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**GENDER DIFFERENCES IN AGRICULTURAL PRODUCTIVITY:
A SURVEY OF EMPIRICAL EVIDENCE**

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ABSTRACT

This paper reviews the econometric evidence on gender differences in agricultural productivity. It provides a methodological overview and a critique of (1) production function-based estimates of technical and labor productivity differences by gender, (2) individual (gender-disaggregated) labor supply and earnings functions and (3) studies of the determinants of technological adoption.

The review finds that (1) in general, male and female farmers are equally efficient as farm managers. Women farmers' lower yields are attributable to lower levels of inputs and human capital than men. However, the use of coefficients estimated from these studies for simulation exercises may not be valid if endogenous input choice is not considered; (2) returns to schooling for both men and women are significant in dynamic agricultural settings where modern technologies have been introduced. Returns to an additional year of women's education range from 2 to 15 percent, which compares favorably with those of men; and (3) farmers with more education are more likely to adopt new technologies. Providing universal primary education also stimulates early adoption by female farmers, whom other women are more likely to imitate. Farmers with more land and farm tools are also more likely to adopt new technologies. To the extent that women farmers may have less education, less access to land, and own fewer tools, they may be less likely to adopt new technologies.

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1. INTRODUCTION

The measurement of gender differences in agricultural productivity is complicated by differences in farming systems and social and cultural institutions.¹ It may be possible to estimate gender differences in efficiency in farming systems where men and women manage separate plots, as in many African farming systems (Boserup 1970), but it is more difficult to isolate managerial efficiency differences in agricultural settings where plots are cultivated jointly by male and female family members and hired labor. In the latter setting, found in the "male" farming systems of Asia and Latin America, the farm

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¹ Ideally, an evaluation of gender differences in agricultural productivity should be based on estimates of total factor productivity, in which an index of output is divided by an index of inputs, aggregated over all types of outputs and inputs, respectively. In addition to aggregation problems, however, lack of gender-differentiated data on inputs and outputs has prevented the use of this approach. Existing studies therefore use partial productivity measures, such as yield and labor productivity, or estimate differences in technical and allocative efficiency.

manager is usually assumed to be the male head of the household, regardless of the actual contribution of women to decisionmaking and farm labor.

In countries where it is possible to identify the gender of the plot manager, direct estimates of gender differences in technical efficiency have been made. The production function studies either estimate male and female production functions separately, or estimate a pooled regression with a dummy variable for the gender of the farm manager (or household head). Coefficients from these production functions have also been used to estimate gender differences in labor productivity. Since labor is usually measured in time units, it is assumed to be homogeneous within a category. However, many of the earlier studies did not consider endogeneity of input choices with respect to farmer characteristics.

Another approach provides an indirect measure of productivity by estimating earnings or wage functions, taking into account the heterogeneity of agricultural labor, since individual characteristics and endowments influence labor market participation, earnings, and wages. These studies often provide evidence on returns to men's and women's schooling. The last category of studies examines gender differences in technological adoption. Since new technologies are adopted due to perceived increases in future income streams, they can also provide evidence on productivity. In fact, technological adoption may be a better long-run indicator of productivity gains than static productivity studies that measure output at a given point in time, and that may be affected by short-run variability.

This paper reviews the empirical evidence on gender differences in agricultural productivity. While there is a huge volume of literature that attempts to document differences in productivity, relatively few control for individual endowments by gender, and even fewer for relationships between individual characteristics (for example, education, gender) and input choice. However, if women systematically had lower levels of education and physical assets, an approach that did not control for endowments would tend to overestimate productivity differences due to gender. The accurate diagnosis of the sources of productivity differences, if they exist, is important in order to identify appropriate policy interventions for increasing women's productivity and welfare.² If productivity differences arise from market failures that constrain women from borrowing to increase their human capital, or to adopt new technologies, there is scope for interventions to remove these barriers. Thus, this review is limited to those studies using regression analysis. A list of the studies, together with information on data sources and method of gender disaggregation, is found in Table 4 in the Appendix.

The remainder of the paper is organized as follows: Parts 2, 3, and 4 discuss the methodology and empirical evidence on (1) production function-based estimates of technical efficiency and labor productivity differences by gender; (2) individual (gender-disaggregated) labor supply and earnings functions; and (3) studies of the determinants

² It is often argued that women's lack of access to resources results in lower productivity or inability to respond to economic incentives. See, for example, the papers in Gladwin (1991).

of technological adoption. Part 4 summarizes the evidence and its policy implications, and suggests directions for future research.

2. PRODUCTION FUNCTION APPROACHES

A production function is a technical relationship between inputs and outputs that specifies the maximum level of output possible, given input levels. Technical efficiency reflects the ability of a manager to produce output, given input levels and technology. Suppose that male and female farmers have the same production technology but male farmers are more technically efficient. For the same level of inputs, say L_0 , the quantity produced by male farmers would be greater than that produced by female farmers (Figure 1). The female farmers' production function may be "inside" the male production function because they use traditional technologies, due to lack of knowledge, lack of access to modern inputs associated with new technologies, or higher costs to adopting the new technologies.

Technical efficiency does not imply allocative or economic efficiency, however. Allocative efficiency means that resources are used so that the value of an additional unit of output (the value of the marginal product) is equal to the cost of an additional unit of input. Given relative prices, the allocatively efficient allocations would be at points A and B for male and female farmers, respectively. Thus, technically inefficient farmers may be allocatively efficient.

Figure 1—Technical and allocative efficiency of male and female farmers

METHODOLOGY

Most of the studies reviewed in this paper follow the primal approach to production analysis and estimate directly the production function of a farm manager i in household j :

$$Y_{ij} = f(V_i, X_i, Z_j), \quad (1)$$

where Y_{ij} is quantity produced, V_i is a vector of inputs used by farm manager i (including land, labor, capital, and extension advice); X_i is a vector of individual attributes, including gender; and Z_j are household- and community-level variables. Correlation of input use with individual and household characteristics can be captured by interaction terms $V_i X_i$ and $V_i Z_j$.

Single-equation estimation of production functions is problematic because the quantities of output and variable inputs are simultaneously determined by the conditions of profit maximization, so the stochastic disturbance term may be correlated with input levels. While it can be argued that households attempt to maximize *expected* profits, the right-hand side variables in single equation estimation should all be exogenous. While (1) correlation between input levels and the error term can be corrected through instrumental variables estimation, and (2) regional effects and heteroscedasticity can be controlled for using fixed effects methods and appropriate estimators of the variance-covariance matrix, respectively, endogeneity of regressors can be avoided by estimating profit or cost functions instead of the production function.

The dual approach to production analysis estimates profit functions as a function of input and output prices, and derives the input demand and output supply functions from the restricted profit function.³ This approach has its advantages when there are multiple outputs and inputs, as in a multicrop farming system. Modeling input choice explicitly also allows for the possibility that farmer characteristics influence the choice of conventional inputs. For example, if more educated farmers are more likely to use modern inputs (for example, fertilizer), than less educated farmers, a production function specification that included both farmer education and fertilizer usage would overstate the contribution of fertilizer and understate that of education. Moreover, this approach also enables the researcher to distinguish productivity differences due to differences in access (which assumes the existence of barriers) or in input choices. The dual approach has only been recently applied to the analysis of gender differences in productivity.

Most of the early empirical work on gender differences in technical efficiency has used the Cobb-Douglas production function:

$$Y = \alpha_0 L^{\alpha_1} T^{\alpha_2}, \quad (2)$$

where Y is output, L is labor input (hired or family), and T is a vector of land, capital, and other conventional inputs. Usually, the equation is estimated by ordinary least squares (OLS) by taking logarithms on both sides; a typical form is as follows:

³ For a thorough discussion and empirical applications of duality approaches in production, see Fuss, McFadden, and Mundlak (1978).

$$\ln Y = \alpha_0 + \alpha_1 \ln L + \alpha_2 \ln T + \beta \ln E + \tau \text{EXT} + \delta \text{GENDER} + \epsilon, \quad (3)$$

where Y , L , and T are as defined above; E is educational attainment or an indicator variable for level of schooling (of the farm manager, or household head, or the members of the household); EXT is an index of extension services; GENDER is the gender of the household head or farm manager; and ϵ is the error term. The coefficient that indicates gender differences in technical efficiency is δ , an intercept shifter. Correlation between the gender of the farmer and other inputs can be captured by interaction terms. Estimates of output per unit of input are usually obtained through yield regressions, where the dependent variable is yield per hectare and the input variables are expressed in terms of inputs per hectare.

Although the Cobb-Douglas functional form is convenient to estimate since it is linear in parameters, it is an unduly restrictive form to impose on the underlying production relationship. It assumes strong separability between inputs and an elasticity of substitution equal to 1.0. Less restrictive approaches using second-order Taylor's expansions to approximate the functional form (for example, the translog, normalized quadratic and Leontief), while widely used in other applications, have rarely been used to analyze gender differences in productivity, with the few exceptions discussed below. Even flexible forms, however, involve a guess about the underlying functional form. Nonparametric approaches may have greater potential by not imposing a specific

functional form, so long as the properties of a production function are preserved.⁴ The implications of neglecting endogeneity of input choice, however, may be more serious than functional form considerations.

EMPIRICAL EVIDENCE

Technical Efficiency Differences

The survey has identified seven studies that estimate differences in technical efficiency between male and female farm managers or household heads using production functions.⁵ The results are summarized in Table 1. Three studies in Kenya found that the gender of the farm manager was an insignificant determinant of output per hectare (Moock 1976; Bindlish and Evenson 1993; Saito, Mekonnen, and Spurling 1994).

A study of 152 maize farmers in Kenya's Vihiga District in the 1970s estimated yield functions for all farms, with a female manager dummy, and for male and female-managed farms separately (Moock 1976). The study found that women, who

⁴ See, for example, Goldman and Ruud (1993).

⁵ The distinction between farm manager and household head may be important in households where the husband may be the titular head but the majority of farm decisions is made by the wife. This is especially true if the husband is absent for prolonged periods due to seasonal migration for wage work, or if the wife in a polygamous marriage has her own farm, both of which are common in Africa.

