Does Better Nutrition Raise Farm Productivity?

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Household-level data from Sierra Leone are used to test whether higher caloric intake enhances family farm labor productivity. This is the notion behind the efficiency wages hypothesis, which has found only weak empirical support. A farm production function is estimated, accounting for the simultaneity in input and calorie choice. Instruments include prices, household demographic characteristics, and farm assets. The latter two sets of instruments are later dropped to explore the robustness of the results to different specifications of exogeneity. The exercise shows a highly significant effect of caloric intake on labor productivity, providing solid support for the nutrition-productivity hypothesis. The marginal effect on productivity falls drastically as calorie consumption rises but remains positive at moderately high levels of intake. One result is a fall in the effective price of food, a decline that is larger for households that consume fewer calories.

I. Introduction

The potential biological relationship that relates current and past nutrition intakes to labor effort per unit of time, or efficiency units of labor, has attracted the interest of economists and nutritionists for some time. Economists have been especially interested in how labor

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markets might adapt to such a relationship, an interest that has spawned the efficiency wages hypothesis. Developed by Leibenstein (1957) and Mazumdar (1959), it has been formalized and extended by Mirrlees (1975), Rodgers (1975), Stiglitz (1976, 1982), Bliss and Stern (1978*a*), and Gersovitz (1983).[1] This hypothesis has been used to explain (1) why constant real wages in the agricultural sector of a developing economy are part of an equilibrium with involuntary unemployment, (2) the distribution of food within a household, (3) household savings decisions, and (4) shadow wage rates. More recently this idea has even been offered as an explanation of involuntary unemployment in industrial countries (Yellen 1984). While other nonbiological relationships between wages and efficiency units of labor have been suggested,[2] it is the relationship between nutrient intake and labor productivity that remains the primary motivation for the efficiency wages hypothesis as it is applied to developing countries.

Despite this body of theory, the empirical evidence on the existence and shape of a function that relates nutritional status to labor productivity is not abundant, and it is especially lacking for farm labor productivity. Nutritionists and doctors have been interested in this question since before the 1920s. Experiments conducted in Minnesota during the 1940s (Keys et al. 1950) have shown that activity levels drop precipitously when males are subjected to dramatic decreases in caloric intakes from moderate to extremely low ones.[3] What is not clear from this evidence is whether people would not adapt fully to long exposure to low caloric intakes without a decrease in productivity.[4] Also it is not clear whether at higher levels of caloric intake, corresponding to a larger proportion of the developing world, similar relationships would hold.

[1] Bliss and Stern (1978*a*, 1978*b*) provide an excellent survey, as do Binswanger and Rosenzweig (1984).

[2] For instance, Stiglitz (1982) hypothesizes a morale effect of higher wages, and Weiss (1980) hypothesizes a potential screening effect, that more productive workers have a higher opportunity cost in self-employment activities.

[3] The daily caloric intakes of 32 men were reduced from 3,500 calories to 1,500 over a 24-week period, then increased to 1,800 calories per day. Productivity was not directly measured, but vastly reduced activity levels were one of the adaptive mechanisms observed.

[4] Sukhatme and Margen (1982) argue that over a range of caloric intakes the efficiency with which calories are used by humans may vary positively with the level of intake, gradually adjusting as intake levels change so as to reequilibrate energy intake with energy expenditure. The best evidence on this comes from the Minnesota semistarvation experiments (Keys et al. 1950), which do show an adjustment in efficiency of calorie use by the body, but not so complete as to prevent a major reduction in activity levels. However, evidence on long-term bodily adaptation at higher levels of caloric intake is extremely scanty.

In other, nonexperimental, evidence from this period Kraut and Muller (1946) report increases in hourly productivity of German coal miners, steelworkers, and workers dumping debris out of railcars when their daily food rations were exogenously increased in special work camps. There were no control groups, however, and no nonlabor inputs (or institutional changes) were measured, so some caution has to be used when interpreting the findings. Also the workers presumably knew they were getting the extra rations, which were not tied explicitly to labor supply, so a morale effect is possible.

More recent empirical evidence trying to answer these general questions has severe methodological problems and shows weak or no patterns.[5] One exception is an experimental study (Wolgemuth et al. 1982) that shows a relationship (significant at the .075 level) between current energy supplementation and output of male Kenyan road construction workers.[6] Other less carefully constructed experimental studies have shown no positive effects of current energy intakes on worker productivity (e.g., Immink and Viteri 1981*a*, 1981*b*).[7]

A number of nonexperimental studies exist, but they are plagued by simultaneity problems. Even the experimental studies report certain evidence that is similarly affected. Typically ordinary least squares (OLS) regressions are run that relate individually measured output, often of sugarcane cutters or road construction workers, to such endogenous variables as calorie consumption (Immink and Viteri 1981*a*, 1981*b;* Immink, Viteri, and Helms 1982), blood hemoglobin levels (Popkin 1978; Wolgemuth et al. 1982), dummy variables for current illnesses (Wolgemuth et al. 1982), or weight-for-height mea-

[5] Spurr (1983), Martorell and Arroyave (1984), and Latham (1985) provide recent surveys of the nutrition literature. Strauss (1985) gives a far more detailed critique than is given here.

[6] Dirt dug per day increased 12.5 percent for workers with high-calorie supplements (1,000 calories per day) vs. workers with low-level supplements (200 calories per day). The presupplementation daily caloric intake of both groups was approximately 2,000 calories. Because less food was consumed at home, the net increase in daily caloric intake was only 500 calories for the highly supplemented group and almost none for the low-level supplemented group. Workers were randomly assigned to groups; however, attrition was high (the sample size is only 47 workers). Since workers knew which group they were in, it is possible that any selectivity in attrition might have been different for the two groups.

