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AUTHOR Appleton, Simon; And Others
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ABSTRACT

Recent survey data from Cote d'Ivoire (Ivory Coast) is used to investigate gender differences in education and labor market participation. Previous researchers believed that women in Cote d'Ivoire were less likely than men to work for wages because of child-related obligations and low wage rates. This report contradicts this commonly held view by finding that women's limited participation is attributed, at least in part, to their having less education. A model based on the simultaneous determination of wages and choice of participation in the labor market is presented and is claimed to be more accurate than the common practice of treating these variables separately. The study also focuses on the differences between males and females in access to schooling. Numerous figures and tables of statistical data appear throughout the report. (DB)

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WORKING PAPER NO. 8
POLICY ANALYSIS

SDA Working Paper Series

Editorial Board

SOCIAL DIMENSIONS OF ADJUSTMENT IN SUB-SAHARAN AFRICA
WORKING PAPER NO. 8

Policy Analysis

Gender, Education, and Employment in Côte d'Ivoire

Simon Appleton, Paul Collier, and Paul Horsnell

The World Bank
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Paul Collier is director of the Unit for the Study of African Economies, Institute of Economics and Statistics, Oxford, England. Simon Appleton and Paul Horsnell are both research officers in the same Unit.

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SDA Working Paper Series Foreword

Integration of social and poverty concerns in the structural adjustment process in Sub-Saharan Africa is a major driving force behind the design of the World Bank's adjustment lending program in the Region. To further the goal, the Social Dimensions of Adjustment (SDA) Project was launched in 1987, with the United Nations Development Programme and the African Development Bank as partners. Since then many other multilateral and bilateral agencies have supported the project financially as well as with advice. The task presents a formidable challenge because of the severity of economic and social constraints in Africa and the intrinsic difficulty of tracing the links between economic policies and social conditions and poverty. It is essential to have a continuous professional dialogue between all concerned parties, so that the best ideas get discussed by the best minds, and become, as quickly as possible, available for implementation by policymakers. This is the aim of the SDA working paper series.

To fulfill its mission, the SDA Project operates on different levels. Conceptually, contributions need to be made which advance our understanding of how the economic crisis in Africa on the one hand and the adjustment response on the other hand affect the living conditions of people. Empirically, major improvements are needed in our knowledge of the social dimensions of life in Africa, how they change, and whether all groups in society participate effectively in the process of economic development. Gaining this knowledge will demand new efforts in data collection and policy oriented analysis of these data. Most importantly, policy actions are needed in the short term to absorb undesirable side-shocks stemming from the adjustment process so that the poor and disadvantaged are not unduly hurt, and in the long term to ensure that these groups fully participate in the newly generated growth. The SDA Project's mandate is to operate, in a concerted way, in all three domains: concepts, data, actions. This working paper series will report progress and experience in all three areas. I encourage every reader's active participation in the series and the work it reports on. It is meant to be a forum not only for exchange of ideas but even more importantly to advance the cause of sustainable and equitable growth in Africa.

Edward V.K. Jaycox
Vice President, Africa Region

The Social Dimensions of Adjustment (SDA) Project Working Paper Series

The SDA Project has been launched by the UNDP Regional Programme for Africa, the African Development Bank, and the World Bank in collaboration with other multilateral and bilateral agencies. The objective is to strengthen the capacity of governments in the Sub-Saharan African Region to integrate social dimensions in the design of their structural adjustment programs. The World Bank is the executing agency for the Project. Since the Project was launched in July 1987, 30 countries have formally requested to participate in the Project.

The Project aims to respond to the dual concern in countries for immediate action and for long-term institutional development. In particular, priority action programs are being implemented in parallel with efforts to strengthen the capacity of participating governments (a) to develop and maintain statistical data bases on the social dimensions of adjustment, (b) to carry out policy studies on the social dimensions of adjustment, and (c) to design and follow up social policies and poverty alleviation programs and projects in conjunction with future structural adjustment operations.

The working paper series "Social Dimensions of Adjustment in Sub-Saharan Africa" aims to disseminate in a quick and informal way the results and findings from the Project to policymakers in the countries and the international academic community of economists, statisticians, and planners, as well as the staff of the international agencies and donors associated with the Project. In the light of the three terrains of action of the Project, the working paper series consists of three subseries dealing with (a) surveys and statistics, (b) policy analysis, and (c) program design and implementation.

The Surveys and Statistics subseries focuses on the data collection efforts undertaken by the SDA Project. As such, it will report on experiences gained and methodological advances made in the undertaking of household and community surveys in the participating countries to ensure an effective cross-fertilization in the participating countries. The subseries would also include "model" working documents to aid in the implementation of surveys, such as manuals for interviewers, supervisors, data processors, and the like, as well as guidelines for the production of statistical abstracts and reports.

The Policy Analysis subseries will report on the analytical studies undertaken on the basis of both existing and newly collected data, on topics such as poverty, the labor market, health, education, nutrition and food security, the position of women, and other issues that are relevant for assessing the social dimensions of adjustment. The subseries will also contain papers that develop analytical methodologies suitable for use in African countries.

Another subseries, Program Design and Implementation, will report on the development of the conceptual framework and the policy agenda for the project. It will contain papers on issues pertaining to policy actions designed and undertaken in the context of the SDA Project in order to integrate the social dimensions into structural adjustment programs. This includes the priority action programs implemented in participating countries, as well as medium- and long-term poverty alleviation programs and efforts to integrate disadvantaged groups into the growth process. The focus will be on those design issues and experiences which have a wide relevance for other countries as well, such as issues of cost-effectiveness and ability to reach target groups.

Contents

1. *Introduction and Summary*

In the Côte d'Ivoire girls are less likely to go to school than boys, women are less likely to work for wages than men. These phenomena are of potential concern to policymakers because of their correlation with social and economic problems. At the social level, there are now good grounds for regarding maternal education as an effective means of improving the well-being of children[1]. Hence, the returns to girls' schooling may be greater than parents realize. At the economic level, lack of female participation in the labor market may indicate a deadweight cost of allocative inefficiency and additionally be an impediment to structural adjustment. Structural adjustment is essentially about resource mobility, the labor market being one of the central processes for this mobility. If women do not participate in the market, the reallocation of their labor may be more difficult.

This paper uses recent survey data to investigate gender differences in education and labor market participation. We find that the two are inter-linked. This in itself is not surprising. Extrapolating from developed country experience we might hypothesize that women participate less in the labor market partly because of child-bearing and child-rearing commitments and partly because of differentially low wage rates reflecting gender discrimination in wage setting. In turn, since the economic return to education depends upon participation and wage rates, the return to, and hence the extent of, female education would be lower than for males. This plausible extrapolation turns out to be false. We present and estimate a model incorporating the simultaneous determination of wages and the choice of participation. We demonstrate that this is methodologically superior to the more common practice of treating these as separable processes, showing the biases involved in that practice. The model is used to show that the limited participation of women in the Ivorian labor market is explained neither by child-related obligations, nor by differentially lower wage rates. It is partly the direct result of limited female education. women are far more likely to participate if they are educated. If the different educational endowments of women and men were fully to explain their differential participation in the labor market then the entire focus of inquiry would shift to the determinants of access to education. We would be able to dismiss the labor market both as an arena of gender discrimination and as an explanation for differential investment in education. However, this is not the case: controlling for education, women are still markedly less likely than men to be in the labor market. This, we suggest, is the result either of gender discrimination at recruitment or of women having systematically lower employment aspirations than men. To the extent that it is the former it may provide an explanation for the preference parents reveal to invest in the education of boys. That is, discrimination in the labor market would give rise to three of the observed gender biases:

1. Controlling for education, women in the labor force are less likely to work for wages than are men;
2. Parents are less likely to invest in the education of girls than in that of boys;
3. Women are less educated and hence are less likely to be in the labor market.

Our dataset enables us only to model the supply side of the labor market. Because we cannot simultaneously model the demand side, we are unable to distinguish between explanations based on discrimination on the part of employers and explanations based on the differential reluctance of women to enter the labor market. Just as the evidence can be interpreted as suggesting discrimination, so it can also be interpreted as showing differences in preferences. This is unfortunate because the two types of explanation have somewhat different implications. Suppose that instead of discrimination by employers, the cause of the observed biases is that women tend to have lower career aspirations than men. Such differential aspirations would account for why women are under-represented in the labor market as reflecting a reluctance to apply for such high status work. Low female aspirations could explain why girls are less likely to go to school in terms of a parental perception that, because of lower aspirations, girls would make less good use of educational opportunities than would boys. Finally, it might account for under-performance of girls once in the schooling system. A finding of the study is that a major reason why girls fail to continue to secondary schooling is that they perform less well in the primary school learning examination used to ration rate school places. If the underlying problem is low female aspirations then the appropriate policy intervention is probably to improve girls' access to and performance in education, whereas if it is discrimination at recruitment the solution must be the labor market. While we cannot distinguish between these explanations, we can, however, demonstrate two results of importance for policy. First, the gender-related differences in education and labor market participation cannot be explained by the arguably benign process of female specialization in child-rearing. It is therefore likely to be due either to discrimination (by employers and/or by parents) or to differentially low work aspirations on the part of women or to both. In either case this would pose a problem worthy of the concern of policymakers. Second, on either interpretation the problem can, as it happens, be diminished by increasing female access to and performance in education. This is because education powerfully reduces—but does not fully eliminate—the differential participation of women and men in the labor market. This can be interpreted either as education differentially raising female aspirations or as employers discriminating less against educated women than against the uneducated.

In Chapter 2 the focus is the Ivorian labor market. Our analysis of urban labor markets finds no evidence for the presence of gender biases in the process of wage determination—indeed Ivorian labor markets are more equitable in returns than markets in developed countries. However, we find strikingly large gender-based differences in labor market participation. Women are far less likely than men to work for wages. The two most superficially likely explanations for this turn out not to hold. First, it might have been that women are less likely to be waged simply because they are less likely to be in the labor force. However, restricting the analysis to adults already in the labor force, the powerful bias in labor market participation remains. Second, it might have been that women with children were less able to work the regular hours often required in wage employment and so participated in the labor force through self employment. This would have explained lower female participation in the labor market in terms of the differential household obligations of women. However, controlling for participation in the labor force, we find that the presence of children had no significant effect upon whether women participated in the labor market. Further, there is no evidence that child bearing reduces the potential duration of female working life relative to males. Evidently, social mechanisms for spreading the burden of child rearing are sufficient for children not to be a barrier to female labor supply to the market.

The question remains, therefore, as to why women have such a markedly lower participation rate in the labor market than men. Part of the explanation is that females are less well endowed by their parents and the government with education, and education increases the likelihood of participation in the labor market for both women and men. However, this is only part of the explanation: controlling for educational attainment and other characteristics, women in the labor force remain markedly less likely to participate in the labor market. Why is this the case?

We suggest that it is the result partly of female aspirations being lower than male, and partly of employers discriminating against women at recruitment, but the evidence for these propositions is necessarily indirect. The first evidence for differentiated aspirations concerns the differential effects of education. Although for both males and females education increases labor market participation, it increases female participation far more substantially than male. Proximately, this is because female participation in the labor market is far more sensitive than male to expected wages, and expected wages are increased by education. One interpretation of this result would be that pervasive societal differences in the expected occupational attainments of men and women give rise to different aspirations which are eroded by female education, either indirectly or because the prospect of higher earnings induces a change in female aspirations. Evidence that females have significantly lower aspirations than males is that they perform less well at school examinations, controlling for other characteristics. Note that in most developed countries females tend to outperform males in examinations.

However, there is also some evidence of discrimination at recruitment. First, we find that *credentialism* operates for women but not for men. That is, with the wage response included in the participation decision, educational diplomas are a significant determinant of female participation but not male. We have found no basis for labor market rationing of males through educational screening. For males the only separable effect of education on participation works through wages. Second, we find that for both women and men a significant effect on labor market participation is the paternal but not the maternal employment history. Parental employment history could potentially influence participation either through raising aspirations or though parental job contacts being a source of job offers. However, if the major effect were through aspirations then we would expect maternal employment history to be a significant influence upon female labor market participation, which it is not. Hence, the most likely interpretation is that fathers in employment help to acquire jobs for their daughters and sons.

The gender differential in access is shown to be confined to the private sector. The public sector is for given characteristics equitable in both returns and access, and hence the three to one plurality of males in the public sector is entirely accounted for by inferior female access to those characteristics. Given we find no evidence of rationing by educational achievement in the public sector, the major reason for lower female participation is low educational levels mapping onto lower wages and therefore onto a lower supply response.

In Chapter 3 our focus is on access to education. Where private processes of information acquisition are gender biased there is a case for an offsetting bias to be created in public processes. Our investigation of the determinants of educational acquisition reveals such a bias in private behavior. Females in the 11 to 18 age group are shown to be far more likely than males to receive no education at all. Since primary education is unrationed, this must reflect lower demand for female education. However, this pattern cannot be attributed to schooling costs, as the gender bias is shown to be greatest in high income households. We demonstrate the presence of an urban bias in the process of female educational acquisition, with the absolute gender bias against females being particularly marked, especially in the predominantly Moslem north. The gender bias is greatest in households with low parental education, suggesting that unless corrected, the inequalities in educational attainment may be self-perpetuating.

The root of this failure in private information processes is likely to be a combination of a *principal-agent* problem and deficient information flows. Given customary marriage arrangements, the increased factor rewards from education are more likely to accrue to the parental household from sons rather than daughters. Further if there is a perceived bias against females in high return activities such as urban labor markets there is again a disincentive to invest in female education as, even without the presence of the principal agent problem, the private marginal benefit of education will be seen as lower for females.

For post-primary education, differential educational achievements are potentially functions of either supply side rationing or differential demand patterns Access to progressive levels of secondary education given the completion of primary education exhibits less gender inequality than primary education Our analysis shows that it is the rationing constraint that drives the process of admission to lower cycle secondary school, and lower female admissions are primarily due to inferior female performance in the primary school leaving examination rather than to patterns of lower demand for places.

We suggest that the poorer academic performance of females may be due in part to the composition of the examinations themselves, and that females not only receive less quantity of education but also poorer quality in a male orientated educational system Alternatively the root cause may be differential aspirations The analysis further shows that females from low income households have a particularly low pass rate at the Certificat d'Etudes Primaires Elementaire (CEPE) level. Again, income inequality and educational inequality are seen to be interlinked and potentially self-perpetuating.

Hence, we find that there are substantial gender-related asymmetries in the Ivorian labor market Quite unlike developed countries these relate not to wage discrimination but to differences in participation in the labor market, which in turn are strongly related to differential achievements in education. This inferior academic performance of females is a combination of both factors correctable within the structure of the educational system and also factors alterable only at a more general level. Chapter 4 provides a brief discussion of the policy implications of the study.

Note

1. See Lockwood and Collier (1988) for a survey of the literature

2. The Labor Market

2.1 Participation and Wage Determination

In this chapter our focus is on the labor market. In Part 2.1 we investigate the determination of participation and wage rates for the labor market as a whole and identify gender differences. In Part 2.2 we push the analysis one stage further, disaggregating the labor market by the extent of unionization and by whether the employer is in the public or private sector.

2.1.1 *Introduction*

We now present and estimate a model incorporating the simultaneous determination of wages and the choice of participation in an African urban labor market. We demonstrate the bias obtained by considering these as two separable processes, and identify the underlying determinants of both mechanisms in the Côte d'Ivoire. Educational, informational, gender and other demographic characteristics are all found to be important explanatory variables, and a strong linkage between the two processes is discovered.

In Section 2.1.2 we formally state the model underlying the subsequent analysis, and describe the estimation methods employed. This is inevitably technical and so the essential concepts are summarised now for those wishing to omit the section. Whether an individual participates in the labor market is viewed as depending upon the wage they can thereby earn (the *actual wage*) being greater than the wage below which they will choose not to sell their labor (the *reservation wage*). Both of these wages depend upon the characteristics of the individual. We will find that gender powerfully influences the reservation wage but has no effect upon the actual wage. Education increases the actual wage relative to the reservation wage and so increases labor market participation. Van der Gaag and Vijverberg (1989) have previously estimated the relationship between characteristics and the actual wage. However, because they considered this as a separate process their results are biased. By restricting their analysis to those in wage employment their estimates are conditional upon the characteristics of those in wage employment. In Section 2.1.3 the model is estimated on a full sample using dummy variables for gender, and in Section 2.1.4 the analysis is repeated on samples disaggregated by gender. While there exists no method that does not simultaneously model the demand side of the labor market that can decompose the effects of differential access to labor markets by characteristics, and the effects the characteristics might have on the reservation wages underlying the choice behavior of agents, we provide quantitative measures of the magnitude of the combined effects of these two influences. Section 2.1.3 contains the definition of the sample used for estimation, and definition and discussion of the variables used. Finally in Section 2.1.5 we provide a brief discussion of non-wage benefits and the comparative legal position of workers by gender in the Ivorian labor market.

2.1.2 *The Modeling of Wage Determination and Labor Market Participation*

We begin from the basis of a standard labor force participation model derived from the neoclassical theory of labor supply. Individuals base their decision to participate in a labor market upon their evaluation of a reservation wage, W^R say, which is determined by the marginal rate of substitution between consumption and leisure at the level of zero consumption. An individual's evaluated reservation wage reflects both a set of observed individual characteristics which we will denote by X, and a set of unobservable characteristics, denoted by ξ. Hence we postulate

$$W_i^R = W^R\left[X_i, \xi_i\right]$$

While W^R is an unobservable variable, we do observe W, the realized market wage of a participating individual. This is also a function of a set of observed characteristics, say Z, and a set of unobservable characteristics, say δ. Labor force participation will then be observed if and only if:

$$W^R\left[X_i, \xi_i\right] < W\left[Z_i, \delta_i\right]$$

However, with W^R unobserved, we can only detail a binary variable, say Y, which takes on a value of unity if an individual is observed to participate in the labor market and zero otherwise, with Y a function of X, ξ, and the level of realisable wages. Hence the two basic equations in a one sector participation/non-participation model are, evaluated on a sample of size n,

$$W_i = W\left[Z_i, \delta_i\right] \qquad (i = 1, \ldots, n)$$

$$Y_i = Y\left[X_i, W_i, \xi_i\right] \qquad (i = 1, \ldots, n)$$

To this point the approach is common and relatively uncontroversial. Estimation of the parameters of the model is seen to involve the estimation of a structural participation equation, and also a market wage equation, with the unobservable vectors of personal characteristics, δ and ξ, being incorporated into the respective error terms. However, the choice of estimation procedures may prove to be non-trivial. The choice of procedure hinges upon the assumptions made on first the errors of the wage equation, secondly the errors of the participation equation, and thirdly on the interrelationship between the two sets of error terms. We estimate the model on data using four separate procedures. All assume normality of the errors in the regression upon the logarithm of observed wages, and we employ two specifications for the participation equation with two modelings of the interrelationship between the errors. We now specify each of these procedures.