Table 1—Production function studies with estimates of male-female differences in technical efficiency: Gender effects from pooled regressions

Study	Sample	Gender Variable	Coefficient	t-Ratio	Dependent Variable Definition	Endogenous Input Choice	Comments
<u>Kenya</u>							
Moock 1976	All farmers N = 152 Males = 101 Females = 51	Female dummy x log of area planted in maize	0.090	1.29	Log of maize output per acre	No	Female dummy, though positive, was not significant by itself and was excluded.
		Female dummy x log of plant population per acre	-0.280*	-1.85			
		Female dummy x log of labor input per acre	0.108***	2.15			
		Female dummy x primary schooling dummy	0.167**	1.98			
		Female dummy x log of extension contact index	-0.028	-1.5			
Bindlish and Evenson 1993	All farmers N = 675 Male heads = 434 Female heads = 241	Female head dummy	-0.022	-0.23	Log (1n) of crop production in 1990	No	Coefficients of age of head, and primary education of head are negative but insignificant.
Saito, Mekonnen, and Spurling 1994	All plots No. of plots = 601 Male plots = 419 Female plots = 182	Male farmer dummy	0.126	1.22	Log of total value of crop production at plot level	No	Coefficients of land, capital, male and female labor, female hired labor, dummies for fertilizer and tractor use, maize/beans dummy, tenure dummy significant and positive. Formal education insignificant. Extension weakly significant.
		Male farmer dummy	-0.017	0.18	Log of total value of crop production at farmer level	No	Coefficients of land, capital, male family labor, female family labor, extension dummy, and maize/beans mixture positive and significant. Formal education and age of farmer insignificant.

(continued)

Table 1 (continued)

Study	Sample	Gender Variable	Coefficient	t-Ratio	Dependent Variable Definition	Endogenous Input Choice	Comments
<u>Burkina Faso</u>							
Bindlish, Evenson, and Gbetibouo 1993	All farmers N = 2,406 Male heads = 2,233 Female heads = 103 Unclear = 70	Female head dummy Proportion of female extension workers in division	-0.146 34.65***	-1.63 2.71	Log (1n) of crop production at the farm level	No	Logarithm of total area cropped and the number of parcels have positive and significant coefficients.
<u>Nigeria</u>							
Saito, Mekonnen, and Spurling 1994	All farmers No. of farmers = 226 Males = 210 Females = 15	Male farmer dummy	-0.130	0.47	Log of total value of production at household level	No	Coefficients of land, female family and hired labor, insecticide use dummy, and age of household were positive and significant.
	All plots Total no. of plots = 1,174 Male plots = 885 Female plots = 289	Male farmer dummy	0.559***	5.59	Log of total value of crop production at plot level	No	Coefficients for land, capital, male and female family labor, male and female hired labor, insecticide use were positive and significant. Dummy for 1-8 years of education and age of head were negative and significant.
<u>Korea^a</u>							
Jamison and Lau 1982	Mechanical farms = 1,363 ^a (90.2% male heads)	Male head dummy	0.95**	2.33	Log of value of agricultural crop output (won)	No	Education measured by the household average, excluding the household head, is significant and positive. Education of household average has stronger effect than education of household head. Land, labor, animal and mechanical power, and fertilizer have significant and positive coefficients.

(continued)

Table 1 (continued)

Study	Sample	Gender Variable	Coefficient	t-Ratio	Dependent Variable Definition	Endogenous Input Choice	Comments
<u>Korea^a</u>							
Jamison and Lau 1982	Nonmechanical farms = 541 ^b (87.6% male heads)	Male head dummy	0.059	0.87	Log of value of agricultural crop output (won)	No	Education measured by the household average, excluding the household head, is significant and positive. Education of household head insignificant and not included. Land, capital, labor, animal power, and fertilizer are positive and significant.
<u>Thailand</u>							
Jamison and Lau 1982	Chemical farms = 91 ^a (97.8% male heads)	Male head dummy	0.076	0.28	Log of output (kg.)	No	In highest R ² regression (not reported here), maximum education of household head has a positive and statistically significant rate of return of 3%. Dummy variables for educational levels positive but not significant, extension negative but insignificant, labor and land coefficients positive and significant.
	Nonchemical farms = 184 ^b (99.5% male heads)	Male head dummy	0.269	1.05	Log of output (kg.)	No	In highest R ² regression (not reported here), maximum education of head has a positive and statistically significant rate of return of 2.4%. A dummy for four years of education is significant and positive; extension significant and positive; labor, land, and coefficients positive and significant.

^a Coefficients are reported from regressions with highest R².

^b Coefficients are reported from regressions with highest R².

*** Significant at 1 percent.

** Significant at 5 percent.

* Significant at 10 percent.

make up a third of the sample, are at least as productive as men: the female farmer dummy, though positive, was not significant by itself in the pooled regression and was dropped from the final specification. Interactions between the female farmer dummy and other inputs suggest that while women benefit less from more densely planted farms, they tend to make better use of labor on maize farms than men. Exposure to the Ministry of Agriculture's extension service is associated with greater technical efficiency if the farmer is male and not too well educated, although there are efficiency gains for women with at least a primary education. The negative coefficient of the interaction term of extension and schooling suggests that both services may be substitutes: less educated farmers tend to substitute extension advice for lack of education. Results from separate regressions for male- and female-managed farms indicate that primary education has a positive and significant effect on yields for women, but a negative and significant effect for men. However, extension contact tends to benefit men but not women, since extension contact with women may have been limited in the 1970s.

Another study for Kenya was based on a survey conducted in three districts (Kakamega, Muranga, and Kilifi) in 1989-90 (Saito, Mekonnen, and Spurling 1994).⁶ Cobb-Douglas production functions with the gross value of maize, beans, and cowpeas per hectare as the dependent variable were estimated at the plot level, further disaggregated by gender of plot manager. Although a farm-level regression more appropriately captures gender differences in managerial efficiency, plot-level regressions

⁶ The authors also perform similar analyses for Oyo State, Nigeria.

were used to take advantage of more degrees of freedom. The regression for all plots shows that, while positive, the male plot manager dummy is insignificant. In separate regressions, capital (value of farm tools and equipment) is a positive and significant determinant of the gross value of output per hectare for male plots but is insignificant for female plots. A tenure dummy, which proxies full exercise of land rights, is not significant for either male or female plots. This may reflect lack of variation in tenure across plots. While the effect of extension is positive and highly significant for male plots, it is insignificant for female plots. This may indicate better use of extension services by men.

These results need to be taken with caution due to methodological problems. The pooling of all plot-level observations without an appropriate fixed or random effects estimation procedure is likely to result in inefficient and inconsistent estimates. Farm-level unobservables would be common to plots farmed by a single farmer, and error terms across plots could be correlated. A more serious omission, however, is due to the computation of gross value of output based only on the value of maize, beans, and cowpea production. Tree crops were not included because their lengthy gestation period makes the assumption of a common production technology unrealistic. However, it cannot be argued that gross value of output based on food crops alone is truly representative of a plot manager's efficiency, especially if men are more likely to cultivate tree crops. Alternatively, since women are traditionally involved in the

cultivation of food crops, they could have developed, over time, a comparative advantage in food crop cultivation.

An evaluation of the effects of the training and visit (T&V) extension system in Kenya offers insights into changes that occurred since the 1970s (Bindlish and Evenson 1993). The T&V system, recently implemented with World Bank support, has altered the traditional British-style extension service, and has employed a substantial number of women as field agricultural extension workers, in contrast to home economics extension where women have traditionally specialized. In Kenya, where there is a strong tradition, particularly among women farmers, to form groups, farmers' groups have also been used as "contact points" by extension workers. The study, based on a national survey conducted in 1989-90, also finds that female heads are equally efficient as male heads (the female dummy is not significant), and that extension, measured by the ratio of field extension workers to farm households, has a positive effect on output and total factor productivity.⁷

Burkina Faso offers an interesting contrast to the above results (Bindlish, Evenson, and Gbetibouo 1993). Regression results show that women farmers are significantly less

⁷ An earlier version of the paper found that female heads were more productive when female extension workers were assigned to the sublocation. However, when assignment of female extension workers was predicted based on locational characteristics, including the percentage of female farmers in the area, the interaction between female extension workers and female farmers became insignificant (personal communication with R. Evenson). This suggests that placement of extension workers may be endogenous, and that female extension workers may in fact be assigned to areas where they can be more productive, i.e. where female farmers are predominant.

productive than men in most crops and have total values of output that are about 15 percent lower. However, the proportion of female field extension workers has a significant and positive effect on output.

The negative female farmer dummy may be due to cultural, religious, and ethnic differences between Burkina Faso and Kenya, rather than to differences in the extension system, which has now shifted over to the T&V system. Fewer farm households are headed by women in Burkina Faso. Moreover, a woman comes under the authority of another male family member when her husband is away. However, the significant proportion of plots managed by women—especially those planted to food crops, sorghum, and millet—may partly explain why the proportion of female extension workers has a positive and significant effect on output. As farmers and plot managers, women benefit from interaction with female extension workers.

A more recent study using detailed agronomic panel data from Burkina Faso suggests that asymmetric roles and obligations within the household may have more serious implications on allocative, rather than technical, efficiency (Udry 1994). Rather than estimate production functions, Udry regressed plot yields on plot characteristics and individual characteristics, including the gender of the plot manager. Plots controlled by women have significantly lower yields than similar plots within the household planted with the same crop in the same year, but controlled by men. The yield differentials are due to significantly higher labor and fertilizer inputs per acre on plots controlled by men. These differences in input intensity between male- and female-managed plots persist even

after land quality, measurement error, or risk management behavior are taken into account, and contradict the assumption of Pareto-efficient resource allocations within the household.

The two Asian countries included in the Jamison and Lau (1982) study of farmer education and farm efficiency offer some counterpoint to the studies on African farming systems. Although the study was primarily intended to measure returns to schooling in an agricultural setting, it also provides some evidence on gender differences (or lack thereof) in technical efficiency in the rice-based "male" farming systems of Thailand and Korea. In contrast to the high percentage of female-headed households or female-managed farms in Africa, only 10.5 percent of the Korean farms are managed by women, and an even lower 1.1 percent of farms in the Thai sample are managed by women.⁸ Since the study does not focus on gender effects, only a dummy for the male household head is the indicator of gender differences in technical efficiency. Results show that the gender of the household head does not significantly affect output in both countries, except in Korean mechanical farms, where men have a 9.5 percent advantage. (No such advantage exists in nonmechanical farms.) While this finding is quite robust across alternative specifications, Jamison and Lau (1982) do not offer any explanation except

⁸ The Korean sample was divided into mechanical and nonmechanical farms based on the use of farm machinery. Only 9.8 and 12.4 percent of mechanical and nonmechanical farms were headed by women, respectively. Similarly, the Thai sample was divided into chemical and nonchemical farms based on the application of inorganic chemicals. Even smaller percentages of 2.2 percent of chemical farms, and 0.5 percent of nonchemical farms, were headed by women.

for a conjecture that there may be possible skill or strength requirements in the use of mechanical power. This explanation may be specific to the type of mechanical traction used, since, in other countries in East Asia (for example, Japan), women have been able to make use of small-scale machines, especially in transplanting and harvesting.