[7] Immink and Viteri (1981*a*, 1981*b*) studied Guatemalan sugarcane cutters, who were provided with a low- and a high-calorie supplement. The different supplements were given to entire villages; thus random assignment of workers to groups was not achieved. Changes in productivity between the two groups showed no differential response, any real differential being swamped by village and seasonal factors. Experimental evidence on iron deficiency as a possible cause of low labor productivity is stronger. Basta et al. (1979) find that anemic Indonesian rubber tree tappers who received an iron supplement were able to catch up in productivity to nonanemic workers, both those receiving a supplement and those receiving a placebo.

surements[8] (six such studies are approvingly cited by Martorell and Arroyave 1984). Using a slightly different twist, Baldwin and Weisbrod (1974) and Weisbrod and Helminiak (1977) regress daily and weekly wages of St. Lucia plantation workers on a set of parasitic disease dummy variables. While the expected positive sign is usually found between productivity and the nutrition or health measures, the attributed causation is in doubt. Unobserved production or earnings function shifters, such as ability or land quality, will also shift food consumption (hence caloric and iron intake) through associated income changes. Health and nutritional status outcomes such as illness and weight for height depend on current and past inputs such as food consumption and time allocation and so will also be endogenous.[9]

The evidence collected to date is overwhelmingly on workers whose individual outputs are easily observed, such as sugarcane cutters or dirt diggers on road construction crews. Labor productivity–nutrition linkages of workers on family farms have been largely unexplored, despite the overwhelming importance of family farms in developing countries.[10] This gap may result from the severe data requirements of such an exercise,[11] including the necessity of having to infer labor productivity instead of directly measuring it. Indeed Bliss and Stern (1978*b*, p. 390) in discussing such a possibility conclude: "We should not be dogmatic. We suggest, however, that an attempt to tease something out of the data, which is much more delicate than the crude production function, with all the problems attendant to that simple exercise, will not be justified." Nonetheless, the gap is sufficiently serious and the hypothesis important enough to warrant more study.

This paper reports an attempt to test and quantify the effects of current nutritional status (measured by annual caloric intake) on annual farm production and, hence, labor productivity using farm

[8] Weight for height is often taken as a proxy for current nutritional status and is sometimes hypothesized to play an independent role in raising biological maximum work capacity (Spurr 1983).

[9] Strauss (1982) and Pitt (1983) provide recent evidence of food consumption responses to income and prices for households in poor countries. Pitt and Rosenzweig (1986) report reduced-form responses of adult illness among Indonesian households to food prices and community socioeconomic variables.

[10] A recent exception, in addition to this paper, is Deolalikar (1984). Viteri (1971) reports that a group of Guatemalan farm workers who had a calorie supplementation for 3 previous years could accomplish standardized farm tasks in less time than an unsupplemented group. This study suffers from the same simultaneity problems as discussed previously. The supplemented workers all came from the same farm in a higher-income region, which paid higher than average wages, while the unsupplemented workers were all from one of the poorer areas in Guatemala.

[11] Not only are traditional farm input-output data needed, but also data on nutrient intakes, possibly anthropometric measurements, and potential instrumental variables. Prices are an obvious source of such instruments, which require intertemporal and/or regional variation.

household level data from Sierra Leone. Farm work there—hoe agriculture—is physically demanding. It is thus a good setting in which to test whether nutrition affects labor productivity. This is done by estimating a farm production function in which caloric intake may enhance efficiency units of labor, while accounting for the simultaneity involved in household choices of calorie consumption and in levels of variable farm inputs. A farm household model is outlined, both to motivate the choice of instruments and to provide first-order conditions with which the parameters can be given a more meaningful interpretation. The instrument set includes prices of commodities consumed, output, and variable farm inputs; quasi-fixed farm inputs; and household demographic characteristics. While treating fixed inputs, and perhaps demographics, as predetermined is common when estimating production, cost, or profit functions (Lau 1978), unobserved farm heterogeneity resulting from land or management quality differences may lead to systematic differences between households in levels of those predetermined inputs. The estimates are therefore examined to see how robust they are in reducing the instrument set, and the hypothesis is tested of no correlation between the quasi-fixed factors and random production function disturbances.

The results show a highly significant and sizable effect of caloric intake on farm output, even after accounting for its endogeneity. The effects are greatly attenuated as calorie levels rise but remain positive over a large range of caloric intake. These measured productivity effects can be interpreted as lowering the shadow price of food below the market price. At the sample mean the reduction for rice (the staple) is computed to be between 20 and 40 percent, rising substantially for low-caloric-intake households. Moreover, both the significance and size of the calorie effects are reasonably robust to the ways in which calories enter the production function, to the inclusion of other human capital related variables, to different assumptions concerning the substitutability of family and hired labor, and to assumptions concerning the exogeneity of certain of the instrumental variables.

II. Model

A farm household model (see Singh, Squire, and Strauss 1986) slightly modified to allow nutrient intakes to affect farm output is used to represent household behavior. Households are assumed to choose a consumption bundle of foods (X_a), nonfoods (X_n), and leisure (X_l), input levels of effective family (L_*^f) and hired (L_*^h) labor, and nonlabor variable farm inputs (V) to maximize their utility subject to a farm production function, time, and budget constraints. Since the caloric consumption that potentially matters is at the individual level, a model explaining food consumption of individuals would be better. However, since the available data are at the household level, this is not pursued.

Farm output, Q, is hypothesized to be a function of effective hours of family (L_*^f) and hired (L_*^h) labor, variable nonlabor inputs (V), fixed capital (K), and land cultivated (A):

$$Q = F(L_*^f, L_*^h, V, K, A). \tag{1}$$

Effective labor, both family and hired, is a function of caloric intake (X_c^f, X_c^h) at the individual level and hours worked (L^f, L^h). Individual-level caloric intake in turn is a function of household food consumption, a function that depends on intrahousehold distribution and biological food-calorie conversion rates. It is the inflow of calories during the current year that is hypothesized to affect annual effective labor. No attempt is made to measure effects of deficiencies that occurred long ago, a stock effect, though to the extent that current and past intakes are correlated the joint effects are being captured. Family and hired labor are hypothesized to have the same effective labor function, although they may be at different points on the function because their intakes are different. In specifying effective labor we follow the efficiency wages literature (Bliss and Stern 1978*a*, 1978*b*) by making effective labor the product of labor hours and a function relating efficiency per hour worked to caloric intake:[12]

$$L_*^i = h(X_c^i)L^i, \quad i = f, h. \tag{2}$$

The efficiency per hour worked function, $h(\cdot)$, is often hypothesized to have a portion that is increasing at an increasing rate followed by a portion increasing at a decreasing rate. It can begin at the origin or from a positive caloric intake. Figure 1 provides an illustration.