The first two approaches are the least computationally demanding, but are restrictive and theoretically ad hoc. They assume that there is zero covariance between the errors of the wage equation and the participation equation, and hence the two functions can be separately estimated. Let the vector Z have k components and the vector X have l, where both Z and X include a constant term. Then we can formulate a wage function in terms of its coefficient vector β, and depending on the functional form employed, a likelihood function for the participation decision in terms of a coefficient vector γ. Hence:

$$\ln W_i = \sum_{j=1}^{k} \beta_j Z_{ij} + e_i \qquad (i = 1, \ldots, n)$$

and letting $S_i = \sum_{j=1}^{l} \gamma_j X_{ij} + \gamma_{l+1} \hat{W}_{ij} \qquad (i = 1, \ldots, n)$

$$L(\gamma) = F\left[Y_i, S_i\right].$$

Using a probit formulation for the participation decision we have the likelihood function, where Φ is the distribution function of the standard normal,

$$L(\gamma) = \prod_{i=1}^{n} \Phi\left[\frac{s_i}{\sigma}\right]^{Y_i} x \prod_{i=1}^{n}\left[1 - \Phi\left[\frac{s_i}{\sigma}\right]\right]^{1 - Y_i},$$

or using a logit specification,

$$L(\gamma) = \prod_{i=1}^{n}\left[\frac{e^{S_i}}{1 + e^{S_i}}\right]^{Y_i}\left[\frac{1}{1 + e^{S_i}}\right]^{[1 - Y_i]}$$

which convert to the respective log-likelihood functions (setting σ to unity),

$$LogL(\gamma) = \sum_{i=1}^{n} Y_i \log \Phi\left[^{S_i}\right] + \sum_{i=1}^{n} [1 - Y_i] \log\left[1 - \Phi\left[^{S_i}\right]\right]$$

and

$$LogL(\gamma) = \sum_{i=1}^{n}\left[Y_i S_i - \log[1 + e^{S_i}]\right].$$

Hence separate estimation involves the estimation of the wage equation by ordinary least squares. From this equation an estimated wage can be imputed for each individual in the sample, and then the chosen form of the log-likelihood function can be maximised. However this approach suffers from the potential of sample selection bias. As the selection rule derived from the participation equation is potentially related to the error term of the wage equation, all coefficients may be biased, and inferences based on them can be misleading. Further this implies that the estimated wage used in the structural decision equation may be incorrect, thus potentially biasing all coefficients in that equation. The computational simplicity of the approach derives solely from the assumption of zero covariance of errors which can never be known a priori. Hence, we present the results of separate estimation solely for comparison with more analytically based procedures.

The first attempts to correct for sample selection bias used a multivariate normal distribution, for example, Amemiya (1974) and Gronau (1974). While tractable, the full likelihood function for modeling the joint distribution of errors is cumbersome. We return to this function in our later discussion of the more general case which does not impose the multivariate normal distribution. Given the demanding nature of the full information approach, two stage limited information approaches for obtaining consistent estimates have been developed following Heckman (1976) and (1979). The method adds a variable to the wage equation which proxies the non-random component of the error term. As was shown by Melino (1982), tests on the significance of this variable are equivalent to the Lagrange Multiplier test of the null hypothesis of no sample selection bias. Indeed, Melino demonstrates that the squared t-ratio of the coefficient on this variable is in fact the relevant Lagrange Multiplier statistic.

Specifically this approach implies the initial estimation of the reduced form participation equation. Thus Y should be modelled as a function of all the components of the union of the sets of variables X and Z, say Q. Identification conditions for the full model were derived by Nelson (1975), and imply that if we do not impose zero covariance of errors as in the separate estimation approach, there must be at least one member of Z not in X. Say Q has K members, multivariate normal distribution for the model implies a probit formulation for the reduced form participation equation, with parameters, say, β. From this estimation the inverse Mills ratio, λ, is defined as;

$$\lambda_i = \frac{\phi\left[-\sum_{j=1}^{K} \beta_j Q_{ij}/\sigma\right]}{1 - \Phi\left[-\sum_{j=1}^{K} \beta_j Q_{ij}/\sigma\right]}$$

Where ϕ is the density function of the standardized normal distribution and Φ is its cumulative distribution function. This is then used as an explanatory variable in the wage equation, and thus the estimated values from this compensated equation can be used in the structural participation equation.

A more general approach is provided by Lee (1982) and (1983). Let the errors of the wage and participation equations be ε_1, and ε_2 respectively with distribution functions $F_1[\varepsilon_1]$ and $F_2[\varepsilon_2]$, and marginal density functions $f_1[\varepsilon_1]$ and $f_2[\varepsilon_2]$. Further let ρ be the simple correlation coefficient between the errors, and Φ and Φ^{-1} respectively the distribution function of the standard normal and its inverse. Then Lee has shown that the full likelihood function may be written as

$$L = \sum_{i=1}^{n} f_1[\varepsilon_1]^{Y_1}\left[\Phi\left[\frac{\Phi^{-1}[F_2[\varepsilon_1]] - \rho\,\Phi^{-1}[F_1[\varepsilon_1]]}{\left[1-\rho^2\right]^{1/2}}\right] \times \right.$$

$$\left.[1 - F_2[\varepsilon_2]]^{[1-Y_1]}\right.$$

If normality for the errors of the wage equation is assumed,

$$\Phi^{-1}[F_1[\varepsilon_1]] = \frac{\varepsilon_1}{\sigma} \text{ and } f_1[\varepsilon_1]] = \frac{1}{\sigma}\phi\left[\frac{\varepsilon_1}{\sigma}\right].$$

For any given form for the participation equation the likelihood function is then still tractable, but involves the computation of the cumulative normal of a function involving an inverse normal for each observation during each pass of the function during an iteration Further, the first order derivatives are distinctly cumbersome.

The generalized two stage procedure involves estimation for the reduced participation equation and then the appropriate variable for inclusion in the wage equation is:

$$\phi\,[\Phi^{-1}F_2[\varepsilon_2]]/F_2[\varepsilon_2]$$

and the coefficient on this variable will be $\sigma\rho$. A probit specification for $F_2[\varepsilon_2]$ then reduces to the Heckman two stage procedure, but the method allows for any specification. Thus we estimate the model using both a logit and a probit specification for $F_2[\varepsilon_2]$ to derive the logit and probit two stage estimates of the model. Hence our four estimation procedures are logit and probit formulations for the participation equation, with and without the imposition of zero covariance between the error terms.

2.1.3 An Application to Urban Areas of Côte d'Ivoire.

In this section we estimate a wage and participation model for urban areas of the Côte d'Ivoire using the procedures and methodology described in the previous section. The data is drawn from the second in the series of the Côte d'Ivoire Living Standards Surveys (CILSS), which was conducted in 1986 on a total sample of 1,601 households. The design, procedures, and implementation for this survey are explained in Ainsworth and Muñoz (1986) and Grootaert (1986).

Respondents to the survey gave information about their main and secondary employment over the previous seven days and the previous year. We define an individual as being in wage employment if their main work over the previous year, i.e. the work they devoted most time to, was as a worker in a non-family business which gave them a monetary reward. The trades and industries in which wage employment were found are shown with frequencies in Table 2.1. The sample was limited to individuals aged 16 or over and under 65 at the time of survey who were not in schooling. The distribution of the percentages of the sample in waged employment thus defined are shown

Table 2.1 Occupations in Waged Employment Sample, CILSS 1986

	Female	*Male*	*Total*
Shopkeeper	3	0	3
Salesperson	3	7	10
Warehouse Man	0	3	3
Buyer for trader	0	2	2
Blue Collar Worker	1	6	7
Laborer	0	29	29
Forester/Lumberjack	0	1	1
Welder	0	1	1
Tailor	3	6	9
Mechanic	0	15	15
Repairman	0	2	2
Mason	0	7	7
Electrician	0	5	5
Painter	0	1	1
Hairdresser	4	0	4
Baker	0	1	1
Butcher	0	1	1
Photographer	1	1	2
Housemaid	9	0	9
Cook	1	2	3
Guard	0	18	18
Doctor	2	1	3
Nurse	2	7	9
Medical Orderly	3	1	4
Lawyer	0	2	2
Journalist	0	2	2
Teacher: Second/Univ.	10	28	38
Teacher: Primary	12	23	35
Principal	0	4	4
Architect	0	5	5
Taxi or Bus Driver	0	24	24
Truck Driver	0	7	7
Train Conductor	0	5	5
Sailor	0	1	1
Pilot or Stewardess	2	4	6
Machine Operator	0	1	1
Transporter	0	1	1
Draftsman	0	1	1
Administrator	0	5	5
Accountant	1	9	10
Cashier	1	0	1
Engineer	0	9	9
Secretary	34	7	41
Telephone Operator	1	0	1
Programmer	0	1	1
Priest	0	1	1
Armed Forces	0	14	14
Artist	0	1	1
Other Technical/Prof.	11	78	89
Other	6	56	62
Total	110	406	516

in Map 1, IBRD #22576, across the 100 clusters surveyed in the CILSS, with figures for urban clusters being boxed, and rural clusters represented by a diamond. The map demonstrates that on the basis of our definition the rural labor market is virtually non-existent, with only two rural clusters directly to the east of Abidjan around the delta of the Bia river, recording participation rates of above 10 percent. We have therefore restricted the analysis to the 43 urban clusters of the survey, where, as is shown in Map 1, only four show participation rates below 10 percent. This produced a final sample of 2,324 individuals of whom 515 were participants in the labor market as defined, a participation rate of 22.2 percent.

Our setup of the model and data is somewhat similar to Grootaert (1988) and can be contrasted to that of van der Gaag and Vijverberg (1989), who estimated wage equations on the 1985 cycle of the CILSS. First, their sample consists solely of wage earners, and hence the lack of any underlying selection rule potentially biases their results. They note their estimates are all conditional on being in wage employment. However this is misleading, a more accurate statement would be that the estimates are conditional on the characteristics of those in employment, which conveys far more of a restriction. In other words, the returns to characteristics estimated by such a method are not the returns relevant to the participation decision by an individual. Secondly, they only use a recall period of seven days, when a period of a year is available. This has several flaws. It potentially ignores relevant information as there need not be a strong correlation between the main job over a seven day period and that over a year. Indeed over a quarter of the sample we have defined as waged reported a different main job for the previous week. There is also the danger of including individuals for whom, seen over a year, wage employment is a minor component of their time allocation.

The dependent variable for the wage equation is defined as the logarithm of the hourly CFA Franc wage. Respondents were given the choice of the time unit for measurement of wages, and these have been standardized to an hourly wage using their information on hours worked per day and days per week where relevant.

The choice of explanatory variables for wages is problematic as some variables, such as experience, are employment specific and hence not observed for nonparticipants. Such variables therefore need to be instrumented and imputed values used for all observations. However instrumentation for experience proved difficult, indeed experience was only well proxied by age, which is a desirable regressor in its own right. Therefore we have used AGEY, the individual's age measured in years, with the important proviso that its effect captures an experience as well as an age effect.

Table 2.2 Mean and Standard Deviation of Variables

Variable	*Mean*	*Standard Deviation*
AGEY	32.400	12.577
SEX	0.466	0.499
CEPE	0.372	0.483
BEPC	0.097	0.296
HIGHER	0.037	0.189
TECH	0.106	0.308
GRP	3.003	2.891
GRS	0.966	1.601
GRS2	0.223	0.732
YRU	0.121	0.774
YRST	0.337	1.024
HEAD	0.270	0.444
MARRIED	0.565	0.496
CHILD	0.568	0.495
ABIDJ	0.472	0.499
NAT	0.814	0.389
FGOVT	0.076	0.265
MGOVT	0.008	0.088

We have also used as explanatory variables the dummy variables SEX, NAT, and ABIDJ, taking the values of unity if an individual is male, Ivorian, and if they are domiciled in Abidjan respectively. Means and standard deviations for these and all variables used are shown in Table 2.2. We have also used a series of educational indicators, the construction of which follows logically from the structure of the Ivorian educational system, and indeed the same variables are used by van der Gaag and Vijverberg.

Schooling in the Côte d'Ivoire begins with a six grade primary system with two grades each at levels defined as preparatory, elementary, and intermediate. At the end of the sixth grade students take the primary school leaving exam, the Certificat d'Etudes Primaires Elementaire (CEPE). Thus we use two variables for primary school education, GRSP representing the number of completed grades of primary school, and CEPE representing possession or not of the CEPE diploma. It should be noted that we can measure grades of primary education and not years, repetition of grades is prevalent and introduces a divergence between the two measures.

Secondary school education consists first of a lower four year cycle leading to the examination for the Brevet d'Etud.. du Premier Cycle (BEPC), and then a further three years before baccalaureat and then the university system. We thus use the variables GRS1, GRS2, YRU to denote grades completed in the first and second cycles of secondary education, and years in university respectively. We use the variable BEPC for possession of that diploma, and HIGHER for baccalaureate and higher degrees, namely Probatoire, License, Masters, and Doctorates. For such individuals who have recorded their highest qualification achieved as either a technical or other diploma, there is no information on which of the CEPE, BEPC, or higher degrees they might have obtained, although we know the highest grade of education they achieved. Hence we have made the assumption that they obtained any diplomas that their educational grade implies they would have obtained. The final variable used in the wage equation, and then only for our two stage procedures is LAMBDA, the sample selection correcting variable as detailed in the previous section, and derived from the reduced form estimation of the participation equation.

It might seem logical also to include as determinants of wages variables denoting technical or professional training. For reference two such variables, TECH, taking the value unity if a technical or professional qualification is held, and YRST, years of technical or professional training are included in Table 2.2. However there are two strong grounds for not including them as explanatory variables. First, such training may be carried out as a result of employment, or as on the job training, and we have no data to distinguish between post- and pre-employment technical training. In the Côte d'Ivoire both methods of training coexist, (Grootaert (1988)), although the balance is moving towards on the job training in both informal and modern sectors, (Bas (1989)). Secondly, even if such a distinction could be made, such measures are still not desirable regressors. Without attempting a simultaneous model of educational acquisition, education has to exhibit strong exogeneity. The decision to acquire technical or professional training pre-employment is a very strong signal by an individual of an intention to seek modern sector employment, indeed over 75 percent of those with technical qualifications in our sample are employed. On the CILSS data the former argument is sufficient to justify exclusion as such training may be employment specific, but the latter argues more generally that the grounds for exogeneity are weaker for technical compared to other forms of education.

For the participation equation we explain WAGED, a binary variable taking the value unity if an individual is in wage employment as defined above. As explanatory variables we use AGEY and its square, SEX, NAT, and ABIDJ all as defined above. We also use the dummy variables, HEAD, MARRIED, and CHILD taking the value unity if the individual is recorded as head of household, married, or has children respectively. During our estimation we used other formulations for CHILD, i.e. numbers of children and

Table 2.3 Côte d'Ivoire 1986 Determinants of Wages and Employment
Parameters of Wage Regression

	(1) and (3)	*(2)*	*(4)*
CONSTANT	3.33386	4.29600	4.38493
	(18.524)**	(15.060)**	(15.448)**
AGEY	0.04595	0.03908	0.03847
	(12.810)**	(10.270)**	(10.110)**
CEPE	0.40808	0.30367	0.29162
	(2.731)**	(2.090)**	(2.004)**
BEPC	-0.06351	-0.05459	-0.05191
	(-0.610)	(-0.530)	(-0.509)
HIGHER	0.13284	0.20493	0.21796
	(0.658)	(1.017)	(1.093)
ABIDJ	0.05195	0.00833	0.00468
	(0.953)	(0.129)	(0.073)
SEX	0.01616	-0.19855	-0.20880
	(0.212)	(-2.245)**	(-2.392)**
GRSP	0.03902	0.02388	0.02286
	(1.596)	(1.013)	(0.966)
GRS	0.16720	0.13817	0.13459
	(5.855)**	(4.840)**	(4.732)**
GRS2	0.21858	0.18500	0.18182
	(4.288)**	(3.622)**	(3.593)**
YRU	0.08161	0.06209	0.05925
	(2.132)**	(1.601)	(1.547)
NAT	0.07345	0.05333	0.04336
	(0.809)	(0.610)	(0.495)
LAMBDA	—	-0.35190	-0.39260
		(-4.277)**	(-4.674)**
R-Squared	0.58604	0.59976	0.60266
Log-Likelihood	-533.16998	-517.91098	-516.04999

* = significant at 10 percent
** = significant at 5 percent
— = not applicable

breakdowns by gender of children, but the results of these formulations proved no more illuminating than the results we report for the simple binary variable case. There is no a priori case for believing that educational variables exert any other effect on the participation equation than their effect working through wages, but we include the variables CEPE and BEPC to see whether for given wages, education tends to reduce or increase the propensity to work. This provides a better approach to the effects of education than estimation that does not include an imputed wage. This amounts to no more than the estimation of a reduced form equation, and normally leads to very powerful educational effects, (for example Behrman and Wolfe (1984a, 1984b)), and usually to highly misleading inferences about the role of human capital. This interpretation of any effects caused by the CEPE and BEPC diplomas can however only be strictly justified if no barriers to entry to the labor market exist, or at least no barriers to entry whose severity is affected by education. In effect our participation equation also includes the selection criteria used by employers in drawing from the population willing to work at the market wage. This means that the effect of a variable on the reservation wage cannot be disentangled from its effect on barriers to entry, except on the basis of a priori judgments. A priori, diplomas are likely to affect the probability of selection in the face of barriers to entry, and their inclusion in the participation equation can therefore be justified on these grounds. We provide measures of the scale of these joint effects in the later discussion, but it should meantime be noted that the discussion of participation involves an underlying choice decision, which in the presence of barriers to entry caused by rationing is an amalgam of both individual and employer decisions.

Information is given in the CILSS about the occupational background of parents. As parental experience may carry information effects, or by the direct operation of a contact system facilitate entry into the labor market, we use the two variables FGOVT and MGOVT. These take the value of unity if respectively the individual's father and mother had employment that was technical, professional, or in the government sector. The final determinant of participation, although of course not in the reduced form participation equation, is the imputed expected logarithmic wage, EXPW, derived from the associated wage determination equation.

The model was run using all four of the estimation procedures described in the previous section. We present the results of all procedures, using in our tables the following nomenclature:

1. logit participation model, separate estimation;
2. logit participation model, two stage estimation,
3. probit participation model, separate estimation,
4. probit participation model, two stage estimation.

The models were run for the sample as a whole, and a disaggregation is carried out in the next section. We present estimates for the entire sample first, showing in Table 2.3 the parameters of the wage equation, and in Table 2.4 the parameters of the participation equation. For estimation methods (2) and (4) the underlying reduced form equations are shown in Table A.1 of the Appendix for completeness, as its coefficients are by themselves uninterpretable. For Table 2.3, as separate estimation implies a simple regression for wages,

Table 2.4 Parameters of Participation Decision

	(1)	(2)	(3)	(4)
CONSTANT	-14.43350 (-12.417)**	-13.77280 (-13.343)**	-7.73896 (-12.740)**	-7.33309 (-13.743)**
AGEY	0.38238 (8.011)**	0.32595 (6.103)**	0.19496 (7.955)**	0.16578 (5.963)**
AGE2	-0.00571 (-9.705)**	-0.00491 (-7.707)**	-0.00299 (-9.924)**	-0.00256 (-7.794)**
HEAD	1.72523 (8.553)**	1.48935 (7.006)**	1.00147 (8.932)**	0.87893 (7.445)**
MARRIED	0.45215 (2.230)**	0.39132 (1.922)*	0.23151 (2.094)**	0.20422 (1.840)*
FGOVT	0.48837 (2.060)**	0.41826 (1.746)*	0.26168 (2.003)**	0.23192 (1.764)*
MGOVT	0.36963 (0.503)	0.30522 (0.415)	0.23012 (0.567)	0.19660 (0.486)
CHILD	0.08799 (0.428)	0.08027 (0.390)	0.03991 (0.354)	0.03504 (0.310)
ABIDJ	0.63753 (4.471)**	0.58989 (4.067)**	0.34198 (4.407)**	0.31717 (4.014)**
SEX	0.95811 (5.291)**	1.03707 (5.702)**	0.54630 (5.541)**	0.58294 (5.902)**
CEPE	0.58295 (2.028)**	0.55647 (1.961)**	0.27873 (1.762)*	0.26984 (1.710)*
BEPC	0.06210 (0.240)	0.01998 (0.077)	0.01811 (0.124)	0.00980 (0.067)
NAT	0.45746 (2.475)**	0.41865 (2.246)**	0.18527 (1.863)*	0.16859 (1.682)*
EXPW	0.88007 (3.624)**	0.96236 (3.806)**	0.51476 (3.779)**	0.54428 (3.861)**
Log-Likelihood	-733.09998	-732.35999	-736.96997	-736.69000

* = significant at 10 percent
** = significant at 5 percent

	Predicted							
	0	1	0	1	0	1	0	1
Actual	1722	87	1716	93	1726	83	1724	85
	214	301	213	302	217	298	214	301

the results are of course the same for methods (1) and (3).