Coefficients estimated from the above studies have been used to simulate the gains from increasing women's levels of physical and human capital.⁹ Estimates of potential productivity gains simulated from the Mook coefficients range from yield increases of 7 percent (if female maize farmers were given sample mean characteristics and input levels) to 9 percent (if they were given men's input levels and other characteristics). Giving all women at least a year of primary education would raise yields by 24 percent, reflecting the gains to providing primary education in a setting where women have very low educational levels. Simulations with the Saito, Mekonnen, and Spurling (1994) coefficients suggest a 22-percent increase in women's yields on maize, beans, and cowpea plots if women farmers were given the human capital and input levels of male farmers.

However, these simulation results need to be interpreted with caution, since they do not reveal how levels of inputs may be raised. To a great extent, differences in input use may be driven by differences in education, since more-educated farmers are more likely to use modern inputs. Moreover, these simulations may also inaccurately depict the gains if a change in input use did occur, since the Cobb-Douglas production

⁹ See, for example, Saito, Mekonnen, and Spurling (1994) and World Bank (1994).

technology assumes constant elasticities, and presupposes that changing the levels of one input does not change the elasticities with respect to other inputs.

An alternative approach could have been to perform a Oaxaca decomposition of the yield differential between male and female farmers (Oaxaca 1973). Although this approach has been used to decompose the wage gap, it can also be applied here:

$$\bar{Q}_m - \bar{Q}_f = \hat{b}_m(\bar{X}_m - \bar{X}_f) + (\hat{b}_m - \hat{b}_f) \bar{X}_f, \quad (4)$$

where \bar{Q}_m and \bar{Q}_f represent mean yields of males and females respectively, \hat{b}_m and \hat{b}_f are estimated output coefficients of male and female farmers, and \bar{X}_m and \bar{X}_f are ^{mean} levels of endowments and inputs of male and female farmers. That is, the overall average male-female yield gap can be decomposed into the portion due to differences in input endowments ($\bar{X}_m - \bar{X}_f$), evaluated using male coefficients; the other portion is attributable to differences in the returns, or output elasticities ($\hat{b}_m - \hat{b}_f$), that males and females get for the same endowment or input application.

Labor Productivity Differences

Coefficients estimated from production functions have been used to measure gender differences in labor productivity.¹⁰ Output elasticities with respect to male and female labor are used to compute marginal products, usually evaluated at the sample mean. The validity of this approach depends on the assumptions embodied in the underlying production function, particularly those regarding the separability of the factors of production, as well as those regarding endogeneity of inputs.

Some studies have used rather restrictive production technologies, for example, a Cobb-Douglas production function with male and female labor further disaggregated into family and hired labor (Saito, Mekonnen, and Spurling 1994). However, the conclusions regarding the relative productivity of male and female labor are suspect because the Cobb-Douglas technology imposes additive separability on the production function. Other studies have used flexible functional forms and tested explicitly for separability of the production function.

¹⁰ Other approaches to measure gender differences in labor productivity have included the number of women of working age (or the proportion of women in the household labor force) in regressions on output as one of the explanatory variables in the production function. However, this is not a good proxy for women's share of labor input in agricultural production if some crops are more intensive in women's time than others, because of the sexual division of labor in agriculture; if the returns to crops grown by males and females vary; and if the value of home production is not measured correctly. Expressing the number of women in adult equivalent units is faulty, because these units are usually based on a food energy consumption measure, not in terms of labor input by task. Examples of the above include Jamison and Mook (1984), von Braun, Puetz, and Webb (1989), and von Braun, de Haen, and Blanken (1991).

For example, Kumar and Hotchkiss (1988) estimate translog production functions for various crops and test for separability between male and female labor in Nepal. Separability implies that the marginal rates of substitution between pairs of factors in the separated group (for example, male and female labor) are independent of the levels of factors outside that group (in their specification, land). The marginal products of men's and women's labor are found to be independently significant for all crops except early paddy, often with different marginal products. The separability test indicates that, except for the dry season crops, men's and women's labor are weakly separable; that is, they have different marginal rates of substitution with land. This suggests that, at least for some crops, there is little justification for aggregating men's and women's labor. For the dry season crops, women's higher marginal product of labor may be due to competing activities of fuel and water collection, which impinge on women's time for agricultural production.

Laufer uses ICRISAT data from six villages in India over a three-year period to estimate generalized quadratic production functions for three major crops with farm-level fixed effects to control for unobserved farm-specific factors (Laufer 1985). Weak separability of labor inputs with respect to all other nonlabor inputs was tested by imposing restrictions on the coefficients.

Likelihood ratio tests indicate that male and female labor are not (weakly) separable in the production of sorghum and rice. Male and female labor are good complements in sorghum and rice, but male labor is more substitutable for animal power

and land. Male labor is complementary with machines (mainly irrigation pumps), while female labor is substitutable. In contrast, male and female labor are found to be good substitutes in legumes, and equally complementary with animal power. In all crops, male and female labor are combined in relatively fixed proportions with land. Male and female labor could be complementary in sorghum and rice, but substitutable in legumes because of the different nature of tasks required to grow the different crops, and the traditional gender division of labor by task. While the marginal product of male labor is greater than that of female labor in all crops, the ratios of marginal products lie within the range of ratios of observed wages paid in the six villages, suggesting that farmers rationally equate relative wages to the ratios of marginal products.

Jacoby's (1992) study of the households in the Peruvian Sierra uses both Cobb-Douglas and sequentially restricted translog production functions to estimate marginal products and test for the substitutability of male and female labor. First, he estimated two Cobb-Douglas production functions for crops and livestock, respectively. He finds that while adult female labor does not have a significantly positive effect on crop output, it is highly significant in the livestock equation, with a coefficient almost twice that of adult males. Adult males have about the same coefficient in both equations. If peasant households optimally allocate time across tasks, the evidence suggests a sexual division of labor, with women spending relatively more of their time than men in livestock production, and men specializing in fieldwork. Second, Jacoby used a sequentially restricted translog production function to test for the nonseparability of male and female

labor. A sexual division of labor implies that labor productivity of men and women will be differentially affected by the presence of other inputs; that is, the marginal rate of transformation between labor of men and women will vary with the levels of other inputs. Since the marginal rate of transformation from a Cobb-Douglas production function depends only on the ratio of male and female labor inputs, the Cobb-Douglas production function imposes separability. Rather than do so, Jacoby sequentially tests for a nonseparable technology. He first tests for strong separability against the joint alternative of weak separability or nonseparability, then tests weak separability against nonseparability alone, through a series of translog regressions.

The conclusion from the sequence of tests is that only the nonseparabilities of male and female labor with respect to land and farm animals are statistically significant. This is attributable to the sexual division of labor in agriculture: since men are involved in tasks such as plowing and transport, farm animal input would tend to be more substitutable with the hours men spend on the farm than with women's hours. The use of more animals for plowing, for example, would replace adult male time spent with the foot plow, but would not cause women to substitute out of weeding. The strong negative interaction between female labor and farm size may also be a consequence of the sexual division of labor: households with little land tend to supplement their income with nonagricultural activities. In near landless households, men tend to spend more time in off-farm work, leaving most agricultural tasks in the hands of women. The elasticity of substitution between male and female labor at the geometric means of the data equals

1.84, which is higher than that allowed by the Cobb-Douglas specification (namely, one), but does not indicate extremely high substitutability.

The ratio of female-to-male marginal products of 0.64 indicates that men contribute more to total farm output at the margin than women, which is reasonably close to the relative wage. This may be due to women's tendency to be casual rather than regular farm workers in the Peruvian Sierra, and to spend more time, on average, in household and nonfarm business activities than in farm work. Jacoby argues that women may sort themselves into these activities because they are "innately" more productive in them than men, and that these productivity differences are accentuated by the acquisition of sector-specific human capital by both men and women. However, since the same agricultural tasks (for example, land preparation) are performed by women in other farming systems, it may not be valid to use the findings from the static measurement of a production technology to generalize about a division of labor that may be jointly determined by culture and comparative advantage. An alternative explanation proposes that there could be a premium related to body size in some agricultural operations that would then be reflected in higher productivity of men relative to women. This is explored in further detail below.

Another study analyzes women's agricultural productivity as a factor influencing men's demand for wives in Côte d'Ivoire, where polygyny is common (Jacoby 1993). Jacoby estimates a conditional profit function as a function of input prices, capital inputs, and male and female family labor inputs. Although the profit function is Cobb-Douglas,

it is modified to allow the labor productivity parameters to depend on permanent farm characteristics, including crop composition. Ordinary least squares estimates show that female labor seems to contribute a larger share to farm profit than male labor, though this difference decreases with first-differenced OLS estimates. After accounting for heterogeneity of labor inputs and using instruments thereof, however, men contribute more to farm profit than do women. When crop composition is taken into account, the contribution of female labor to profit is negatively related with cocoa, coffee, cotton, rice, and maize. Crops traditionally grown by women, such as tubers, interact positively (or only weakly negatively) with female labor. While the results seem to support specialization by crop according to comparative advantage, in the longer run, crop composition (and the gender division of labor by crop) is also endogenous. Indeed, recent evidence from Africa suggests that the gender division of labor by crop is changing, as women increasingly cultivate cash crops (Saito, Mekonnen, and Spurling 1994).

3. WAGE AND EARNINGS FUNCTIONS

METHODOLOGY

Estimates of labor productivity based on coefficients from production functions typically refer to the productivity of an additional unit of labor. Thus, they abstract from the heterogeneity of the agricultural labor force. The link between individual characteristics and wages is better explored by the literature on wage and earnings functions, which include (1) determinants of labor market participation (or, more generally, time allocation) of men and women and (2) determinants of wages or earnings.

An individual's decision to participate in the labor market is determined by trade-offs between income, leisure, and home production activities, subject to a full income constraint that takes into account an individual's income earning opportunities, given a time endowment. An individual will participate in the labor market if the market wage is greater than or equal to his or her reservation wage, which is determined by individual characteristics (for example, education and experience), household characteristics (for example, landholdings and assets), and market conditions. For women especially, the choice is often not simply between market wage participation and leisure, but allocation among market work, leisure, and home production activities. For members of agricultural households, the allocation may be among wage labor, own farm labor, home production activities, and leisure. The literature on bargaining models of the household and models of marital formation also suggests that individual unearned incomes and

spouse's characteristics affect time allocation decisions.¹¹ Community characteristics such as household proximity to services, presence of factories or small-scale industries, and characteristics of the agricultural production environment such as irrigation and seasonality also affect participation in wage labor activities.