Competitive markets are assumed to exist for all commodities. In principle, provided that a nutrition-productivity relationship exists and perfect information on it exists for both employees and employers, wage per effective hour (the efficiency wage), not clock hour, would be taken as given by family and hired laborers. The associated full income constraint can be written as

$$\begin{aligned} w_*^f h(X_c^f)T + (p_a Q - w_*^f L_*^f - w_*^h L_*^h - p_v V) + E \\ = p_a X_a + p_n X_n + w_*^f h(X_c^f) X_l, \end{aligned} \tag{3}$$

[12] For simplicity different types of family or hired labor, such as male adult and female adult, are aggregated. In principle each might have a different function relating efficiency per hour worked to caloric intake.

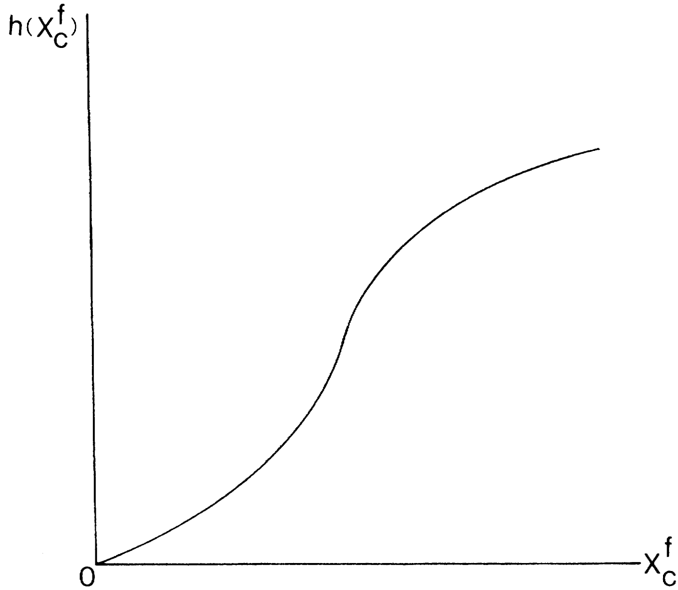

FIG. 1.—Prototype efficiency labor function

where w_*^f is the wage per hour of effective labor hired out—thus $w_*^f h(X_c^f)$ is the wage per clock hour, assuming for simplicity the same efficiency labor function to apply to labor hired out and to family farm labor—w_*^h is the hired-in wage per effective hour, the p's are prices, T is total household clock time available,[13] and E is any non-labor, nonfarm income. The term in parentheses represents the returns to quasi-fixed farm inputs, or profits.

From the first-order conditions it is clear that the real marginal price of foods is less than the market price to the extent that on-farm (and off-farm) labor productivity varies positively with caloric intake.[14] Also the marginal value product of efficiency labor, family or hired in, is equated to the efficiency wage for labor hired out or hired in. Other conditions are standard.

For the purpose of estimating the farm production function, the agricultural household model provides a set of variables that may be

[13] Following Grossman (1972), time available could be modeled as non–sick time available, where morbidity depends in part on nutrient intake. This is not followed here because the necessary data are unavailable.

[14] With interior solutions this condition is

$$\frac{\partial U}{\partial X_a} - \lambda p_a\left[1 - L^f \frac{\partial F}{\partial L_*^f}\frac{dh}{dX_a} - \frac{w_*^f}{p_a}(T - X_l - L^f)\frac{dh}{dX_a}\right] = 0,$$

where $U(\cdot)$ represents the utility function and λ the Lagrange multiplier.

taken as exogenous or at least predetermined to the household, hence that are candidate instrumental variables. These variables can be grouped into prices, farm assets, and quasi-fixed household characteristics. Prices of outputs and inputs and quantities of fixed inputs are often taken as uncorrelated with stochastic disturbances when estimating cost or profit functions of pure firms (e.g., Lau 1978). The farm household model suggests that prices of consumption commodities and household characteristics, such as size and age composition, that might be taken as predetermined and also might enter household utility are additional candidates. However, if there exist unobservable household or farm characteristics, such as management skills or land quality, that persist over time, these will arguably affect the accumulation of farm assets and certain household characteristics, thus making those variables inappropriate as instruments. In view of this potential difficulty, exogeneity of instruments is tested and the robustness of estimates to choice of instruments is examined.

III. The Data and Study Setting

The data are from a cross-section survey of households in rural Sierra Leone taken during the 1974–75 cropping year (May–April). Sierra Leone was divided into eight geographical regions chosen to conform with agroclimatic zones, and they were used to stratify the sample. Within these regions, three enumeration areas were randomly picked and households sampled within them. Households were visited twice in each week to obtain information on production, sales, and labor use, among other variables. Half the households were visited twice during one week per month to obtain market purchase information.

The data set contains details on outputs, family and hired labor use (there is not much use of nonlabor variable inputs in Sierra Leone), capital stock, land use, and household characteristics. It also provides estimates at the household level of food consumption from both market purchases and home production of 196 different foods (see Strauss [1982, 1984] for details of variable construction). From these data, estimates of household caloric availability have been constructed using food composition tables (Food and Agricultural Organization 1968). This data set also has regional average price and wage data with sufficient variation to have supported estimation of a moderately large (seven commodity groups) complete demand system (Strauss 1982). It is then a good data set with which to estimate farm household-level production functions, including a measure of caloric availability, having good data on outputs and inputs as well as data on the type of instrumental variables required for estimation.

