Gender and Education Impacts on Employment and Earnings in West Africa: Evidence from Guinea*

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I. Introduction

Labor markets in developing countries are likely to be heterogeneous in the sense of having segments that differ in terms of the factors affecting entry and earnings. One potentially important distinction is between self-employment and wage employment; another is between the private sector and the public sector. In this study we use survey data from Conakry, the capital of Guinea, to investigate the structure of the urban labor market in a West African setting. We address three questions which, largely due to a lack of comprehensive household survey data compared with other regions, have not been adequately investigated in the African context. First, with regard to overall labor market structure, do labor market segments differ with respect to the determinants of entry and earnings? Second, do investments in schooling yield equivalent returns in different sectors of the labor market? Third, do women enjoy the same access as men to employment in all sectors of the labor market, and do they receive earnings comparable to men with similar backgrounds?

We investigate these issues by modeling employment outcomes and earnings of men and women in three labor market alternatives (in addition to nonparticipation): self-employment, private sector wage employment, and public sector wage employment. The few prior studies of labor markets in Africa using micro data generally have not considered the full range of alternatives and, hence, provide only a partial picture of the labor market. S. Appleton et al. (for Côte d'Ivoire) and P. Glewwe (for Ghana), for example, consider the distinction between private and public wage employment but not the choice of self-employment versus wage employment.[1] The self-employed, however, make up a large part of the African urban economy. W. Vijverberg models the choice of, and earnings in, nonagricultural self-employment as an alternative to nonparticipation and wage employment in a sophisticated analysis for Côte

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d'Ivoire that also considers migration choices.[2] His study does not, however, distinguish between public and private wage employment.

The questions addressed by this study are important from a policy perspective. This is the case particularly for the many countries in Africa, including Guinea, that are implementing policy reforms aimed at economic liberalization and expansion of the private sector. A prominent aspect of reform in Guinea and several other countries has been a reduction in the size of the civil service.[3] The consequences of such measures both for economic growth and for equity will depend in large part on how the labor market functions—in particular, on the access of men and women to wage and self-employment in the private sector and on the returns to schooling and experience in private sector employment. Improving economic opportunities for women in Africa is also an important objective. Little information currently exists, however, on gender differences in employment determinants and earnings and on how policy can effectively influence labor market outcomes for women. Finally, this study is one of the few for Africa that examines the determinants of earnings of the self-employed. Since this is a sector in which both women and the poor are likely to be found, policies targeting the self-employed are potentially effective means of raising incomes of particularly vulnerable groups.

The remainder of this article is organized as follows. In section II the empirical methodology is discussed. Section III describes the data. Sections IV and V present the results of the analyses of sector of employment allocation and earnings. The article concludes in section VI with a discussion of the findings, emphasizing their implications for policy.

II. Empirical Methodology

In the first part of this study we estimate a model of labor market participation and sector of employment determination. A plausible view of the labor market in Conakry would allow for three distinct employment alternatives: self-employment, wage employment in the private sector, and wage employment in the public sector. We begin with this general specification and apply statistical tests to see if these divisions are warranted. In view of the number of discrete choices involved, the multinomial logit model provides the most suitable econometric approach to estimating the probabilities that an individual will be found in each sector. The model is derived from the assumption of utility maximization through the theory of probabilistic choice as developed by D. McFadden.[4] Briefly outlined, utility conditional on the choice of each labor market alternative j (including nonparticipation) is specified in linear form:

$$V_{ij} = B_j X_i + u_{ij}, \tag{1}$$

where V_{ij} is the indirect utility function of individual i in labor market sector j and is a linear function of education, age, income, and other factors (X_i); β_j is a vector of parameters indexed on sector; and u_{ij} is the stochastic component of utility capturing unmeasured determinants of choice. The individual is assumed to choose the sector k ($k = 1, 2, 3, 4$) for which V_{ij} is highest. Thus the probability that sector j is chosen by individual i, using equation (1), is

$$\begin{aligned} P_{ij} &= P_r(V_j > V_k) \text{ for all } j \neq k \\ &= P_r(B_j X_i + u_{ij} > B_k X_i + u_{ik}) \\ &= P_r(B_j X_i - B_k X_i > u_{ik} - u_{ij}). \end{aligned} \tag{2}$$

Assuming the u_{ij}'s are distributed independently and identically Gumbel, their differences ($u_{ik} - u_{ij}$) have a logistic distribution, and the probabilities take the multinomial logit form, which is straightforward to estimate. Since the formulation of the choice problem in (2) implies that only differences in utilities derived from different sectors matter for the determination of choice probabilities, some method of normalization is required. We normalize by setting the parameter vector associated with nonparticipation equal to zero. The estimated coefficients β_j are then interpreted as the effects of a variable on the utility of being in employment alternative j relative to the utility from the base category of nonparticipation.[5]

Next we estimate earnings functions for men and women in each sector of employment. Since the determination of participation in a given sector is unlikely to be random, the familiar problem arises of sample selection bias in the earnings equation estimates. This occurs when unmeasured characteristics of an individual influence both the wage and the sector selection process. If these characteristics are correlated with the right-hand-side variables in the earnings function, the estimated coefficients on these variables will be biased. To control for potential selection bias, we use the two-stage method developed by L. F. Lee. Selectivity terms (inverse Mill's ratios) are calculated from the estimated probabilities of employment for each worker in a sector and included in the earnings function, which takes the following Mincerian form:

$$\ln W_{ij} = \gamma_j Z_i + \tau_j \lambda_{ij} + \epsilon_{ij}, \tag{3}$$

where $\ln W_{ij}$ is the natural log of wages or earnings of individual i in sector j; Z_i is a vector of explanatory variables containing human capital measures such as level of education and experience; λ_{ij} is the selectivity term; and ϵ_{ij} is a disturbance term with an expected value of zero. Ordinary least squares estimation of (3) will yield consistent parameter estimates of γ_j. In addition, a significant estimate of τ_j, the coefficient on the selectivity term, indicates the presence of selectivity.[6]

A further issue in the estimation of earnings functions that contain selectivity terms, such as this one, is identification. It is preferable that the data contain one or more variables that affect sector allocation decisions but do not directly affect earnings. In our data set, nonlabor income and demographic characteristics of the household serve as identifying variables for the λ_{ij}.

III. Data

The data used in this study are taken from a survey of 1,725 households, conducted in Conakry, Guinea, in 1990. The survey contains detailed information on respondents' labor market activities and earnings as well as a wide range of other data on the structure of the household, education, and other social and economic characteristics. For the self-employed, data on enterprise revenues and costs over the week and month preceding the survey were collected, making it possible to calculate enterprise profits or losses. After deleting observations with missing data, the sample consisted of 3,566 men and 3,306 women 15 years of age and older. Means and standard deviations of the explanatory variables for men and women are given in table 1.

Labor force participation, defined to include individuals currently working as well as those searching for work, is much higher for men than for women in Conakry. For women 15–65 years old the participation rate is only 40%. For men between 30 and 50 years of age participation is almost universal, but it is very low for men younger than 30. For those 21–30 years old, it is just 40%, although many ''nonparticipating'' men in this age group are in training as unpaid apprentices. Job opportunities appear to be limited in Conakry, reflecting the stagnation of Guinea's economy prior to the initiation of economic and institutional reforms in 1984 and slow per capita growth combined with a rapid rise in Conakry's population in the years since then.

Table 2 shows, for men and women, the portion of the labor force found in each sector of the labor market as well as mean years of schooling and occupational experience by sector. With regard to sector placement, there are major differences between men and women.[7] Although men in the labor force are dispersed relatively equally across the three sectors of private wage, public wage, and self-employment, women are predominantly engaged in self-employment. Women who are self-employed, moreover, tend to be from households in the lower deciles of the per capita household income distribution. Only slightly more than one-fifth of employed women are wage employees, the majority in the public sector. With regard to education, mean years of schooling vary widely among sectors, pointing to large differences in skill requirements in different parts of the labor market. In the public sector, both men and women tend to be well educated, averaging 10.2 and 11.9 years of schooling, respectively. Among the self-employed, in contrast, average

TABLE 1

MEANS AND STANDARD DEVIATIONS OF VARIABLES USED IN THE ANALYSIS

Variables	Men: Mean	Men: Standard Deviation	Women: Mean	Women: Standard Deviation
Age (%):				
15–20	.234	.423	.256	.437
21–30	.300	.458	.326	.469
31–40	.214	.410	.221	.415
41–50	.124	.330	.106	.308
51–65	.100	.301	.067	.251
Older than 65	.028	.164	.023	.151
Education level completed (%):				
None	.436	.496	.663	.477
Primary school	.348	.477	.243	.429
Secondary school	.141	.348	.073	.261
University	.076	.263	.021	.142
Married (%)	.442	.497	.644	.479
Number in household:				
Own children <2	.191	.436	.229	.431
Other children <2	.332	.636	.359	.628
Own children 2–5	.333	.694	.380	.625
Other children 2–5	.729	1.056	.857	1.113
Children 6–14	2.123	1.910	2.402	2.000
Males and females 15–20	1.676	1.659	1.666	1.606
Men 21–64	2.339	1.687	1.965	1.536
Women 21–64	1.826	1.391	2.242	1.518
Men >64	.123	.343	.112	.325
Women >64	.086	.305	.114	.350
Unearned income (Guinean francs)	$.194 \times e^{+6}$	$.546 \times e^{+6}$	$.182 \times e^{+6}$	$.547 \times e^{+6}$
Migrated since 1985 (%)	.183	.387	.194	.395
Ethnic group (%):				
Soussou	.461	.499	.493	.500
Fulani	.264	.441	.246	.431
Malinke	.204	.403	.202	.402
Other	.070	.255	.059	.236
Household receives electricity (%)	.881	.324	.882	.322
Resides near city center (%)	.124	.329	.136	.343
Interviewed in April–September (%)	.426	.495	.444	.497
No. of observations	3,566		3,306	

levels of education are very low—just 3 years for men and 2 years for women. Among private sector wage employees, the educational attainment of men and women diverges sharply, with men averaging 5 years and women averaging more than 9 years of schooling.