An individual i 's ($i = m, f$, for male and female, respectively) participation in an activity k ($k = w, l, h$ for work, leisure, and home production, respectively) is given by

$$I_k^{*i} = \tau_{0i} + X_i \tau_{1i} + Z \tau_{2i} + e_i, \quad (5)$$

where I_{ik}^* is a vector of binary dependent variables; $I_{ik}^* = 1$ if individual i participates in the k th activity, 0 otherwise; X is a vector of individual characteristics that influences an individual's time allocation; Z is a vector of household and market factors affecting time allocation decisions; τ is the vector of coefficients to be estimated, and e is an error term. The participation equations are usually estimated using a maximum likelihood probit technique and the results used to correct for sample selectivity in the wage regression.¹²

Much of the empirical work on earnings functions follows Mincer (1974), drawing from Becker's (1964) work on human capital and dynamic human capital accumulation

¹¹ See, for example, Manser and Brown (1981), McElroy (1990), and McElroy and Horney (1981).

¹² See Heckman (1983) and Maddala (1983).

models (Yoram Ben-Porath 1967). The Mincerian earnings equation, taking sample selection into account, is given by

$$\begin{aligned} \ln w_i &= \alpha_0 + \beta S_i + \Gamma_1 E_i + \Gamma_2 E_i^2 + \mu_i & \text{if } I_i = 1 \\ I_i^* &= \tau_{0i} + X_i \tau_{1i} + Z_i \tau_{2i} + e_i & , \\ I_i &= 1 \text{ if } I_i^* > 0 \text{ and } I_i = 0 \text{ if } I_i^* \leq 0 \end{aligned} \quad (6)$$

where $\ln w_i$ is the natural logarithm of wages (or earnings) of the i th individual, observed only if the individual participates in the labor market; S_i is the number of years of schooling (or indicators of levels of schooling); E_i is the number of years of work experience, which enters as both linear and quadratic terms; and μ_i is the stochastic disturbance term. It is usually hypothesized that $\Gamma_1 > 0$ and $\Gamma_2 < 0$, and β is interpreted as the rate of return to an additional year of schooling. Estimates of rates of return to schooling are useful in evaluating the desirability of expanding educational opportunities in rural areas.

EMPIRICAL EVIDENCE

Table 2 presents estimated returns to schooling derived from wage functions for men and women in rural areas. In most of these studies, wages are expressed as a function of individual characteristics (education, experience), family characteristics (number of male and female workers, landholding, farm equipment, housing

Table 2—Estimates of returns to schooling for rural men and women

Study	Estimation Method	Schooling Category	Males	Females
<u>India</u>				
Rosenzweig 1980	Ordinary least squares	Years of schooling	0.10 (0.83)	0.02 (1.30)
Mukhopadhyay 1991	Semilog with selectivity correction for wage earner status	Years of schooling	0.0161*** (3.39)	0.0352*** (3.21)
<u>Sri Lanka</u>				
Sahn and Alderman 1988	Semilog with selectivity correction for labor force participation	Primary school	0.103** (1.90)	0.178** (2.32)
		Grade 6-10	0.145*** (2.61)	0.189** (2.18)
		General Certificate Exam	0.331*** (5.45)	0.571*** (6.77)
		University/Post- graduate	0.793*** (6.30)	1.060*** (6.28)
<u>Philippines</u>				
Behrman and Lanzona 1989	Semilog with selectivity correction for contractual and fixed wages	Years of schooling Wet Season	0.084*** (6.2)	0.067*** (4.0)
		Dry Season	0.105*** (6.4)	0.070*** (4.0)
<u>Peru</u>				
Khandker 1990	Ordinary least squares	Primary	0.05 (1.56)	0.05 (0.63)
		Secondary	0.06** (2.26)	0.10 (1.11)
		Postsecondary	0.21** (2.29)	0.20 (1.30)
	Maximum likelihood	Primary	0.06 (1.63)	0.08 (0.83)
		Secondary	0.09*** (2.60)	0.13 (1.11)
		Postsecondary	0.26*** (3.51)	0.27 (1.02)
	Household fixed-effect	Primary	0.11*** (3.20)	0.37*** (5.73)
		Secondary	0.17*** (3.46)	-0.02 (0.26)
		Postsecondary	0.42*** (3.27)	0.26 (1.66)

*** Significant at 1 percent.

** Significant at 5 percent.

* Significant at 10 percent.

indicators), and community characteristics (distance to the village, presence of rural industry, average wages and prices, seasonality dummies, village dummies). In situations where entry into wage labor may be a function of individual characteristics, a probit for labor market participation was estimated and used as a selectivity correction. Some of the family characteristics, particularly those related to wealth and nonlabor income (land, farm equipment, unearned income), or previous decisions (marital status, number of male and female workers in the household) are used as identifying variables in the probit equation. Although the scope of these studies may be broader than estimating wage functions, only this aspect is discussed here.

In rural India, returns to education are low because wages depend mainly on local market characteristics rather than personal characteristics like education. Rosenzweig (1980) tests predictions about labor supply behavior of landless and landholding households derived from a neoclassical utility-maximizing model based on competitive assumptions. In the model, if schooling augments efficiency, there may be a greater demand for own-farm labor, and the response of market labor supply to educational levels in landholding households will be algebraically less than that in landless households. The same result also holds for differential experience, if such experience is relevant to managerial efficiency only on a household's own land. The model also predicts that household members on farms with more productive assets will participate less in the labor market, since higher asset levels increase the demand for labor time in farm production, and through the income effect, increase the demand for leisure.

These predictions are tested using data from the 1970-71 round of a three-round national sample survey of rural households in India. Selectivity may not be important in this sample since, due to geographic immobility of rural households and the nature of rural occupations, wage rates are influenced more by community characteristics than by personal attributes, after taking gender into account.

The coefficient on years of schooling from a simple earnings function suggests that the rate of return to an additional year of schooling for men is 6 percent. This drops to 3.9 percent when village-level variables (weather, presence of factory and small-scale industries, village size, distance of residence from village, and presence of an agricultural development project) are added to the regressors. When the district-level average male wage is included, however, the coefficient on schooling drops to 1 percent and is no longer significant. The rate of return to an additional year of schooling for women is in the neighborhood of 2 percent, but the coefficient is not significant in any of the regressions. The size of the village, distance of the household to the village, and the average district-level female wage are significant determinants of female wages. The results suggest that labor is not perfectly mobile geographically in rural India, and wage rates are not greatly affected by human capital attributes in nonsalaried, private-sector occupations that characterize rural labor markets.

The lack of significance of schooling in the more fully specified equations of nonsalaried nongovernment workers does not mean that schooling does not increase earnings in India. It is highly correlated with salaried or government jobs, whose

computed mean wage rates are higher than those observed in the sample of rural workers used.

In contrast, a more recent study in West Bengal, India, finds positive and significant returns to male and female schooling. Mukhopadhyay (1991) estimates sample selection-corrected wage functions for rural men and women age 15 to 65. The selectivity equations suggest that education, age, land size, and assets reduce participation in the hired agricultural labor force. A Muslim man is likely to work as an agricultural wage worker, but a Muslim woman is not, since Muslim women are "proscribed" from working in the field as farm labor in West Bengal. Technology variables also decrease the probability of women's participation in the market for agricultural labor.

Daily wages of men and women are significantly affected by age (which captures returns to experience) and schooling. The returns to experience seem to be greater for men, while the private rate of return to female schooling (3.5 percent) is larger than the return to male schooling (1.6 percent). Moreover, women with larger areas of land apparently are able to command higher wages as agricultural workers. The coefficients of the inverse Mills ratio from the selectivity equation are significant but of opposite sign for both men and women, suggesting that low wage male workers are likely to work in wage employment, while high wage female workers are more likely to work in agricultural wage labor.

The positive and significant returns to schooling from this study, based on a 1990 data set, are a contrast to the insignificant rate of return to both male and female

schooling in the earlier study based on a 1970-71 survey. It is possible that the diffusion of the "green revolution" technology in the 1970s and 1980s may have created an environment where returns to schooling can be significant. Control of the flow of irrigation water, and the timely application of fertilizer, insecticide, and other chemical inputs associated with the seed-fertilizer technology may have raised skill requirements for agricultural laborers, such that more educated laborers may receive higher wages.

Another view from South Asia is offered by a study of wage determinants in Sri Lanka, a country which has high enrollment rates and educational attainment for both men and women relative to countries at similar per capita income levels. Sahn and Alderman (1988) correct for self-selection into the labor force and infer that labor force participation is influenced by personal and household characteristics (such as age, education, landownership, marital status, the presence of young children and the elderly in the household, the number of males and females in the household, distance to markets, and nonlabor income). The selectivity-corrected wage equations show that in rural areas, for both men and women, schooling has a positive effect on wages of both men and women, with gross rates of return increasing with the level of education. Estimates of internal rates of return for continuing one's education to pass the General Certificate of Exams are twice as high for women than for men in rural areas (14.4 compared to 7.4 percent), although the returns for graduating from university are about the same for men and women (13 to 14 percent).

Consistent with the hypothesis that technological change in agriculture increases returns to schooling, Behrman and Lanzona (1989) find substantial returns to education for men and women in five Philippine rice villages, where modern rice varieties were planted in 87 percent and 79 percent of rice area in the wet and dry seasons, respectively. They estimate selectivity corrections for fixed-rate and piece-rate wages to control for the higher work intensity of the contractual (piece-rate) arrangement that is reflected in higher mean wages.

Adult schooling is a significant determinant of reported wages in both seasons, after controlling for hourly or piece-rate contracts. Male rates of return to an additional year of schooling are 8 percent and 10 percent in the wet and dry seasons, respectively. The rate of return to female education is 7 percent in both wet and dry seasons; however, the difference between the sexes is not statistically significant. At the sample means, the estimates imply that an additional year of schooling increases the daily wage for men by about 0.33 pesos per day in the dry season (from a mean predicted wage of 10.36 pesos per day) and 0.42 pesos per day in the wet season (from 16.03 pesos per day), with similar increases for women of about 0.28 and 0.25 pesos per day (from means of 8.10 in the dry season and 10.05 in the wet season). The additive terms for individual types indicate that women receive significantly lower wages than men (about 48 percent lower) with equal schooling and age in the wet season, though there is no significant difference in the dry season. The selectivity controls for fixed and contractual wages are also

significant, which implies that those who participate in fixed wage contracts are less productive.

