction on (6). Notice also that it is the sum of the current and lagged ln *Y* on the right-hand side of both the poverty and wage regressions. Thus, one collects two terms in ln *Y* in the steady-state.

Table 4 Elasticities of rural poverty to average farm yield

	Short-run				Long-run			
	Direct effect	Via wage rate	Via food price	Total short-run	Direct effect	Via wage rate	Via food price	Total long-run
Head-count index	–0.119	–0.014	–0.044	–0.177	–0.239	–0.237	–0.400	–0.875
	(67.3)	(7.9)	(24.8)	(100.0)	(27.3)	(27.0)	(45.7)	(100.0)
Poverty gap index	–0.219	–0.026	–0.065	–0.309	–0.437	–0.434	–0.591	–1.462
	(70.7)	(8.3)	(21.0)	(100.0)	(29.9)	(29.7)	(40.4)	(100.0)
Squared poverty gap index	–0.296	–0.035	–0.082	–0.413	–0.592	–0.587	–0.748	–1.928
	(71.7)	(8.4)	(19.9)	(100.0)	(30.7)	(30.5)	(38.8)	(100.0)

dependence in real wages and the relative price of food (Table 3). Total elasticities are about five times higher in the long run.

WHY ARE ELASTICITIES HIGHER FOR HIGHER VALUES OF α?

It can be seen in Tables 2 and 4 that the (absolute) elasticities of the poverty measures to farm yield gains are higher for higher values of the parameter α. To see what this means, note that the Foster-Greer-Thorbecke poverty measures can also be written in the form

$$P_1 \;=\; P_0(1-\mu^p/z); \qquad P_2 \;=\; P_1\left[1+\mu^p/z \;+\; \frac{(\;\;^p/z)^2}{1-\mu^p/z}\right], \tag{9}$$

where μ^p and $\;^p$ 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 ($1 - \mu^p/z$ falls). Furthermore, inequality among the poor—as measured by the CV ($^p/\mu^p$)—must be decreasing as average farm productivity increases.[14] Thus, with growth, there is also an improvement in distribution among the (changing number of) poor. Note too that although we do find significant yield and wage effects on relative poverty, overall distribution neutrality (in that relative poverty measures are unchanged) is not inconsistent with higher impacts on absolute poverty as α rises. For example, all consumption levels could increase at the same rate while the mean consumption of those below the poverty line also rises; whether it does or not depends on the density of consumptions in a neighborhood of the poverty line.[15]

5. ANTECEDENTS

Much of the scholarly debate on agricultural growth and rural poverty in India started with 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. Subsequent papers questioned (among other things) sensitivity to changes in the period of analysis (Griffin

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[14]Since μ^p is increasing with yield, a higher elasticity for P_2 than P_1 must imply that ^{p} is falling.

[15]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 $f(z)/P_0 < {}^{\mu g}/z$ where $f(z)$ is the density of consumption at the poverty line z.

and Ghose 1979; Saith 1981). Ahluwalia (1985) and Bell and Rich (1994) returned to the Ahluwalia regressions, adding data for another year (1977/78), and broadly confirmed his conclusions.[16]

With so few observations, and relatively little sustained agricultural growth over the period, the Ahluwalia series was clearly not ideal for this purpose (Srinivasan 1985). Our doubling of the number of observations over the original Ahluwalia series, and the substantial growth that has occurred since the late 1970s, has allowed us a more powerful test for the effects of agricultural growth on rural poverty.

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 included the Sen (1976) index, but did not draw out implications concerning the depth of impacts on the poor. Our results indicate differences in the impact of agricultural growth between different measures, arising from changes in distribution below the poverty line. We find no support for claims that productivity gains in Indian agriculture by-passed the poorest.

Past work has not tested for pro-poor distributional effects (which, as we have noted, is a different thing to testing impacts on higher-order poverty measures). Our approach, using simulated measures to isolate distributional shifts, has an antecedent in the method used by Datt and Ravallion (1992) to decompose changes in poverty into growth

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[16]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.

and redistribution components. We cannot explain the distributional component of the head-count index, but we do find that wages and yields can help explain the distributional components of the higher order measures. Higher wages and yields entail an improved distribution from the point of view of poverty, although the bulk of their effect (in terms of the elasticities in Table 2) is clearly through the growth component.