These sectoral differences in education level are reflected in the occupational distributions of men and women in each sector. In public em-

TABLE 2

Mean Levels of Schooling, Occupational Experience, and Capital Expenditures by Sector of Employment and Sex

Sector	Men	Women	All
Self-employment:			
Number	547	759	1,306
Schooling (years)	2.9	1.8	2.2
Experience (years)	10.7	7.2	8.7
Capital expenditures (10,000 Guinean francs/year)	1.39	.108	.641
Private wage employment:			
Number	544	70	614
Schooling (years)	5.1	9.3	5.5
Experience (years)	8.5	5.0	8.1
Public wage employment:			
Number	472	141	613
Schooling (years)	10.2	11.9	10.6
Experience (years)	12.5	9.3	11.8

Note.—Capital expenditures for self-employed only.

ployment, where the workforce is well educated, a large share of the jobs (47% for men and 59% for women) are in professional or managerial occupations. Among workers in the private wage sector, where schooling tends to be higher for women, slightly over half of the women are in professional, managerial, or clerical jobs compared with just 16% of men. Unlike men, very few women in the private wage sector are found either in unskilled work or in skilled trades. In self-employment, women are almost exclusively found in retail activities (e.g., food commerce), compared with only half of the men. Measured in terms of revenues, profits, or enterprise capital expenditures, female-operated enterprises tend to be much smaller than those owned by men.

These descriptive results suggest that the sectoral distinctions used in this analysis correspond to real differences in skill levels and occupation. There are also clear indications of heterogeneity by gender in the labor market. We next explore these issues further using the multivariate methodology described above.

IV. Participation and Sector of Employment

Parameter estimates from multinomial logit models of sector of employment for men and women are presented in table 3. Since the estimates themselves do not indicate the effect of a change in an independent variable on the probability of entering a sector, we have calculated these effects from the parameters and the data and report them beneath the parameter estimates. To determine if the sectoral decomposition of the labor market underlying the multinomial logit model is justified, we constructed Wald tests of the equality of the slope parameter vectors associated with each pair of sector choices (e.g., male self-employment and private wage employment). For both men and women, the null hypothesis was easily rejected at the .001 level of significance in every case except women for private wage and public wage employment, where equality was rejected at the .05 level. Equality of schooling effects specifically, represented in the models by dummy variables for completed primary, secondary, and university education, was also rejected for each pair of sectors for both men and women. The tests thus confirm that the urban labor market is heterogeneous in the sense that the determinants of entry into different segments of the labor market are not the same.

The estimated impacts of schooling on sector allocation shown in table 3 are in line with the descriptive statistics presented above. For both men and women, more education reduces the likelihood of being self-employed while it strongly increases the likelihood of being in the public sector; education clearly is the key to civil service employment. For private wage employment, the effects of schooling differ for men and women. More education reduces the probability that a man will enter the private wage sector, although the effects are smaller in absolute value than for self-employment. For women, on the other hand, being better educated strongly increases the probability of being a private wage employee. As noted in the previous section, the types of occupations open to, or chosen by, women in this sector are in general quite different than those for men, for whom an education is clearly not a requirement of entry.

Previous studies in developing countries, including some using African data, also find that education is important in determining in which portion of the labor market an individual works. Consistent for the most part with our results, Vijverberg found that for men and women in Côte d'Ivoire the probability of being a wage employee rises with education level while the probability of nonagricultural self-employment falls with additional schooling. Glewwe found that in Ghana schooling is positively associated with entry into wage employment and, among wage employees, that those with a better education are more likely to be in the public than in the private sector.[8]

Age is represented in the models by dummy variables for different age categories.[9] Sector participation probabilities of men and women generally follow an inverted U-shaped profile with age. Consistent with the descriptive statistics presented above, adults older than 30 are more likely to be employed in any activity than are those younger than 30. Marriage is also generally strongly and positively related to sector entry for both men and women. The impact for women is noteworthy because studies of female labor force participation typically find that married women are less likely to work than unmarried women. Presumably, this reflects a higher reservation wage resulting from access to their spouses' incomes. Our results, in contrast, are consistent with nonpooling of

TABLE 3

MAXIMUM LIKELIHOOD ESTIMATES OF MULTINOMIAL LOGIT SECTORAL CHOICE MODELS (Changes in Probabilities in Brackets)

	Men			Women		
Variables	Self-Employment (1)	Private Wage (2)	Public Wage (3)	Self-Employment (4)	Private Wage (5)	Public Wage (6)
Intercept	−1.881	−1.601	−4.386	−2.642	−5.982	−8.979
Completed education (relative to none):						
Primary	−.802***	−.261*	.508***	−.568***	1.764***	2.775***
	[−.0991]	[−.0115]	[.0635]	[−.1145]	[.0317]	[.0751]
Secondary	−1.221***	−.218	1.566***	−.801***	2.399***	3.925***
	[−.1585]	[−.0406]	[.1875]	[−.1663]	[.0534]	[.1764]
University	−2.327***	−.704***	1.689***	−1.297***	3.699***	5.517***
	[−.1976]	[−.0822]	[.2410]	[−.2324]	[.1193]	[.3622]
Age (relative to < 21):[a]						
21–30	. . .	. . .	. . .	.899***	1.559***	2.694***
				[.1118]	[.0181]	[.0195]
31–40	1.116***	1.424***	2.455***	1.276***	2.035***	4.517***
	[.0346]	[.0927]	[.1374]	[.1627]	[.0223]	[.0800]
41–50	1.136***	1.333***	3.047***	1.582***	2.598***	4.924***
	[.0216]	[.0531]	[.2236]	[.2102]	[.0358]	[.0964]
51–65	.0003	.535***	1.677***	1.155***	1.447*	4.058***
	[−.0366]	[.0363]	[.1036]	[.1483]	[.0117]	[.0603]
Over 65	−.706*	−1.009***	−1.459	−.173	1.325	−6.506
	[−.0410]	[−.0615]	[−.0286]	[−.0170]	[.0177]	[−.0021]
Married	1.938***	1.665***	1.709***	.818***	−.242	.571*
	[.1426]	[.1090]	[.0684]	[.1147]	[−.0098]	[.0124]
Own children <2	.254*	.244*	.246	.047	−.043	−.166
	[.0125]	[.0124]	[.0073]	[.0083]	[−.0005]	[−.0048]
Other children <2	−.017	.080	.073	.179**	.211	−.094
	[−.0063]	[.0078]	[.0040]	[.0269]	[.0038]	[−.0044]
Own children ages 2–5	.266***	.012**	.227**	.206***	.072	−.005
	[.0226]	[−.0142]	[.0117]	[.0308]	[.0007]	[−.0016]
Other children ages 2–5	−.056	−.163**	−.176*	.015	−.132	−.113
	[.0038]	[−.0121]	[−.0083]	[.0034]	[−.0023]	[−.0028]
Children ages 6–14	−.080*	−.196***	−.064	.055*	−.048	.130*
	[.0000]	[−.0179]	[.0019]	[.0076]	[−.0015]	[.0034]