The most striking aspect of Table 2.3 is that sample selection bias is a problem for this data, and the results of a straightforward wage regression are highly misleading. LAMBDA is highly significant for both the logit and probit reduced form generations of it, judged either on a straightforward t-test or on Melino's (1982) derived Lagrange Multiplier test for sample selection bias. This strong effect leads to considerable changes in the sizes and significance of coefficients, the most striking change caused by the adjustment for sample selection bias being the effect of gender, positive but insignificant in the unadjusted formulation, but significantly negative in the adjusted formulations. There is shown to be a wage premium for a female of some 20 to 21 percent over a male with the same characteristics. Hence any bias against females is seen as being due to inferior female access to those characteristics which augment wages, and in access to the labor market itself. We expand on this issue by a decomposition by gender in the next section. However it is worth noting at this point that our results lend credence to two views of the urban labor market. First, there is evidence of non-homogeneity of labor by gender, and secondly the positive bias to women, once characteristics are adjusted for, supports an hypothesis that the level of female participation in the urban work force is suboptimal relative to the given level of male participation.

For the estimation of the returns to education to have any relevance, they must be calculated as the returns observed by an individual in the general population. Results obtained conditional on wage employment, i.e. not taking sample selection bias into account, cannot theoretically be meaningful inputs into, say, a model of education acquisition. Table 2.3 gives an indication of the magnitude of errors obtained if we do infer from returns based on simple regression Comparing the results of the simple regression with the sample selection compensations, we see that the estimated premia obtained through possession of the CEPE diploma falls from 4 percent to 29 or 30 percent. Likewise the return to a completed grade of lower cycle secondary school falls from 16.7 to 13.8 percent, and for a grade of upper secondary from 21.8 to 18.2 percent. Aggregated over four years of lower secondary and three years of upper secondary the accumulated divergence is in excess of 20 percent. This demonstration of the magnitude of the error that can be made by ignoring the selection rule casts a shadow of doubt on most of the studies of the returns to education cited in Psacharopoulos (1985), particularly for countries with low levels of labor force participation and hence potentially greater sample selection bias.

With the acute sample selection bias exhibited by the simple regression, we can only draw meaningful conclusions on the basis of (2) and (4), which tell the same story. There are no wage premia on the basis of location or nationality. Significant premia exist for females, as noted above, and for the possession of qualifications. Likewise significant returns exist to completed grades of both lower cycle and upper cycle secondary education, with a range of 13.5 to 13.8 percent for the former and 18.2 to 18.5 percent for the latter, while no significant returns are observed in primary education or university once the compensation for diplomas achieved is made. Age increases wages at the rate of 3.8 or 3.9 percent per year, with, as cautioned above, this effect also including the proxy of the effect of a year's extra experience. Formulations with age also entering quadratically were tested, but gave no significance for the quadratic terms. The combined effect of age and experience on wages is therefore presented as log-linear.

Table 2.4 demonstrates how the errors made by ignoring sample selection bias are compounded if the imputed expected wage obtained from a wage equation is then used in a participation equation. For both the logit and probit specifications of the structural participation decision, imputation of the uncompensated wage, respectively models (1) and (3), leads to underestimation of the direct effect of wages, and considerable changes in the magnitude of other significant determinants. There is no necessary reason for the corrected wage versions to produce a greater proportion of observations whose participation behavior is predicted correctly by the model, the major effect of the correction is the reallocation of explanatory power across the set of determinants. Indeed, as evidenced by the predicted versus actual crosstabulations shown in the table, the corrected logit predicts five fewer observations correctly compared to the separately estimated model, while the effect of the compensation on the probit formulation is a gain of just one. The predictive power of each is impressive, explaining some 87 percent of all observations, 58 percent of participants and 95 percent of non-participants.

At the 10 percent level of significance, the sets of significant and insignificant determinants of participation are identical across the four formulations of the model. We thus discuss the quantitative effects common to the formulations, and then present measures of the quantitative differences between the corrected and the uncorrected imputed wage versions, and between the alternative specifications of the distribution of the error term.

The justification for the concern in deriving a meaningful imputed logarithmic wage is shown in the significance of the linkage between wages and participation with the anticipated positive coefficient. The truncation of the female labor supply relative to male implied by the results of the wage determination exercise is confirmed, with males exhibiting a greater propensity to participate. Given that our base group of non-participants may include individuals whose participation is precluded by social or cultural barriers, or by direct discrimination in gaining entry to the labor market, it is not possible to disentangle the issue of whether women have inferior access or merely higher reservation wages. However, our result is suggestive of the hypothesis of inferior access and discrimination, either from society or the labor market itself, even though our wage equations imply that no such discrimination exists within the labor market in the return to female labor.

Other than the effects of gender and wages, Table 2.4 also shows that participation is affected by a mixture of demographic, informational, and educational variables. There is a significant quadratic effect in age, peaking across the models in the range of 32 to 34 years. Participation is shown to be higher in Abidjan compared to other urban areas, and also among the Ivorian population compared to international migrants, although the combined effects of location and nationality are still less than the effect of gender. The three household composition and type variables, HEAD, MARRIED, and CHILD convey the result that having children has no effect on participation, while household heads are more likely to be participants. Further, on this combined sample of males and females, married individuals show a greater propensity for labor force activity than those unmarried.

Even allowing for the premia obtainable from education in the returns to labor, the CEPE dummy is still positively significant, i.e. educated individuals show a greater propensity to participate beyond that induced by higher wages. However this increased propensity is not further enhanced by post primary education, the BEPC dummy is insignificant as were all other measures of post-primary education used in the model.

The effect of paternal education is shown to be a significant positive factor in participation. This effect is potentially a combination of both an access effect with the possibility of manipulating existing family links within the labor market, and also informational. Maternal participation, which is at a very low level, is seen to have no significant effect.

To assess the quantitative effect of the determinants of participation we choose the following procedure. For a given characteristic, say X, which increases the probability of an individual being a labor force participant, we determine the percentage difference, *P*, in the absolute wage necessary to induce an equal propensity to participate in an individual without X compared to an individual with X and otherwise identical characteristics. Hence in the absence of any barriers to entry caused by non-possession of X, this percentage is simply the percentage difference between the individuals' reservation wages. If γ_1 is the coefficient on X, and γ_2 the coefficient on the imputed logarithmic wage, then some manipulation yields:

$$P = 100\left[e^{\gamma_1/\gamma_2} - 1\right].$$

Hence such a measure is independent of the absolute level of wages or the common vector of other characteristics. In part (a) of Table 2.5 the measure has been calculated for the four models considered, and the significant variables in Table 2.4, other than age, which is considered below, and of course the imputed wage.

To first compare the models, for all variables but SEX, the introduction of a corrected imputed wage leads to considerable changes in magnitude. Further, the difference between the logit and probit formulations in their uncorrected form is on average far less apparent in their corrected forms. The distortion necessarily involved in separate percentage estimation is seen to be greatly magnified if quantitative rather than qualitative inferences are made. As cautioned above, inferences can only be validly drawn on the basis of (2) and (4), the other two approaches are included merely as reference to the loss function involved in using the less general and encompassed models.

We can now return to the above discussion of the cause of the sexual bias observed in participation. If this was indeed merely the result of females evaluating higher reservation wages, then Table 2.5 implies that, controlling for all other characteristics, female reservation wages would be at nearly three times the level evaluated by males. This would not appear to be credible, and indeed implies that barriers to entry are considerable. While disentanglement of the effects of differing reservation wages and barriers to entry is impossible, a belief that females face no barriers to entry is shown to be equivalent to the belief that female reservation wages are three times the level of male reservation wages. Likewise it can be inferred that non-Ivorians also face inferior access. The majority of non-Ivorians are male immigrants from neighboring states, migrating on economic grounds, and a priori

Table 2.5 Percentage Wage Differentials to Equalize Propensities

	(1)	(2)	(3)	(4)
(a) Characteristics				
HEAD	610.17	370.01	599.72	402.71
MARRIED	76.16	50.17	56.79	45.52
FGOVT	74.18	54.44	66.26	53.12
ABIDJ	106.35	84.59	94.32	79.09
SEX	197.03	193.77	189.00	191.84
CEPE	93.94	78.29	71.85	64.18
NAT	68.17	54.50	43.32	36.31
(b) Age (Compared to maximum of quadratic)				
Max (Years)	33.48	33.19	32.60	32.38
Age 20	138.67	143.00	151.54	105.60
Age 30	8.18	5.33	4.01	2.69
Age 40	31.71	26.67	37.42	31.41
Age 50	487.00	322.59	480.19	330.79

they would be expected to have relatively low reservation wages. The most overt demonstration of this is provided by Joshi et al. (1976), who suggest that non-Ivorians face unequal access to the transmission of information. In the wake of *sans-travail* demonstrations, non-Ivorians are generally hired only when the Office de la Main d'Oeuvre is unable to provide Ivorian staff of the type requested, and non-Ivorian Africans are reluctant to enter the labor exchange offices.

Acquisition of education might be expected on balance to increase an individual's reservation wage, whereas the estimated values for P for non-possession of the CEPE diploma are positive and of a large magnitude. There are two possible explanations. First the decision to acquire primary education might be endogenous with the labor force participation decision. However this is not consistent with the lack of significance in the participation equation of possession of the BEPC diploma, taken on average four years later. The more plausible explanation is that in a rationed labor market the CEPE diploma acts as a selection device, and hence the uneducated not only have lower expected wages but also inferior access.

The magnitude of the effect of headship on participation is only explainable in terms of a full theory of the intra-household distribution of economic status and responsibility. However, consider a model in which headship entails the economic responsibility to generate at least a household subsistence level of monetary income. Then if labor market participation was the only method of cash generation, in a household with no other members in employment the value of P implied by lack of headship would be infinite. In a model where headship entailed the power to control the entire household pattern of labor and heads assigned greater weight to their own leisure, the value of P would be expected to be negative. Hence the high value for the headship characteristic is at least consistent with economic responsibility falling mainly on household heads, or with heads acting altruistically in their valuation of the relative marginal utilities of leisure of household members.

A further possible explanation is that there is a degree of endogeneity in headship. If economic status is a determinant of intra-household political status a degree of endogeneity may exist, at least on the grounds that the timing of intergenerational transfer of headship may be influenced by the economic status of the potential successor. Alternatively, combining the competing auspice theory of Caces et al. (1985) with the cooperational conflict models of Sen, greater economic status may encourage participants to opt out of the structure of the original household and form a competing auspice, loosely tied but economically distinct. While the grounds for the endogeneity of marriage are weaker, the same theory of economic responsibility shifting due to possession of a characteristic is supported by the apparent lower reservation wages of married individuals, given that there are few reasons for believing that marriage facilitates overcoming barriers to entry other than through the extension of the potential information network.

Paternal employment history may well help to reduce barriers to entry by informational content or direct contacts, and also parental example may induce a greater inclination for work thus reducing the reservation wage. Both sets of factors support the sign of P for the FGOVT characteristic. Finally the relative scarcity of employment possibilities outside Abidjan promotes a high positive value for ABIDJ, which may possibly be increased by a greater choice of other activities (particularly rural and semi-rural activities) in urban areas outside Abidjan tending to increase reservation wages.

The effects of age are quantified in part (b) of Table 2.5, where the percentage wage increment necessary to induce an equal propensity for participation for individuals aged 20, 30, 40, and 50 compared to an individual with the maximum propensity implied by the age quadratic. The expression for P is a defined above with the distinction that γ_1 is now defined as the difference between the maximum value of the quadratic and the value for the age of the individual under consideration, holding other characteristics constant. The results demonstrate that the combination of reservation wages and unequal access shows a very diverse pattern by age with dramatic patterns of increase for the two extremes of the age scale.

2.1.4 Gender Aspects of Labor Market Activity

In the previous section we discovered significant gender effects in both the process of wage determination, with a premium for females once sample selection was fully incorporated, and also in participation with a lower propensity for labor market activity exhibited by females. We now replicate this methodology and model with a decomposition by gender.

The spatiality of labor force participation by gender is shown in Map 2, IBRD # 22577 for females and Map 3, IBRD # 22578, for males. The market for female labor is non-existent in all but four of the fifty-seven rural clusters in the 1986 CILSS, and in none of these does the participation rate reach 10 percent. Participation rates are seen to vary considerably in urban areas. Of the forty-three urban clusters, ten show no participation, and a further fifteen rates of less than 10 percent, while the rate surpasses 40 percent in three clusters in Abidjan. Map 2 therefore suggests an almost totally urban basis for female labor market participation, with an apparent concentration in Abidjan compared to other urban areas.

Comparison with Map 3 demonstrates that female participation is considerably lower than male participation, as was suggested by the analysis of participation in the previous section. For males the rural market, while truncated, is widespread, and particularly evident in the East Forest area of the country, while relatively high participation rates are observed in the urban clusters, with almost a half of those clusters showing rates above 40 percent.

While there is some presumption that an analysis of the rural market for male labor might be possible, we confine the following analysis to urban areas to focus on the comparative gender issues. With participation and variables defined as in the previous section, this produces samples of 1,286 females and 1,124 males. As suggested above, the participation rates are very disparate—36 percent for males and only 8.6 percent for females.

The aggregated analysis revealed a wage premium for females compared to males with equal wage augmenting characteristics, and it was suggested that the root cause of any perceived or real discrimination against women in the returns to labor would then be based on differential access to such characteristics, and in access to the labor market. The extent of this differential access is shown in Table 2.6, where tabulations are shown by gender for the sample population and also for the subset in waged employment. Part (a) of the table represents the percentage of individuals in the group under consideration possessing a set of educational characteristics, and also average years of education broken down into primary, post-primary (non-technical), and technical.

The distribution of educational achievements and years of schooling across the sample population demonstrates how poor female access is to these characteristics compared to males. The average male receives 62 percent more primary education measured in terms of grades completed and is 102 percent more likely to have achieved possession of the CEPE diploma. The extent of this differential access increases at post-primary level, with males receiving 214 percent more post-primary education in terms of grades, being 237 percent more likely to achieve possession of the BEPC diploma, and 313 percent more likely to achieve baccalaureate or higher qualifications. Inferior access to females is further replicated for technical training and qualifications, although as noted above such measures may be endogenous with respect to the participation decision. The CILSS contains information on self reported reading, writing and numeracy skills. Even at this most basic educational level, while accepting the

Table 2.6 Characteristics of Wage Earners in Relation to Total Population
Percentages, unless noted otherwise

	Population		*Waged*	
	Females	*Males*	*Females*	*Males*
(a) Educational Characteristics:				
CEPE	25.20	50.80	79.10	69.20
BEPC	4.60	15.50	27.30	25.20
HIGHER	1.50	6.20	13.60	11.90
TECH	7.20	15.00	47.30	33.10
Read	39.40	66.90	87.30	80.30
Write	38.50	64.40	87.30	77.80
Numerate	40.10	70.20	90.00	84.00
Grades:				
Primary	2.33	3.77	5.15	4.54
Post Primary	0.63	1.98	3.65	3.34
Years Technical	0.26	0.50	1.55	1.05
(b) Other:				
FGOVT	7.30	7.60	25.50	11.60
MGOVT	0.80	0.70	3.60	1.50
Married	59.20	52.50	63.60	73.80
Age (Years)	31.80	34.70	31.90	35.50
Abidjan	44.30	48.30	71.00	59.00
Ivorian	84.10	78.40	89.10	82.50
Log Hourly Wage	—	—	6.03	6.07

— = not applicable

shortcomings of such self assessed measures, the illiterate and innumerate population is dominated by females.

Moving to the sub-sample of waged individuals provides a further illustration of the danger of making inferences on the basis of a sub-sample without due reference to the sample selection rule. Not only does the clear picture of inferior female access to education fail to emerge, it is reversed. Employed females tend to have more education than employed males, measured in terms of grades or diplomas, with the difference being particularly marked for primary and technical education, and also for the self assessed basic educational indicators. There may be two factors driving such a distortion from the characteristics of the sample population. First, as is evidenced by Table 2.1, there are sections of the labor market where the required skill level is expected to be low, in which there is little or no female participation. Secondly, given the truncation of female labor supply relative to male supported by the results of the previous section, any drawing from the sample population into the waged sample which is based primarily on education, could produce such a configuration of characteristics as that observed.

While part (b) of Table 2.6 is intended primarily for the purpose of reference, and no inferences can be made that will necessarily survive the empirical modeling performed below, a few patterns are worth noting. First, the inference of Maps 2 and 3 that the female labor market is heavily concentrated in Abidjan relative to the male labor market is confirmed. Secondly, there is a very marked pattern in the distribution of FGOVT, with the presumption that females tend to make more use of the informational or access content of paternal employment histories than males. Finally, on average females do receive a lower wage than males, even with their higher average educational characteristics. That the previous section still managed to find a positive wage premium for females within this pattern when separate estimation techniques could not, provides an indication of how powerful the effects of the sample selection adjustment are on this dataset.

We now run the one sector participation of the previous section on two samples, the disaggregation being by sex. The two underlying reduced form equations are presented in Tables A.2 and A.3 of the Appendix, and the wage equations are shown in Table 2.7 for males and Table 2.8 for females, where the nomenclature for models follows that of the previous section. Given the predominance of males in the workforce it is not surprising that the pattern of results shown in Table 2.7 is similar to that of Table 2.3. Males receive some 4.1 percent increase per year due to the combination of age and experience, significant differential effects exist for primary diplomas and years of secondary education, and the selection compensation variable is highly significant. No premia exist in wages on grounds of location or nationality.

The relatively smaller sample size for females might a priori be expected to produce greater sample selection bias than that exhibited by the male sample. However, LAMBDA is insignificant for the probit and logit reduced form generators. An explanation is provided by unequal access, not to the capacity to realize the wage augmentations of characteristics, but to those characteristics themselves. The key lies in education While Table 2.8 can only disentangle the effects of diplomas and years of education for the BEPC diploma, (negatively significant), and grades of secondary education, no educational variables can be discarded as likelihood ratio tests reveal joint significance for the singularly insignificant educational variables. As the purpose of the wage equations is to provide the best measure of expected wages this collinearity is not a problem in itself, and indeed the specification bias resultant in dropping variables to give significance elsewhere is undesirable as it necessarily entails discarding relevant information. In total, while the differential effects of years of education and diplomas achieved may not necessarily be separated, education does affect wages. The differences between the educational characteristics of waged and unwaged females are extreme, very low levels in the unwaged sample and almost universal non-zero levels in the waged sample. The greater the extent to which wage affecting variables effectively partition the two samples, the less will be the bias caused by sample selection.