The major weaknesses in the data are the absence of individual-level data on caloric intake and the absence of measures of nutritional outcomes, especially anthropometric or clinical measures. Anthropometric and clinical variables would be useful to distinguish different possible effects on productivity of long-term (chronic) and short-term (acute) deficiencies. Also body size may have an independent effect from current nutrient intake on labor productivity (Spurr 1983). Ideally the dietary information one would like would include actual intakes for individuals.

The measure available in the Sierra Leone data is availability, not intake. The two may differ systematically, especially if food waste is positively related to income levels. However, intake data are difficult to obtain accurately. Recall methods have potential inaccuracies and, in addition, may be unrepresentative of average annual intake if the data come from one or two interviews during the year, as is common with food recall surveys. The Sierra Leone data were collected throughout the year, twice weekly for production-related variables and twice during one week per month for the market purchase information. It is not obvious whether more measurement error is introduced by using annual household availability data or individual intake data measured infrequently. Clearly, though, the best data would be frequently measured intakes at the individual level. Since such data are not available for this sample, the household-level calorie variable has to be converted into an average per family worker.

Two methods are used to make this conversion to see how robust the results are. At one extreme one could assume that food is shared equally among family members and divide household availability by household size. This seems unreasonable, though, so another assumption used is that individual food consumption is proportional to approximate caloric "requirements" for a moderately active person of a given age and sex.[15] This allows adults to get a higher share than under the equal distribution assumption, though perhaps not as high as they in fact receive.

Both of these methods assume that intrahousehold allocation of

[15] Estimating caloric "requirements" is very imprecise because of wide interindividual variation in activity levels and digestive efficiency. The weights in this study are taken from Food and Agricultural Organization (1957):

	Age			
Sex	0–5	6–10	11–15	16+
Male	.2	.5	.75	1.0
Female	.2	.5	.7	.9

Data were unavailable to correct for differential requirements of pregnant or lactating women.

foods does not vary systematically with socioeconomic factors, such as income or assets, or relative wage rates. There is very little evidence on this question, and none for Sierra Leone. What scanty evidence exists suggests that food sharing between workers and nonworkers may be greater for wealthier households, as might occur if returns from calories are decreasing.[16] In this case the conversions will overstate intakes of workers from higher-income households while understating intakes of workers from poorer households. This should reduce the estimated calorie coefficient.

For hired laborers annual caloric availability data are not directly available. Two approaches are pursued for estimation: the hired labor's calorie consumption is simply omitted, and a regional proxy variable is formed. Given the parameter normalization that is used in estimating the efficiency hours function (see eq. [5]), omitting hired labor's calorie consumption is equivalent to assuming equal consumption at the sample mean by all hired laborers. Forming a regional proxy allows for interregional variation in this measure. Since workers who hire themselves out are identified in the data, this proxy can be calculated as a weighted average of daily caloric availability per consumer equivalent (or per capita) of all households in a region. The weights used are the proportion of total regional hours hired out that comes from each household. This reduces the weighted-average caloric intake for hired laborers beneath the simple regional average since poorer households, which are larger, also tend to provide a proportionately greater amount of labor sold out.

If predicted household and hired labor caloric intakes covary positively within regions and if calories do indeed enhance labor productivity, then estimates of the household's calorie coefficient(s) will be biased upward.[17] Such a positive intraregional sorting of hired and household labor by nutrient intake might arise either if household and hired labor are complements in production, which seems unlikely, or because of a management enhancement effect of current or

[16] A study of nutrient consumption of households in Laguna, Philippines (Fabella 1982) reports a negative association between both husband's and wife's caloric intake relative to their children's, and husband's wage rate. This suggests a more egalitarian distribution with higher income. Behrman and Deolalikar (1985) report negligible responses of separate caloric intakes of adult males, adult females, boys, and girls to assets in a set of Indian households. If anything there was a slightly greater response to assets for the children, also consistent with a positive equality-wealth relationship. In a different context both Rosenzweig and Schultz (1982) and Sen and Sengupta (1983) find that landless households in India exhibit greater male-female child mortality rate differentials than do landed households. Other asset variables add very little explanation, however.

[17] Although this strictly applies to only a linear model, the logic seems applicable in this nonlinear model as well. Note that the correlation has to be within regions for the argument to apply to the estimates that use the regional proxy for hired labor's calories.

past nutrition, with management and hired labor being complementary.[18] If positive nutritional sorting did exist, then wage rates should vary positively with current nutrient intake, since there would have to be an incentive for better-fed employers to attract better-fed workers. However, wage differentials need not imply positive sorting. Unfortunately, there is no direct evidence on either the sorting or wage differential question for this sample.[19]

What is known about rural labor markets in Sierra Leone bears indirectly but inconclusively on this question. Sierra Leone is characterized by active rural labor markets (see Spencer [1979] for details) with approximately 15 percent of labor hours hired in. Much hiring is reciprocal, with payment either in cash or in kind (including meals in the field eaten with household workers). Payment in meals could reflect a recognition of nutritional-productivity effects, but it is also consistent with other hypotheses, such as economizing on travel time to and from fields, which are often far from homes. Landless laborers are virtually nonexistent in Sierra Leone, as in much of West Africa, so hired workers are themselves farmers who work only limited amounts of time (under 15 percent of labor supply) for hire. They tend not to work for the same household over long periods of time but move from one farmer's fields to another's (Spencer 1979). Most hired laborers, roughly 87 percent, are paid by the day. Payment by task is not the norm, but is confined to male laborers engaging in brushing, tree felling, or swamp digging, all very physical activities. Wage rates (including in-kind payments) vary by season, by sex, and by region, but not by job performed. Thus if better-fed workers work at more demanding tasks, which are paid better, this does not show up in the data.