Turning to the explanatory variables, past studies have followed Ahluwalia in using farm output per head rather than per acre, which we have used. The log of output per person is simply the sum of the logs of output per acre and acres per person, and we initially included both variables. The fact that we can reject the null that they have the same coefficients means that farm output per person is not the relevant variable for predicting poverty. Although output per person and output per acre are highly correlated (the correlation coefficient is 0.89 over the 35 annual observations), the latter variable is clearly the better predictor of how much the poor share in agricultural growth. In interpreting this finding, it should be recalled that we do not have either farm or nonfarm employment on the RHS (we return to this point). Controlling for wages and prices, it may well be that yield per acre is the better predictor of the underlying employment effects of agricultural growth; for example, one would expect it to better reflect the potential for multiple cropping in a given year, a key determinant of the demand for agricultural labor. Our finding that lagged output matters, and it has a similar effect as current output, echoes Ahluwalia (1985).

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. Bardhan (1984, Chapter 14) found a significant poverty-reducing effect of higher real agricultural wages in cross-sectional data for West Bengal. We know of only one other study that looked at the impact of the real wage rate on the evolution over time of rural poverty in India, namely van de Walle (1985), who used the earlier Ahluwalia poverty measures for 1959–71. However, 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.

The level of the nominal price index has figured prominently in past work. Saith (1981), Narain (see Desai 1985) and others (Mathur 1985; Gaiha 1989) added the level of the CPI to the original Ahluwalia (1978) model, and this sparked further debate.[17] 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 that 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 suspect that the significant price level effect identified in the literature reflects an omitted variable bias, the key omitted variable being the real wage

[17]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.

rate. While we have allowed effects of price levels, this proved to be insignificant once the real wage was added.[18] (Notice also that we included current and lagged price levels in the Wald test for omitted variables reported in Table 2.)

However, we do find effects of inflation on poverty via wages and food prices. The inflation rate (as distinct from the price level) also appears to be an important omitted variable in past work on the determinants of rural poverty in this setting. For example, if one reestimates our wage regression, dropping the rate of inflation and adding the log of the CPI, then the latter has a significant negative coefficient (–0.17 with a t-ratio of 3.49); yet this is clearly spurious, for as soon as one returns the rate of inflation to the same regression, the effect of the price level vanishes (the coefficient drops to 0.04, with a t-ratio of 0.93, while the inflation rate is highly significant (a coefficient of –0.70, with a t-ratio of 7.49).[19]

We remain somewhat puzzled as to why there is such a strong adverse effect on the poverty measures of higher food prices. Following the argument in the introduction, we expected this to be a distributional effect. Our tests do not support that view (Table 2), though data limitations (notably the use of a single fixed-weight deflator) entail that we can only expect to pick up any distributional effects on the income side, leaving out any effects via demand behavior. Nonetheless, mean rural consumption is strongly negatively

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

[19]One can also find a highly significant, but spurious, effect of the relative price of food on real wages if one does not control for the rate of inflation.

correlated with the relative price of food (the simple correlation coefficient over the 24 survey rounds is 0.82.) That is surprising, given that (in this largely closed food economy) the rural sector as a whole produced more food than it consumed. However, it should be noted that there was no significant trend over the whole period in the relative price of food; its ability to help explain the poverty measures appears to be due to a correlation with the fluctuations in poverty over time. The explanation may lie in common shocks to food supply and/or savings behavior.

It has been argued that nonagricultural employment was also an important factor in reducing rural poverty in India. Presumably this was to some extent the result of agricultural growth (Hazell and Haggblade 1993), though it has also been argued that there was an independent effect (Sen and Ghosh 1993; Sen 1996). Testing for an independent effect of nonfarm employment is difficult, given that most NSS rounds did not collect data on rural employment, and there is only sporadic data from other sources.[20] We tested (current and lagged) nonagricultural net domestic product (urban plus rural) per head of rural population as an additional explanatory variable for the absolute poverty measures in Table 2, but it was insignificant and had little effect on other coefficients. We repeated this test dropping the real agricultural wage rate from the poverty regressions, to

[20]Sen (1996) presents regressions for the head-count index (using an earlier version of our data set that we provided) in which variables for "nonagricultural employment" and "commercialization" appear on the right-hand side. However, no information is supplied in his paper as to where these data come from, and no reply was received to our letter asking for sources. We assume that the nonagricultural employment series was created by some sort of interpolation. Substantial interpolation would have been required, given the sparse employment data.

see if there is an indirect effect via wages; again, nonagricultural output per capita was (highly) insignificant.

We have examined this issue further at the state level in India, and found that deviations from trend in (state-specific) nonagricultural output do help explain the fluctuations over time in the poverty measures. However, we find that differences between states in the trend of nonagricultural output growth had no power to explain differences in the trend rate of poverty reduction, once one controls for the trend rates of farm yield growth and initial conditions related to physical and human infrastructure (Datt and Ravallion 1997b).