The wage premium found for females in the previous section by the use of a simple gender dummy variable, is only a valid conclusion if all other characteristics were not affected by gender The disaggregated wage equations help to provide a better measure of the direction and degree of any discrimination in wages. Imputing the mean characteristics of waged males into the female wage equation produces an estimated wage 2.57 percent lower than that generated for the same characteristics by the male equation. However, imputing the mean characteristics of waged females gives a wage 0.74 percent higher in the female wage equation than the male. If we follow Greenhalgh (1980) in regarding the average of the two as the best measure of discrimination, then discrimination against females in wages is only of the order of 0.92 percent. Imputing the mean population characteristics, for males the estimated wage from the female equation is 2.49 percent lower than that given by the male equation, but for mean female population char-

Table 2.7 Côte d'Ivoire 1986 Determinants of Male Wages and Employment
Parameters of Wage Regression

	(1) and (3)	(2)	(4)
CONSTANT	3.28167 (17.042)**	4.02739 (16.788)**	4.04446 (16.581)**
AGEY	0.04766 (11.913)**	0.04127 (10.337)**	0.04127 (10.222)**
CEPE	0.46389 (2.726)**	0.35719 (2.196)**	0.35017 (2.137)**
BEPC	0.02753 (0.226)	0.05795 (0.484)	0.05595 (0.471)
HIGHER	0.07603 (0.326)	0.17996 (0.780)	0.17991 (0.787)
ABIDJ	0.01428 (0.195)	-0.03976 (-0.558)	-0.04086 (-0.573)
GRSP	0.03952 (1.461)	0.03362 (1.321)	0.03427 (1.332)
GRS	0.14972 (4.515)**	0.12389 (3.840)**	0.12309 (3.814)**
GRS2	0.20661 (3.471)**	0.15854 (2.663)**	0.16042 (2.722)**
YRU	0.08605 (1.907)*	0.06876 (1.520)	0.06937 (1.557)
NAT	0.09992 (0.974)	0.05123 (0.526)	0.04321 (0.440)
LAMBDA	—	-0.40913 (-4.814)**	-0.42087 (-4.749)**
R-Squared	0.58547	0.60668	0.60691
Log-Likelihood	-424.70001	-407.98001	-407.87000

* = significant at 10 percent
** = significant at 5 percent
— = not applicable

Table 2.8 Côte d'Ivoire 1986 Determinants of Female Wages and Employment
Parameters of Wage Regression

	(1) and (3)	(2)	(4)
CONSTANT	3.69173 (9.283)**	3.96211 (4.506)**	4.60526 (5.121)**
AGEY	0.03808 (4.540)**	0.03554 (3.257)**	0.02938 (2.647)**
CEPE	0.27496 (0.869)	0.25932 (0.855)	0.22230 (0.739)
BEPC	-0.33340 (-1.659)*	-0.34263 (-1.779)*	-0.36464 (-1.896)*
HIGHER	0.36764 (0.912)	0.35919 (0.936)	0.34370 (0.896)
ABIDJ	0.24708 (1.703)*	0.23037 (1.[illegible]8)	0.19373 (1.339)
GRSP	0.02475 (0.420)	[illegible]96 (0.217)	-0.01071 (-0.167)
GRS	0.22171 (3.962)**	0.21305 (3.622)**	0.18861 (3.131)**
GRS2	0.21980 (2.205)**	0.21679 (2.281)**	0.21121 (2.230)**
YRU	0.07235 (1.008)	0.06713 (0.960)	0.05618 (0.800)
NAT	-0.08635 (-0.419)	-0.08997 (-0.459)	-0.10005 (-0.510)
LAMBDA	—	-0.07011 (-0.340)	-0.24658 (-1.120)
R-Squared	0.62719	0.62759	0.63141
Log-Likelihood	-102.88000	-96.53100	-95.96400

* = significant at 10 percent
** = significant at 5 percent
— = not applicable

acteristics the female equation produces a wage 5.01 percent higher, an average discrimination in favor of females of 1.26 percent.

Therefore, while the large and significant discrimination in favor of females found in the last section is now seen as the result of a too highly aggregated wage equation, we can still stress that there is no evidence of discrimination against women in the process of wage determination in the Côte d'Ivoire. In fact, the Ivorian wage determination system appears to be far more equitable across genders than is implied by empirical research in American labor markets (Cain (1986)).

Tables 2.9 and 2.10 show the parameters of the participation decision resultant on using the expected wages derived above for respectively males and females, where again the nomenclature for models follows that employed in the previous section. For both genders there is a significant wage response, a quadratic age response and also headship implies a greater propensity to participate. There is no effect attached to maternal employment history, and the strong positive effect of marriage found in the previous section is lost in the disaggregation. However gender differences emerge in the other determinants of participation used in the estimation.

The informational content and access improving potential of paternal employment history represented by FGOVT is shown to be confined to females. Given that informational paucity is unlikely to follow gender divisions, this implies that paternal employment has mainly access effects, perhaps combined with some

Table 2.9 Parameters of Male Participation Decision

	(1)	(2)	(3)	(4)
CONSTANT	-12.10840 (-8.632)**	-11.45340 (-9.566)**	-6.85640 (-8.886)**	-6.33872 (-9.863)**
AGEY	0.39175 (6.662)**	0.34830 (5.091)**	0.21433 (6.787)**	0.19163 (5.109)**
AGE2	-0.00585 (-8.357)**	-0.00523 (-6.617)**	-0.00327 (-8.782)**	-0.00291 (-6.812)**
HEAD	1.61258 (5.868)**	1.44663 (4.946)**	0.93239 (5.787)**	0.84453 (4.944)**
MARRIED	0.20104 (0.643)	0.18397 (0.588)	0.12513 (0.700)	0.11458 (0.641)
FGOVT	0.21507 (0.698)	0.19377 (0.627)	0.09455 (0.535)	0.09244 (0.523)
MGOVT	-0.15740 (-0.155)	-0.13055 (-0.129)	-0.00525 (-0.009)	0.01237 (0.020)
CHILD	0.52482 (1.863)*	0.47460 (1.680)*	0.31669 (1.903)*	0.28673 (1.722)*
ABIDJ	0.71139 (4.193)**	0.67551 (3.[illegible]33)**	0.38381 (3.995)**	0.36579 (3.761)**
CEPE	0.28771 (0.785)	0.28494 (0.785)	0.10720 (0.515)	0.14603 (0.707)
BEPC	-0.14[illegible]41 (-0.454)	-0.14507 (-0.462)	-0.10611 (-0.576)	-0.07622 (-0.419)
NAT	0.64111 (2.976)**	0.60795 (2.769)**	0.31150 (2.560)**	0.29987 (2.418)**
EXPW	0.62304 (2.014)**	0.66693 (2.048)**	0.39845 (2.239)**	0.37994 (2.044)**
Log-Likelihood	-480.23001	-480.17001	-483.17999	-483.63000

* = significant at 10 percent
** = significant at 5 percent

	Predicted							
	0	1	0	1	0	1	0	1
Actual	603	76	602	77	606	73	603	76
	128	277	127	278	130	275	126	279

lowering of the reservation wage. In other words, if males have relatively free access to the labor market while barriers exist for females, the access improvement occasioned by paternal contacts is a more valuable commodity for females.

The effect of children is to lower female participation, but not significantly, while male participation rates are raised. For females, this implies that the household structure is able to absorb the requirements of child care without necessarily causing career breaks, while for males the presence of children magnifies the economic responsibility they are allotted within the household. While fertility and female labor force participation are often found to be inversely related in industrial economies (Papanek (1976)), our results are fully consistent with previous studies in less developed economies (surveyed in Chant (1989)), where no convincing relationship has been found.

Participation rates are significantly higher for males domiciled in Abidjan, and for Ivorians, while no such pattern emerges for females. The effects of education are however the greatest difference between genders. For males, education has no effect on participation other than the effect working through wages. How-

Table 2.10 Parameters of Female Participation Decision

	(1)	(2)	(3)	(4)
CONSTANT	-18.43070 (-8.096)**	-17.92020 (-8.261)**	-8.89480 (-8.258)**	-8.76419 (-8.293)**
AGEY	0.47725 (4.600)**	0.43031 (4.066)**	0.20713 (4.194)**	0.20038 (4.015)**
AGE2	-0.00660 (-4.655)**	-0.00592 (-4.151)**	-0.00299 (-4.468)**	-0.00289 (-4.291)**
HEAD	1.5023[illegible] (3.401)[illegible]	1.35071 (3.045)**	0.82389 (3.397)**	0.79915 (3.296)**
MARRIED	0.45794 (1.375)	0.40083 (1.199)	0.21982 (1.252)	0.21204 (1.208)
FGOVT	0.80522 (2.196)**	0.72673 (1.965)**	0.51893 (2.707)**	0.50470 (2.630)**
MGOVT	0.36360 (0.356)	0.32990 (0.322)	0.11597 (0.214)	0.11568 (0.214)
CHILD	-0.41002 (-1.300)	-0.36393 (-1.149)	-0.23301 (-1.441)	-0.22409 (-1.385)
ABIDJ	0.28892 (0.989)	0.27955 (0.955)	0.22137 (1.476)	0.22111 (1.469)
CEPE	1.47911 (3.209)**	1.49643 (3.301)**	0.62253 (2.615)**	0.63533 (2.680)**
BEPC	0.[illegible]7540 (2.432)**	0.92746 (2.284)**	0.53247 (2.302)**	0.53052 (2.287)**
NAT	0.69576 (1.520)	0.66513 (1.453)	0.20624 (1.016)	0.20072 (0.990)
EXPW	1.16534 (3.274)**	1.22344 (3.333)**	0.65455 (3.280)**	0.64964 (3.234)**
Log-Likelihood	-225.28000	-224.99001	-228.99001	-2[illegible]9.16000

* = significant at 10 percent
** = significant at 5 percent

	Predicted							
	0	1	0	1	0	1	0	1
Actual	1114 63	16 47	1114 63	16 47	1116 67	14 43	1116 67	14 43

ever the possession of both diplomas significantly increases the rate of female participation. Hence either education leads to significant reductions in the reservation wage, an effect we quantify below, or, more likely, education acts as the selection device for female labor in the presence of barriers to entry. This has the further implication that even though no significant sample selection bias was found in the female wage equation, the results of such an equation can still not be generalized to the general female population. In other words, in the presence of barriers to entry which are reduced by education, the major return to education may not be in its mapping to income through wages, but rather in the ability of education to facilitate the realization of the capacity to utilise that mapping.

We use the procedure outlined in the previous section to calculate the necessary percentage wage augmentations needed to equalize the propensity of an individual without a characteristic to that of an individual with the characteristic under consideration and otherwise equal attributes, for the results of Tables 2.9 and 2.10. These are shown in Table 2.11, where the differentials are calculated for the sample selection

Table 2.11 Percentage Wage Differentials to Equalize Propensities

	Logit		*Probit*	
	Males	*Females*	*Males*	*Females*
(a) Characteristics:				
HEAD	775.03	201.63	823.31	242.17
MARRIED	31.76*	38.77*	35.29*	38.59*
FGOVT	33.71*	81.12	27.54*	117.47
CHILD	103.73	-25.73*	112.70	-27.17*
ABIDJ	175.35	25.67*	112.70	40.55*
CEPE	53.30*	239.78	46.87*	165.91
BEPC	-19.55*	113.42	-18.18*	126.29
NAT	148.82	72.23*	120.18	36.20*
(b) Age: (Compared to maximum of quadratic)				
Maximum (Years)	33.30	36.34	32.93	34.67
Age 20	300.20	264.49	259.55	160.41
Age 30	8.90	21.49	6.77	10.18
Age 40	42.22	6.68	46.70	13[illegible]
Age 50	791.26	146.54	832.52	184.5[illegible]

* Denotes cases where variable is insignificant in the relevant participation equation.

corrected logit and probit versions by gender. As noted above these figures can be interpreted as measures of the affect of the characteristics on the reservation wage, in the absence of barriers of entry.

The strong effect of education on female participation is now quantified, the joint effects of CEPE and BEPC, beyond effects working through wages, imply that if possession of these characteristics does not facilitate entry into the labor market by overcoming barriers to entry, then reservation wages are 78 percent lower for females with the diplomas compared to those without both. The differentials also show a considerably flatter pattern by age for females.

2.1.5 Non-Wage Benefits and the Status of Employment

We now give a short account of the gender decomposition of characteristics reflecting non-wage benefits and the status of employment. A breakdown of characteristics by gender, showing percentages unless stated, is shown in Table 2.12, where the most notable feature is in fact the uniformity by gender.

Individual affiliation to a union is information unavailable in the CILSS, but the presence of a union at the place of work is recorded. Hence this measure of unionization at least proxies access to unionization or access to the free-riding benefits or otherwise of unionization. Table 2.12 demonstrates that there is no difference by gender in access to unionization. We attempt to capture the legal protection in employment with whether employment is subject to the Ivorian minimum wage laws, (not necessarily that a minimum wage is received), and whether a formal wage contract has been signed. Again, women do not appear to have a lower legal status or access to legal protection, and indeed are more likely to be contracted, probably reflecting the lack of a female equivalent to the more casual and likely to be uncontracted male laborer market. While there are fewer female receipts of non-wage benefits such as paid sick leave and holidays, pensions and subsidized medical care, these differentials are considerably less important than the difference in receipt of social security. Thus the intra-labor market discrimination appears to be less than the discrimination in the official benefit system. Further, such non-wage benefits may in part be functions of years of tenure, some 30 percent higher for the average male, and thus are not necessarily symptomatic of discrimination by gender alone.

Table 2.12 also throws further light on the differential effect by gender of parental employment histories described in the previous sections. Only 3.4 percent of males and 3.6 percent of females have a relative as an employer or director at work, and even fewer are following the employment histories of their parents directly. Hence the effect of paternal employment history would again appear to work through the use of a generalized contact system.

The percentage of waged individuals using their own tools at work is presented as a check of whether there might be a prevalent underlying capital restraint on participation, and a figure of 9.1 percent for both males and females suggests that this is unlikely to be a significant factor.

We have not included hours of work as an endogenously determined variable in the above analysis as the distribution of hours worked suggests that the participation decision in the vast majority of choices is

Table 2.12 Gender Differences in Access to Unions and Legal Protection

	Females	Males
Union at Work	60.0	61.2
Wages Subject to SMIG	65.5	65.1
Signed Contract, Spec. Wages	45.5	40.7
Paid Holidays	68.2	72.3
Paid Sick Leave	55.4	63.2
Retirement Pension	48.2	55.3
Free or Subsidized Medical Care	49.1	55.0
Receive Social Security	25.4	36.3
Use Own Tools at Work	9.1	9.1
Either or Both Parents Did Job	2.7	2.0
Employer a Relative	3.6	3.4
Years of Employment	7.6	10.3
Hours per Day	7.5	8.1
Days per Week	5.7	5.9

a simple minimum hours or more, or nothing choice. This is illustrated by the high mean values of hours worked per week, 42.8 for females and 47.8 for males.

2.2 Sectoral Choice and Labor Market Participation

2.2.1 *Introduction*

The model of labor force participation presented in Section 2.1.2 made no reference to the sectoral composition of the labor market, and the wage function estimated was therefore an aggregate function. Implicitly, therefore, individuals were modelled as reacting to the observation of the average wage commanded by their characteristics, and to an average disutility of work. We now extend that analysis to cases in which individuals perceive segmentations within the labor market and their participation decision is simultaneous with sectoral choice, affected by the variety of wage offers open to them and to differences in their evaluation of the disutilities of work across the sectors.

The focus of research involving endogenous sectoral choice has fallen mainly on differences in returns between sectors, and thus on the estimation of the sectoral wage functions. In particular when corrections are made for sample selection bias, they have been made with reference only to workers in other sectors, and thus all results obtained are conditional upon the characteristics of those in employment. In other words, studies that tackle the endogeneity of sectoral choice avoided by single equation methods using sectoral dummies, have taken the participation decision as given. The bulk of these studies have concentrated on either union versus non-union wage differentials e.g. Robinson and Tomes (1984), Robinson (1989), or on public sector versus private sector differentials, e.g. van der Gaag and Vijverberg (1988) who estimate a switching regression model on data drawn from the 1985 cycle of the CILSS. In Blank (1985), the focus is on sectoral choice between public and private sector, although while wages affect the choice no adjustment is made for sample selection bias. A more general approach has been provided by Gyourko and Tracy (1987), who allow for endogenous choice for public sector and unionized status, although, as we argue below, their treatment of the role of wages in this decision is perverse. Grootaert (1988) adopts an approach similar to our own in that he models a sequential decision process in which participation in the labor market is treated as a decision prior to that of the choice of sector.

The remainder of Section 2.2 is structured as follows. In Section 2.2.2 we utilize single equation methods that make no reference to either the participation decision or the endogeneity of sectoral choice to provide an overview of the sectoral decomposition of the Ivorian labor market. Section 2.2.3 provides a methodology that takes both of these factors into consideration, and this model is estimated in Section 2.2.4 for a labor market comprising a public and a private sector. In Section 2.2.5 non-homogeneity of the private sector is introduced with that sector disaggregated on the basis of the access employees have to unionization. Finally, in Section 2.2.6 the trace between two cycles of the CILSS is analyzed.

2.2.2 *A Single Equation Non-Adjusted Model*

To provide an overview of the issues involved in sectoral decompositions of the Ivorian labor market, we utilize the single equation techniques pioneered by Smith (1976a, 1976b, and 1977), and developed by inter alia Mellow (1982) and Asher and Popkin (1984). To generalize the developed form of such single equation methods, assume we have J sectors in the disaggregated labor market, denoted by the dummy variables D_1 to D_J. There are K breakdown dummy variables X_K, as well as a set of L other determinants of wages in the vector of characteristics Z. In this formulation the set of determinants may include employment specific characteristics as no reference to, or compensation for, a base group of non participating individuals is made. However, sectoral specific variables, other than terms involving the sectoral dummies, are of course ineligible determinants of wages. The method then involves the estimation by ordinary least squares of the following equation, assuming the J^{th} sectoral dummy is used as a base;

Table 2.13 Relative Wages as Implied by One Equation Model

	Private Sector	*Public Sector*
Female		
Non-Educated		
Non-Ivorian		
Non-Union	56.019	63.477*
Unionized	80.349	64.228
Ivorian		
Non-Union	62.687***	61.331***
Unionized	90.172	62.057***
Educated		
Non-Ivorian		
Non-Union	74.748***	90.715
Unionized	107.213	91.789
Ivorian		
Non-Union	83.866*	87.649
Unionized	120.320*	88.686
Male		
Non-Educated		
Non-Ivorian		
Non-Union	52.830***	71.757
Unionized	75.776*	72.422
Ivorian		
Non-Union	59.289***	69.156**
Unionized	85.039	69.974***
Educated		
Non-Ivorian		
Non-Union	70.494***	102.288
Unionized	101.111	103.499
Ivorian		
Non-Union	79.111***	98.831
Unionized	113.471*	100.000

*** = significantly different from base group at 5 percent
** = significantly different from base group at 10 percent
* = significantly different from base group at 20 percent

$$\ln w_i = \alpha + \sum_{l=1}^{L} \beta_l Z_{il} + \sum_{j=1}^{J-1} \gamma_j D_{ij} +$$

$$\sum_{j=1}^{J} \sum_{k=1}^{K} \delta_{jk} [D_{ij} \times X_{ik}] + U_i \qquad (i = \ , ..., n)$$

Hence incremental scores can be estimated for the $2JK$ combinations of sectors and the vector of breakdown variables, $\overline{X}$, as:

$$S[D_j,\overline{X}] = \gamma_j + \sum_{k=1}^{K} \delta_{jk} X_k \qquad (j = 1, ..., J \text{ and } \supset \overline{X})$$

Normalizing on a particular sector and vector of characteristics, say D^* and $\overline{X}^*$, the wage for all possible combinations of sectors and characteristics expressed in percentage terms of the base group wage is given by:

$$100e^{[S[D_j,\overline{X}] - S[D^*,\overline{X}^*]]}$$

Tests of the significance of wage differentials are therefore tests of the equality of the incremental scores, and therefore may be conducted as F tests on the coefficients of the underlying single equation. For example a test of wage discrimination between a group possessing the $k1$ and $k2$ characteristics in $\overline{X}$ in sector $j1$ and a group possessing the $k3$ characteristic in sector $j2$ reduces to simply a test of the hypothesis:

$$\gamma_{j1} + \delta_{j1k1} + \delta_{j2k2} - \gamma_{j2} - \delta_{j2k3} = 0.$$

We estimate this model using the breakdown variables sex, education (namely possession of the CEPE diploma), nationality and unionization, with two sectors defined by private and public employment. With no information on the union affiliation of the individual, unionization is defined as the presence of a union at the workplace. The components of Z are those explanatory variables defined in the previous chapter, barring SEX, CEPE, and NAT [illegible]ch as breakdown variables enter in the interactive [illegible]s with the sectoral dummies. In addition we use the variable YEARS and its square, defined as years of experience, and TECH and YRST, defined in the previous chapter but excluded from that exercise due to their potential endogeneity.