Men and women ages 15–20	−.016	−.166***	−.025	−.062*	−.015	.144*
	[.0048]	[−.0178]	[.0028]	[−.0102]	[−.0005]	[.0044]
Men ages 21–64	−.073	.006	−.094*	.051	−.128	−.224**
	[−.0058]	[.0058]	[−.0059]	[.0095]	[−.0020]	[−.0062]
Women ages 21–64	−.105*	−.073	−.101	−.169***	.012	−.197*
	[−.0062]	[−.0019]	[−.0037]	[−.0243]	[.0014]	[−.0045]
Men >64	−.177	.346*	−.037	.049	−.058	−.961*
	[−.0303]	[.0471]	[−.0082]	[.0135]	[.0016]	[−.0269]
Women >64	−.034	.308	.065	.105	.413	.255
	[−.0160]	[.0349]	[−.0023]	[.0129]	[.0070]	[.0052]
Unearned income	$-.275 \times e^{-6}$	$-.221 \times e^{-6}$	$-.814 \times e^{-7}$	$-.378 \times e^{-6}$**	$.307 \times e^{-7}$	$-.569 \times e^{-7}$
	$[-.217 \times e^{-7}]$	$[-.153 \times e^{-8}]$	$[.113 \times e^{-7}]$	$[-.568 \times e^{-7}]$	$[.202 \times e^{-8}]$	$[-.659 \times e^{-9}]$
Migrated since 1985	.304*	.290*	−.414**	−.236*	−.342	−.748**
	[.0303]	[.0335]	[−.0436]	[−.0301]	[−.0037]	[−.0164]
Ethnic group (relative to Soussou):						
Fulani	.650***	.075	.125	−.348***	−.210	−.188
	[.0646]	[−.0185]	[−.0065]	[−.0511]	[−.0024]	[−.0024]
Malinke	.445***	.082	.345**	−.430***	−.210	−.330
	[.0348]	[−.0155]	[.0160]	[−.0618]	[−.0018]	[−.0058]
Other	.499***	.057	.396	−.204	−.463	.262
	[.0404]	[−.0217]	[.0197]	[−.0323]	[−.0081]	[.0109]
Household receives electricity	−.058	−.055	.404*	.247*	.333	.684
	[−.0126]	[−.0146]	[.0326]	[.0318]	[.0036]	[.0146]
April–September	−.060	.048	−.019	.460***	.224	.086
	[−.0076]	[.0082]	[−.0014]	[.0691]	[.0026]	[−.0012]
Resides near city center	−.007	−.507	−.008	.362***	.004	.127
	[.0160]	[−.0518]	[.0117]	[.0564]	[−.0016]	[.0011]
Ln likelihood		−2,916			−2,090	
χ^2		2,485			1,106	
No. of observations	547	544	472	759	70	141

NOTE.—Base employment category is nonparticipation. For continuous explanatory variables, the partial derivative of the probability of being in a sector with respect to a change in the variable is shown. For dichotomous variables, the change in probabilities is shown. This is calculated as the difference in the probability of being in the sector when the variable equals 1 and when it is zero. The figures shown are the average effects for the male and female samples.

[a] Excluded group for men is under 31.

* Significant at the 1% level.

** Significant at the 5% level.

*** Significant at the 10% level.

incomes within the household—a plausible situation in this African context.[10] Alternatively, however, married women may be more likely to work because they are able to secure capital through their husbands to start up small enterprises or use their husband's connections to obtain wage employment. With regard to the estimates for men, note that marital status may not be an exogenous determinant of male labor force behavior. Instead, men may get married once they have secured a livelihood.

A number of other demographic factors can be expected to influence labor force participation and sector allocation, particularly for women, who must balance domestic responsibilities with the need to augment family income. We focus here on the effects of children; discussion of the estimates for other demographic variables can be found in Glick and Sahn (see n. 6). Young children exert theoretically contradictory effects on a woman's reservation wage, hence on her participation. Children require care and supervision, but they also increase the need for market goods. In the present context, an additional (own) child younger than 2 years neither reduces nor increases the probability of a woman's participation in any sector, but having an additional child age 2–5 has a strongly significant positive impact on the probability of being self-employed. These estimates suggest that the care of older children, but not that of infants, may be combined with some self-employment activities. An additional other (not one's own) child age 2–5 has no effects, but other children younger than 2 raise the likelihood of participation of women in self-employment, presumably reflecting substitution in market work for mothers of infants who withdraw from the labor force. For men (who engage in very little child care), the association of young children of one's own and entry is positive for all sectors, but as with marriage, being a father of young children may be an outcome of having employment rather than a determinant.[11]

In accord with expectations from theory, the coefficients on exogenous (nonlabor) income, which is expected to raise the reservation wage, are generally negative, although significant only for women in self-employment. Having migrated to Conakry in the 5 years preceding the survey is positively associated with men's entry into self-employment and private wage employment but negatively associated with public sector employment. The latter is not unexpected, since new arrivals are less likely to have contacts that would provide access to scarce public sector jobs.

The results also shed light on a number of factors that have a particular influence on women's participation in self-employment (table 3, col. 4). Several of these reflect aspects of the urban infrastructure. Residing in a *sous-prefecture* located near the commercial center of the city encourages female self-employment activity, probably because of the lower costs of transportation to markets where retail activity is conducted.

Electricity in the household also encourages women to become self-employed, presumably by making production in the home or working during evenings easier.[12] Women are more likely to be engaged in self-employment activity from April to September, which is the rainy season. Although this period corresponds to a general reduction in commercial activity in Conakry, women may become more active in household enterprises during this time to substitute for men who return to their villages for agricultural work.

Ethnic group membership also has impacts. Fulani and Malinke women are less likely to be self-employed than are women in the excluded Soussou category. Tradition and social networks may explain in part the predominance of Soussous in self-employment. Also, women who belong to other ethnic groups (which are well-off relative to Soussous in terms of measures such as per capita household expenditures) may have the option of avoiding what might be considered low-status work.[13]

Finally, note that our models include only general education and not vocational schooling or apprenticeship training. In alternative specifications of the logit models (available from us on request), apprenticeship training was associated with entry into both self-employment and private wage employment for men and women and public employment for men, while formal vocational schooling was associated with entry into private and public wage work for men and women. However, these variables were not included in our preferred specification because it is likely that they are endogenous to sector choice, since many individuals no doubt choose such training in order to enter employment in a given field or a particular sector of the labor market.

Gender and Sector Entry Probabilities

The multinomial logit estimates imply that differences in sector allocations of men and women can be attributed in part to gender differences in background characteristics, especially schooling. For example, women, who have lower mean schooling than men (3.7 compared with 6.5 years), are underrepresented in the public sector, where jobs tend to require some education. In this section we control for differences in background to assess the impact of gender itself on entry into different sectors of the labor market. Table 4 shows the predicted probabilities of sector participation by sex and education level calculated from the male and female logit estimates in table 3 and setting the other explanatory variables to the mean values for the entire (male and female) sample. The table, therefore, shows the sector entry probabilities for men and women with similar characteristics.

The predicted overall probabilities of working (the sum of the sector probabilities) are substantially lower for women than for men with the same characteristics at all levels of schooling except university (just 2%

TABLE 4

Predicted Sector Entry Probabilities of Men and Women, by Level of Education

Sex/Level of Education Completed	Self-Employment	Private Wage	Public Wage
Men:			
None	.248	.180	.038
Primary	.131	.164	.074
Secondary	.078	.155	.193
University	.028	.104	.239
Women:			
None	.222	.006	.001
Primary	.132	.036	.024
Secondary	.099	.064	.072
University	.043	.165	.250

of all women). For self-employment specifically, entry probabilities are quite similar for men and women. The difference in overall participation probabilities, therefore, is due to women's much lower chances of entering wage employment. This applies to both public and private sectors. Men are far more likely than women of identical backgrounds to be in public wage employment at all levels of education except university. The probabilities for less than primary, completed primary, and secondary schooling are 4%, 7%, and 19%, respectively, for men compared with 0.1%, 2.4%, and 7.0%, respectively, for women. Lower educational attainment thus only partly explains the low presence of women in the public sector workforce.

Differences in private wage entry probabilities for men and women with similar characteristics are also large, especially at lower skill levels. Men who have not completed primary school and those who are primary school completers have 18.0% and 16.4% probabilities of entry, compared with just 0.6% and 3.6%, respectively, for women. Thus for women with little schooling (who lack the qualifications for public employment), own-enterprise activities are virtually the only source of labor income. We discuss the possible reasons for the gender gaps in public and private sector wage employment in the conclusion. For now we note that studies previously cited on Côte d'Ivoire and Ghana also found that women are less likely than men to be in wage employment, controlling for education and other factors.

V. The Determinants of Earnings

Estimates from earnings equations for men and women in self-employment, private wage employment, and public wage employment are presented in table 5. In the case of wage employees, the dependent variable is the natural log of the sum of monetary and in-kind compensation received in the last week divided by hours worked. For the self-employed, it is the natural log of hourly net enterprise earnings, calculated as weekly enterprise revenues minus costs divided by total hours worked. We estimated the self-employment earnings functions in two ways: on the sample of single-person enterprises (90% of all enterprises), and on the sample that includes multiperson enterprises, using the characteristics of the most educated family worker to represent these enterprises.[14] Since the results proved generally to be quite robust to the choice of sample, we present only the estimates for the full sample of enterprises. The wage and self-employment earnings equations were corrected for potential selectivity bias, employing the method described above, using the sectoral choice model estimates to calculate the selectivity factors.[15] With the exception of women in the private wage sector, discussed below, the coefficients on the lambda terms were insignificant, indicating an absence of sample selectivity.