Our analysis at the state level also allowed for an independent effect of government spending, but did not find a reasonably significant effect on either the trend rates of poverty reduction or deviations from trend (though the effect was mildly significant in the latter case for the head-count index). However, we did find evidence of an indirect effect via higher farm yields. To test this, Datt and Ravallion (1997b) regressed the agricultural yield on the other explanatory variables in their model, including public spending at state level. The latter had a significant positive impact; yield had an elasticity of 0.29 (t-ratio=3.18). This suggests that state development spending has helped reduce rural poverty through its impact on average farm yield. There could well have also been effects via rural nonfarm employment; we did find significant effects on nonagricultural output. However, we were also unable to reject the hypothesis that growth in agricultural yields and nonagricultural output Granger-caused higher development spending; lagged agricultural yield and nonagricultural output had significant positive effects on state development spending. Further work is needed to disentangle these effects.

There has also been concern in India that the employment elasticity of agricultural growth has been falling, and (hence, it is argued) that agricultural growth is now less poverty reducing than it used to be.[21] What do our data and models suggest? Although there is not a suitable time series of agricultural employment, we can directly test for a trend decline in the farm-yield elasticity of our poverty measures. To do so, we added an interaction effect between time (mid-point of the NSS round, in years) and the wage-weighted farm yield variable (Table 2). The interaction effect was insignificant for all three poverty measures, both absolute and relative. We do not find a trend decline in the yield elasticity of rural poverty in India.

Aside from the specification of the RHS variables, studies in the literature have differed in their assumptions (often implicit) about dynamics, and the properties of the error term. Virtually all of the poverty regressions in this literature are static. The only 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 (1983). The shortage of surveys for estimating the poverty measures will entail a small-

[21]For a recent argument, see Sen and Ghosh (1993), although their own results do not appear to be conclusive on this point.

sample bias in dynamic models. We tested for an independent dynamic effect, by adding the lagged poverty measure to the regressions in Table 2.[22] This was insignificant. Note, however, that there are still dynamic effects via wages and prices in our model. We also found that if we dropped the relative price of food from the right-hand side, then the poverty regressions called for either a lagged dependent variable or an autoregression (AR1) correction to the error term. Significant dynamics in a poverty regression that excludes this variable can thus be attributed to omitted variable bias.[23]

Serial dependence in wages and prices is more likely to be a structural feature, reflecting stickiness in wage and price adjustment. Some past time series models of real wage determination have been static.[24] Omitting lagged wages in our model leads to substantial overestimation of the short-run effect of agricultural productivity on wages and underestimation of the long-run elasticity, and (of course) the residuals were highly serially dependent (a Durbin-Watson statistic of 0.77). 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 also considerably higher than

[22]The test allowed for uneven spacing by estimating the model under a common factor restriction. This is equivalent to estimating a static model with an AR1 error term, in which the usual autoregression coefficient is raised to the power of the elapsed time between surveys.

[23]In an earlier working paper version (summarized in Ravallion and Datt 1996b), we had omitted the relative price of food from the poverty regressions, and so were drawn toward a dynamic specification. Comments on that working paper by Sen (1996) led us to test the effect of adding the relative food price. Note, however, that the conclusions we had drawn in Ravallion and Datt (1996b) about dynamic responses of poverty measures to agricultural growth remain valid, given the dynamic behavior of both food prices and wages in response to agricultural growth.

[24]For example, Acharya and Papanek (1992) estimate models of agricultural wage determination for India using time series data, but do not allow for dynamic effects.

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 (real wages and farm yields are cointegrated in our data) entails that his model does not have a long-run solution, and it appears to impart an upward bias on the short-run yield effect on wages.

It is of interest to see what a poverty regression consistent with the Ahluwalia specification looks like on our new data. The Ahluwalia specification gives

$$\ln P_{0t} = \underset{(2.45)}{12.94} - \underset{(3.95)}{0.72}(\ln YPC_t + \ln YPC_{t-1}) - \underset{(0.37)}{0.001}\,t \,, \tag{10}$$

where YPC is agricultural output per capita. The R^2 is 0.76, 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.[25] Notice that Ahluwalia's specification gives a higher elasticity to agricultural output, namely –1.44, as compared to our long-run estimate of –0.88 (Table 4). The main reason for this appears to be our use of output per acre rather than output per person; when we rerun the Ahluwalia specification using our yield variable, the long-run elasticity becomes –0.92 (t=5.24). The Narain-Saith specification gives

$$\ln P_{0t} = \underset{(2.80)}{96.92} - \underset{(2.73)}{0.82}\ln YPC_t + \underset{(2.30)}{0.57}\ln CPI_t - \underset{(2.54)}{0.047}\,t \,. \tag{11}$$

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[25]The Durbin-Watson statistic is 0.76, and the Breush-Godfrey test with two lags indicates significant serial correlation at the 1 percent level.