Khandker (1990) uses household survey data from the 1985-86 Peruvian Living Standards Survey (PLSS) to estimate differences between males and females in labor market participation, productivity (as measured by wages), and returns to schooling. A probit for labor market participation was estimated separately for males and females in rural areas, with unearned income, landholding, and marital status as identifying variables. The wage equations with selectivity correction were estimated using maximum likelihood techniques. While schooling and experience are insignificant determinants of rural women's wages, the maximum-likelihood results indicate that schooling has high rates of return for men in rural areas. The insignificance of the regressors in the female rural wage equation is due to the high standard error of the wage regression. There are relatively few women working in the wage sector in rural areas, and since most of them work as teachers or clerks, the variation in wages is small.

Household fixed effects estimation is used to control for the effects of unobserved household characteristics. Rates of return from ordinary least squares, maximum likelihood, and fixed effects estimation are summarized in Table 2. The correction for unobserved household heterogeneity increases the returns to male education at all levels but decreases returns to women's education at secondary and postsecondary levels. Rates of return to male schooling from the maximum likelihood estimates are 6 percent at the primary level (compared to 8 percent for women); 9 percent for secondary schooling

(compared to 13 percent for females); and 26 percent for postsecondary (27 percent for females). The fixed effects estimates yield an 11 percent rate of return to male schooling at the primary level (compared to 37 percent for females); 17 percent at the secondary level (compared to -2 percent for females); and 42 percent at postsecondary levels (and 26 percent for females). This disparity indicates that, in rural areas, parents may have reasons for investing less in daughters than in sons, although, overall, the private rate of return is higher for women.

The above studies suggest that returns to schooling for both men and women, although higher in nonagricultural occupations, are significant in dynamic agricultural settings where modern technologies have been introduced. In contrast, where labor market participation of women is limited or constrained by cultural factors, returns to female schooling are low, and women receive significantly lower agricultural wages than men.

However, low female agricultural wage rates may not be reflected in rates of return to schooling. In many countries, women earn about half the male agricultural wage, but most daily wage earners in agriculture are uneducated. The difference between male and female wages could reflect productivity differences, since physical size may affect earnings. While size may not affect certain work (for example, transplanting, weeding, and farm management), it could have a premium in, for example, land preparation. A study of a rural public works program in India finds that while being female per se is not a deterrent to participation, taller individuals are more likely to participate, with men

being taller than women on the average (Deolalikar and Gaiha 1992). To the extent that height is an indicator of long-term nutritional status, and that, in some societies, boys are favored in nutrition and health outcomes, gender differentials in agricultural wages could reflect life-cycle effects of intrahousehold resource allocation.¹³

4. GENDER DIFFERENCES IN TECHNOLOGICAL ADOPTION

METHODOLOGY

Studies of the adoption of agricultural innovations at the farmer level are concerned with analyzing the determinants of the degree of use of a new technology in a long-run equilibrium when the farmer has full information about the new technology and its potential (Feder, Just, and Zilberman 1985, 256). The adoption of a nondivisible technology (such as farm machinery) is usually measured in terms of a dichotomous variable, whereas the adoption of divisible technologies (such as modern varieties or variable inputs) is indicated by the intensity of input use (for example, quantity applied, or quantity applied per hectare) or the percentage area using the new technology.

To estimate the determinants of the intensity of adoption, given the decision to adopt or not to adopt a technology, the appropriate procedure would be to use a two-stage procedure, and first estimate the probability of adoption, using a probit regression. In the

¹³ These differences in wages, and differences in work energy intensity, could also affect the current allocation of nutrients within the household.

second stage, the intensity of input use could be estimated either by ordinary least squares, with the selectivity correction from the first equation, or both equations could be estimated simultaneously by maximum likelihood, similar to the estimation of earnings functions with corrections for participation discussed above.

Simultaneous equations considerations are also important in modeling the adoption of agricultural technology because (1) the impact of extension on technological adoption may be influenced by farmer characteristics; (2) lagged awareness and adoption of technologies may influence current adoption decisions; and (3) input use is often interdependent, as in the adoption of a package of technologies (Birkhaeuser, Evenson, and Feder 1991). These issues can be addressed in two-stage (and often, iterative) estimation techniques, in which lagged cumulative adoption or awareness probabilities are computed. For example, the probability of receiving extension advice can be modeled as a function of farmer and extension system characteristics in the first stage. The first stage results can then be used to correct for sample selection in the awareness (or adoption) equation. Alternatively, if panel data exist, the lagged cumulative predicted adoption probabilities of other households can be used as a proxy for copying effects.

Interdependent input adoption decisions can be addressed by treating the demand for new technologies as one of the input demand functions in the dual approach. Given a cost function or a restricted profit function, differentiation with respect to the input price yields an input demand function that is a function of input and output prices and

fixed inputs. These equations can be estimated jointly with cross-equation restrictions to take advantage of increased efficiency of estimation.

EMPIRICAL EVIDENCE

This section reviews the evidence on differential adoption of agricultural innovations by gender; selected coefficients are summarized in Table 3.

Appleton et al. (1991) estimate logit equations to examine gender effects in investment decisions in coffee, cocoa, and livestock adoption in Kenya, Tanzania, and Côte d'Ivoire. They employ an iterative technique to estimate the effects of sequential and late copying on adoption. Three possible gender effects are considered: (1) differences between male and female household heads or decisionmakers; (2) differential effects on adoption of the availability of male and female labor; and (3) gender-specific copying effects. The first is captured by the gender dummy, the second by the number of (adult) males and females in the household, and the third by measures of previous and current adoption by other farmers, disaggregated by gender.

Table 3—Gender-related determinants of technology and crop adoption

Study	Dependent Variable	Household Demographic Variables		Human Capital Variables		Extension, Adoption, and Copying Variables		Wage and Price Variables	
		Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a
<u>Kenya</u>									
Appleton et al. 1991	Coffee adoption (logit)	Dummy for female decisionmaker	-6.2* (1.7)	Primary education dummy	-0.47 (0.7)	Percentage of households growing the crop	
		Number of adults	0.22 (1.5)	Female decision-maker education	4.82 (1.5)	Same sex	2.1 (1.1)		
				Age	0.065 (0.6)	Opposite sex	0.3 (0.2)		
				Age squared	-0.001 (1.0)	Early adopters Same sex	7.0* (1.8)		
						Opposite sex	1.0 (0.5)		
	Livestock investment (logit)	Dummy for female head	-0.39 (1.3)	Primary education dummy	0.58 (1.9)*	Proportion of households in cluster already owning livestock	1.49*** (2.4)
				Age	0.055 (1.1)				
				Age squared	-0.00035 (0.7)	Average expected probability of livestock adoption by other households in cluster	-.27 (0.3)		
<u>Tanzania</u>									
Appleton et al. 1991	Livestock investment (logit)	Dummy for female head	-1.2 (1.6)*	Primary education dummy	0.53 (1.1)	Proportion of households in cluster already owning livestock	6.0*** (4.5)
				Age	0.17 (1.5)				
				Age squared	-0.0019* (1.6)				

(continued)

Table 3 (continued)

Study	Dependent Variable	Household Demographic Variables		Human Capital Variables		Extension, Adoption, and Copying Variables		Wage and Price Variables	
		Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a
<u>Côte d'Ivoire</u>									
Appleton et al. 1991	Coffee adoption (logit)	Dummy for female head	-1.00 (1.0)	Primary education dummy	0.018 (0.0)	Dummy for cooperatives	0.46 (0.9)
		Number of males	0.0078 (0.1)	Age	-0.064 (0.9)	Percentage of households already growing coffee	1.75** (2.2)		
		Number of females	-0.0539 (0.4)	Age squared	0.0004 (0.5)	Dummy for cooperatives	0.452 (1.3)		
	Cocoa adoption (logit)	Dummy for female head	0.099 (0.2)	Primary education dummy	0.405 (1.4)	Percentage of households already growing cocoa	1.844*** (2.7)
		Number of males	0.124 (1.2)	Age	0.0049 (0.1)	Average expected probability of cocoa adoption by other households in cluster	1.269 (0.8)		
		Number of females	-0.046 (0.5)	Age squared	-0.00018 (0.3)				
<u>Zambia</u>									
Jha, Hojjati, and Vosti 1991	Fertilizer use (probit)	Female head dummy	-0.600 (-1.623)	Education of head	0.071 (1.235)	Extension advice dummy	0.491 (1.201)
		Dependency ratio	0.008 (0.040)	Age	-0.040*** (-2.735)	Predicted use of hybrid maize	-2.824* (-1.763)		
						Predicted use of oxen	-2.750 (-1.290)		
	Ox cultivation (probit) (1=ox, 0=hoe)	Female head dummy	-0.657* (-2.095)	Education of head	0.016 (0.365)	Extension advice dummy	-0.193 (-0.543)
		Dependency ratio	0.228 (1.230)	Age	-0.031*** (-3.142)	Predicted use of fertilizer	-0.847* (-1.754)		
						Predicted use of oxen	0.450 (0.880)		

(continued)

Table 3 (continued)

Study	Dependent Variable	Household Demographic Variables		Human Capital Variables		Extension, Adoption, and Copying Variables		Wage and Price Variables	
		Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a
<u>Zambia</u> Jha, Hojjati, and Vosti 1991	Hybrid maize use (probit)	Female head dummy	-0.297 (-1.105)	Education of head	0.075* (1.832)	Extension advice dummy	0.578 (1.622)
		Dependency ratio	-0.100 (-0.546)	Age	-0.017* (-1.911)	Predicted use of fertilizer	-0.336 (-.697)		
						Predicted use of hybrid maize	-1.176** (-2.569)		
	Cotton, soybean, or sunflower growing (probit)	Female head dummy	-0.013 (-1.477)	Education of head	-0.009 (-.230)	Extension advice dummy	0.499 (1.610)
		Dependency ratio	0.084 (0.516)	Age	-0.013 (-1.477)	Predicted use of fertilizer	-0.121 (-0.287)		
						Predicted use of oxen	0.188 (0.507)		
	Plant nutrients used per fertilized hectare (kg.)	Female head dummy	-1.138 (0.13)	Education of head	2.718* (1.658)	Extension advice dummy	3.684 (0.294)
		Dependency ratio	-0.991 (-0.133)	Age	-0.155 (-0.414)	Predicted use of hybrid maize	62.660 (1.610)		
						Predicted use of oxen	-42.946 (-0.962)		
	Percent area fertilized	Female head dummy	-0.067 (-1.338)	Education of head	0.007 (0.809)	Extension advice dummy	-0.015 (-1.338)		
		Dependency ratio	-0.022 (-0.609)	Age	0.0002 (0.085)	Predicted use of hybrid maize	-0.025 (-0.134)		
						Predicted use of oxen	-0.049 (-0.224)		