Our first step was to use Wald tests to determine whether the overall earnings structures differ between sectors of the labor market, that is, to see if the disaggregation of earnings functions by sector is justified.[16] For men, the equality of earnings determinants (comparing all slope coefficients except the lambdas) is rejected at the 0.001 level for self-employment and private wage employment and for self-employment and public wage employment, and at 0.05 for private wage and public wage employment. For women, equality is rejected at the 5% level for self-employment and public employment and at 10% for self-employment and private wage employment, but cannot be rejected for private and public wage employment. With respect to the last case, note that the cell sizes are small (137 and 69 women, respectively, in public and private wage sectors), resulting in a test with low power. Overall, the determinants of earnings do vary by sector, providing further evidence of heterogeneity in Conakry's urban labor market.

We are also concerned specifically with differences by sector in the returns to schooling, represented in the earnings functions by dummy variables for completed primary, secondary, and university education (for self-employed women, secondary and university are lumped together). In no case for either men or women could the equality of the education coefficients as a group be rejected for any pair of sectors, and only in a single case (men's private and public wage employment) do the differences in any of the individual schooling level estimates come near to being significant at conventional levels. Thus the returns to schooling in different labor market sectors are statistically similar, although as we discuss below, the point estimates vary considerably between sectors in a number of cases.

We now examine the schooling estimates in more detail, starting

TABLE 5

SELECTIVITY CORRECTED LOG HOURLY EARNINGS EQUATIONS

	MEN			WOMEN		
VARIABLE	Self-Employment	Private Wage	Public Wage	Self-Employment	Private Wage[a]	Public Wage
Completed education (relative to none):						
Primary	.2737	.2494	.2878	.2518	.1190	.5947
	(1.55)	(3.11)***	(3.33)***	(1.93)*	(.46)	(2.15)**
Secondary[b]	.6865	.5737	.3329	.3292	.3319	.8184
	(2.36)**	(5.29)***	(2.81)***	(1.33)	(1.13)	(2.36)**
University	1.3159	.9019	.6211	. . .	.7791	1.2887
	(2.80)***	(6.13)***	(4.58)***		(2.17)**	(3.01)***
Occupational experience	.0182	.0142	.001	.0582	.0167	.0197
	(2.56)**	(3.59)***	(0.20)	(3.24)***	(.68)	(2.17)**
$(\text{Experience})^2$	. . .	. . .	. . .	−.0017	. . .	. . .
				(2.57)**		
Age (relative to <31):						
31–40	.2702	.0955	.0492	.3325	−.1987	.0828
	(1.60)	(.93)	(.31)	(3.02)***	(.84)	(.43)
41–50	.5843	.2579	.2346	.2890	.0606	.2036
	(3.08)***	(2.35)**	(1.24)	(2.11)**	(.17)	(.80)
51–65[c]	.3387	.1471	.0896	.1456	−.1668	.2554
	(1.55)	(1.20)	(.53)	(.93)	(.26)	(.71)
Older than 65	.8622	−.0679	.4846	. . .	. . .	. . .
	(2.12)**	(.30)	(.87)			

Business capital/10,000 Guinean francs:	.0714	. . .	. . .	.2054	. . .	. . .
	(8.90)***			(3.79)***		
April–June	−.1767	.0514	.0559	.0360	−.0247	.0161
	(1.19)	(.56)	(.80)	(.32)	(.09)	(.14)
July–September	−.4221	.1205	.0776	−.2364	.2276	.1286
	(2.82)***	(1.42)	(1.13)	(2.06)**	(.83)	(1.06)
October–December	−.8395	−.1820	.1601	−.3165	.1536	−.0116
	(5.65)***	(2.06)**	(2.38)**	(2.58)***	(.54)	(.09)
Ethnic group (relative to Soussou):						
Fulani	−.1457	−.0177	.0856	.0311	.1510	−.0246
	(.93)	(.23)	(1.30)	(.30)	(.56)	(.21)
Malinke	.2454	.0207	−.0420	.2312	.0441	−.2164
	(1.45)	(.24)	(.67)	(1.94)*	(.18)	(1.88)*
Other	.3620	.1857	−.0154	−.0373	.6925	.1231
	(1.51)	(1.42)	(.15)	(.20)	(1.45)	(.77)
λ	−.0879	.0527	−.0993	.1857	. . .	.1773
	(.43)	(.44)	(.87)	(1.15)		(.95)
Intercept	5.9949	5.2140	5.4453	4.8343	5.3134	4.4068
	(16.00)	(23.92)	(16.14)	(19.67)	(19.39)	(6.81)
Adjusted R^2	.306	.136	.184	.095	.023	.128
F	13.49	6.66	8.03	5.46	1.13	2.42
No. of observations	454	540	470	637	69	137

NOTE.—*t*-statistics are in parentheses.
[a] OLS estimates.
[b] Includes university for women in self-employment (one observation).
[c] Includes over 65 for women.
* Significant at the 1% level.
** Significant at the 5% level.
*** Significant at the 10% level.

with the results for men. In self-employment, the coefficients on schooling are positive but are only significant for secondary and university education. The point estimates for secondary and university education (which only a small portion of self-employed men possess) are 0.69 and 1.32, implying hourly earnings that are twice and almost three times higher, respectively, than for men who have not completed primary school.[17] It should be noted that the estimates may understate the total benefits of schooling in enterprises, since schooling also appears to have indirect effects on profits through its impact on the use of capital inputs.[18]

Among male wage employees, the return to primary schooling appears to be slightly higher in the public sector than in the private sector, but the situation is reversed at higher levels, with much lower returns to postprimary schooling in the public sector. The estimates imply benefits (relative to the base of no schooling or less than primary) from completed secondary and university education of 77% and 146% in the private wage sector, compared with only 50% and 105% in the public sector. Although we cannot reject equality of the private and public sector coefficients for secondary and university schooling ($p = .14$), the large differences in the point estimates are striking. They suggest a compression or narrowing of the civil service pay structure, a phenomenon reported for many African countries.[19] Among better skilled male wage employees, therefore, those in the public sector appear to be at a disadvantage.[20]

For women in self-employment, completed primary schooling raises hourly enterprise profits by around 30% relative to having less than a primary education. Although the coefficient on primary schooling is only significant at the 10% level ($t = 1.93$), restricting the sample to women in single-person enterprises (all but 5% of the total) yields a larger and more precise estimate ($t = 2.22$). Thus there are returns in women's small enterprises to basic skills acquired through a primary education. On the other hand, there appears to be little incremental benefit to secondary or greater schooling—the coefficient is not much larger than that for primary school and is not significant. The small number of women in this subsample with a secondary education or better may make it difficult to estimate the effects of this schooling. Still, the contrast with the large increase in self-employment profits accruing to men with a secondary education is noteworthy and suggests that postprimary schooling may not be very useful for the types of activities in which self-employed women are involved.[21]

For women in the public sector the returns to education at all levels are substantial: the rather flat wage-education relation observed for men is not found for women. The schooling coefficients for women are roughly double those for men in the public sector for each level of education, though the differences are not jointly significant ($p = .51$). Note, however, that the higher incremental schooling effects for women do not imply that absolute wage levels are higher for women than for men. As discussed below, predicted public sector wages of men actually exceed those of women for all levels of education.

The selectivity-corrected earnings regression for women in the private wage sector yielded several highly implausible parameter estimates, particularly for schooling, for which the coefficients were actually negative. These results apparently reflect collinearity of the schooling covariates with the selectivity term, which was highly significant. This in turn suggests that the equation is not properly identified.[22] We therefore present instead in table 5 the results of the model without the selectivity correction with the caveat that there is likely to be some bias through sample selection in these OLS estimates. All the education parameter estimates in this regression are positive, but they are not individually significant except for university education, and they are jointly only marginally significant ($p = .15$). The returns to schooling also appear low compared both with those for women in the public sector and for men in the private sector. In assessing the magnitudes and the significance levels of the estimates, however, it should be kept in mind that the sample consists of only 69 observations.

Turning to the other covariates in the earnings equations, returns to experience are captured by a continuous variable for years of occupational experience as well as a series of dummy variables for age level to capture general experience. The age dummies are also included to control for cohort effects on earnings, which may be important in view of the decline in the quality of the educational system during the later years of the Sekou Touré regime. Age is associated with higher earnings in private wage employment for men and in self-employment for both men and women. Occupational experience in self-employment raises earnings for both men and women, further underscoring the benefits of human capital in small enterprises. For men, the effect of experience is basically linear (with experience raising earnings by about 2% per year), while for women a quadratic specification gave the best fit. The benefits from experience for women begin at about 5% per year and peak at about 17 years. As with secondary schooling in self-employment, the difference in the effects of experience for men and women may reflect differences in the types (and perhaps the scale) of enterprises they operate. The nature of women's small-scale retailing activities may impose limits on the benefits of experience beyond a certain point.