Unionization is in fact compulsory for Ivorian public sector workers, yet 23.5 percent of public sector workers said there was no union at their workplace. The scale of this response implies it is more than mere miscoding, and indicates the presence of areas within the public sector where the effect of labor associations is either invisible or neglible even though workers are nominally unionized. It may be indicative of a relatively casual section of public sector workers, and would help explain the pattern of public sector recruitment observed in Section 2.2.6.

The results of this regression and the means and standard deviations of the variables over the sample of waged individuals are shown in the Appendix Table A.4. To illustrate the implied relative wages we have taken a base group of male public sector educated and unionized individuals and followed the

Table 2.14 Characteristics of Workers According to Sector and Unionization

	Private Sector		*Public Sector*	
	Non-Unionized	*Unionized*	*Non-Unionized*	*Unionized*
Sex (% female)	25.90	14.20	1[illegible].90	24.90
Ivorian	69.80	74.30	93.60	96.00
Has CEPE diploma	47.50	67.30	77.40	86.10
Has BEPC diploma	10.80	21.20	29.00	37.30
Years of Education	5.20	7.40	7.80	10.30
Age (years)	31.70	35.60	36.40	36.00
Log hourly wage	5.30	6.20	6.11	6.50
Hours per day	7.96	8.20	7.91	7.77
Days per week	5.90	5.80	5.81	5.82
Subject to SMIG	31.90	73.40	54.80	87.10
Has signed contract	21.80	49.70	21.00	58.70
Paid holidays	31.90	84.10	79.00	90.00
Paid sick leave	27.30	71.70	72.60	77.10
Retirement pension	23.00	61.10	71.00	65.60
Free/Subsidized medical care	16.80	49.60	64.50	79.60
Receive Social Security	9.30	45.10	33.90	44.70
Use own tools	6.80	3.50	8.10	14.90
Years of employment	5.50	9.60	9.50	10.10
Number in category	139.00	113.00	62.00	201.00

procedure outlined above. The resultant relative wages are shown in Table 2.13.

The general pattern demonstrated in Table 2.13 combined with exhaustive tests of significance of the differentials against all the other fifteen possible base groups (not shown) is how comparison between groups is complicated by unionization. Wages are significantly increased by unionization in the private sector but not the public sector. The combination of factors is such that male public sector wages are significantly higher than private non-unionized wages in all categories, bar non-educated Ivorians, but not different from private unionized wages, while for females public sector wages are significantly lower than private unionized wages in all categories bar educated non-Ivorians and not different from private non-unionized wages. Hence even though there is no significant discrimination in wages by gender in either sector, a pattern emerges of female public sector wages following a lower private sector counterpart. Hence if the affect of unionization in the private sector is simply a result of the difference between a formal and an informal private sector, we have the implication of formal sector wages being higher in the private sector than the public for females and no different for males, even though there is no significant intra-sector discrimination in wages.

The above framework is, however, highly restrictive. Returns to the characteristics in the sector Z are constrained to be equal across sectors, and no account is made of the potential bias arising from self-selection. In the next section, we provide a methodology for overcoming such shortcomings, and examine the determinants of wages and participation in multi-sectoral models. We first present descriptive statistics on the public-private, union-non-union decomposition used above. These are presented in Table 2.14, where figures are percentages unless stated.

Table 2.14 is highly suggestive of marked intersectoral differences in access, and also in the non-wage benefits of employment. The bimodal nature of the female labor force is shown by the concentration of females in both the lowest paid and least educated sector, public unionized. Non-Ivorians are heavily concentrated in the private sector, but other than nationality there is considerable heterogeneity between the characteristics of the private non-unionized and the private unionized sector. The non-unionized sector draws more heavily on females and the younger and relatively less educated, and provides strikingly fewer non-wage benefits with little contracted labor or protection under wage laws. The sharpest difference is shown by the receipts of social security with 91 percent of the non-unionized sector being unwilling or institutionally unable to receive benefit compared to 55 percent in the unionized sector.

It appears therefore that the presence of a union in the workplace provides a possible disaggregation of the private sector into a formal and an informal sector, and gives considerable evidence of an heterogenous sector.

The differences are less marked within the public sector where the unionization rate, noting that this is defined by the presence of a union at the workplace, rather than by participation, is 76 percent compared to 45 percent in the private sector. Unionization thus defined is associated with relatively higher levels of educational attainment, greater recourse to protection under contracts and wage laws and an absolute wage which is on average 48 percent higher than for non-unionized public sector workers. The difference in non-wage benefits is considerably less marked than in the private sector, indeed the non-unionized workforce is shown to have greater access to pension rights than the unionized workforce.

2.2.3 *Estimation of Endogenous Sectoral Choice*

We now provide a methodology to model multi-sectoral participation decisions with a modification of the approach used in the previous chapter. Assume that the labor market has J sectors. Individuals then evaluate a series of J reservation wages as a function of a vector of observable individual characteristics, X, as well as unobservable characteristics ξ. Hence:

$$W_{ij}^{R} = W^{R}\,[X_i, \xi_i] \qquad (j = 1, \ldots, J)$$

Reservation wages are allowed to vary between sectors for several reasons. The disutility of work and the perceived status from employment may vary, and individuals may place differing values on any non-wage benefits derived from employment by sector. Likewise if participation is affected by capitalized rather than contemporaneous wages and wage differentials, and earnings and tenure profiles vary by sector, we assume that the reservation wages pick up differences in the perceived capitalization. We return to this last point at greater length in the next section.

In total the above factors allowing the reservation wage to vary by sector imply that participation decisions are not simply based on the highest wage offer available to the individual in question. An individual may quite rationally choose to participate in a sector which provides a lower wage than that offered in another sector.

We observe a market wage in each of the J sectors, a function of a vector of observable characteristics Z, and also a set of unobservable characteristics incorporated into an error term. We define δ_j as the difference between the reservation wage and the market wage in a particular sector j. Then participation in that sector is observed for individual i if and only if:

$$\delta_{ij} < 0 \text{ and } \delta_{ij} < \min_{k \neq j} \delta_{ik}$$

However, with the reservation wages being unobservable, we can only detail the variable Y which takes $J + 1$ values, say $Y = 0$ for non-participants and $Y = j$ for participants in sector j, $(j = 1, \ldots, J)$. We can proceed as follows. The differentials δ_j are a function of Z, market wages and an error term which we denote as η_i. With δ_0 set to zero define ε_j as:

$$\varepsilon_{ij} = \min_{k \neq j} \delta_{ik} + \eta_j$$

where

$$\delta_{ij} = \sum_{m=1}^{M} \gamma_{mj} X_{ij} + \gamma_{(m+1)j} W_j + \eta_j$$

assuming the vector Z has m elements. Hence substitution yields that $Y = j$ if, and only if,

$$\varepsilon_{ij} = \sum_{m=1}^{M} \gamma_{mj} X_{ij} + \gamma_{(m+1)j} W_{ij}$$

The model can now follow the analysis of Lee (1983). The form of the likelihood function is directly determined by the distribution of the residuals η_{ij} defined above. A multinominal logit form of the selection equation is derived by assuming that these residuals η_{ij} are identically distributed and also independent following a type I extreme value distribution with their cumulative distributions up to an arbitrary constant k being given by

$$F[\eta_j < k] = e^{[-e^{-k}]}.$$

Then the probability that

$$\varepsilon_{ij} < \sum_{m=1}^{M} \gamma_{mj} X_{ij} + \gamma_{(m+1)j} W_{ij},$$

i.e. that individual i participates in sector j, is given by a multinominal logit formulation as is demonstrated by Domencich and McFadden (1975).

As in the previous section we now have a full likelihood function which is theoretically tractable, but again involves many computations of the cumulative normal of a function involving inverse normals evaluated at conditional probabilities. However a two-stage method yielding consistent estimates is available, although even this proves to be non-trivial as the covariance matrix yielded by OLS estimations of the wage functions including the sample selection correction variables proves to give incorrect standard errors and must be adjusted. This two-stage procedure, or in fact three stage procedure as we investigate the determinants of participation as well as those of wages, involves the following steps. First, the reduced form participation equation is estimated as a multinominal logit, with Y modelled as a function of the union of the vectors X and Z. From this we derive the J predicted probabilities for each individual, P_{ij}. Then as shown by Lee, the relevant j wage regressions are given by

$$\ln W_{ij} = \sum_{k=1}^{K} \beta_{jk} Z_{ij} + \sigma_j \rho_j \lambda_{ij} + U_{ij}$$

where

$$\lambda_{ij} = \frac{\phi[\Phi^{-1}[P_{ij}]]}{P_{ij}},$$

assuming the vector Z has K members and where σ_j is the standard error of the wage equation for the jth sector, ρ_j the correlation coefficient between the errors of the relevant wage equation and the participation, and ϕ and Φ are respectively the density and distribution function of the standard normal.

The approach so far yield consistent estimates for the β_{jk} and the $\sigma_j \rho_j$. However, the standard errors given by OLS are not correct, and the necessary correction is demonstrated by Greene (1981). Finally the participation equation is estimated as multinominal logit using the vector X and the estimated wages. The following two sections employ this procedure to estimate first a bi-sectoral and then a tri-sectoral model of the Ivorian labor market using the same data as in the previous chapter.

Finally it should be noted that the sectoral demand for labor functions are implicitly assumed to enter into the selection equation, which therefore is seen as the summarization of two processes. There is the decision of individuals to seek employment in a particular sector and the decision by employers to draw the individuals from the queue of those wishing to be employed. Thus, as in the previous chapter, it is impossible to disentangle the role of the determinants of participation in reducing reservation wages compared to their effect on barriers to entry. The sectoral demand for labor functions may be explicitly incorporated even though we cannot directly observe the employers' decision not to draw individuals from a queue, as is the case in the CILSS, following Abowd and Farber (1982).

However the explicit incorporation of labor demand decisions poses serious problems of logical consistency and identification in a model involving non-participants, even if disappointed applicants are observed. The relevant wage variable for the participation decision in a particular sector should then be determined not only by the sectoral wages, but also as a function of the probability of gaining entry to the sector. Identification of the reduced form equation will then generally imply the need for a restriction that the covariance between the errors of the individual choice and the employer choice equations be zero. Removal of this restriction involves the assumption that individual choice behavior is not affected by relative changes in the probability of employment between sectors.

2.2.4 *A Bi-Sectoral Model of the Ivorian Labor Market*

We can now estimate the model outlined in the previous section on the 1986 CILSS data, using a framework where individuals face the choices of non-participation, public sector or private sector employment. For estimation of the two wage functions we use the explanatory variables employed and defined in the previous chapter, plus the selection variable derived from the underlying multinomial logit reduced form equation. However, some discussion is necessary to justify the choice of determinants of participation and sectoral choice.

It is customary in a participation equation to include expected wages as a determinant, and also in sectoral choice models to include expected wage differentials. However Gyourko and Tracy (1987) depart from this consensus and omit differentials in their four sector model for those observed to be participants. This is a far more radical departure than they signal it to be, and deserves some elaboration. The Gyourko and Tracy argument is that first, worker choice is a function of the capitalization of such differentials, and if tenure/earnings profiles vary between sectors current wage differentials may be poor proxies. Their second argument is that as using wage differentials implies that factors influencing wages must enter into the reduced form equation, leading to cases of spurious correlation, for which they give an example of the consequences of using a percentage unionization rate as an explanatory variable in a model explaining unionization.

We consider these points in ascending order of importance. Their second point is minor, such variables as percentage unionization rates simply cannot be used in the model they propose, or equivalents in more general models. If beyond that they are implying that correlations occurring in reduced form equations render the results uninterpretable, then the answer is of course, the reduced form is only used to generate selection variables to estimate wages which are then used in a structural explanatory model.

Gyourko and Tracy's first point is also flawed. Their own response to their argument is to drop wage terms entirely from their model of sectoral choice. The illogicality of this is seen by turning their argument around. Suppose contemporaneous wage differentials are a good proxy for capitalized differentials. Then, given their set of explanatory variables for choice includes some variables in their wage equations and excludes others, all they estimate is a reduced form equation prone to all the problems they hope to avoid by excluding differentials, and beyond that a reduced form which is also misspecified. Without the endogenous modeling of sectoral job loss functions, which in itself is fraught with difficulties as only pure involuntary retrenchment should be included, there is only one sensible solution to the problem they highlight. Wage levels and differentials must be included, and this approach can be justified for all possible scenarios. If workers do respond to contemporaneous differentials rather than capitalized there is clearly no problem, nor is there if there is a strong correlation between the two. If there is not, and this is course untestable, then the Gyourko and Tracy argument is that workers who participate may rationally decide to enter a sector where their contemporaneous wage is lower than in some other sector. However, given that the underlying reservation wages in our model are allowed to vary according to which pairwise comparisons to non-participation are being made, the problem is obviated as this situation can still occur.

Specifically, as long as the determinants of tenure do not include any variable left out of the choice equation, then any differences between capitalized and contemporaneous differentials is picked up in the differential reservation wages, which of course may differ due to other reasons than tenure profiles. Hence even with the implicit Gyourko and Tracy assumptions of zero expected correlation between the two types of wage differentials when agents do not respond at all to contemporaneous differentials, dropping wage differentials from the model is an inferior and encompassed strategy.

Gyourko and Tracy also exclude age from their model on the grounds that it should only be included if agents show a switch in behavior over time. This is testable by using age as a determinant, and in our model involving non-participation, we do, from the results of the previous chapter, expect such a shift between non-participation and participation in some sector, although there is no a priori expectation of the inter-sectoral pattern. In total we use the same explanatory variables in the choice equation as in the one sector model of the previous chapter, plus expected wage variables. As decisions are a function of differentials and also levels, as non-participation is a choice, we use two variables, PUBWAGE representing the expected logarithmic public sector wage, and DIFF being the difference between the expected logarithmic private and public sector wages. Given the high correlation that might be expected between the two measures we acknowledge that while completely compensating for wage effects decomposition of the two effects may not necessarily be possible.

The results of the two sector wage determination exercise are shown in Table 2.15, where the OLS results are also included for reference. The selection variable, LAMBDA, is insignificant for the private sector but significant for the public sector. Thus there is no significant difference between the wage earned by an average representative of those employed in the private sector and an individual drawn from the population at random, while there is a significant 27 percent difference for a representative of the public sector over the randomly drawn individual. This figure is derived by multiplying the coefficient LAMBDA in the public sector wage regression by the average value of LAMBDA for workers in that sector. There is no pattern by gender by either sector, but location plays an important role in the private sector where wages are 23 percent higher in Abidjan, compared to the insignificantly different from zero premium of 6.9 percent for other urban areas in the public sector. The lack of significance of gender in either equation implies that the significant premium for females found in the aggregated wage equation of the previous section may simply be a sectoral composition effect. It also agrees with the general result of the simple one equation Smith model presented in Section 2.2.2, which found higher wages for females in the private sector and for males in the public sector, as do the OLS equations in Table 2.15. The corrected version produces the same result in the private sector but the reverse, while still insignificant in the public sector. Taking these results with those of the one sector model the hypothesis of wage discrimination against females is clearly disproved. There is also a pattern by nationality, a 33.8 percent premium for non-Ivorians in the public sector and a 19.1 percent premium for Ivorians in the private sector, and while very close neither effect is significantly different from zero at the ten percent level of significance.

The variability of wages due to age and education is shown to be greater in the private sector. The combined effect of age and experience is to increase wages

Table 2.15 Wage Determination in Bi-Sectoral Choice Model

	OLS		Two-Stage	
	Private Sector	*Public Sector*	*Private Sector*	*Public Sector*
CONSTANT	2.98619	3.99052	3.40770	5.48573
	(12.020)**	(12.716)**	(8.661)**	(8.304)**
AGEY	0.05124	0.03491	0.05068	0.02199
	(9.616)**	(6.615)**	(9.502)**	(3.039)**
CEPE	0.35506	0.23348	0.33122	0.09397
	(1.830)*	(0.890)	(1.703)*	(0.355)
BEPC	-0.12168	-0.01116	-0.12382	-0.00780
	(-0.580)	(-0.099)	(-0.591)	(-0.070)
HIGHER	-0.17326	0.15530	-0.13699	0.1985[illegible]
	(-0.397)	(0.722)	(-0.314)	(0.930)
ABIDJ	0.32482	-0.16510	0.23494	-0.06900
	(3.096)**	(-2.005)**	(1.905)*	(-0.769)
SEX	-0.03397	0.07395	-0.14761	-0.01136
	(-0.287)	(0.766)	(-1.025)	(-0.112)
GRSP	0.02633	0.09046	0.02388	0.06083
	(0.886)	(1.986)**	(0.803)	(1.310)
GRS	0.19028	0.11443	0.18372	0.07858
	(4.162)**	(3.211)**	(4.004)**	(2.072)**
GRS2	0.21596	0.20115	0.22810	0.15232
	(2.191)**	(3.583)**	(2.309)**	(2.595)**
YRU	0.24121	0.08355	0.23690	0.05844
	(2.313)**	(2.158)**	(2.275)**	(1.478)
NAT	0.15005	0.08958	0.19069	-0.33794
	(1.336)	(-0.465)	(1.646)	(-1.580)
LAMBDA	—	—	-0.18220	-0.36649
			(-1.380)	(-2.564)**
R-Squared	0.5765	0.5382	0.5798	0.5501

* = significant at 10 percent
** = significant at 5 percent
— = not applicable

by 5.1 percent per annum compared to 2.2 percent in the public sector, and a steeper post-primary education schedule is also observed in the private sector.

The parameters of the choice equation are shown in Table 2.16, with the parameters for a base group of non-participants shown in the first two columns, and the implied private versus public sector choice estimates shown in the third column. This column implies that the choice between sectors is essentially economic. The only other variables that influence sectoral choice are location, gender and nationality, factors that are more likely to show significance due to the presence of barriers to entry. Females find access to the private sector more difficult than to the public sector, compared to males, as do agents in Abidjan compared to other urban areas. Further non-Ivorians have a lower propensity for public sector employment compared to Ivorians.

The result that females are relatively more favored in the public sector rather than the private supports the results of van der Gaag and Vijverberg's switching regression model estimated on the 1985 CILSS. However the descriptive statistics presented in Section 2.2.2 show a clear gender pattern within the private sector when a decomposition by unionization is made. Closer investigation of this motivates the further decomposition of the labor market carried out in the next section.

Table 2.16 Côte d'Ivoire 1986 : Bi-Sectoral Choice Model

	(1)	(2)	(3)
CONSTANT	-13.96140	-23.21940	9.25800
	(-8.631)**	(-11.960)**	([illegible].328)**
AGEY	0.32050	0.35490	-0.03440
	(5.193)**	(4.674)**	(-0.395)
AGE2	-0.00469	-0.00474	0.00005
	(-5.883)**	(-5.015)**	(0.046)
HEAD	1.35799	1.35497	0.00302
	(5.393)**	(4.616)**	(0.009)
MARRIED	0.23396	0.48512	-0.25116
	(0.962)	(1.776)*	(-0.[illegible]24)
FGOVT	0.23900	0.41404	-0.17504
	(0.836)	(1.401)	(-0.561)
MGOVT	0.28607	0.55265	-0.26658
	(0.333)	(0.657)	(-0.334)
CHILD	-0.12650	0.10436	-0.23086
	(-0.521)	(0.367)	(-0.743)
ABIDJ	1.48423	0.64101	0.84321
	(5.187)**	(1.935)*	(2.327)**
SEX	1.07297	0.38945	0.68352
	(4.765)**	(1.437)	(2.223)**
CEPE	0.25750	0.60821	-0.35072
	(0.785)	(1.567)	(-0.848)
BEPC	-0.26707	-0.02101	-0.24606
	(-0.841)	(-0.067)	(-0.760)
NAT	0.33220	2.30423	-1.97202
	(1.371)	(5.924)**	(-4.970)**
PUBWAGE	0.84986	1.96929	-1.11943
	(2.541)**	(5.426)	(-2.955)**
DIFF	0.95687	0.79615	0.16072
	(1.730)*	(1.289)	(0.249)
Log-Likelihood -989.68			

* = significant at 10 percent
** = significant at 5 percent
Code. [illegible]ivate/None, (2) Government/None, [illegible]vernment/Private.