IV. Functional Form Specification and Results

The agricultural production function estimated is a Cobb-Douglas function with effective family labor, effective hired labor, capital, and land as inputs (see App. for variable definitions). The production elasticities are allowed to vary linearly with the percentage of cultivated land that is upland. This is an attempt to capture differences in land quality between swamps and uplands and may also capture

[18] It seems likely that past nutritional intake, together with education and other human capital investments, would have more of an allocative impact than current intake. However, current intake is likely to proxy for past intake as well.

[19] Using a sample of Indian agricultural workers, Deolalikar (1984) reports a weak correlation between earnings and worker weight for height from a two-stage least-squares (2SLS) regression with demographic variables as instruments.

some output composition effects since swamps tend to produce rice in pure stands while uplands tend to be intercropped (Spencer and Byerlee 1977, p. 18). This specification gives rise to the estimating equation

$$\begin{aligned}\log Q = \beta_1 &+ (\beta_2 + \beta_3 U)[\log L^f + \log h(X_c^f)] \\ &+ (\beta_4 + \beta_5 U)[\log L^h + \log h(X_c^h)] \\ &+ (\beta_6 + \beta_7 U)\log K + (\beta_8 + \beta_9 U)\log A + \beta_{10} U + \epsilon,\end{aligned} \tag{4}$$

where $U \equiv$ upland as a percentage of cultivated acreage, the β's are parameters, and ε is an independent, identically distributed (iid) error term with zero mean and constant variance.

The specification reported here for the efficiency per hours worked function is quadratic in daily calories per consumer equivalent (or per capita), normalized so that the function value equals one when calories consumed equals the mean for the sample:

$$h(X_c^i) = 1 + \alpha_1\left(\frac{X_c^i}{\bar{X}_c^f} - 1\right) + \alpha_2\left[\left(\frac{X_c^i}{\bar{X}_c^f}\right)^2 - 1\right], \quad i = f, h. \tag{5}$$

This specification is reasonably flexible, even allowing for a range of negative productivity effects at high levels of food intake. It does not allow for both convex and concave portions, but it is likely that observed values would be on the concave portion of the curve since that is the more relevant economic region. The normalization allows a ready interpretation of the computed value of $h(\cdot)$ at different caloric consumption levels as the efficiency of a labor hour relative to that for the sample representative worker. It has the further advantage that $h(\cdot)$ equals one if the calorie coefficients are zero, so the usual agricultural production function is a special case of the one hypothesized here. Other functional forms for $h(\cdot)$ were used in estimation including a cubic function, which showed very little statistical improvement over the quadratic and log-reciprocal and log-log functions.[20] In addition, a Cobb-Douglas specification was estimated in which family and hired labor are permitted to be perfect substitutes, but with different efficiency weights. All estimates show the same broad patterns.

The basic set of instrumental variables used appears in appendix table A1, along with their summary statistics.[21] The regional average

[20] All were normalized in the same way as the quadratic. The log-reciprocal specification, $\log h = -\alpha[(\bar{X}_c^f/X_c^f) - 1]$ reported in Strauss (1984), forces $h(\cdot)$ to be sigmoid in shape.

[21] These instruments are output price, rice price, root crop and other cereal price, oils and fats price, fish and animal product price, miscellaneous foods price, nonfoods price, male adult wage, wage squared, hired labor calorie consumption, hired labor calorie consumption squared, capital stock, upland, land, capital × upland, land × upland, household size, and number of adults.

caloric consumption of hired labor and its square are included in this instrument set, which is equivalent to assuming a pool of labor available to the household for hiring that the household cannot affect. This assumption is relaxed later.

Taking the logarithm of a quadratic function introduces nonlinearity in both parameters and variables into the estimating equation, for which nonlinear two-stage least squares (NL2SLS) is used (see Amemiya 1983).[22] Estimates for the production function (eq. [4]) using the basic instrument set are provided in table 1. The first column gives a linear two-stage least-squares estimate of the Cobb-Douglas function when no calorie variable is included, the family and hired labor variables being treated endogenously. Column 2 contains the NL2SLS estimates with the quadratic effective labor hours function, while the third column repeats the estimation after the jointly insignificant upland and land-upland interaction variables are dropped.[23] Column 4 uses the per capita calorie availability measure for both household and hired labor, but is otherwise the same regression as column 3.

In all three cases in which they are included both the calorie and calorie squared coefficients are significant at more than the .01 level, with calorie consumption contributing positively to output (see table 3). Coefficients of other production function inputs are all significant at the .025 level (col. 3), in contrast to the estimates that omit the calorie variables for which only the family labor and upland variables have asymptotic standard normal statistics of over one.

It is possible that the calorie variables are picking up the effects of other human capital variables. This is explored by repeating the regressions and entering years of English and Islamic education (most respondents had none) into the family effective labor function as well as household head's age and age squared. The coefficients of these human capital variables are completely insignificant, while the calorie coefficients remain highly significant. The remaining coefficients are quite close in magnitude to those reported in table 1.

The fact that only a very crude proxy, percentage upland, is available for land quality could also bias upward the calorie coefficients. Another variable that is related to land quality and available in the data is the average age of bush on fallowed land. To the extent that better-quality land is cultivated more extensively, one would expect that less time in fallow would be allowed, so that a lower average age

[22] Quadratic terms and interactions of exogenous variables can be used as instruments in NL2SLS in addition to levels. The only such term used in this study is wage squared. Other terms resulted in a numerically singular cross-product matrix of instruments. The Davidon-Fletcher-Powell algorithm as available in the Fair-Parke program (see Fair 1984) was used to minimize the objective function.

[23] The Wald test statistic (χ^2 with 2 df) is 1.86.