In the public sector, occupational experience has a significant but small (about 2% per year) impact on women's wages. For men, the coefficient is close to zero and insignificant. This results in part from the inclusion of age and in part from the selection correction term. Even excluding these variables from the regression, the effect remains very small compared both with women in the public sector and men in self-employment. In the private wage sector, the estimated annual returns to experience for men and women are fairly similar (about 1.4% and 1.7%) but only significant for the former. Overall, the effects of occupational experience on wages in both private and public employment in Conakry are on the low side compared with other African estimates and indicate that employees at the high end of the seniority scales have suffered relatively more from inflation in recent years.[23]

Enterprise capital is represented in the self-employment earnings regressions by the reported value of capital purchases made for the enterprise over the past year. It was necessary to use this proxy for capital stock because the survey lacked a reliable measure of the actual capital stock in each enterprise. Annual capital expenditures will be an accurate representation of actual capital stock only under some fairly restrictive assumptions, a limitation that should be kept in mind when interpreting the estimates.[24] Businesses run by men employ far greater amounts of capital than those run by women (see table 1), reflecting differences in the scale and types of operations run by men and women and perhaps also constraints on women's access to credit. However, capital inputs have positive and strongly significant effects on enterprise earnings for both men and women (table 5). In fact, the magnitude of the coefficient for women is more than twice that for men, and the difference is statistically significant ($p = .014$). Also with regard to self-employment, enterprise profits for both men and women show marked seasonal variability compared with wages, probably reflecting the involvement of many small enterprises in marketing agricultural products.

The earnings regressions in table 5 include neither apprenticeship training nor formal vocational education in professional or technical schools, both of which are common in Guinea. We have not included these variables because they are more likely than regular schooling to be statistically endogenous to earnings. If individuals choosing such training tend to benefit more from it in terms of increased productivity and pay than those who do not choose it, the estimated impacts will be biased upward. Nevertheless, the effects of both apprenticeship training and formal vocational education are of significant policy interest, so dummy variables for having had these forms of training were included in alternative specifications of the earnings functions (reported in Glick and Sahn; see n. 6). The generally insignificant coefficients on these variables suggest that for the most part apprenticeship training and formal vocational education do not confer direct productivity benefits. The only exceptions are an improvement in men's private sector wages as a result of vocational schooling and a very large (more than 90%) increase in hourly profits accruing to women in self-employment who have had apprenticeship training. Recall, however, that both apprenticeship and vocational training backgrounds show strong associations with entry into specific sectors. This indicates that they may function primarily to permit access to specific sectors or professions, although some caution is in order in

TABLE 6

Male-Female Hourly Earnings Differentials by Sector

	Self-Employment	Private Wage	Public Wage
Mean ln male earnings	6.269	5.705	5.848
Mean ln female earnings	5.478	5.752	5.664
Overall ln earnings differential	.790	−.046	.184
Contribution of characteristics	.350	−.176	.042
Education	.036	−.182	−.037
Experience	.087	.048	.014
Age	.069	.016	.039
Capital	.098	...	...
Other variables	.060	−.058	.027
Residual	.440	.129	.141

making such an inference, since these variables are potentially endogenous to both sector choice and earnings. Among the self-employed, apprenticeship training for women is much rarer than for men but appears to offer women a path out of traditional commerce activities, which may be why it is associated with higher earnings. About half of the women with such a background are engaged in trades or manufacturing, compared with no more than 4% of all self-employed women.

Comparison of Male and Female Earnings

We turn next to comparisons of male and female earnings in each sector of the labor market. As shown in table 6, mean earnings in two of the three cases differ substantially between men and women. For self-employment, the gap is very large: the antilog of the difference in log mean earnings (equal to the ratio of male and female mean earnings) is 2.2, indicating that average hourly profits for men exceed those for women by 120%.[25] Among public sector employees, average male earnings exceed female earnings by about 20%. In private wage employment, in contrast, wages are actually slightly higher for women (reflecting the superior schooling of female private sector employees), but the difference is not significantly different from zero.

Do these gender differences in earnings reflect differences in productivity-enhancing characteristics such as schooling and experience, or are they caused by discrimination in pay (or in access to better-paying occupations)? Discrimination, if it occurs, may be reflected in differences in returns to men's and women's characteristics or it may enter additively, that is, independently of returns. For self-employment, a Wald test indicates that returns to characteristics jointly differ between men and women (the null hypothesis of equality of slope coefficients excluding lambda is rejected at the 0.05 level), although the equality of schooling coefficients specifically could not be rejected. For both public and private wage employment, the equality of the complete vectors of male and female slope coefficients, and of schooling coefficients alone, could not be rejected. This does not imply equality of wages for men and women with equivalent backgrounds; instead, there appears to be primarily an additive gender impact. In regressions for the private and public wage sectors pooling male and female workers, coefficients on dummy variables for being female were negatively signed, though significant only for the public sector.[26]

To assess the relative contributions to male-female earnings gaps of characteristics and discrimination (the gender effect), we use the decomposition procedure originally developed by R. Oaxaca. The difference in average male and female (log) wages is decomposed into two parts: one due to differences in the average levels of schooling, experience, and other covariates of men and women, and the other due to differences in pay for men and women with the same values of these characteristics. The second part arises from differences in the coefficients in the earnings equations (both intercept and slopes) for men and women and is usually interpreted as a measure of wage discrimination, though it may also reflect unobserved differences between men and women that affect earnings. To save space, we do not present the derivation of the decomposition here; the variant of the technique we adopt for this analysis is developed in D. Neumark and in R. Oaxaca and M. Ransom.[27]

Below the total earnings differentials in table 6 we show the portions due to differences in male and female characteristics and to gender (the residual). The former is disaggregated into contributions of specific factors. For self-employment, about 45% (i.e., .35/.79) of the difference in male and female mean log earnings is due to differences in characteristics. Men's greater occupational experience (10.7 vs. 7.2 years) and their larger amounts of capital together account for about half of the characteristics portion of the differential. Fifty-five percent of the total difference in male and female self-employment earnings reflects the residual gender effect. Obviously, for this sample the residual cannot be attributed to discrimination in pay by employers. Instead, self-employed women may be confined to activities (mostly in retailing, as described above) that are relatively less profitable. Male-female differences in self-employment occupations may themselves reflect discrimination against women, though not in the labor market per se. Women may be unable to secure credit to start up more lucrative businesses, and as discussed below, they are disadvantaged in access to apprenticeships that facilitate entry into skilled trades.

In the public sector, only about 25% of the male-female wage differential is accounted for by differences in male and female characteristics, leaving 75% unexplained by observed factors. Thus women working for the government are also paid less than are men with equivalent backgrounds, although the residual, like the overall gap, is modest compared with self-employment.[28] Note that gender differences in schooling serve to reduce somewhat the overall gap because women working in the public sector are slightly better educated than men. For the private wage sector, both the male-female overall earnings differential and the portion of the differential due to differences in characteristics are negative. This means that the differences in the mean characteristics of men and women—in particular, women's greater schooling—are responsible for the negative male-female wage gap. On the other hand, the residual is positive (and similar in size to that for the public sector). Thus women in the private wage sector also receive less compensation than men with the same characteristics, reducing the amount by which the mean wage of women exceeds that of men. The private wage decomposition results should be treated with caution, however, as most of the parameter estimates in the private wage regression for women underlying the decomposition, as well as the estimated impact of gender, were not statistically significant.

As with self-employment, the residual differences in male and female wages in the private and public sectors may reflect the different occupational structures of men and women—that is, there may be occupational discrimination rather than wage discrimination within occupations. Women in wage employment are highly concentrated in a few fields such as teaching, nursing, and secretarial work, which alone account for 41% of all female wage employees.[29] Although these are relatively skilled occupations, pay and advancement within them may be limited compared with professional and clerical occupations dominated by men. It is possible in principle to extend the earnings decomposition analysis to distinguish between occupational discrimination (unequal access to an occupation for women and men with the same qualifications) and pay discrimination.[30] However, initial attempts to incorporate occupation effects by including occupation dummies (e.g., professional and managerial, clerical services, etc.) in the wage equations and decompositions yielded few notable results. This is probably because the standard occupation classifications are too broad to capture several relevant male-female occupational differences in our sample. For example, almost half of the female professional and managerial employees in the private and public wage sectors are nurses or teachers, while men are dispersed more evenly among job types. However, inclusion of these finer occupational distinctions in the separate earnings functions for men and women underlying the decomposition analysis is generally not feasible, because few men hold ''female'' jobs (e.g., nursing) and few women hold other jobs that are dominated by men. Thus we are not able to directly incorporate these job distinctions in the analysis of pay gaps in private and public wage employment. Still, the very different occupational distributions of men and women point to occupational segregation as a factor behind the residuals in the male-female wage decompositions.