(continued)

Table 3 (continued)

Study	Dependent Variable	Household Demographic Variables		Human Capital Variables		Extension, Adoption, and Copying Variables		Wage and Price Variables	
		Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a
<u>India</u>									
Mukhopadhyay 1991	Proportion of rice area under high-yielding varieties	Men's schooling	-0.00102 (.20)	Ln agricultural predicted male wage	-.177* (1.68)
				Women's schooling	-0.00103 (.19)			Ln agricultural predicted female wage	.143*** (2.00)
<u>The Philippines</u>									
Behrman and Lanzona 1991	Percentage area direct seeded (wet season)	Men's schooling	10.87* (1.7)*	Rice price	100.86*** (5.4)
				Women's schooling	9.33*** (3.16)			Fertilizer price	-3.59*** (5.4)
				Men's age	13.13*** (2.9)			Predicted male wage	-191.18** (2.5)
				Women's age	-3.59*** (4.5)			Predicted female wage	-65.06* (1.9)
				Age squared of men	-0.13*** (2.5)				
	Percentage area under modern varieties (wet season)	Men's schooling	17.38*** (4.2)	Rice price	55.29*** (4.70)
				Women's schooling	8.33*** (5.0)			Fertilizer price	-0.38 (0.90)
				Men's age	12.85*** (4.4)			Predicted male wage	-193.55*** (3.90)
				Women's age	-4.53*** (8.8)			Predicted female wage	-118.47*** (5.40)
				Age squared of men	-0.13*** (4.1)				

(continued)

Table 3 (continued)

Study	Dependent Variable	Household Demographic Variables		Human Capital Variables		Extension, Adoption, and Copying Variables		Wage and Price Variables	
		Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a	Definition	Coefficient ^a
<u>The Philippines</u>									
Behrman and Lanzona 1991	Percentage area using tractors (wet season)	Men's schooling	5.40** (2.4)	Rice price	0.50 (0.10)
				Women's schooling	-0.78 (0.90)			Fertilizer price	4.04*** (17.50)
				Men's age	1.45 (0.9)			Predicted male wage	-26.98 (1.0)
				Women's age	-0.89*** (3.2)			Predicted female wage	-36.94*** (3.10)
				Age squared of men	-0.01 (0.6)				
<u>Guatemala</u>									
von Braun, Hotchkiss, and Immink 1989	Export crop adoption (probit)	Total labor available in household	0.085 (-1.544)	Years of schooling of household head	0.323 (0.692)
		Share of female labor in total labor of household	-1.004** (-2.016)	Age	-0.008 (-0.833)				

*** Significant at 1 percent.
 ** Significant at 5 percent.
 * Significant at 10 percent.

The authors distinguish two types of copying effects. Sequential copying is measured by the coefficient of the percentage of households in the cluster who have adopted earlier. Simultaneous copying is based on the estimated probability that other households will adopt.

Simultaneous copying is estimated using a multistage procedure. In the first stage, the probability of a household *i*'s adoption is modeled as a function of a set of explanatory variables. The estimated coefficients are then used to predict each household's adoption probabilities, and these estimated probabilities are aggregated over all the other households in a cluster. In the next stage, the average of the estimated adoption probability of other households is added as a regressor, and the adoption logit is reestimated, providing a first estimate of household *i*'s propensity to adopt as a function of other households' adoption probability. This procedure is repeated until the estimated value of the adoption propensity converges. Gender-specific copying effects are proxied by entering separately the proportion of growers headed by a person of the same sex as the household under consideration, and the proportion of growers of the opposite sex.

Application of the above technique to the analysis of coffee adoption in Kenya finds that female decisionmakers are less likely to grow coffee than male decisionmakers, although the negative coefficient of the female dummy is only weakly significant. More educated females, however, are more likely to grow coffee. While simultaneous copying from growers of the same or opposite sex is not important, early adoption by growers of

the same sex has a positive and significant effect on the probability of coffee adoption. Thus, household decisionmakers are more likely to copy from an adopter of the same sex.

Since the importance of the gender effect cannot be judged directly on the basis of the estimated coefficients, the authors use these estimates in a series of simulations. In one simulation, the 74 female-headed households are assumed to be male-headed, and the number of households predicted to adopt increased from 24 to 91. While it is easy to argue that coffee would be more widely adopted if female-headed households were as likely to adopt as male-headed households, the simulation masks the unobservables that the female dummy captures.

An additional set of simulations, which simulates the percentage increases in average probabilities of adopting coffee due to a 10-percent increase in each of the explanatory variables, is more informative.¹⁴ Increases in women's education have greater effects on coffee adoption than increases in land size. The elasticity of the adoption probability with respect to education is 0.6 for early adoption and 1.4 for late adoption, while corresponding elasticities for land are only 0.2 and 0.6, evaluated at the means of all variables in the group of female decisionmakers.

In contrast to coffee adoption, gender differences in livestock adoption in Kenya are insignificant. Better educated farmers, farmers with more land, or those who have previously grown coffee, had a farm, or a bank account in 1975 were more likely to

¹⁴ Personal communication from Kees Burger and Jan Willem Gunning, 12 March 1992.

invest in livestock. In addition, the sequential copying variable (the percentage of other households in the cluster who previously owned cattle) is significant, but the simultaneous copying variable, a proxy for other households' current adoption, is not. In Tanzania, female-headed households are less likely to have livestock, but households with more land are more likely to have cattle. Previous adoption by other households is also an important influence on a household's probability of adoption. In Côte d'Ivoire, while farmers with larger land areas are more likely to adopt either coffee or cocoa, and the existence of a larger percentage of other households who already grew these crops increase the likelihood that a particular household would grow coffee or cocoa, gender was an insignificant determinant of the probability of coffee or cocoa adoption.

An evaluation of the training and visit (T&V) extension system in Kenya estimated the probabilities of awareness and adoption of specific technologies, namely, spacing, improved seed use, top dressing, chemical use, and stalk borer control (Bindlish and Evenson 1993). The availability of information from a 1981/82 Rural Household Budget Survey, in addition to the 1989-90 extension survey, made the construction of lagged and cumulated endogenous variables possible; these included lagged field extension worker to farm household ratios, lagged sublocation staff dummies, and three cumulated logistic variables to capture "learning from neighbor" effects of lagged advice, awareness, and adoption.

Results from the probit regressions suggest that household size and age of the household head were generally unimportant determinants of awareness or adoption,

although more schooling increased the likelihood of adoption. Farmers from female-headed households were as likely to become aware of and adopt technologies as those from male-headed households, and had a higher probability of adopting complex practices such as top dressing, chemical use, and stalk borer control. The significant coefficients of the lagged adoption variables (and the insignificant coefficients of the lagged awareness variables) suggest that neighbors influence adoption more through observable adoption of technology than through awareness of the technology.

An instrumental variables approach was used to predict fertilizer usage, cultivation by oxen, use of hybrid maize, and use of cotton, soybeans, and sunflower by smallholders in Eastern Province, Zambia (Jha, Hojjati, and Vosti 1991). Probit regressions for the adoption of the above technologies were estimated using maximum likelihood. Predicted values were then used in a second-stage regression. The inverse Mills ratio from the fertilizer use equation was used as a selectivity correction in second-stage regressions of percent area fertilized and nutrients used per fertilized hectare.

Although the female head dummy is negative in all of the regressions, it is significant only in the equation for ox cultivation. Education of the household head does not significantly affect the likelihood of adoption, but older household heads seem less likely to adopt new technologies. Extension advice is not significant, and the predicted values of hybrid maize use, fertilizer use, and ox cultivation have unexpected negative signs, except in the nutrient application equation. As expected, better infrastructure positively affects fertilizer use and nutrient application intensity. The only variables that

appear to affect adoption consistently are land area (positive), the cooperatives dummy (positive), and the location dummies.

Differences of male and female farmers in technological adoption cannot be identified in family farming situations where males and females jointly cultivate agricultural land and pool the incomes therefrom. Instead, in such countries as India, the Philippines, and Guatemala, various studies on the adoption of modern technologies or high-valued crops examine the differential effects of male and female schooling, experience (proxied by age), male and female wages, and household composition.

In West Bengal, India, environmental factors such as the suitability of land for high-yielding variety (HYV) rice, proportion of irrigated cultivable land, and the yield ratio of HYV to traditional varieties are the most important determinants of the percentage area planted to high-yielding rice varieties (Mukhopadhyay 1991). Although HYVs have been argued to be riskier than traditional varieties, recent evidence suggests that the additional risk is relatively small, and given land quality, assured irrigation, and expected higher profits does not significantly affect the decision to adopt. Surprisingly, neither male nor female schooling significantly affects HYV adoption. The size of land owned and value of assets are also unimportant, consistent with findings that the HYV technology in rice is scale-neutral. Higher female wages and lower male wages, however, are positively related to the adoption of HYV technology. Since the new technology uses a higher proportion of male-to-female labor, an increase in male wages would increase cost and decrease the proportion of land under modern varieties.

Conversely, an increase in women's wages would decrease the area under traditional varieties and encourage HYV adoption.

In the Philippines, land size, irrigation, and tenure status are important determinants of modern technology adoption (Behrman and Lanzona 1989). In contrast to India, men's and women's schooling has a significant and positive effect on the percentage area directly seeded and planted to modern varieties (MVs), while only men's schooling significantly affects tractor use. While tractor use is invariant to male age, direct seeding and modern variety use exhibit diminishing returns to male age, a proxy for experience. Older women are less likely to adopt any of the modern technologies. As expected, higher rice prices and lower input costs (fertilizer price, male and female wages) increase the percentage of area planted to modern varieties. The negative impact of both male and female wages is probably due to increased demand for both male and female labor compared to traditional varieties, since MVs require more weeding (a traditionally female task) due to increased fertilizer use, and the increased yields require more harvest and postharvest labor. Furthermore, MVs increase the demand for hired labor, not only because of increased seasonal demand, but also because family labor declines absolutely, as women in farm households shift to more lucrative enterprises and provide supervision rather than labor in farm production.