2.2.5 *The Model with Private Sector Decomposition*

The one equation model and descriptive statistics presented in Section 2.2.2 of this chapter, suggested that important differences might emerge, particularly in the private sector, when unionization was considered. As the unionization variable available in the CILSS records the presence of a union in the workplace rather than individual affiliation, there can be some doubt as to what this decomposition really records. On the one hand we may pick up benefits to labor from unionization itself, with non-affiliated agents in a workplace where there is a union deriving benefits as free-riders. Alternatively this definition may be a proxy that provides an effective division between heterogenous components of the labor market, for instance between a formal and an informal sector. Thus effects attributed to unionization could be in fact sectoral differences.

In this section we use this unionization definition to divide the private sector used in the previous section into two separate sectors. Thus we estimate a model where individuals face the choice of non-participation against participation in one of three sectors, private non-unionized, private unionized, or public.

It might seem logical to apply the decomposition to the public sector as well, to provide four sectors, in fact using the same divisions as analyzed in Gyourko and Tracy (1987), who however do not have non-participants as a base group. We do not apply this extra decomposition for a variety of practical and theoretical reasons. First, the relatively small sample of public sector non-unionized individuals places great demands on the model, and as noted in Section 2.2.2 all public sector workers are theoretically unionized. However, such a model was still estimated and provided no significant differences in wages within the public sector, the same result as was suggested by the Smith model of Section 2.2.2, nor were there significant differences in the participation model.

Further, as we observe unionization in such a general way with no information on affiliation rates within a unionized sector, we cannot confidently present such a model as one of endogenous unionization and sectoral choice, but rather as one of endogenous sectoral choice where the presence or otherwise of unions defines sectoral boundaries. The issue is therefore whether we are observing three simultaneous decisions, i.e. participation, sector, and unionization, or simply the first two of these. Given our one step removed observation of unionization the balance must be to favor the latter interpretation. Hence we are imposing an homogeneous public sector and attempting to capture the a priori greater heterogeneity within the private sector by a decomposition according to access to unionization.

Using this tri-sectoral model of the labor market and the same variables as in the previous section, the wage equations estimated using the selection compensation variables derived from the reduced form multinomial logit are presented in Table 2.17. The difference in the underlying reduced form selection equation results in only minor changes in the public sector wage equation compared to that shown in Table 2.15. However, the two private sector equations show a marked differ-

Table 2.17 Côte d'Ivoire 1986: Tri-Sectoral Choice Model: Wage Equation

	Private Unionized	*Private Non-Unionized*	*Public Sector*
CONSTANT	4.38557	3.01637	5.47915
	(6.583)**	(5.128)**	(8.527)**
AGEY	0.04586	0.04296	0.02203
	(6.251)**	(5.246)**	(3.097)**
CEPE	0.06387	0.45220	0.09436
	(0.239)	(1.672)*	(0.357)
BEPC	-0.13399	-0.08044	-0.00867
	(-0.440)	(-0.266)	(-0.078)
HIGHER	-0.16162	-0.40522	0.21254
	(-0.365)	(-0.387)	(0.994)
ABIDJ	0.00613	0.27985	-0.05745
	(0.030)	(1.740)	(0.631)
SEX	-0.23342	-0.06914	-0.00902
	(-1.147)	(-0.346)	(-0.090)
GRSP	0.08521	-0.01999	0.06107
	(2.027)**	(-0.489)	(1.319)
GRS	0.12700	0.19831	0.07918
	(2.214)**	(2.811)**	(2.1[illegible]
GRS2	0.19925	0.17915	0.14937
	(1.485)	(1.025)	(2.546)**
YRU	0.24164	0.38637	0.05616
	(2.577)**	(1.270)	(1.417)
NAT	0.19330	0.13294	-0.33951
	(1.264)	(0.763)	(-1.596)
LAMBDA	-0.38468	0.16623	-0.37109
	(-2.267)**	(0.709)	(-2.645)**
R-Squared	0.6928	0.4743	0.5508

* = significant at 10 percent
** = significant at 5 percent

ence. LAMBDA is significantly different from zero in the unionized private sector, and implies that the average employee in this sector commands a wage 27 percent higher than that that would be given to an individual drawn at random in that sector. The equivalent figure for the public sector employee is also, as in the previous section, 27 percent. Evaluated across the entire sample of both participants and non-participants, the average expected absolute wage in the private unionized sector is 0.6 percent higher than in the public sector. In the non-unionized sector LAMBDA is insignificant, with a positive coefficient, i.e. wages are lower for the average individual in that sector compared to that offered to the randomly drawn individual by that sector. The regressions identify significant returns to the primary diploma and grades of lower secondary in the non-unionized private sector, and to primary and lower secondary grades as well as years of university education in the unionized private sector.

There are again no significant distinctions by gender in any sector, and the wage premium for Abidjan found for the private sector in the previous section is now seen to be exclusively confined to the non-unionized component of that aggregate.

The associated participation equation for this tri-sectoral model is shown in Table 2.18, where rebasing gives the coefficients shown in columns 4 to 6 thus presenting the complete set of pairwise comparisons derived from the multinomial logit estimation. The estimated public sector wage is entered as PWAGE, and DIFF1, and DIFF2 are respectively the differentials of PWAGE from respectively the unionized and non-unionized private sector estimated wages.

Table 2.18 confirms the general determinants of participation found in the previous chapter. Age and headship both exert a strong influence on participation but not sectoral choice. The effect of paternal occupation is insignificant. This effect was found to be confined to females, however the small sample sizes make decomposition of multi-sectoral models by gender impossible. Table 2.18 shows this effect to be strongest in the public sector and the private unionized sector, hence there is a presumption that the strong informational benefits of paternal occupation found for females in the previous chapter, are mainly concentrated in these sectors.

The presence of children provides a significant explanation of participation in the private unionized sector compared to non-participation, and in both the private unionized and public sectors compared to the private non-unionized sector. The relative favoring of women in the public sector compared to the private found in the previous section is only significant compared to the private unionized sector. However, gender is still a significant determinant of participation in both private sectors, with females finding access more difficult than males, while no such significant effect prevails in entry to the public sector. The result that nationality significantly affects the chances of participation in the public sector, while providing no significant barrier in the private sector found in the bisectoral model is seen to also apply in this case of a disaggregated private sector. There are no differences in the relative case of access to the two components of the private sector on the grounds of nationality, while, as evidenced by Table 2.14, the public sector is almost exclusively Ivorian. Participation overall is generally more likely in Abidjan, with the private unionized sector being particularly biased towards the capital.

Table 2.18 Côte d'Ivoire 1986: Tri-Sectoral Choice Model: Participation Equation

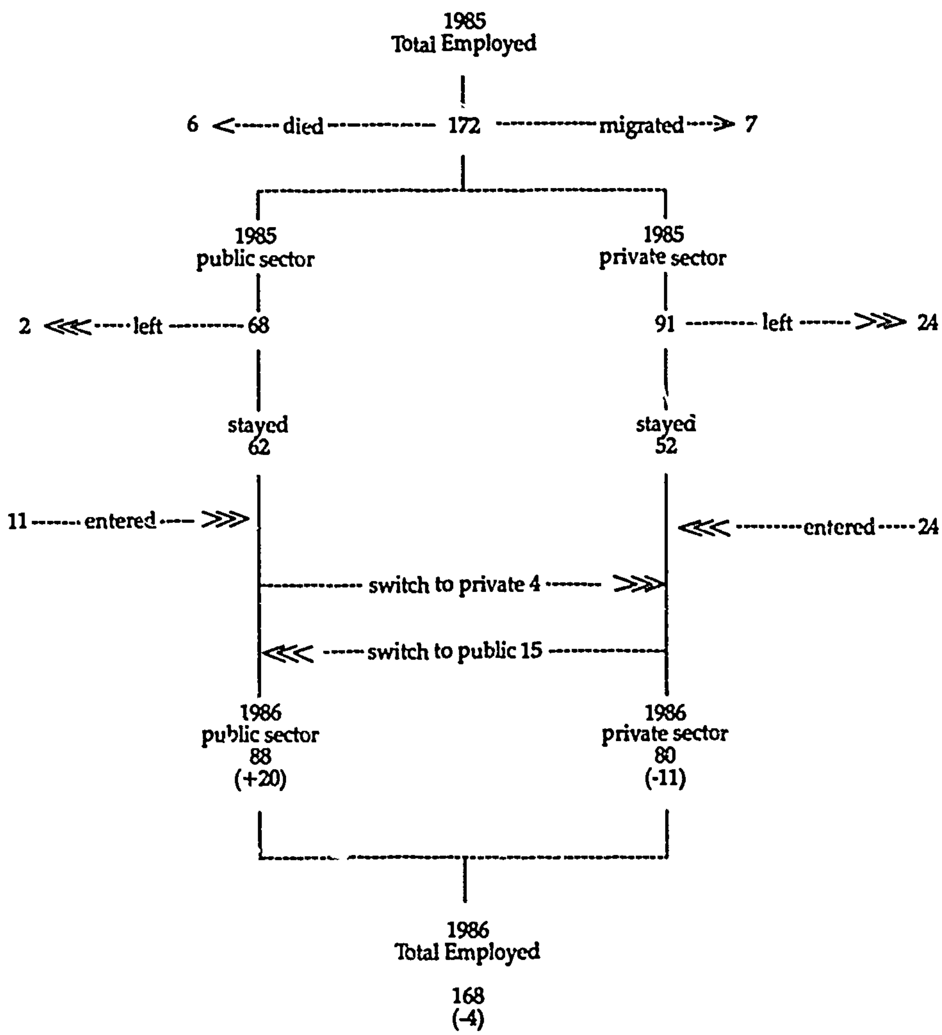
	(1)	(2)	(3)	(4)	(5)	(6)
CONST	12.79380 (-4.761)**	-19.39030 (-6.377)**	-25.64500 (-11.177)**	-6.59652 (-1.827)*	-12.85120 (-4.246)**	-6.25465 (-2.067)**
AGEY	0.33752 (4.336)**	0.30851 (3.068)**	0.37218 (4.695)**	-0.02901 (-0.241)	0.03465 (0.333)	0.06367 (0.562)
AGE2	-0.00485 (-4.600)**	-0.00458 (-3.635)**	-0.00488 (-5.005)**	0.00028 (0.178)	-0.00002 (-0.017)	-0.00030 (-0.213)
HEAD	1.19827 (3.973)**	1.57687 (4.148)**	1.49530 (4.951)**	0.37860 (0.873)	0.29703 (0.795)	-0.08157 (-0.199)
MARRIED	0.07546 (0.264)	0.45447 (1.226)	0.49887 (1.792)*	0.37900 (0.886)	0.42341 (1.187)	0.04441 (0.113)
FGOVT	0.12149 (0.332)	0.46671 (1.268)	0.47620 (1.580)	0.34522 (0.760)	0.35471 (0.872)	0.00949 (0.026)
MGOVT	0.15957 (0.133)	0.20956 (0.203)	0.30302 (0.347)	0.05000 (0.038)	0.14345 (0.199)	0.09346 (0.102)
CHILD	-0.44737 (-1.569)	0.65808 (1.700)*	0.39129 (1.312)	1.10545 (2.522)**	0.83866 (2.272)**	-0.26679 (-0.639)
ABIDJ	1.12541 (2.546)**	2.31110 (3.946)**	1.27445 (2.887)**	1.18568 (1.758)*	0.14903 (0.267)	-1.03665 (-1.688)*
SEX	0.74509 (2.724)**	1.41317 (3.761)**	0.35387 (1.259)	0.66408 (1.532)	-0.39522 (-1.111)	-1.05931 (-2.563)**
CEPE	[illegible] (0.228)	0.34103 (0.665)	0.38721 (0.905)	0.23849 (0.394)	0.28467 (0.525)	0.04618 (0.085)
BEPC	-0.52017 (-1.007)	-0.35794 (-0.686)	-0.51841 (-1.242)	0.16224 (0.253)	0.00176 (0.003)	-0.16048 (-0.325)
NAT	0.37498 (0.941)	0.60846 (1.504)	2.61143 (6.169)**	0.23348 (0.479)	2.23645 (4.414)**	2.00297 (4.225)**
PWAGE	0.64150 (1.238)	1.35869 (2.577)**	2.26188 (5.607)**	0.71719 (1.123)	1.62038 (2.969)**	0.90319 (1.819)*
DIFF1	1.27318 (0.996)	1.03353 (0.683)	3.03556 (2.198)**	-0.23965 (-0.137)	1.76237 (1.071)	2.00203 (1.198)
DIFF2	0.42064 (0.938)	0.15714 (0.309)	-0.61865 (-1.186)	-0.26350 (-0.436)	-1.03929 (-1.671)*	-0.77579 (-1.301)
Log-Likelihood	-1,122.5					

* = significant at 10 percent
** = significant at 5 percent
Code: (1) Private non-Union/None, (2) Private Union/None, (3) Government/None, (4) Private Union/Private non-Union, (5) Government/Private non-Union, (6) Government/Private Union.

2.2.6 *The Labor Force Trace*

The 1986 cycle of the CILSS used the same sampling frame as in 1985, with all the households in half the clusters being reinterviewed, and a new random sample drawn in the other clusters. This feature enables us to trace individuals between surveys, and we have used this trace to examine the changes in the labor market over the year. Applying the definition of waged employment used in Part A to the 1985 CILSS produced a sample of 172 individuals who could be traced onto the 1986 survey, or at least whose fate could be verified by another household member. Of these, six had died and seven migrated, all of whom had been employed in the private sector. The sample of waged individuals in the 1986 survey who could be traced back to the 1985 survey was 168, although the age range in the definition used before began one year

Figure 2.4 Changes in Employment 1985-86



later to maintain comparability. Hence overall employment had increased by two, allowing for deaths, and at first sight the labor market appears static. However this conceals a pattern of fluidity within the labor market, as is evidenced by Figure 2.1. The sample of those alive and domiciled in the same place in both years, gives a sample of 68 public sector employees and 91 private sector employees for 1985. The first major surprise is that only two individuals left the public sector to become unwaged, with four switching into the private sector. The public sector therefore retained 62 employees out of 68, and the active labor force as a whole retained all but two. There is therefore no evidence of large scale retrenchment in the public sector from 1985 to 1986. Further there appears to have been no effective freeze on hiring, the public sector absorbed 11 individuals who were unwaged in 1985, and 15 who switched from the private sector. Overall we record public sector employment in 1986 as being some 18 percent higher than in 1985. This survey evidence is contradicted by official estimates of public sector employment and we are not in a position to reconcile this discrepancy. Since the sample size of those in public sector employment is small, it is tempting to dismiss the discrepancy as being attributable to sample bias. However, the nature of the problem is likely to be deeper than this since the survey records 26 public sector entrants during 1986 at a time of a hiring freeze. Either these were not genuine entrants, which casts more general doubt upon the validity of the data, or the hiring freeze was less than fully effective. It is possible that there is a grey area at the edge of public sector employment. As noted in Section 2.2.2 all public sector employees are officially unionized.

Yet 23.5 percent do not record the presence of a union at their workplace, and this includes the majority of the new entrants.

The picture is rather different in the private sector. Only 52 out of 98 employees (including those who migrated) are retained in the sector. There is also a higher inflow rate from those previously unwaged than in the public sector, but most strikingly nearly one in six of of 1985 private sector employees switch into the public sector, with a net loss of 11 individuals as four move in the opposite direction. Overall the private sector contracts, in terms of its employment level, by some 12 percent. Thus in this trace sample public sector employment moves into the majority between 1985 and 1986. This is confirmed by the general samples, van der Gaag and Vijverberg calculate that 41 percent of employment is in the public sector in 1985, with our different definition giving a marginally higher level. However this definition in 1986 gives a proportion of 51 percent over the whole traceable and non-traceable sample. Overall the story is therefore not one of continued public sector restraint following structural adjustment, but rather an unbound public sector making up for the years of frozen hiring, probably by casual appointments.

The story is further clarified by consideration of wage changes. The sample of those in constant employment proved too small to provide a satisfactory breakdown of wage changes, particularly when sample selection adjustment is required, but the averages are themselves illuminating. Taking average changes in the log wage and converting to absolutes, shows that the nominal public sector wage in 1986 was 95.7 percent of its 1985 level, and the private sector wage at 90.7 percent of its 1985 level, with inflation running at some 2 percent. Hence in a period of real wage decreases, total employment was static while public sector employment increased sharply. An explanation is that the relaxation of the structural adjustment controls reduced the barriers to entry to the public sector. With a queue of individuals wishing to work in the public sector at market wages but unable to do so, the effect of reduction of barriers proved greater than the disincentive effect of lower wages. Further with an increase in the relative public sector wage compared to the private sector, net switching into the public sector is observed.

The male wage for those in constant employment fell to 93.6 percent of its 1985 level and the female wage to 93.8 percent. The above fluidity in employment therefore showed no clear gender pattern. In the unionized sectors of the economy wages fell to 93.9 percent of the 1985 level, and in the non-unionized sector to 92.8 percent. Unionization, while as in the previous chapter observed at one step remove, did not protect wage levels. Holders of the CEPE diploma experienced a fall in wages of 3.3 percent, while those without educational diplomas found wages falling by 12.2 percent.

3. *Gender Differences in Access to Schooling in the Côte d'Ivoire*

3.1 Introduction

The previous chapter has demonstrated that the root cause of inferior female access to waged employment, and therefore to inferior labor mobility, lies in education. This chapter analyzes the underlying gender differentials in access to education.

Attention is focused on entry to each of the three main stages of schooling in the Côte d'Ivoire: primary school, lower secondary school, and upper secondary school. Entry at each stage is analyzed conditional on individuals having completed the previous stage. thus those who did not complete primary school are not included in the subsample used to estimate the determinants of entry to lower secondary school, and those who did not complete lower secondary school are excluded from the analysis of entry to upper secondary school. This approach reflects the sequential nature of schooling decisions and allows one to analyze separately gender differentials at each stage of schooling. The approach also uses the information that is available about those who are still at school and thus whose completed schooling is unknown. Unfortunately, the analysis of entry to primary school suffers from the problem of censoring since many children who have not enrolled in school are between the observed minimum and maximum ages for entry. To avoid treating such children as if they would never enrol, only those above the age of the oldest child in the first year of primary—ten years old—were included in the sample used to analyze primary school enrollment.

The basic data about the school enrollment of those included in the various subsamples used for analysis is presented in Table 3.1.

It is clear from Table 3.1 that there are marked gender differentials in access to education at all three stages of schooling. Indeed, the figures show that, compared with boys, girls are:

i) 91 percent more likely not to go to primary school.
ii) 37 percent more likely not to go to lower secondary school condiuonal on completing primary school.
iii) 14 percent more likely not to go to upper secondary school conditional on completing lower secondary school.