The few previous studies comparing male and female earnings in

Africa do not uniformly show a female disadvantage in pay after controlling for background factors, in contrast to more consistent evidence (though from even fewer studies) of gender differences in entry into different labor market sectors. The results of W. Milne and M. Neitzert for Kenya and Vijverberg for Côte d'Ivoire suggest, as the results of this study do, an earnings disadvantage to women in wage and self-employment. On the other hand, Glewwe does not find evidence of wage discrimination against women in Ghana and even finds a premium to being female in the public sector in some model specifications. Similarly, J. Armitage and R. Sabot, using data on the private and public wage sectors in Tanzania and Kenya, find evidence of a wage premium to men only in the public sector in Tanzania.[31] The lack of a consistent finding of bias against women in pay suggests that economic conditions, labor market policies, and sociocultural factors that affect wage determination differ among countries. However, the results may also differ because the methodological approaches in these studies vary widely, perhaps most importantly with respect to the presence or absence of controls for sample selectivity.

VI. Discussion and Implications

This study of Conakry, Guinea, represents one of the few efforts to date to analyze the determinants of labor force participation, sector of employment, and earnings in a sub-Saharan African economy. Estimates from multinomial logit models indicate that the determinants of entry into various sectors of the labor market are quite different, supporting the hypothesis of heterogeneity in the urban labor market. Education, in particular, plays an important role in allocating labor force participants among self-employment, private wage employment, and public employment alternatives. We found that the overall structure of earnings also differed significantly among sectors, although the equality of returns to schooling in different sectors could not be rejected. Education raises the earnings of men and women in all sectors of the labor market, confirming in a very low-income urban setting the benefits to investments in the schooling of boys and girls.

Our analysis of self-employment earnings is particularly noteworthy in view of the fact that the informal sector (consisting in large part of small enterprises like those in our sample) is expected to be the primary source of job growth in Africa for the foreseeable future. Earnings of self-employed men and women in Conakry increase with level of education, experience, and, for women, apprenticeship training, as well as with inputs of physical capital. Thus investments in schooling can be expected to contribute to future productivity growth in small informal enterprises as well as in the formal sector. For female enterprises, however, the benefits from education and experience appear to decline after a certain level, suggesting that there will be limits to growth in incomes for women in self-employment if they remain primarily involved in traditional commerce activities.

In addition to education, gender was found to play an important role in determining both labor market participation patterns and earnings in Conakry. Women are far less likely than men to be wage employees, which for the public sector at least partly reflects men's greater educational attainment. Controlling for differences in schooling and other characteristics, however, women are still much less likely than men to enter either public or private wage employment. With respect to earnings, the returns to schooling are substantial for both men and women and are similar statistically. In part because they have more schooling and experience, men tend to earn more than women do. Decompositions of the earnings differentials, however, indicate that in every sector of the labor market men earn more than women even when these factors are controlled, reflecting either gender discrimination in pay or differences in the occupations that men and women tend to enter. In sum, with respect both to entry into certain portions of the labor market and earnings, women appear to fare less well than men. In the remainder of this discussion, we focus attention on the reasons for these disparities and on policies that might reduce gender inequalities in economic opportunities and incomes.

The determinants of male and female labor market outcomes are to be found both in the operation of the labor market itself and outside the labor market, in parental investments in schooling for boys and girls. One labor market barrier women may face is discrimination in hiring. There is evidence from a number of West African countries that employers are less willing to hire women than men or are willing to hire them only in traditionally female occupations in which pay and prospects for advancement are poor.[32] In Guinea, female vocational education and training graduates with training in nontraditional (i.e., ''male'') occupations find it difficult to obtain jobs in these fields.[33] Employer preferences for men are likely to be grounded both in realistic appraisals of particular costs associated with hiring women (their higher absenteeism and disruptions or expenses related to maternity and child rearing) and in simple biases or stereotyping.[34] Policies to address these factors can take a number of forms, including government provision of child care services, strong enforcement of antidiscrimination laws, and financial incentives to employ women.[35]

Since the private sector in Guinea is expected to assume a more important role in the future, employment practices in this sector bear particular scrutiny. Our finding of very low private wage entry probabilities for women relative to men points to barriers to women's employment in the private wage sector, particularly at low skill levels. There is also evidence that women released from public sector employment during Guinea's civil service reform were less able than men to find new wage employment in the private sector.[36]

In addition to demand side factors that reduce women's access to wage employment, constraints or cultural attitudes that affect the supply side are undoubtedly also important. These include women's domestic responsibilities, their spouses' (or their own) traditional attitudes about work outside the home, and culturally conditioned low career aspirations. Each would serve to reduce female participation in the labor market or encourage entry into less formal self-employment activities. Direct information on the importance of these factors is lacking, however, and obtaining such knowledge would be a valuable objective for future research.

With regard to investments in schooling, raising levels of human capital is essential both for overall poverty alleviation in Guinea and to close the gender gap in economic opportunities and earnings. This study indicates that better schooling for girls will eventually increase female entry into formal or wage employment and raise women's incomes in both wage and self-employment. An ambitious education policy in the past several years has been successful in raising primary school enrollments in Guinea, reversing the declines that occurred during the 1980s. Primary school enrollment rates remain among the lowest in the world, however, and girls are especially disadvantaged. Girls make up only 32% of total primary school enrollments, and the gender disparity in enrollments increases sharply at higher levels.[37] Moreover, the quality of education received by girls who do go to school may suffer in comparison with that of boys because of problems such as poor attendance and an unfavorable classroom atmosphere.[38] In recognition of these problems, the government of Guinea has recently taken several steps toward improving girls' access to schooling.[39]

The results of this study also suggest that increasing female enrollment in vocational education would improve women's access to employment, particularly wage employment. Currently, only a small percentage of enrollees in postprimary vocational training centers in Guinea are girls. In addition, both the number of institutions and the types of programs or areas of study open to girls in Guinea are very limited compared with those for boys.[40] Thus the segregation of women into specific wage occupations reflects to an extent the organization of the training system. Expanding girls' access to training in traditionally nonfemale occupations should ease their entry into potentially more lucrative fields. However, given that female graduates with training in ''male'' occupations in Guinea encounter difficulty getting hired, such an intervention should not proceed without also addressing labor market factors such as employer attitudes.[41]

This problem points to a more general concern with policies to expand schooling for girls, which also fail to account for labor market conditions. Because of a lack of jobs open to women, improvements in women's education may result in an oversupply of educated female labor and high levels of unemployment. This is not purely a gender issue; unemployment rates are high for educated young men in Guinea (as throughout Africa) and will probably remain so for some time given the probability of weak growth of the formal sector. But women also face gender-related barriers to employment. Labor market opportunities may be especially limited for women with primary but not higher schooling, who may shun self-employment but lack the relatively good access to professional wage jobs enjoyed by the small elite of women with a secondary and university education. Indeed, overall female labor force participation in Guinea exhibits the same U shape with respect to education level observed in many other developing countries—it is relatively high for women with no schooling, low for primary schooling, and high for secondary and higher schooling.[42]

In addition, until barriers in the labor market to women's employment are reduced, it may prove difficult to convince parents to educate their daughters. Parents in Guinea view education largely as a means of gaining future support for the family.[43] Since women are less likely to be working than are men with equivalent schooling, parents may correctly perceive the returns to investing in girls' schooling to be low compared with boys'.[44] Changes in labor market practices to improve access for women thus are complementary to efforts to raise investments in girls' human capital.

Our results also indicate that certain policies oriented largely toward self-employment can improve women's opportunities and incomes. One of these is encouragement of apprenticeship training for girls. Only 8% of girls ages 15–20 in our sample are reported to be working as apprentices, one-third the percentage of boys. Such training, however, is associated with female entry into self-employment and the private wage sector as well as with higher earnings of women in self-employment; the latter occurs in part because women with backgrounds as apprentices enter male-dominated activities such as skilled trades. Also, in view of the high returns to capital found in female enterprises, the development of credit institutions that are responsive to the needs of self-employed women represents a potentially effective means of raising the incomes of poor women.

Notes

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1. S. Appleton, P. Collier, and P. Horsnell, ''Gender, Education, and Employment in Côte d'Ivoire,'' Social Dimensions of Adjustment Working Paper no. 8 (World Bank, Washington, D.C., 1990); P. Glewwe, ''Schooling, Skills,

and the Returns to Education: An Econometric Exploration using Data from Ghana,'' Living Standards Measurement Working Paper no. 76 (World Bank, Washington, D.C., 1990).

2. W. Vijverberg, ''Educational Investments and Returns for Women and Men in Côte d'Ivoire,'' *Journal of Human Resources* 28 (1993): 933–74.

3. Some 32,000 workers were released from the public payroll in Guinea during 1985–89. D. Sahn and B. Mills, ''Reducing the Size of the Public Sector Work Force: Institutional Constraints and Human Consequences in Guinea,'' *Journal of Development Studies* 31 (April 1995): 505–28.

4. D. McFadden, ''Conditional Logit Analysis of Qualitative Choice Behavior,'' in *Frontiers of Econometrics,* ed. P. Zarembka (New York: Academic Press, 1973).