Finally, a study of the determinants of export crop adoption by smallholders in the Western Highlands of Guatemala finds that the probability of export crop production is positively related to farm size, and negatively affected by the share of women's labor in

the household, off-farm income, and traditional attitudes of the household head toward growing maize for food (von Braun, Hotchkiss, and Immink 1989). Age and education of the household head do not significantly affect the likelihood of growing export crops. Estimates of marginal effects show that an increased share of women's labor in total labor significantly reduces the probability of growing export vegetables, controlling for the total labor force of the household. This could be because adoption of the new crop is primarily a male decision, consistent with larger differentials between male and female education in Guatemalan Indian agricultural areas compared to the rest of mestizo Latin America.

5. SUMMARY AND RESEARCH ISSUES

SUMMARY AND POLICY IMPLICATIONS

Six out of seven studies on differences in technical efficiency between male and female farmers found insignificant dummies for the gender of the farm manager or household head. That is, female farmers are equally efficient as male farmers, once individual characteristics and input levels are controlled for. The only exception was the study on Burkina Faso, where the female farmer dummy was negative and significant. In this setting, women may be more constrained by cultural factors from having more active roles, and levels of education and technical development are lower. More recent work on Burkina Faso suggests that lower input intensities on women's plots, which

result in lower yields, result from asymmetric roles and obligations within the household—casting doubt on the assumption of Pareto efficiency.

The neglect of endogenous input choice does not consider possible relationships between farmer characteristics and input use, and thus does not address how actual levels of input use are to be increased. To the extent that better educated farmers are more likely to adopt modern inputs, these studies underestimate the consequences of underinvestment in women's education in rural societies. Moreover, even land size, which is usually assumed to be exogenous, is correlated with land quality. In societies where land is allocated to women based on marriage, land size is unlikely to be exogenous.

Despite the mixed evidence on technological adoption by gender, most of the technology adoption studies reviewed find that better educated farmers, regardless of gender, are more likely to adopt new technologies. Increasing the educational level of female farmers by giving them universal primary education has higher marginal effects on the probabilities of adoption than increasing the educational level of male farmers, due to the generally lower levels of female education in most rural areas. Previous awareness and adoption of modern technology, particularly by farmers of the same sex, also increased the probability of current adoption. The significance of gender-specific copying effects highlights the need not only for female extension agents to work with female farmers, but also for contact farmers to be women, if women farmers are more likely to copy innovations from other women. Most of the studies reviewed also suggest

that farmers with larger areas cultivated and higher values of farm tools are more likely to adopt new technology. To the extent that women farmers may have less education, less access to land, and own fewer tools, they may be less likely to adopt new technologies.

It is difficult to generalize whether an additional unit of male or female labor is absolutely more productive because the gender division of labor varies widely across crops, tasks, and farming systems. Labor scarcity, and not gender, is more strongly linked to higher marginal products of labor. In Nepal, the higher marginal product of women's labor in dry season crops is probably due to the demands of traditional tasks, which reduce time in agricultural production (Kumar and Hotchkiss 1988).

Although returns to schooling for both men and women appear to be higher in nonagricultural occupations, they are significant in dynamic agricultural settings where modern technologies have been introduced. In contrast, in areas where labor market participation of women is limited or constrained by cultural factors, returns to female schooling are low, and women receive significantly lower wages than men. These generally lower returns for females in these settings may be one reason why parents invest less in their daughters' education.

These wage studies, however, may not reveal some important sources of the male-female agricultural wage gap. In many countries, women earn about half the male agricultural wage, but most daily wage earners in agriculture are uneducated. The difference between male and female wages could reflect size-related productivity

differences, since physical size may affect earnings in some agricultural tasks. To the extent that size, proxied by height, is an indicator of long-term nutritional status, and that, in some societies, boys are favored in nutrition and health outcomes, gender differentials in agricultural wages could reflect life-cycle effects of intrahousehold resource allocation (Behrman 1994).

METHODOLOGICAL ISSUES

Rigorous measurement of gender differences in agricultural productivity is still a relatively new field of research. A number of conceptual and methodological issues should therefore be addressed in future work.

Gender Differences and Publication Bias

The above review has found that when differences in yields, earnings, or technological adoption are attributed to differences in individual characteristics (human or physical capital) and input levels, gender per se is often an insignificant determinant of agricultural productivity. This may be surprising to those who expect intrinsic gender differences. However, any review of the comparative literature may be biased, since it is more likely that studies that report differences in productivity will be submitted for publication, and, given that, are more likely to be accepted (Haddad, Alderman, and

Hoddinott 1994). Such "publication bias" is common to all scientific disciplines, but has been found to exaggerate the prevalence of studies that show differences between cognitive styles of men and women.¹⁵ When such studies of gender differences in cognitive processes are more carefully analyzed, a large number do not reject the null hypothesis of no difference in cognitive outcomes, nor adequately control for characteristics of the male and female samples.

While this may seem to be a philosophical point, there is a real danger in attributing differences in agricultural productivity solely to gender. In so doing, researchers and policymakers are not able to identify whether there are systematic differences in the distribution of underlying variables—many of which are subject to policy intervention—between men and women. By addressing differences in the characteristics that contribute to lower yields, earnings, or technological adoption of female farmers, appropriate agricultural policy interventions can be better designed. Closing the gaps in educational attainment and relieving exogenous constraints in access to markets and resources are a case in point.

Gender Dummies

Most of the studies estimating gender differences in technical efficiency use a dummy variable for female (or male) household head, farmer, or decisionmaker. While

¹⁵ On publication bias, see Begg and Berlin (1988). On publication bias in studies of cognitive differences between men and women, see Fausto-Sterling (1992).

headship may count, a dummy variable masks intrahousehold decisionmaking processes. What makes a female household head different from a male head? Does a simulation in which female-headed households are assumed to be male-headed make sense? There is a lot of institutional texture that is hidden by the use of headship dummies—one does not actually know what male and female *members* of the household do. The existence of female-managed plots in male-headed households is a case in point. It may be more appropriate to disaggregate by gender of the plot manager, rather than by gender of household head, particularly in societies where individuals in households farm their own plots (Saito, Mekonnen, and Spurling 1994; Udry 1994).

Biases in Valuation of Output

Greater disaggregation of decisions and tasks may provide more insight into household decisionmaking processes, but this is constrained not only by the lack of gender-disaggregated data, but also by conceptual difficulties in defining what to measure. Even among carefully conducted econometric studies, the definition of output can undercount women's production. With the exception of the studies of monocrop farmers, the studies reviewed used gross value of output, either in absolute terms or per land area. If women grow a larger share of subsistence crops, it is possible that regressions on gross value of output may understate female farmer's technical efficiency.

While this issue can be avoided by dealing exclusively with monocrops, it is an oversimplification of the diverse farming systems in which women operate. One appropriate procedure to use in a multicrop setting would have been a profit function approach, in which prices of inputs and outputs are explanatory variables.

Another controversial issue is the valuation of home production, which none of the studies mentioned here addresses. It has been argued that since women attend to domestic chores while performing agricultural labor, their productivity in agriculture is lower. However, this judgment is based on a conventional output measure, which disregards the value of home production. This measurement issue has yet to be satisfactorily addressed.¹⁶

Simultaneous Equations and Selectivity Considerations

With a few exceptions, most of the studies reviewed do not properly model simultaneous agricultural decisions. This is especially true in the technological adoption equations, where input use logits are estimated without taking into account their interdependent nature. There are very few studies in which crop choice or fertilizer use is first modeled, and then intensity of input use is then analyzed, conditional on crop choice or decisions regarding use or nonuse of fertilizer. The inclusion of endogenous

¹⁶ Among the few studies that estimate the productivity of women's work at home is Singh and Morey (1987).

explanatory variables such as crop mix and off-farm income in adoption equations could create simultaneous equations bias.

Life-Cycle Implications of Intrahousehold Allocations

The above review has not analyzed the processes that determine the allocation of human and physical capital to men and women. Educational attainment, land ownership, and long-term nutritional status of men and women are the outcomes of their parents' or kinship groups' allocation decisions. Increasing evidence suggests that these decisions may involve nonunitary preferences, or even noncooperative behavior. Whatever the process that underlies these decisions, parental allocation decisions will have long-term implications on the productivity of children. Underinvestment in girls' education could, for example, lead to lower probabilities for female farmers to adopt new technologies; inheritance laws which favor boys would imply that men would have greater inherited landholding sizes; allocation of nutrients towards boys would have long-term effects on height and productivity.

The importance of addressing intrahousehold allocation issues in policy formulation has been discussed elsewhere; this point is reiterated here.¹⁷ Further research should be directed not only towards understanding the results of intrahousehold allocations, but, more importantly, to understanding the processes of decisionmaking in agricultural households, by male and female heads, and family members. There is much

¹⁷ See Haddad, Alderman, and Hodinott (1994) for a review.

to be learned from insightful fieldwork and from contributions of anthropologists and other social scientists.

APPENDIX

Table 4—List of studies reviewed

Category	Area and Date of Data Collection, Sample Characteristics, and Crops	Method of Gender Disaggregation	Comments
Production-Function Studies			
Male-Female Differences in Technical Efficiency			
<u>Kenya</u>			
Moock 1976	Kenya, 1971-72. Farms in Vihiga Division that received loans for the purchase of hybrid maize seeds and fertilizers and comparison farms that were not loan recipients. 152 farmers, 101 males and 51 females. Maize.	Separate farms of men and women; regressions performed on male and female farms separately and for whole sample. Female dummy interacted with other variables used in pooled regression.	Cobb-Douglas production function; crop yield.
Bindlish and Evenson 1993	Seven districts in six provinces, 1989-90 long rains season; resurvey of 1981-82 rural household survey for baseline information; 675 farmers, 434 male household heads, 241 female household heads. Maize and other crops.	Regression performed on pooled male and female farms with female head dummy.	Value of total output per farm.
Saito, Mekonnen, and Spurling 1994	1989-90, 720 households in three districts. Maize and other crops.	Regressions performed on plot-level and household-level data with male farmer dummy for the pooled regression; regressions also performed separately for male-female plots and male-female farmers.	Cobb-Douglas production function; total value of output.
<u>Burkina Faso</u>			
Bindlish, Evenson and Gbetibouo 1993	401 villages from 12 CRPAS; panel data 1984-1991, 2,406 households with production data; 17,000 plots; 650 male managers and 519 female managers. Maize, sorghum, millet, cotton, rice, groundnuts.	Regressions performed on farm-level data for male and female farms, with female head dummy.	Value of total output per farm.
Udry 1994	Burkina Faso ICRISAT survey; 1981-85 panel study of 150 households, 6 villages, 3 agro-climatic zones. Study uses data from 1981-83; 4,655 cultivated plots	Yield regressions for plots controlled by men and women, with gender dummy	Not a production function study, but yield regressions for plots controlled by men and women for the same crop in the same year.