These summary statistics should be handled with care since they are each based on different samples and age cohorts. However, the figures do seem to indicate that girls face additional difficulties in getting onto each rung of the academic ladder to those faced by boys, although they become more successful in surmounting those obstacles at higher stages of schooling. This does not necessarily imply that the admissions procedures to higher stages of schooling are themselves superior in terms of gender equity than are those to the lower stages because the pattern could simply be due to a sample selection effect. To see how this could arise, consider a very simple model of entry to school. individuals have a single propensity to go to school and enrollment into higher levels of schooling is restricted to those with progressively higher propensities to go to school. Also assume that this propensity is on average higher for boys than for girls but is otherwise normally distributed with equal variances for both sexes. In such a model it is clear that entrance procedures to higher stages of schooling do not use different criteria to those used to determine entry to lower stages and thus these procedures are no superior in terms of gender equity. Nonetheless, if we look at the subsamples of those who have entered previous stages of schooling and see what proportion of them progress further, we will be excluding those with low propensities to go to school—more of whom are girls—and thus the differences between the mean

Table 3.1 Figures for Examination Passes and School Entry for the Samples Used

(i) Sample used for analysis of primary school entrance; 11-18 year olds

	No Primary	*Primary*
Total	668 (27.46%)	1,765
Girls	426 (36.47%)	742
Boys	242 (19.13%)	1,023

(ii) Sample used for analysis of lower secondary school; 10-18 year olds who have completed primary school: numbers by qualification and enrollment

	Lower Secondary Schooling		
	State	*Primary*	*None*
Passed CEPE:			
Total	302 (45%)	95 (14%)	76 (11%)
Girls	95 (35%)	43 (16%)	26 (10%)
Boys	207 (51%)	52 (13%)	50 (12%)
Failed CEPE:			
Total	16 (2%)	7 (1%)	171 (26%)
Girls	5 (2%)	3 (1%)	90 34%)
Boys	11 (3%)	4 (1%)	31 (20%)

(iii) Sample used for analysis of upper secondary school; 30 year olds and younger who have completed lower secondary school

	Upper Secondary Schooling		
	State	*Private*	*None*
Passed BEPC:			
Total	159 (44%)	33 (9%)	73 (20%)
Girls	41 (39%)	8 (8%)	25 (24%)
Boys	118 (47%)	25 (10%)	48 (19%)
Failed BEPC:			
Total	13 (4%)	9 (3%)	72 (20%)
Girls	6 (6%)	4 (4%)	22 (21%)
Boys	7 (3%)	5 (2%)	50 (20%)

propensities to enrol of the boys and girls remaining in the subsamples will be progressively reduced as we consider higher stages of schooling. For example, if entrance exams are important in determining entrance, lower gender differentials in pass rates at higher levels may make it seem as if those exams for higher stages of schooling are less unfavorable to girls when in fact it is merely that only very able girls have been able to reach the position where they can sit the exams. Such effects are partly controlled for in the subsequent analysis since we include as determinants of enrollment many characteristics such as household income per capita. Hence, if the effects of gender per se, as represented by a dummy variable for being female, are found to vary at different stages of schooling, it can not be because those girls who are included in the samples used to analyze higher levels of schooling have atypically high incomes. However, we have no data on *ability* per se—such as I.Q. test results—and thus the possibility of a change in the effects of the gender dummies may be explicable in terms of a sample selection effect working through differences in ability.

One of the interesting features of the data to emerge from Table 3.1 is that although there is a marked gender differential in the proportion of primary school leavers who go to lower secondary school, there is no such differential in the proportions of successful CEPE candidates who enrol in lower secondary school; indeed the differential in this subgroup favors girls. This has considerable implications for policy since in the Côte d'Ivoire state secondary school places are rationed by academic performance. Marks obtained in the primary leaving exam, the Certificat d'Etudes Primaires Elementaire (CEPE), are used as selection criteria by secondary schools (see Glewwe (1988)). Similarly, successful entrants to state upper secondary schools are chosen from those who passed the lower secondary school leaving school exam, the Brevet de Fin d'Etudes du Premier Cycle (BEPC). Consequently, one interpretation of the observation that there are significant gender differentials in CEPE exam performance but not in lower secondary school enrollment conditional on such performance is that girls have inferior access to lower secondary schooling to boys, not because their parents have a lower demand for the education of girls than boys but because the system of admission to state schools works against them. To investigate this question of whether gender inequalities in access to the two cycles of secondary school are due to demand factors or to the rationing system, it was decided to analyze separately:

i) who passes CEPE and BEPC exams;
ii) whether those who pass the primary/lower secondary school leaving exam then enrol in lower secondary/upper secondary school.

Both questions were analyzed using a standard logit model, which can be interpreted using a latent variable formulation. Assuming a nought-one dependent variable, Y_i, there is a latent variable Y_i^* which is a linear function of a vector of determinants X_i and a stochastic term ε_i:

$$Y_i^* = \beta' X_i + \varepsilon_i$$

such that:

$Pr(Y_i=1) = Pr(Y_i^* > 0)$

This yields the logit model if we assume that the stochastic term has a Type I extreme value distribution. In the context of the determinants of exam success, if Y_i is one if the individual passes the exam, the latent variable has a natural interpretation as the exam mark of the individual transformed in some way so as to make the pass mark equal zero. The interpretation of the latent variable when the model is applied to enrollment conditional on passing the exam is less straightforward since enrollment may still be determined by a combination of rationing and choice. Interpretation would be easiest if it is assumed that possession of the leaving exam of the prior stage of schooling is necessary and sufficient for eligibility for entry to the next step of schooling. Under this assumption, the analysis of exam success will explain the allocation of state school place rations whilst the determinants of school enrollment conditional on passing the relevant exam will reflect only demand factors. Together the two processes—rationing and individual choice—would sequentially determine secondary school enrollment. In particular, the latent variable in the model for entry to school conditional upon passing the exam could be interpreted following McFadden (1973) as the outcome of individuals maximising stochastic utility functions; thus if:

$$V_{ij} = \alpha_j' X_i + \nu_{ji} \qquad j = 0, 1$$

where V_{ij} is the indirect utility function from choosing outcome $Y_i = j$, then the latent variable Y_i^* will be equal to the indirect utility function from enrolling normalized around that from not enrolling, that is to say:

$$Y_i^* = V_{i1} - V_{i0}$$

and thus $= \beta = \alpha_1 - \alpha_0$.

However, it is not strictly true that passing the CEPE/BEPC exams is necessary and sufficient for eligibility for a government secondary school place. That success in the relevant exam is not always required for entrance to a state secondary school can be seen from Table 3.1. Nonetheless, since only 3-4 percent of the subsamples used did enter without the appropriate qualifications, and given what Glewwe has reported, we feel confident at dismissing the small numbers who manage to enter state secondary schools without passing the leaving exam of the prior stage of schooling as either. reporting errors; rare exceptions being made to entry procedures; or individuals entering less demanding non-academic *technical* or *vocational* schools. That passes in the relevant leaving exam of the previous cycle are not always enough for individuals to be *unrationed* is suggested by Glewwe who reports that in some parts of the Côte d'Ivoire a *good mark* in the CEPE rather than a simple pass is necessary to meet the selection criteria for lower secondary school in some areas of the country. Consequently, whilst for most of this chapter the results of logit for enrollment in secondary school conditional upon passing the leaving exam of the prior stage of schooling will be interpreted as reflecting demand rather than rationing factors, it should be borne in mind that this may not always be true.

As can be seen from Table 3.1, private secondary schools service between a fifth and a quarter of all secondary school students. Given the existence of a rationed state secondary school sector, one might expect a priori that the private sector would cater for those who are rationed out of the state school system. Consequently, pursuing the initial hypothesis that girls in particular are rationed out of the state lower secondary schools, one would expect proportionately more of them than boys in private lower secondary schools and indeed this is confirmed in Table 3.1. However, the very small numbers of those who fail the CEPE (or the BEPC) who are observed in private schools seems to imply either that rationing of state school places is more severe than was suggested above (that is to say merely passing the exam is not enough to be included in the ration) or that the private sector does not play a marked role in catering for those students rationed out of the state school sector. This issue will be considered later in the section on the results of the model for access to lower secondary school. The reason why the private sector does not educate many of those rationed who fail the CEPE may be a complementarity between the returns from schooling and academic ability: perhaps given the substantial fees charged by private schools, only gifted children (and thus children eligible to enrol in state schools) for whom the returns to schooling are particularly high are likely to enrol. Alternatively, there may be reasons why private schools wish to ration places by academic criteria: for example, if they believe less able students have negative external effects on the more able or if parents—unable to observe the quality of the intake of private schools—measure a school's quality by the average exam performance of its pupils. Nonetheless, it would be desirable to model the behavior of those rationed out of the state sector, but the proportion of CEPE failures who go to private schools is too small to permit interesting results. Instead, only the choices of successful exam candidates will be analyzed. Such individuals face a choice between state, private or no lower/upper secondary schooling which can be modelled using a simple multinomial logit model.

Table 3.2 Variables Used in the Analysis

Dependent Variables:	
PRIM	Equals 1 if the child has received any primary schooling, 0 otherwise
CEPE	Equals 1 if the child has passed the CEPE, 0 otherwise
LSEC	Equals 1 if the child has entered public lower secondary school, 2 if has entered private lower secondary, 0 otherwise
BEPC	Equals one if the child has passed BEPC, 0 otherwise.
USEC	Equals 1 if the child has entered public upper secondary school, 2 if has entered private secondary school, 0 otherwise
Hypothesised Explanatory Variables:	
AGE1, AGE2	Quadratic specification of age
LIVVALPC	Value of household livestock per capita, measured in CFA francs but deflated by 1000
INCPC1, INCPC2, INCPC3	Household income per capita with quadratic and cubic terms are used, the minimum income per capita observed in the sample (-185,300) is subtracted from each income to avoid negative valued
CONPC1, CONPC2, CONPC3	Cubic specification of household consumption per capita, measured in CFA francs but deflated by 10,000
ESTCEPE1	Probability logit reported in Table 3.6 predicts for passing CEPE
ESTCEPE 2	Probability logits reported in Table 3.7 predicts for passing CEPE
The following are nought-one dummy variables which equal 1 if:	
FEMALE	Child is female
NONIVORIAN	Not of Ivorian nationality
NORTH	Household member sampled in clusters 93 or 95-100; these are rural areas in the Northern Savannah area
WEST	Household member sampled in rural clusters to the west of the Bandama River, excluding those included in the NORTH category
URBAN	household member sampled in clusters 1-43; these are all in towns or cities
ABIDJAN	Household member sampled in clusters 1-21; these are in Abidjan
FGOVT	Father's main occupation is wage employment for the government
FPRIVT	Father's main occupation is wage employment for the private sector
DADWAGED	Father's main occupation is wage employment
DADPRIM	Father received some primary schooling
MUMPRIM	Mother received some primary schooling
BOTHPRIM	Both parents received some primary schooling
DADCEPE	Father obtained CEPE
MUMCEPE	Mother obtained CEPE
BOTHCEPE	Both parents obtained CEPE
DADSEC	Father received some secondary schooling
DADBEPC	Father obtained BEPC
MUMBEPC	Mother obtained BEPC

3.2 Explanatory Variables

Table 3.2 defines the variables used in the reported results. The explanatory variables were included on the assumption, discussed above, that the equations for entry to the two cycles of secondary school conditional upon exam success, at least partly, reflected demand factors in the same way as the logits for entry to the unrationed primary sector do. Following Becker (1975) and treating education as an investment in human capital, one would predict the expected returns and costs of schooling to be the key determinants of demand; but unfortunately, measuring both sets of factors is problematic. Expected returns are unobserved and furthermore, conventional wage functions to estimate returns usually rely heavily upon experience and schooling as the explanatory variables: neither of which can be used to analyze prior schooling decisions. Furthermore, the pecuniary costs of schooling are not observed for those who do not attend school and the opportunity costs in terms of child labor are not recorded in the survey (see Grootaert (1988) for estimates of the opportunity costs). Consequently, most of the explanatory variables included in this work are at best proxies for costs and returns and indeed may proxy both, making interpretation problematical. A brief rationale for the inclusion of some of the explanatory variables follows.

i) HOUSEHOLD INCOME AND WEALTH. Household income will be a determinant of the demand for education if, as seems likely, capital market imperfections prevent pupils from securing loans to fund their schooling on the strength of the future earnings premia their schooling will give to them. Further, since education may often be in part a gift to a prospective student from other household members other than the child it will be bounded by their budget constraint like other consumption goods and services. Household wealth is likely to be a determinant of the demand for education for similar reasons.

In all the equations, likelihood ratio tests were carried out to see if household income, consumption or food consumption (all measured per capita) were the most significant sets of economic variables. Consumption was considered as an alternative to income because it was thought possible th. : *income* effects are better measured by consumption due to the smaller reporting errors likely in that statistic (see p.20 of the note by the researchers at Warwick University, Johnson et al (1989)) and possibly due to its being a better reflection of permanent income (although the capital markets in the Côte d'Ivoire are unlikely to be perfect). Food consumption per capita was considered since it may be a better identifier of poverty than income or consumption per capita.

To allow for non-linear effects of income upon demand, the economic variables were entered in quadratic and cubic forms, with likelihood ratio tests being used to test for more parsimonious forms. Since some i .comes were negative, the lowest income was subtracted from each income to avoid problems in interpreting the squared terms.

One final problem that should be noted is that, due to data limitations, the income and related variables are those observed at the time of the survey, which is usually not the time at which the schooling decision being analyzed was made. The approach of analysing whether an individual is enrolled or has ever been enrolled in a stage of schooling contrasts with the approach of Glewwe (1988) (and also with Gertler (1988)) which analyzes decisions about whether to spend th year of the survey in that stage of schooling. The pro em with analysing only current decisions is that it involves assuming that enrollment in a stage of schooling is determined in the same way as are premature drop-outs from that stage. this seems implausible, partly because enrollment into secondary school is rationed but decisions on whether to withdraw are not affected by such considerations. However, even if both decisions were demand determined, the appropriate econometric model may still differ, for example, on a simple version of the screening hypothesis, where sitting the exam is the only benefit from schooling, a person will only drop-out if their circumstances unexpectedly change for the worse. This example may be extreme, but the mechanism probably underlies some of the observed clustering of schooling attainment around the ends of each cycle. Human capital theory could account for the same phenomena if human capital is acquired in *lumps*. The drawback with the method adopted here is that it requires cu ent values of explanatory variables to proxy their value at the time of the schooling decision being analyzed. This will be a source of measurement error and, as will be discussed later, may cause serious exogeneity problems for the sample of under 30 year olds used to examine upper secondary school entry

ii) PARENTAL EDUCATION. Aside from influencing parental income, parental education may affect the demand for schooling by a number of channels. Having educated parents may increase the amount of human capital a child can acquire at school; one simple example being through literate parents reading to their children. A variety of attitudinal mechanisms could also be postulated: educated parents may know more about—or value more highly—the benefits of schooling.

The precise form in which parental education enters the final equations was largely that suggested by the data. separate cumulative dummy variables for the mother's and father's education, together with an interaction term, were used in preliminary estimates

iii) REGIONAL DUMMY VARIABLES. Since the sample contained both urban and rural clusters, an urban dummy variable was used, as was a further variable for those living in Abidjan. The rural part of the sample was divided into east and west of the Bandama River ince this may capture some of the ethnic differences within the Côte d'Ivoire. This division is inevitably rather arbitrary but was prompted by comments by Clignet and Foster (1966) that as well as coming into contact with European influences first, tribes west of the River were organized along matrilineal lines of descent whilst those to the east were patrilineal. A dummy variable for the few clusters in the very North of the country was created since this area is relatively underdeveloped compared to the rest of the country and also differs from other areas by being predominantly Muslim. Interpreting these regional dummy variables is difficult: for example, the urban dummy variable may pick up differences in returns to education due to greater access to the formal labor market; lower opportunity costs of child labor due to less reliance on production from family land, and provision of better school facilities.

iv) OTHER VARIABLES. Dummy variables for parental occupation were used since these may proxy differential access to the formal labor market and hence affect the derived demand for schooling. The probability that the individual would pass the leaving exam of the previous cycle was also entered, partly as a rate of return proxy under the hypothesis that returns to schooling increase with academic ability. However, these predicted examination performance variables could also be interpreted as reflecting any rationing of successful candidates on the basis of their exam results.

v) VARIABLES REJECTED FROM THE ANALYSIS. A number of variables were rejected on the basis of preliminary estimates: notably several dummy variables for parental occupation and variables for the numbers of younger and older brothers and sisters. Land holdings were also insignificant in all cases. A dummy variable for whether any member had worked on family land in the past year was used as a proxy for significant costs to child labor but rejected as insignificant in all cases. Where a variable defined in Table 3.2 does not appear in a particular equation, it is either because it is inappropriate a priori or due to its having been rejected on the basis of likelihood ratio tests.

vi) THE DETERMINANTS OF EXAM SUCCESS. The selection of explanatory variables for exam success was rather more speculative than the selection of those for school entry. Following the argument given above that parental education increases the amount of human capital acquired from school, one would predict parental education to improve exam performance. Furthermore, a stronger economic position of the household might be expected to increase the expenditure on items of educational value, as well as providing a good environment for learning. Regional dummy variables were also used to capture differing school quality across the country.

3.3 Primary School Enrollment

A binary logit was used to analyze the determinants of whether a child had received any primary schooling. A sample of people of the ages of 11 to 18 was used, consisting of household members and their children living elsewhere. If any of the hypothesised explanatory variables were missing for an observation, that observation was rejected. The lower age limit was set to avoid right censoring of observations since the oldest child observed in the first year of primary

Table 3.3 Logit Estimates of Determinants of Enrolling in Primary School: ages 11-18 years

Variable	*Coefficient*	*T-ratio*	*Mean of X*	*Std.D.of X*
CONSTANT	-5.323	-2.220**	1.000	0.000
FEMALE	-1.162	-10.800***	0.480	0.500
AGE1	0.7980	2.399**	14.289	2.264
AGE2	-0.2928E-01	-2.559**	209.284	65.599
NON-IVORIAN	-1.367	-8.640***	0.096	0.295
NORTH	-1.406	-6.580***	0.055	0.228
URBAN	0.4410	3.661***	0.449	0.498
DADPRIM	1.569	8.046***	0.259	0.438
MUMPRIM	1.310	3.216***	0.110	0.313
FGOVT	0.8749	2.045**	0.076	0.266
FPRIVT	0.4095	1.982**	0.132	0.338
INCPC1	0.6102E-01	3.575***	37.419	20.100
INCPC2	-0.5854E-03	-2.555**	1,803.997	3,121.496
INCPC3	0.1628E-05	2.025**	124,660.391	572,84[illegible]
LIVVALPC	-0.5283E-02	-3.643***	8.333	38.290
Log-Likelihood	-1,113.8			

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability

	Predicted		
	TOTAL	*Enrol*	*Don't Enrol*
TOTAL	2,433	1,995	438
Actual			
Enrol	1,765	1,613	152
Don't Enrol	668	382	286

** = significant at 5 percent
*** = significant at 1 percent

Table 3.4 Effect on Probability of Primary School Enrollment of the Nought-One Dummy Variables: Full Sample and Own Gender Only Results (percentage)
(Evaluated at means of other explanatory variables; gender specific means in part b)

(a) When estimated for full sample:

	Probability of Non-Enrollment
No Change, all variables at means	15
Girl	25
Boy	9
Ivorian	14
Non-Ivorian	38
Northern Home	45
Urban Home	12
Other Home (ie, Rural Non-Northern)	17
Mother some Primary	5
Mother no Primary	17
Father some Primary	5
Father no Primary	21
Father not Employee	17
Father Private Employee	12
Father Government Employee	8

(b) When estimated for single sex samples:

	Probability of Non-Enrollment	
	Girls Only	*Boys Only*
No Change, all variables at means	24	4
Ivorian	22	3
Non-Ivorian	44	16
Northern Home	62	15
Urban Home	19	2
Other Home	26	5
Mother some Primary	7	2
Mother no Primary	28	4
Father some Primary	9	1
Father no Primary	32	5
Father not Employee	27	11
Father Private Employee	16	11
Father Government Employee	16	0

school was aged 10. The upper age limit was chosen rather arbitrarily so as to capture behavior that was recent enough to be of interest for policy analysis and to try to limit the extent of the exogeneity problems discussed above concerning the household income variable.