5. The multinomial logit formulation has the limitation that the ratio of the probabilities of choosing two alternatives is independent of the number of alternatives (the ''independence from irrelevant alternatives'' property). We attempted instead to estimate the less restrictive nested multinomial logit model, using alternative nestings of choices (e.g., self-employment and nonparticipation vs. private and public wage employment). In each case, however, the term approximating the correlation of errors fell outside the unit interval, indicating misspecification (see G. S. Maddala, *Limited Dependent and Qualitative Variables in Econometrics: Econometric Society,* Monograph no. 3 [Cambridge: Cambridge University Press, 1983]). Therefore, we report only the estimates from the nonnested logit models below.

6. The selection correction term λ_{ij} equals $\phi(\Phi^{-1}[P_{ij}])/\ P_{ij}$, where $\phi(.)$ and $\Phi(.)$ are the standard normal density and cumulative distribution functions and P_{ij} is the predicted probability of observing individual i in sector j, given by the multinomial logit model. For more details, see P. Glick and D. Sahn, ''Labor Force Participation, Sectoral Choice, and Earnings in Conakry, Guinea,'' Cornell Food and Nutrition Policy Program Working Paper no. 43 (Cornell University, Ithaca, N.Y., 1993); and L. F. Lee, ''Generalized Econometric Models with Selectivity,'' *Econometrica* 51 (1983): 507–12.

7. The table excludes the small percentage of the employed (less than 5%) who report working in more than one job and whose sector assignment would therefore be ambiguous. Given this ambiguity, as well as the small number of observations involved, we also dropped these workers from the econometric analysis of sector determination in the following section.

8. See also H. Alderman and V. Kozel, ''Formal and Informal Sector Wage Determination in Urban Low-Income Neighborhoods in Pakistan,'' Living Standards Measurement Study Working Paper no. 65 (World Bank, Washington, D.C., 1989); T. H. Gindling, ''Labor Market Segmentation and the Determination of Wages in the Public, Private-Formal, and Informal Sectors in San Jose, Costa Rica,'' *Economic Development and Cultural Change* 39 (1991): 585–605; and S. Khandker, ''Labor Market Participation of Married Women in Bangladesh,'' *Review of Economics and Statistics* 69 (1987): 536–41.

9. The model for men does not include a dummy for age 21–30, so the excluded category is age 15–30. The very small number of men age 21–30 in public sector employment precluded obtaining a reasonable estimate for this covariate.

10. A sizable anthropological literature for Africa, and West Africa specifically, indicates that men and women within households do not pool income or make expenditure decisions jointly. See, e.g., E. Fapohunda, ''The Nonpooling Household: A Challenge to Theory,'' and M. Munachonga, ''Income Allocation and Marriage Options in Urban Zambia,'' in *A Home Divided: Women and Income in the Third World,* ed. D. Dwyer and J. Bruce (Stanford, Calif.: Stan-

ford University Press, 1988); J. Guyer and P. Peters, ''Introduction,'' *Development and Change* 18, no. 2 (1987): 197–214, and references therein.

11. Since some of the demographic variables (marriage, children, number of adults) may be simultaneously determined with an individual's participation, it would be preferable to use instrumental variable methods to control for endogeneity. However, our data set (like most data sets) lacks adequate instruments for these variables. An alternative is to drop potentially endogenous covariates, although this will yield the correct reduced form only if the eliminated variables are in fact endogenous; if they are not, there is a risk of omitted variable bias from the correlation of omitted exogenous variables with the included variables. In alternative specifications of our logit models excluding all demographic variables, the magnitudes of the coefficients on a number of the remaining variables do change substantially, but for schooling (the main policy variable of interest) the basic pattern across sectors and by sex was unaltered.

12. Both location and receiving electricity may not be exogenous to labor market choices. However, this is likely to be less of a factor for women than for men.

13. This explanation is not incompatible with the fact that men of all other ethnic groups are actually more likely than Soussous men to enter self-employment. Men's enterprises tend to be much larger than women's, and Soussou men may lack the savings, or be unable to secure the credit, to start their own businesses.

14. P. Moock et al. also use this approach to handling multiperson enterprises in their study of informal enterprises in Peru. It is implicitly assumed, among other things, that the individual with the most schooling is the primary decision maker. P. Moock, P. Musgrove, and M. Stelcner, ''Education and Earnings in Peru's Informal Enterprises,'' Policy, Planning and Research (PPR) Working Paper WPS 236 (World Bank, Washington, D.C., 1989).

15. To construct the selectivity terms for inclusion in the earnings equations, we first reestimated the logit models after dropping 326 employed individuals from the sample (out of a total of 2,533 employed in the sample) for whom reliable earnings data were not available. Most of these were self-employed and included individuals engaged in fishing or agriculture (whose estimated net revenues tended to be implausibly low or negative); workers in multiperson family enterprises other than the most educated worker; individuals in enterprises with negative net revenues (only 1% of enterprises excluding fishing and agriculture), for whom the semilog specification could not be used; and women engaged in home activities partly for sale and partly for household consumption, for which cost data were not collected. Alternatives to dropping these observations from the first stage estimation, such as considering them as a separate category in the multinomial logit models or counting them among the nonparticipants, yielded results for both the selection terms themselves and the other regressors that were barely distinguishable from the earnings function estimates using the chosen approach.

16. The general form of the test statistic used to compare vectors or subvectors of estimated coefficients for different sectors is $(\beta_i - \beta_j)'(V_i + V_j)^{-1}(\beta_i - \beta_j)$, where β_i and β_j are the parameters for sectors i and j and V_i and V_j are their corresponding variance submatrices. The statistic is distributed as chi-square (j) under the null, where j is the number of restrictions. This test assumes $\text{cov}(\beta_i, \beta_j) = 0$ but does not impose equality of the variances of the disturbances for the two groups, as a Chow test (based on pooling the samples and interacting each regressor with a sector dummy) would do.

17. In the semilogarithmic specification, the percentage change in earnings due to the presence of the characteristic represented by a dummy variable is $100^*\{\exp(c) - 1\}$, where c is the coefficient on the variable.

18. Schooling and capital expenditures are positively correlated for male and female enterprises, indicating that educated managers have easier access to credit or better knowledge of how to employ capital in their enterprises. Alternatively, the better educated may be able to save and invest more because their profits are higher. For both men and women, dropping the capital variable from the self-employment earnings regressions results in increases in the magnitudes and *t*-statistics of all the schooling coefficients, suggesting that part of the overall impact of schooling comes from its association with the level of capital inputs used in the enterprise.

19. As real earnings of public employees at all skill levels were being eroded by inflation throughout the region through the early to mid-1980s, adjustments in nominal pay structures tended to favor those at lower levels, thus narrowing the range of earnings. See D. L. Lindauer, O. A. Meesok, and P. Suebsaeng, ''Government Wage Policy in Africa: Some Findings and Policy Issues,'' *World Bank Research Observer* 3 (1988): 1–25. Note that evidence of wage compression is usually gleaned from examination of official government salary scales and establishment surveys rather than through estimation on micro survey data controlling for sector selection, as done here.

20. Comparisons of expected wages of men in the public and private sectors calculated from the regression estimates (thus accounting for base earnings as well as incremental impacts of schooling and other covariates) indicate premiums to working in the private sector of about 18% and 23%, respectively, for men with secondary and university education, confirming the pattern suggested by the schooling coefficients (Glick and Sahn [n. 6 above]). This is the case despite several large civil service pay increases, primarily through increases in base salaries, granted by the government in the late 1980s at the same time as the public sector workforce was being reduced (Sahn and Mills [n. 3 above]). Since the survey was conducted, a huge increase of 145% in civil service base pay in 1991 (partly the result of a national strike) has likely substantially altered the balance between public and private sector pay.

21. Note, however, that postprimary schooling may have an impact on women's earnings through its effect on the level of capital in the enterprise (see n. 18 above). In his study of Côte d'Ivoire, Vijverberg also finds that years of women's primary schooling but not postprimary schooling raises enterprise profits, although unlike our findings in this study, he finds no effects at all for male schooling. Few other studies of self-employment earnings in developing countries exist, and these do not always find impacts of women's (or men's) formal schooling. See S. Khandker, ''Earnings, Occupational Choice, and Mobility in Segmented Labor Markets of India,'' World Bank Discussion Paper 154 (World Bank, Washington, D.C., 1992); S. Teilhet-Waldorf and W. Waldorf, ''Earnings of Self-Employed in an Informal Sector: The Case of Bangkok,'' *Economic Development and Cultural Change* 31, no. 3 (1983): 585–607; and Moock et al. Note that in the first two studies, either sample sizes or the variation in schooling are quite small, making it less likely that significant education effects would be found.