The R^2 is 0.69, though there is still significant positive serial correlation in the residuals. 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, price level, and relative price of food to our preferred model in Table 2, we can do a nested test against the Ahluwalia-Narain-Saith models. All three variables were highly insignificant (individually and jointly); a Wald test gave a Chi-square of 1.14. Our model fits the data better.

6. CONCLUSIONS

We have estimated the effects of farm yield growth in rural India on various poverty measures, real agricultural wages, and the relative price of food, using data spanning the period 1958–94. We find that higher real wages and higher farm yields reduced absolute poverty, and with about the same elasticity. A composite variable—wage earnings per acre times average product per worker—has remarkably strong predictive power for various measures of absolute poverty. When we turn, instead, to measure relative poverty—by fixing the poverty line as a constant proportion of the current mean, we find significant, but much smaller effects, below the poverty line. We also find that the poor gained in absolute terms from lower relative prices of food. However, this effect is not evident in measures of relative poverty. Our results suggest that the bulk, but not all, of the gains to the poor from higher farm yields and higher real wages were via rising average living standards rather than improved distribution. Even so, the gains to the poor from higher average yields were not confined to those near the poverty line, but reached deeper.

We also find evidence of important indirect channels linking average farm productivity to living standards of the rural poor. There is evidence that real agricultural wages responded positively to higher farm yields, presumably through effects on labor demand, such as due to multiple cropping. There is also a strong link through food prices. While the impact of agricultural growth on food prices is quantitatively small, even small food price changes can have large effects on absolute poverty. Inflation also had adverse effects on the poor, via its short-term effect on real wages and the food prices.

Neither the real wage rate nor the price of food adjusted instantaneously to higher yields. The combined effect of this wage and price stickiness is that the short-run gains to poor people of higher farm productivity are far lower than the long-run gains. Indeed, the short-run effects operating via wages and prices are minor compared to those through other channels. But in the long run, the dynamic general equilibrium effects account for about 70 percent of the steady-state elasticity of absolute poverty to a farm yield increase. Overall elasticities in the long run are five times higher than the short-run values, and long-run elasticities are about one for the head-count index of poverty, rising to about two for the squared poverty gap index.

APPENDIX

DATA SOURCES

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 50th round for July 1993–June 1994. In keeping with data available 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 50 (July 1993–June 1994). 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 (India 1979), and was later also endorsed by the Expert Group on Estimation of Number and Proportion of Poor (India 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). Using the formulae in Datt and Ravallion (1992), point estimates of poverty measures are calculated using data on mean per capita consumption and the estimated parameters of fitted Lorenz curves. Either the Beta Lorenz specification (Kakwani 1980) or the General Quadratic specification (Villasenor and Arnold 1989) were used, depending on which fitted best; both satisfied the theoretical conditions needed for a valid Lorenz curve in all survey rounds. A complete series of the poverty measures can be found in Özler, Datt, and Ravallion (1996), Datt (1997) and World Bank (1997).

THE PRICE INDICES

The CPI for agricultural laborers (CPIAL) 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 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. (The NSS values firewood consumption from own-production at local market prices. Also, see Minhas et al. 1987 for further discussion.) To deal with this problem, we estimated a new deflator, replacing the firewood subseries 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); see Datt (1997) for further discussion of this adjustment to the CPIAL.

The index of the relative price of food is defined to be the ratio of the food component of the CPIAL to the above (adjusted) CPIAL for all commodities.

AGRICULTURAL WAGES

Our data on nominal daily male 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)* [India 1991], and *Agricultural Wages in India* reports since 1984–85. (A complete time series is not available for women.) The primary source of all these data is the Ministry of Agriculture's annual publication, *Agricultural Wages in India* (AWI). These data were aggregated to derive all-India wage rates using state- and year-specific weights, as described in Özler et al. (1996).

AGRICULTURAL OUTPUT AND AREA

These data are collated from various issues of the publication, *Area and Production of Principal Crops in India*, produced by the Ministry of Agriculture. The data are in the form of three annual indices: (1) 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. (Alternatively, one can use value added in agriculture, though this makes little difference, since it is highly correlated with output; the correlation coefficient is 0.97 in logs); (2) 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; (3) the index of net sown area also with the same base period. All three indices refer to the agricultural year from July to June.

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 that coincides with the mid-point of the corresponding agricultural year.

MATCHING NON-NSS DATA WITH NSS ROUNDS

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 (1) 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 (2) a weighted average of the values for agricultural years overlapping with the survey period of that round.

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