TABLE 1

AGRICULTURAL PRODUCTION FUNCTIONS: QUADRATIC EFFECTIVE LABOR FUNCTIONS

Variable	(1)	(2)	(3)	(4)*
Constant	−4.21	−.23	1.22	1.39
	(−1.7)	(−.1)	(1.2)	(1.2)
Effective labor function:				
Calories†	...	1.35	1.33	1.12
		(4.2)	(4.5)	(4.4)
Calories squared†	...	−.41	−.39	−.30
		(−3.2)	(−2.8)	(−4.4)
Family labor†	1.61	1.19	.95	.91
	(4.6)	(4.4)	(5.2)	(4.7)
Family labor × upland†	−1.89	−1.04	−.53	−.67
	(−3.4)	(−2.3)	(−2.2)	(1.9)
Hired labor†	−.27	−.49	−.49	−.41
	(−.9)	(−1.7)	(−2.1)	(−2.3)
Hired labor × upland†	.48	1.03	.99	1.15
	(.9)	(2.1)	(2.7)	(2.7)
Capital	.02	.23	.40	.40
	(.1)	(1.1)	(2.7)	(2.3)
Capital × upland	.004	−.38	−.59	−.63
	(.01)	(−1.4)	(−2.9)	(−2.8)
Land	.2	.36	.26	.28
	(.9)	(1.7)	(2.5)	(2.3)
Land × upland	.2	−.14	...	...
	(.6)	(−.5)		
Upland	11.69	3.00	...	...
	(2.8)	(1.1)		
Function value	2.60	2.80	3.33	4.59
Regression standard error	.59	.54	.51	.56
R^2	.35	.47	.52	.42
Minimum χ^2 test statistic‡	7.47 [4]	9.60 [7]	12.80 [9]	14.64 [9]

NOTE.—Asymptotic standard normal statistics in parentheses; degrees of freedom are in square brackets.
* Calories per person used instead of per consumer equivalent.
† Endogenous variable.
‡ Defined as $\mathbf{e}'\mathbf{Z}(\mathbf{Z}'\mathbf{Z})^{-1}\mathbf{Z}'\mathbf{e}/\hat{\sigma}^2$, where $\mathbf{e}$ is the vector of estimated residuals, $\mathbf{Z}$ is the matrix of instruments, and $\hat{\sigma}^2$ is the estimated regression variance.

of bush would result. When this variable is entered linearly into an effective land function, similar to the effective labor function, its coefficient is just significant at the .1 level, but once again the other coefficients do not change very much.[24]

The estimates in table 1 all use farm assets, household size, and number of adults as instrumental variables. If there exist time-persistent household effects that are unobserved and are correlated with these variables, then these estimates would be inconsistent. Such

[24] The calorie and calorie squared coefficients are 1.37 and −0.39, respectively, with standard normal statistics of 5.1 and −3.1.

household effects, or heterogeneity, might include managerial ability. Even without this heterogeneity the household size and number of adults variables could possibly be endogenous since households with higher incomes might attract more family members to live with them. Since extended families are important in Sierra Leone, this should be considered. In addition, the proxy variable for hired labor's caloric intake, the regional average variable, and its square were included in the instrument set. This may be objectionable since their inclusion would correspond to the underlying model only if hired labor's calorie consumption were exogenous to the hiring household, which may not be the case.[25]

The existence of correlation between the instruments and residuals is tested by using a generalized method of moments specification test (Hansen 1982; Newey 1983), which examines how close the cross-products of instrumental variables and residuals are to zero when evaluated at the estimated parameter values.[26] If the instruments are truly uncorrelated with the unobserved disturbances, then these cross-products of estimated residuals with instruments ought to be close to zero.[27] This specification test is general in that rejection can occur for more than one reason, for example, endogeneity of instruments or omitted variables. Results from these tests are reported in the last row of table 1 labeled "minimum χ^2 test statistic." The statistics are asymptotically distributed as χ^2 variables with degrees of freedom equal to the number of overidentifying instruments. The statistics from columns 2 and 3, the major ones of interest here, have probability values of roughly .2.[28] They are, therefore, not significant at standard levels. Nevertheless, it is still of some interest to examine the robustness of the results to a reduced instrument set.

Table 2 contains reestimates of column 3 from table 1 while systematically dropping groups of instruments. When the regional hired

[25] Dropping these variables from the instrument list does not solve the problem that hired labor's caloric intake is assumed to be homogeneous within regions, whereas it is almost certainly not. The issue here is still whether systematic sorting of well-nourished hired labor with well-nourished hiring households exists.

[26] The statistic is $\mathbf{e}'\mathbf{Z}(\mathbf{Z}'\mathbf{Z})^{-1}\mathbf{Z}'\mathbf{e}/\hat{\sigma}^2$, where $\mathbf{e}$ is the vector of estimated residuals, $\mathbf{Z}$ is the matrix of instrumental variables, and $\hat{\sigma}^2$ is the estimated regression variance. Since the numerator is the minimized value of the objective function used by NL2SLS, this statistic is readily computed from standard computer output. In contrast, a Durbin-Wu-Hausman specification test based on the difference between estimates using the full and reduced instrument sets is complicated to compute for this case. This is because the covariance matrix of the difference is not simply the difference of the two covariance matrices, as in more typical examples, but also depends on the covariance between the two estimators.

[27] Of course, it is necessary that the number of instruments exceed the number of parameters for the cross-products not to be set to zero by the estimation procedure.

[28] The analogous test statistics from the log-reciprocal specifications also have probability levels near .2.

TABLE 2

AGRICULTURAL PRODUCTION FUNCTIONS WITH VARYING INSTRUMENT SET

Variable	(1)	(2)	(3)	(4)
Constant	1.28	1.09	.38	.26
	(1.1)	(.8)	(.2)	(.2)
Effective labor function:				
Calories	1.58*†	1.57*†	1.05*†	.74†
	(8.9)	(8.3)	(1.4)	(2.5)
Calories squared	−.49*†	−.44*†	−.20*†	...
	(−8.4)	(−3.0)	(−.4)	
Family labor	.96†	1.10†	1.00†	.97†
	(4.9)	(4.8)	(3.4)	(3.2)
Family labor × upland	−.52†	−.47†	−.22†	−.23†
	(−1.7)	(−1.4)	(−.4)	(−.4)
Hired labor	−.47†	−.67†	−.81†	−.81†
	(−2.0)	(−2.5)	(−2.3)	(−2.2)
Hired labor × upland	.92†	.86†	1.27†	1.31†
	(2.1)	(1.8)	(1.9)	(1.9)
Capital	.33	.38	1.16†	1.27†
	(1.8)	(2.0)	(2.2)	(2.9)
Capital × upland	−.55	−.51	−1.83†	−1.88†
	(−2.3)	(−2.0)	(−2.5)	(−2.6)
Land	.29	.32	.002†	−.05†
	(2.4)	(2.6)	(.01)	(−.2)
Function value	3.10	1.33	.02	.09
Regression standard error	.57	.61	.66	.68
R^2	.41	.31	.19	.15
Minimum χ^2 test	9.7	3.55	.05	.2
statistic‡	[7]	[5]	[1]	[2]