22. As noted, our multinomial logit specification includes household composition and exogenous income variables that do not enter the wage equations and can therefore be used to identify the selectivity term. However, the coefficients on these variables are not significant in the vector of parameters relating to the choice of private wage employment for women (table 3, col. 5), although for the other parameter subvectors (which also enter into the calculation of the lambda term) they are significant in a number of cases. As Maddala and others have noted, without identification by exclusion restrictions, the selectivity correction term is simply a nonlinear transformation of the variables appearing in

the second-stage (wage) equation. This explains the high correlation with the education covariates.

23. Vijverberg (n. 2 above) for Côte d'Ivoire, J. Armitage and R. Sabot for Tanzania and Kenya, and Knight and Sabot for Tanzania (manufacturing sector only) all report returns to occupational or employment experience that are generally higher than ours. J. Knight and R. Sabot, ''Labor Market Discrimination in a Poor Urban Economy,'' and J. Armitage and R. Sabot, ''Discrimination in East Africa's Urban Labor Markets,'' in *Unfair Advantage: Labor Market Discrimination in Developing Countries,* ed. N. Birdsall and R. Sabot (Washington, D.C.: World Bank, 1991). Note that our occupational experience estimates would be somewhat higher if age, which is correlated with experience, were excluded. Even then, however, the experience impacts tend to be lower than in these studies.

24. It is assumed that purchases of capital goods are made to exactly replace the depreciated value of existing capital and that the rate of depreciation was the same across all firms. Note also that enterprise capital may be to an extent endogenous to profits. Instrumenting this variable is difficult, however, because it is hard to find variables that affect the amount of capital used but do not affect profits directly as well.

25. Men's average hourly self-employment profits are also markedly higher than average wages for men in the private wage and public sectors. A possible reason for this, which is discussed at length in Glick and Sahn (n. 6 above), is that enterprise managers underestimate costs. If the underestimation of costs (or overestimation of profits) is related to enterprise size, the male-female earnings differential will be overstated because women's enterprises are smaller.

26. Strictly speaking, the lack of rejection of equality of the male and female slope coefficients for the private and public wage sectors implies that the male and female samples can be pooled and a sex dummy included to capture gender effects. However, in view of the rather large differences between many of the point estimates in the male and female wage regressions (especially for the public sector), we have chosen to present the results of separate male and female equations for private and public wage sectors as well as self-employment.

27. The decomposition of earnings differentials used in these studies is $\ln W_m - \ln W_f = \beta^*(X_m - X_f) + X_m(\beta_m - \beta^*) + X_f(\beta^* - \beta_f)$, where β_m and β_f are the vectors of estimated parameters from the male and female earnings regressions, and X_m and X_f are the mean values of the explanatory variables for males and females; β^* is the parameter vector from a wage equation estimated on the pooled sample of men and women and is an estimate of the wage structure that would prevail in the absence of discrimination. The first term on the right hand side represents the portion of the overall wage differential accounted for by differences in male and female characteristics. The second term represents the ''male advantage'' from discrimination (the amount by which men with mean male characteristics are paid in excess of the nondiscriminatory wage), and the third term represents the ''female disadvantage'' from discrimination (the amount by which women with mean female characteristics are paid less than the nondiscriminatory wage). The sum of the last two terms equals the total differential from discrimination. D. Neumark, ''Employers' Discriminatory Behavior and the Estimation of Wage Differentials,'' *Journal of Human Resources* 23 (1988): 279–95; R. Oaxaca and M. Ransom, ''On Discrimination and the Estimation of Wage Differentials,'' *Journal of Econometrics* 61 (1994): 5–12.

28. Calculations of predicted public sector wages using the male and female parameter estimates and the data means for the full sample of men and women indicate that at each level of schooling women can expect to earn less

than men with equivalent backgrounds do. This occurs despite the higher incremental schooling effects observed above for female public wage employees.

29. This concentration into a small number of traditionally ''female'' occupations of women in the wage work force occurs throughout the Third World. See R. Anker and C. Hein, ''Why Third World Employers Usually Prefer Men,'' *International Labor Review* 124, no. 1 (1985): 73–90.

30. To do this would involve estimating the probabilities of entry of men and women into each occupation and estimating earnings functions by sex for each occupation. The approach imposes potentially severe identification requirements because entry into an occupation should be treated as endogenous to earnings. See T. Schultz, ''Labor Market Discrimination: Measurement and Interpretation,'' in Birdsall and Sabot, eds. (n. 23 above). Our initial experiments with occupation effects were more modest and involved adding occupation dummies (treated as exogenous) to the earnings functions and decompositions.

31. W. Milne and M. Neitzert, ''Kenya,'' in *Labor Markets in an Era of Adjustment,* vol. 2, ed. S. Horton, R. Kanbur, and D. Mazumdar (Washington, D.C.: World Bank, 1994); Vijverberg; Glewwe (n. 1 above); Armitage and Sabot.

32. See E. Date-Bah, ''Sex Segregation and Discrimination in Accra-Tema: Causes and Consequences,'' in *Sex Inequalities in Urban Employment in the Third World,* ed. R. Anker and C. Hein (New York: St. Martin's Press, 1986); C. Di Domenico, ''Male and Female Factory Workers in Ibadan,'' in *Female and Male in West Africa,* ed. C. Oppong (London: George Allen & Unwin, 1983); M. Peil, *Cities and Suburbs: Urban Life in West Africa* (New York: Africana Publishing Co., 1981).

33. Y. Cogne, ''La participation des jeunes filles dans l'enseignement professionnel en Guinée'' (The participation of young girls in professional education in Guinea) (Conakry, Guinea: République de Guinée, Ministère de l'Enseignement Pré-Universitaire et de la Formation Professionnelle, 1994).

34. This general conclusion emerged from a range of country studies, including several in West Africa, surveyed in Anker and Hein (n. 29 above). Conditions in Guinea are likely to be similar, but it is also necessary to gain more direct information on the Guinean case.

35. Guinea's *Code de travaille* prohibits discrimination in hiring on the basis of sex, but its effectiveness in practice is another matter.

36. B. Mills and D. Sahn, ''Is There Life after Public Service? The Fate of Retrenched Workers in Conakry, Guinea,'' Cornell Food and Nutrition Policy Program Working Paper no. 42 (Cornell University, Ithaca, N.Y., 1993). This study was based on a survey of retrenched public sector workers, conducted in conjunction with the household survey used for the present article.

37. In 1993 girls represented only 25% of college (lower secondary) students, 20% of *lycee* (upper secondary) students, and just 6% of university students. World Bank, *Developing Girls' Education in Guinea: Issues and Policies,* Report no. 14488-GUI (Washington, D.C.: World Bank, 1995).

38. Girls perform substantial amounts of household work that may interfere with their attendance and schoolwork, and they also often face a hostile classroom atmosphere because of the attitudes of teachers and male students; consequently, girls drop out and repeat grades more frequently than boys do and (if they remain in school) they have higher failure rates than boys do on exams to qualify for the next level (World Bank [n. 37 above]). In the labor market, a perception of lower quality of girls' schooling may be a reason for employers' reluctance to hire women. For earlier cohorts of students, policies under the First Republic (1958–84) that encouraged girls' education by, among other things, lowering achievement requirements for girls to enter secondary school and set-

ting a quota for female students in the university may also have created perceptions among employers of lower quality of female graduates.

39. Female teachers have been recruited and sent to rural areas, grade repetition and student pregnancy policies have been relaxed, and greater emphasis has been given to school construction in rural areas. The last of these will improve rural enrollments overall but particularly for girls, since their enrollment is more strongly discouraged by distance to schools (World Bank).

40. Guinea's vocational education system is described in Cogne. For a study of Ghana that also emphasizes the limited training opportunities for girls, see C. Robertson, ''Formal or Informal Education: Women's Education in Accra, Ghana,'' *Comparative Education Review* 28, no. 4 (1984): 639–58.

41. Note that although the training system may be used to improve gender equity in access to jobs, our earnings equations estimates indicate that the direct productivity benefits to such training are limited. Obviously, the system also needs to be improved to better impart needed skills.

42. Summing the predicted probabilities of employment in table 4 in each sector, the probability that a woman works is .23 if she has less than a primary education, .19 if she completed primary school, .23 if she completed secondary school, and .46 if she completed university. For a comparative study of female schooling and employment, see A. Smock, ''Sex Differences in Educational Opportunities and Labor Force Participation Patterns in Six Countries,'' in *Comparative Education*, ed. P. Altbach, R. Arnove, and G. Kelly (New York: Macmillan, 1982).

43. A. Sow, ''Enquète sur la scolarisation des filles en milieu rural: Rapport de synthèse'' (Survey of girls' schooling in rural areas: Synthesis report) (Conakry: République de Guinée, Ministère de l'Enseignement Pré-Universitaire et de la Formation Professionnelle, 1994).

44. Not only are the potential benefits lower; households also face high opportunity costs of educating girls. This is because girls' contributions to household work (which are very important in rural areas) must be reduced if they are to attend school. Parents in Guinea cited this factor as a major reason for their reluctance to send girls to school, in addition to factors such as long distances to schools and coeducation of boys and girls (World Bank).