(continued)

Table 4 (continued)

Category	Area and Date of Data Collection, Sample Characteristics, and Crops	Method of Gender Disaggregation	Comments
Production-Function Studies			
Male-Female Differences in Technical Efficiency (continued)			
<u>Nigeria</u>			
Saito, Mekonnen, and Spurling 1994	1989-80, Oyo State, 250 randomly selected households, 1,174 plots.	Regressions performed on household-level data for all farmers, with male head dummy; regression also on plot-level data for all plots, with male farmer dummy and male-female plots separately.	Cobb-Douglas production function; total value of crop production per household and per plot.
<u>Korea</u>			
Jamison and Lau 1982	Korea, 1973, subsamples of a national survey of 2,254 farmers in nine regions of South Korea. 1,904 farms of which 1,363 used mechanical power and 541 did not use mechanical power. Rice and other crops.	Regressions performed on farm-level data with a dummy for male household head.	Cobb-Douglas production function; total value of production.
<u>Thailand</u>			
Jamison and Lau 1982	Thailand, 1972-73. Reanalysis of a stratified random sample of farm households from 22 villages on the Chiang Mai Valley. Rice. Selected farms: 275 farms, 91 chemical, 184 nonchemical.	Regressions on farm-level data with dummy for male household head.	Cobb-Douglas production function; total value of production.

(continued)

Table 4 (continued)

Category	Area and Date of Data Collection, Sample Characteristics, and Crops	Method of Gender Disaggregation	Comments
Production Function Studies			
Male-Female Differences in Labor Productivity			
<u>The Gambia</u>			
von Braun, Puetz, and Webb, 1989	1985-86, 10 villages in Jahally-Pacharr. Smallholder Rice project, 186 household (compounds), upland cereals (millet, sorghum, maize), groundnuts, rice.	Production function for major field crops had two gender-related variables; number of women in household of working age per adult equivalent, and share of fields of the crop under women's control.	Determinants of land productivity in agriculture.
<u>Rwanda</u>			
von Braun, de Haen, and Blanken, 1991	1985-86, in high altitude zone of Northwestern Rwanda, stratified random sample of 192 households. Maize, sorghum, and other crops.	Dependent variable is net returns per day of family labor available for agricultural and home goods production; gender variable was the share of women of working age in total number of persons of working age in the household.	Determinants of land productivity in agriculture.
<u>Burkina Faso</u>			
Ram and Singh 1988	1980, 105 families from seven villages of three selected areas of Mossa plateau.	Labor input (hours worked) disaggregated for males and females.	Earnings function estimated with net farm income as dependent variable, and land area, schooling of head (or sum of years of schooling of household members), male and female hours worked, value of farming capital, and dummy for tractor use as regressors.
<u>Kenya</u>			
Saito, Mekonnen, and Spurling 1994	1989-90, 720 households in three districts of Kenya. Maize and other crops.	Productions included dummy for male farmer in pooled regression, also estimated separately for males and females.	Cobb-Douglas production function, total value of crop production per household and per plot.

(continued)

Table 4 (continued)

Category	Area and Date of Data Collection, Sample Characteristics, and Crops	Method of Gender Disaggregation	Comments
Production Function Studies			
Male-Female Differences in Labor Productivity (continued)			
<u>Nigeria</u>			
Saito, Mekonnen, and Spurling 1994	1989-90, Oyo State, 250 randomly selected households, 1,174 plots. Maize and other crops.	Labor inputs were disaggregated into male and female labor input, for both family and hired labor.	Cobb-Douglas production function, total value of crop production per household and per plot.
<u>Nepal</u>			
Jamison and Moock 1984	Nepal, 1977-78, 792 households in 12 panchayats in two districts in the terai region. Rice and wheat.	Production functions included the proportion of female labor.	Cobb-Douglas production function.
<u>India</u>			
Laufer 1985	1975-78, ICRISAT survey. Six villages in South Central arid India. Sorghum, legumes, rice.	Labor inputs disaggregated into labor of men, women and children (hours).	Generalized quadratic production function with farm-level fixed effects.
<u>Peru</u>			
Jacoby 1992	1985-86, Sierra region highlands of Peru, Peruvian Living Standards Survey, 1,549 agricultural households chosen from original 5,000. Outputs of crops and livestock aggregated using prices.	Labor input disaggregated into adult male labor, adult female labor, teenager labor, and child labor.	Separate Cobb-Douglas production functions for crops and livestock; translog production function with sequentially restricted parameters, translog with sample selection correction for male and female's probability of working in the farm.
<u>Côte d'Ivoire</u>			
Jacoby 1993	1985-88 Côte d'Ivoire Living Standards Survey, from which 1,229 households were included for this study	Labor input disaggregated into male and female labor in hours per year.	Agricultural profit functions estimated as a function of land area, value of capital goods, male and female labor, and interactions of male and female labor with crop composition dummies

(continued)

Table 4 (continued)

Category	Area and Date of Data Collection, Sample Characteristics, and Crops	Method of Gender Disaggregation	Comments
Wage and Earnings Functions			
<u>India</u>			
Rosenzweig 1980	1970-71, national sample survey of rural households conducted by National Council of Applied Economic Research, heads of households and wives from landless and landholding households.	See-specific labor supply functions, with predicted wages.	Male and female market labor supply functions estimated for landless and landholding households using ordinary least squares-instrumental variables and tobit.
Mukhopadhyay 1991	1989; 1930 farm households and 11,575 persons in the state of West Bengal, India.	Sex-specific labor supply functions, with predicted wages.	Male and female labor supply functions estimated with predicted wages; wage functions estimated separately for men and women with selectivity corrections.
<u>Sri Lanka</u>			
Sahn and Alderman 1988	1980-81 Labor Force and Socioeconomic Survey; nationally and sectorally representative; rural sample with 5,450 males and 5,314 females.	Sex-specific wage equations, with selectivity correction	Wage functions estimated separately for men and women with selectivity corrections.
<u>The Philippines</u>			
Berhman and Lanzona 1989	1985-86, five rice-growing villages in two Philippine provinces, time use data for 473 men, 390 women, and 455 children.	Male and female dummies, schooling of adult males and females as separate regressors.	Semilog wage equations in the dry and wet seasons were estimated with selectivity corrections for price-rate and fixed wage contracts.
<u>Peru</u>			
Khandker 1990	1985-86, Peru Living Standards Survey, 51,000 households in Lima, other urban areas, and rural areas.	Participation and wage functions estimated separately for male and female workers ages 14 to 60.	Ordinary least squares, maximum likelihood, and fixed-effects methods used to estimate wage equations with selectivity corrections for men and women in Lima, other urban areas, and rural areas. Only the results for rural areas are reported here.

(continued)

Table 4 (continued)

Category	Area and Date of Data Collection, Sample Characteristics, and Crops	Method of Gender Disaggregation	Comments
Technology and Crop Adoption			
<u>Kenya</u> Appleton et al. 1991	1982, 441 households in Central and Nyanza Provinces, Kenya; Coffee and other crops, livestock.	Logit regression with a dummy for female household head, adjusted for absent husbands who may be decisionmakers. Percentage of growers currently adopting, and early adopters were disaggregated by gender to test for gender-specific copying effects.	Logit equation for early and late adopters, taking into account sequential and simultaneous copying effects, were estimated. Gender-specific copying effects were also included.
<u>Tanzania</u> Appleton et al. 1991	1983, 498 households in Dodoma, Iringa, Kilimanjaro and Ruvuma provinces, Tanzania.	Logit regressions for livestock adoption included a dummy for female household head.	Logit for livestock adoption estimated with simultaneous copying variable (not differentiated by gender).
<u>Kenya</u> Bindlish and Evenson 1993	Seven districts in six provinces, 1989-90 long rains season; resurvey of 1981/82 rural household survey for baseline information. 675 farmers, 434 male household heads, 241 female household lands. Maize and other crops.	Probit regression with dummy for female household head.	Multinomial probit estimates of determinants of sublocation and staffing advice; binary probit estimates of the probabilities of awareness and adoption of six technology fields. Previous advice, previous awareness, and previous adoption are captured by cumulative logistic terms defined over the cluster.
<u>Zambia</u> Jha, Hojjati, and Vosti 1991	1985-86, 330 smallholders in 10 agricultural districts of Eastern Province, Zambia. Hybrid maize.	Probit regressions with female household head dummy.	Probit regressions for fertilizer use, animal traction vs. ox cultivation, hybrid maize use, and use of cotton, soybean, and sunflower were estimated using instrumental variables. Predicted probabilities of usage were used in second stage estimates. Selectivity correction from the fertilizer use equation was included in regression of plant nutrients per fertilized hectare and percent area fertilized.

(continued)

Table 4 (continued)

Category	Area and Date of Data Collection, Sample Characteristics, and Crops	Method of Gender Disaggregation	Comments
Technology and Crop Adoption (continued)			
<u>India</u> Mukhopadhyay 1991	1989, 1930 farm households and 11,575 persons in the state of West Bengal, India. Rice.	Market wages of men and women, and average schooling of men and women included in equation for adoption of high yielding variety technology.	Proportion of rice area under high yielding variety estimated using maximum likelihood tobit with predicted values of male and female wages, estimated ratio of modern variety to traditional variety yields, and estimated ratio of yield variances.
<u>The Philippines</u> Behrman and Lanzona 1989	1985-86, five rice-growing villages in two Philippine provinces; 125 households in the wet season, 111 households in the dry season, with time use data for 473 men, 390 women, and 455 children.	Predicted wages, adult schooling, and age of men and women were included as determinants of modern technology use.	Reduced form regression of percentage area using modern technology (direct seeding, modern varieties, tractors), as a function of prices, predicted wages of men, women, and children, men's and women's schooling, assets, age of men, women, and children, a dummy for peak season, and interaction between land tenure, irrigation, area, and rice and fertilizer prices.
<u>Guatemala</u> von Braun, Hotchkiss, and Immink 1989	1983 and 1985, 400 small-farm families in the Western Highlands of Guatemala. Export vegetables (snow peas, broccoli, cauliflower, parsley), maize, traditional vegetables.	Probit estimates of determinants of export crop adoption included the share of female labor in total labor of the household.	Probit estimate of export crop adoption as a function of farm size, land quality, off-farm income total labor and share of female labor, household head's age, education, and attitude towards maize production, and village dummies.

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