NOTE.—Asymptotic standard normal statistics in parentheses; degrees of freedom are in square brackets. Instruments dropped from col. 1 are regional hired labor calorie consumption and its square; from col. 2, hired labor calories, household size, and number of adults; from cols. 3 and 4, hired labor calories, capital, land, capital × upland, land × upland, household size, and number of adults.

* Jointly significant at the .01 level.

† Endogenous variable.

‡ See table 1 for formula.

labor calorie variable and its square are dropped from the instrument set (col. 1), the coefficients are not substantially changed, although the fit worsens somewhat. The specification test of orthogonality between residuals and instruments still has a probability value of .2. These estimates are extremely close to those that drop the hired labor calorie variable, which are therefore not reported here. The equation fit worsens still more when the household demographic variables are in addition dropped as instruments (col. 2); however, the coefficient estimates are almost identical to those in column 1, with the calorie coefficients retaining significance at less than the .01 level. Note now that the specification test statistic has become quite insignificant, its probability value rising to .6. Columns 3 and 4 drop the farm asset variables as well. The calorie coefficients remain jointly significant at the .01 level or less (Wald statistic of 9.52), as also evidenced by the linear specification, although the terms in the quadratic specification now lose individual significance. While the magnitudes of the calorie coefficients change for the quadratic $h(\cdot)$ function, the elasticity of $h(\cdot)$ with respect to family calories actually rises a little, compared with the base line estimates from table 1, from .55 to .65 when both farm and household assets are dropped.[29] The land coefficient becomes insignificant and its magnitude drops considerably when the farm asset instruments are omitted. Apparently the remaining instruments predict little of the variation in land input, as evidenced by the large drop in R^2. The hired labor and capital output elasticities change by only a small amount. Dropping both farm asset and demographic variables from the instrument set has lowered the specification test statistic to well under .5.[30]

In sum, the household calorie consumption seems a statistically significant determinant of farm output. While the statistical evidence of possible endogeneity of farm assets and household demographic measures is very weak, even allowing explicitly for that possibility, calorie consumption remains quite significant. How important this relation is in economic terms is the next question to be discussed.

V. Implications

To interpret the coefficients the implied output elasticities and marginal products are first considered. Table 3 reports them using the estimates from column 3 of table 1. Other specifications provide broadly similar patterns. The estimates show roughly constant returns to scale. Interestingly, the 2SLS estimates without the effective labor function (col. 1) imply a returns to scale of .8. The largest change in output elasticities occurs for family labor, which drops to .42. Apparently, with other inputs held constant, households demanding more family labor have a lower equivalent caloric intake per consumer, which biases family labor's coefficients downward.

The marginal products of family and hired labor are fairly close and not significantly different (the standard error of the difference is .45). Both are very close to the sample mean real wage, which is .29.

Family caloric intake has a sizable, statistically significant, output

[29] At the sample mean this elasticity equals $\alpha_1 + 2\alpha_2$ (see eq. [5]), where α_1 is the coefficient on calories and α_2 the coefficient on its square.

[30] Estimates were also made by dropping the farm asset variables while retaining household size and number of adults. Coefficient estimates and their standard errors are very close to those of cols. 3 and 4.

TABLE 3

OUTPUT ELASTICITIES AND MARGINAL PRODUCTS AT SAMPLE MEAN

Input	Elasticity	Marginal Product
Household calorie consumption	.33	.19
	(.11)	(.07)
Household labor hours	.60	.32
	(.18)	(.10)
Hired labor hours	.13	.40
	(.15)	(.44)
Capital	.03	2.06
	(.10)	(6.63)
Land	.26	85.40
	(.10)	(34.49)

NOTE.—Asymptotic standard errors in parentheses. Computed using estimates from col. 3 of table 1.

elasticity of .34.[31] The magnitude of this elasticity varies widely from low-consumption to high-consumption households. As the level of caloric intake reaches 4,500 per day, which is roughly the average intake of the upper third of the sample, the output elasticity falls to only .12. However, at a daily intake per consumer equivalent of 1,500 calories, which corresponds to the average for the lower third of the sample, the output elasticity rises to .49. This figure is remarkably close to the calorie output elasticity of .5 found for Kenyan road construction workers, with an average daily intake of 2,000 calories, in the experiment of Wolgemuth et al. (1982) (see n. 6).

The estimated efficiency units of labor function is plotted in figure 2. As indicated, $h(\cdot)$ reveals the relative efficiency of an hour of labor when compared with labor that consumes calories equal to the sample mean. The function reaches a peak at a daily intake per consumer equivalent of 5,200 calories, and thereafter calories have a negative impact on effective labor. The corresponding value of $h(\cdot)$ is 1.2. Roughly 12 percent of the sample (15 households) have an estimated daily caloric intake per consumer equivalent above this level. This is an extremely large intake level for calories to have a positive effect; however, the effective labor function is flat by the level of 4,500 calories per day ($h[\cdot]$ is 1.17), which is roughly the average intake of the upper third of the sample. Indeed this function rises very gently after 3,750 calories (h being 1.1). The flattening of the effective labor function is also apparent by noting the decline in the elasticity of $h(\cdot)$ with respect to calories from .55 at the sample mean intake to .23 at 4,500 calories per day. For households with low levels of calorie con-

[31] Calorie elasticities and marginal products from the log-reciprocal specifications of the effective labor function are lower, .18 and .10, respectively.

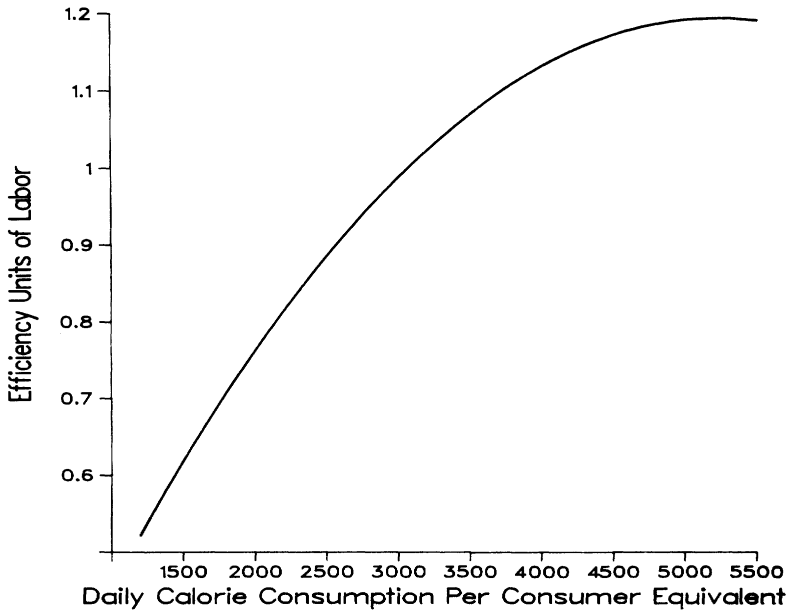

FIG. 2.—Estimated efficiency labor function

sumption per consumer equivalent $h(\cdot)$ rises much more steeply. At an intake of 1,500 calories the calorie elasticity is .75. The level of $h(\cdot)$ is roughly .6, implying that the hourly efficiency of family labor is on the order of 60 percent of the efficiency of a family worker from a representative family.

A different effect may be seen by looking at the first-order condition for food consumption (see n. 14). An increase in caloric intake per consumer equivalent is equivalent to a proportionate reduction in the effective price of food. Taking rice, the staple food in Sierra Leone, and ignoring the effect of higher caloric intake on clock hour wages or on total non–sick time available to the household, these results suggest that at the sample mean a percentage increase in rice consumption will reduce the effective price of rice by 42 percent.[32] Again those percentages vary by level of caloric intake, being in the range of 90 percent for an intake of 1,500 daily calories per consumer equivalent and 15 percent at 4,500 calories. Now clearly these figures are large, especially for the poorer households, although other specifications of $h(\cdot)$ result in somewhat smaller magnitudes.[33] How-

[32] This is calculated assuming a conversion of 3,743 calories per kg of rice, converting this annual figure to a daily per consumer equivalent and multiplying by the marginal product of family calories from table 3.

[33] The log-reciprocal specification of $h(\cdot)$ results in a percentage decline of 22 percent, which still seems large.

ever, they are suggestive, and given the reasonable robustness of these empirical results, these effects should not be dismissed.

VI. Conclusions

It would appear that current nutrient intake, proxied by calories, does raise current farm labor productivity in rural Sierra Leone. These effects seem very strong at low intake levels, dropping off substantially as intake levels rise, but still with some effect at moderate intake levels. As noted, agricultural labor in Sierra Leone is characterized as physically demanding, so these results are not implausible. The effect explored here, however, is a pure worker effect. To the extent that allocative effects of better nutrition are important, the results have understated the impact of better nutrition on output supply.

A number of questions about the nature of the nutrition-productivity linkage remain unanswered, partly because individual-level nutrient intake and anthropometric data were unavailable. The analysis has proceeded on the assumption that current, annual caloric intake directly causes higher productivity. However, current calorie flows are probably correlated with accumulated stocks, such as measured by height or weight, which may have independent effects on productivity. More generally, health may have an impact on productivity and also be correlated with current caloric intake. Thus it is not clear from these estimates how much low-nutrient intake during childhood affects labor efficiency versus current intake or related health outcomes. For policy design this would be useful to know. Individual-level data on nutrient intake and anthropometric or clinical health variables might help economists answer these questions. It is also plausible that the impact of nutrient intake differs by male, female, or child labor and that it has a different impact on home production than on farm production or market earnings. Finally, other studies will have to establish how strong the nutrition-productivity links may be in developing countries with either a greater capital intensity of agriculture production or higher income levels or both.

Appendix

TABLE A1

SAMPLE SUMMARY STATISTICS

	Mean	Standard Deviation
Endogenous variables:		
Farm output quantity index (kg)	2,295.2	1,844.4
Daily family calories per consumer equivalent	3,061.0	1,811.4
Daily family calories per capita	2,434.7	1,610.9
Hours of family labor	3,898.2	2,122.0
Hours of hired labor	816.5	620.8
Exogenous variables:		
Daily hired labor calories per consumer equivalent	2,788.4	1,242.7
Output price index*	.27	.06
Rice price index*	.24	.05
Root crop and other cereal price index*	.58	.46
Oils and fats price index*	.66	.16
Fish and animal product price index*	.56	.31
Miscellaneous foods price index*	.60	.19
Nonfoods price index*	.64	.09
Male adult wage (leones per hour)	.08	.03
Capital stock (in leones)	34.4	31.6
Land cultivated (in acres)	6.8	4.5
Upland as a percentage of land cultivated	.63	.37
Household size	6.3	3.7
Persons 11 years and older	4.4	2.2
Average age of bush in fields (in years)	7.8	6.8
Number of consumer equivalents	4.7	2.4
Years of English education of household head	.4	1.5
Years of Islamic education of household head	1.6	4.1
Age of household head	50.9	15.0

* Leones per kg. For definitions of commodity groups see Strauss (1982), table A.1.

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