The final form for the regression when run over the whole of the sample is presented in Table 3.3. All variables included in the final form were significant at 5 percent. Under the logit model, the size of the effect of each the explanatory variables on the probability of not enrolling varies with what base probability is chosen. The convention adopted here was to evaluate the effect of varying an individual regressor at the mean values of the other regressors. The results of such calculations are given in Table 3.4. Such simulations are illustrative and should be interpreted with care: for example, due to the non-linear form of the logit, the probability of not enrolling at the mean of all the regressors is 15 percent when in fact the observed proportion who have not enrolled in primary school in the sample is 27 percent.

As can be seen from Tables 3.3 and 3.4, there is a sizable and significant negative effect on the probability of going to primary school from being female. At the mean of the other explanatory variables, girls are predicted to be two and a half times more likely not to be enrolled. How other explanatory variables affect this overall gender differential was investigated by splitting the sample by gender; the results being recorded in Table 3.5. The likelihood ratio test for whether the sample should be split in this way gave a statistic of 28.88 which is significant at the 5 percent level with 14 degrees of freedom. Hence the test rejected the hypothesis that all of the explanatory variables had the same coefficients for the female sample

Table 3.5 Logit Estimates of Determinants of Enrolling in Primary School ages 11-18 years, by sex

Variable	Boys only Coefficient	Boys only T-ratio	Girls only Coefficient	Girls only T-ratio
CONSTANT	-5.523	-1.464	-6.294	-1.988**
AGE1	0.8410	1.633	0.8318	1.890*
AGE2	-0.3003E-01	-1.692*	-0.3113E-01	-2.056**
NONIVORIAN	-1.765	-7.613***	-1.005	-4.669***
NORTH	-1.314	-4.902***	-1.544	-4.231***
URBAN	0.6507	3.264***	0.3474	2.222**
MUMPRIM	0.7456	1.194	1.659	3.105***
DADPRIM	1.4733	4.350***	1.593	6.720***
FGOVT	15.6512	0.011	0.6835	1.524
FPRIVT	-0.2379E-01	-0.070	0.6897	2.673***
INCPC1	0.3315E-01	0.538	0.4998E-01	2.339**
INCPC2	0.8316E-04	0.066	-0.5513E-03	-1.912*
INCPC3	-0.2304E-06	-0.030	0.1626E-05	1.547
LIVVALPC	-0.6745E-02	-3.720***	-0.2450E-02	-0.984
Log-Likelihood	(Boys Only)	-484.66	(Girls Only)	-614.70

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability

Boys Only Predicted

	TOTAL	Enrol	Don't Enrol
Actual			
Enrol	1,023	972	51
Don't Enrol	242	155	87

Girls Only Predicted

	TOTAL	Enrol	Don't Enrol
TOTAL	1,168	862	306
Actual			
Enrol	742	636	106
Don't Enrol	426	226	200

* = significant at 10 percent
** = significant at 5 percent
*** = significant at 1 percent

as for the male sample. It should be noted that gender inequalities can vary with a variable despite the coefficients in Table 3.4 not differing between the sexes. This is because, being less certain to enrol for whatever reason, girls are generally more *marginal* cases and thus—according to the formulation of the logit model—more susceptible to given changes in the determinants. Finally, it should be noted that there is a certain ambiguity in the concept of gender differential: here we will mainly refer to the *absolute differential* meaning the difference in the probabilities for the two sexes. However, it may also be of interest to look at the ratio of the two probabilities, which we . ll term the *relative differential*. One problem with the latter measure is that it is sensitive to whether one looks at the ratio of probabilities of receiving schooling or the ratio of probabilities of not receiving it. In other words, a given change in an explanatory variable such as income per capita may tend to increase the proportions of girls to boys both in and out of school.

According to Table 3.4a, non-Ivorians are more than twice as likely not to be enrolled in primary school, assuming mean of other observed characteristics. This may reflect the disruption to schooling due to migration or indeed the school systems of their home countries rather than that in the Côte d'Ivoire, but the significance and magnitude of the differential is such as to be a cause for concern. As Table 3.4b shows, girls are particularly disadvantaged if they are non-Ivorians since this worsens their chances of enrollment by 20 percentage points compared with a 13 point loss for boys. That is to say, absolute gender differentials are greater amongst non-Ivorians. With relative differentials the story is less clear, since the proportions of girls to boys predicted to be in school and to be unenrolled are higher for non-Ivorians. This pattern of gender differentials—that both absolute differentials and relative differentials in the probabilities of enrollment vary positively with factors that increase the probability of enrollment; whereas the relative differential in the probabilities of non-enrollment—is repeated elsewhere in Table 3.4b. The only exceptions to this being the effects of paternal occupation and, interestingly, maternal education—the latter being predicted to reduce relative gender differentials in both the schooled and unschooled populations.

The only regional dummy variables to be significant enough to be retained in the final results were those for the urban-rural divide and for living in the far North of the country. In both cases, the effects on absolute gender differentials are sizable. At the (gender-specific) mean of other variables, coming from the far North reduces the probability of going to primary school by 0.46 percent if one is a girl and by 0.1 percent if one is a boy; whereas residing in a town increases girls' chances of enrollment by 0.7 percent and boys' by 0.3 percent.

As described in the section on explanatory variables, likelihood ratio tests were carried out to see which measure of a household's economic position—income per capita, total consumption per capita, and food consumption per capita—was the most appropriate as a determinant of demand for primary schooling. The test results are reported below and unambiguously pointed to the retention of income per capita only:

Retaining income per capita only: 9.6 (not reject at 5 percent)
Retaining consumption per capita only: 21.2 (reject at 5 percent)
Retaining food consumption per capita only: 34.4 (reject at 5 percent)

Subsequent likelihood ratio tests found that a cubic form for the effect of household income per capita could not be rejected in favor of a linear or quadratic one.

The effect of household income per capita on the probability of girls and boys not going to primary school when calculated at the means of the other regressors is plotted in Figure 3.1 using the results of the model when estimated separately for each sex. Thus calculated, higher income has a dramatic, although diminishing, effect in reducing non-enrollment for most of the ranges of income of relevance. At very high incomes—around 500,000 CFA Francs and above—the income effect becomes perverse for girls. Further, the relative gender differential in the probability of primary school non-enrollment when calculated at the mean of the other variables increases with income, as is plotted in Figure 3.1b. Since this pattern of relative differentials is consistent for the entire range of incomes, it cannot be wholly explained by the non-monotonicity in the income effect for girls. If one is prepared to use these cross-section findings to extrapolate about the effects of future changes in income, it appears that, without the recent introduction of compulsory primary schooling, general economic development by itself would have left the unschooled as an increasingly female dominated group.

Parental education also has strong effects: according to the calculations in Table 3.4 if either parent has some primary schooling, the probability of not enrolling evaluated at the mean of the other regressors is reduced almost fourfold. In absolute terms, these effects are particularly important for girls: having some parental education generally increases girls' chances of enrollment by around 20 percentage points compared with increases for boys of only a few points. As previously remarked, maternal education has a particularly strong effect in reducing gender differentials, the coefficient being insignificantly different from zero for boys. It is interesting that in preliminary estimates, cumulative dummy variables for the parents having completed primary school and having attained higher levels of schooling were all wholly insignificant. These findings about the effects of parental education on child enrollment seem explicable if it is assumed either:

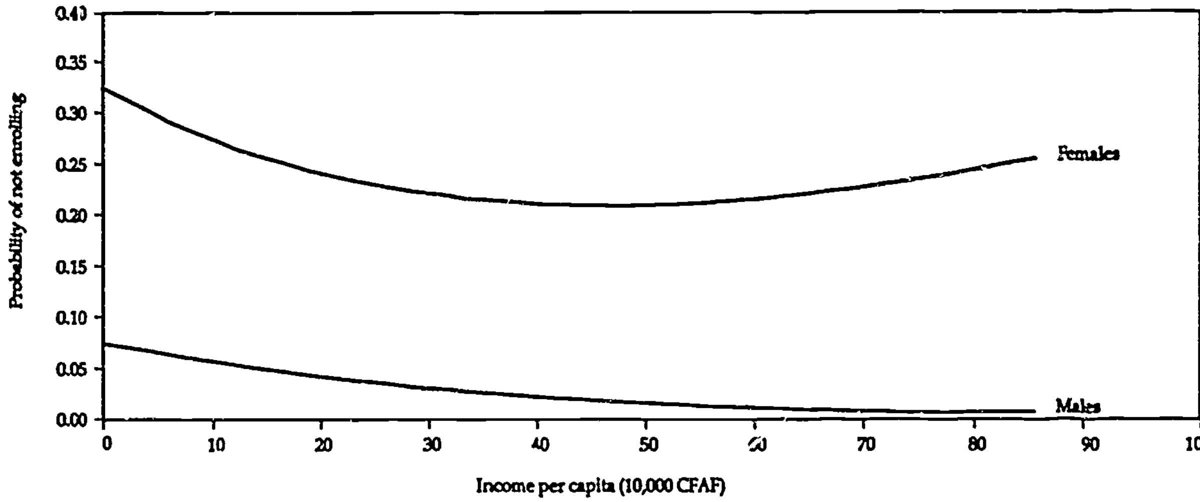
1) That the marginal effects of education on parents' ability to help their children profit from schooling decrease after the early grades of schooling. This attempt to rationalise the finding within the *human capital* framework seems to some extent implausible since it is often assumed that at least four years of schooling are necessary for basic literacy and numeracy.

or

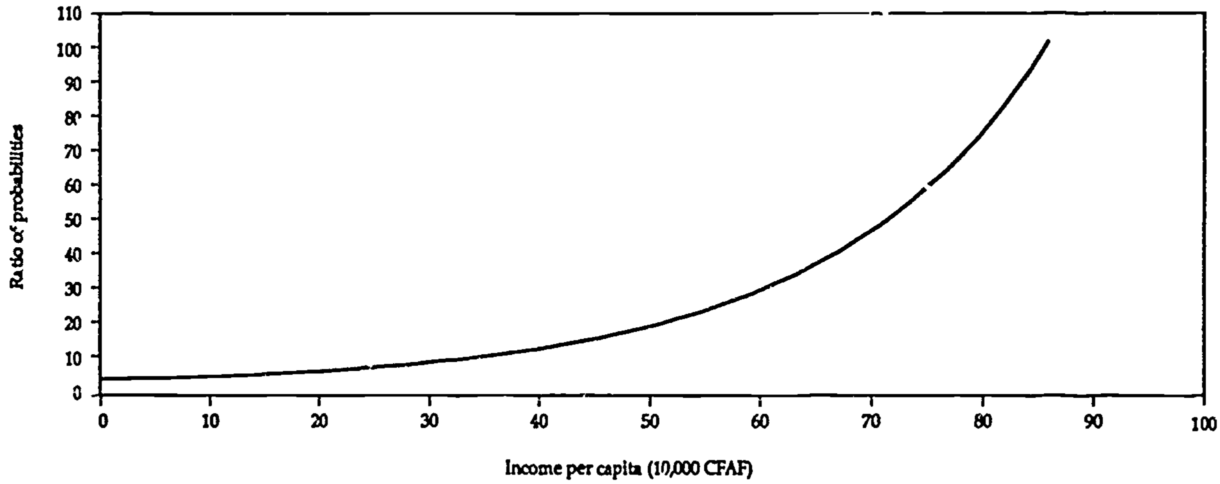
2) That the familiarity with primary schooling that parents have gained from some contact with it themselves is of particular importance in leading them to send their children to school: parents understand what is involved in the schooling process; feel it is less alien and so forth.

Figure 3.1 Income and Primary School Non-Enrollment*

a) Probability of Non-Enrollment by Gender



b) Ratios of Probabilities of Non-Enrollment: Girls to Boys



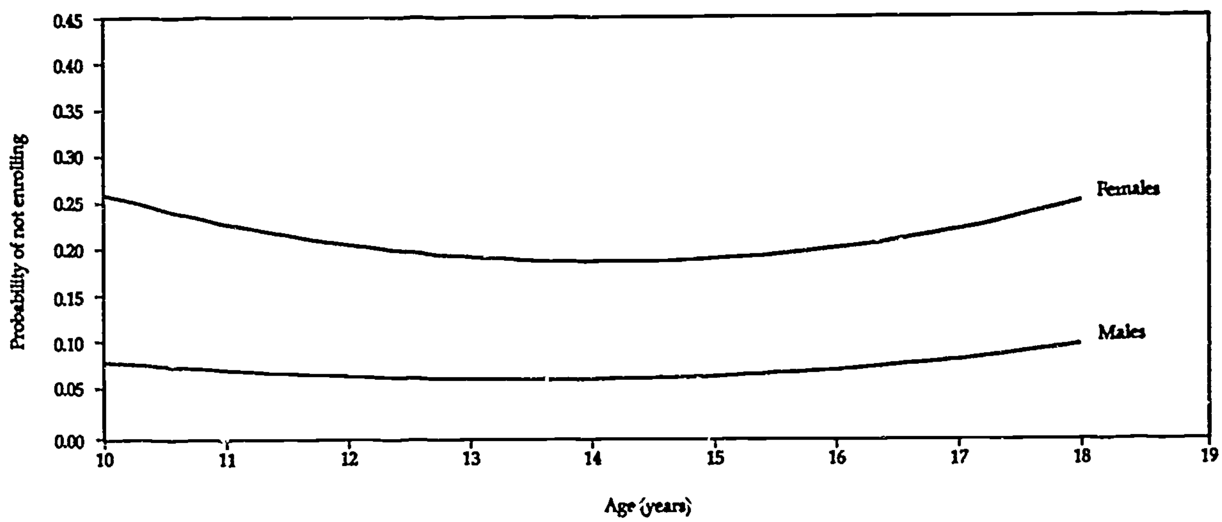
*Probabilities evaluated at the means of other regressors.

Figure 3.2a plots the gender specific effects of age on primary enrollment and reveals that younger children have less chance of enrolling than those toward the middle of the sample age distribution. This finding could reflect some residual right censoring of school attainment but given that no one in the sample aged 11 or over was observed entering primary school, this seems unlikely. The alternative explanation is that primary enrollment rates have been dropping in the early 1980's (World Bank (1988)). Once again, absolute gender differentials are least when total enrollment is high, although relative gender differentials in the probability of not enrolling in primary school, as plotted in Figure 3.2b, appear to have worsened over time, giving some credence to the earlier speculation about the effects of general improvements in income.

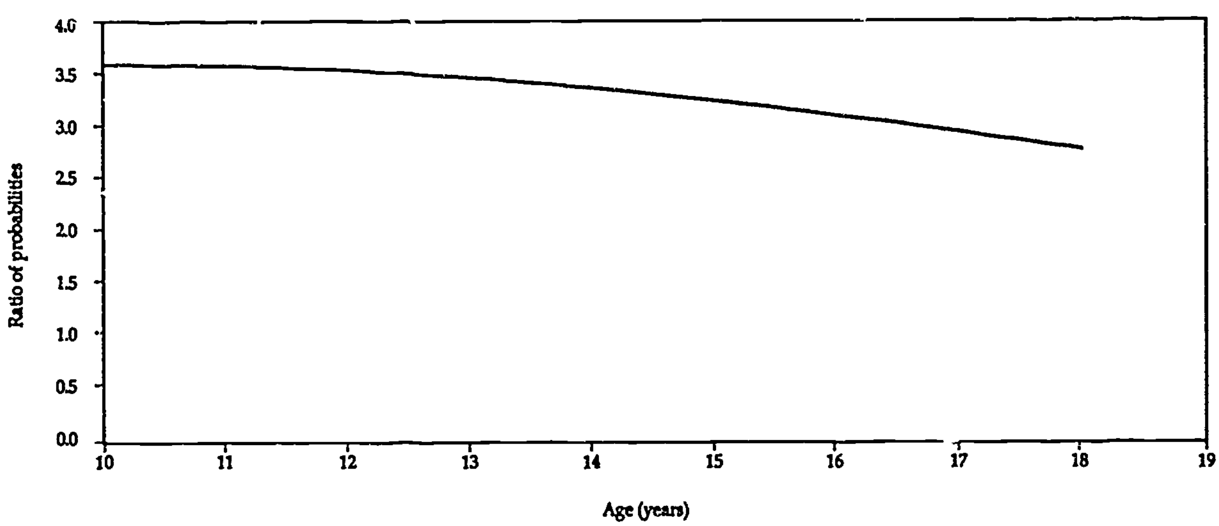
Livestock per capita has a perverse effect in the model but one which is only significant for boys. At

Figure 3.2 Age and Primary School Non-Enrollment*

a) Probability of Non-Enrollment by Gender



b) Ratio of Probabilities of Non-Enrollment: Girls to Boys



*Probabilities Evaluated at the Means of Other Regressors.

the mean probability, an increase in livestock per capita holdings equal to one standard deviation of their distribution across the sample increases the probability of non-enrollment by 4 percent. This finding could be rationalized as reflecting either:

1) Investment in an alternative asset to the human capital of one's offspring. The effect could be strongest for boys since their education may be more akin to an investment good for many families than is the schooling of girls.

or

2) A relatively high valuation of child labor due to the need for them to look after the livestock. This effect could be strongest for boys if tending livestock is more *boys' work* than girls.

Having a father who is in wage employment significantly increases the probability of being enrolled in primary school. Since we attempt to control for household income and parental education, this effect may reflect such fathers valuing education more on account of it being, as we have seen, an important determinant of access to the work they do. This effect is particularly strong for those with fathers in government employment. Splitting the sample, it can be seen that the effects of paternal wage employment are only well determined for girls.

3.4 Entry to Lower Cycle Secondary School

As explained in Section 3.1, entrance to each cycle of secondary school is modelled in two stages: the first investigates the determinants of success in the leaving exam of the previous stage of schooling and the second analyzes entry conditional upon such success. The sample used to investigate entry to lower secondary school consisted of those children aged 10 to 18 who were known to have completed primary school. The lower age limit is unimportant since no one below 10 was observed in secondary school, whilst the upper age limit was chosen to allow a reasonable sample size, to include decisions that took place fairly near to the time of the survey (some 18 year olds were just in the first year of lower secondary school at the time of the survey) and again to avoid the serious endogeneity problems connected with income mentioned in Section 3.2.

Passing the CEPE

To try to identify socioeconomic determinants of performance in the primary leaving exam, a logit for passing the exam was estimated over 10 to 18 year olds who were known to have completed primary school. As was suggested earlier, the logit model may be appropriate here, not only for its desirable statistical properties in handling discrete data, but also because the latent variable underlying it can be interpreted as the exam marks received in the CEPE transformed in such a way as to be zero at the critical passmark. The final form of the logit is given in Table 3.6.

The coefficient on the dummy variable for being female is significant and negative, implying that girls are less likely to pass the exam given the same measured characteristics as the boys. This corresponds with the findings of Grisay (1984) that girls in primary school in the Côte d'Ivoire perform less well than boys in all subjects and at all levels of schooling. As Grisay notes, this finding, whilst often discovered in developing countries, contradicts that generally found in developed countries, where girls typically perform better in some subjects such as languages and also benefit from a faster rate of maturity than boys. Grisay proposed several explanations of her findings in the

Table 3.6 Logit Estimates of Determinants of Passing the CEPE Exam, Conditional upon Completing Primary School: ages 10–18 years

Variable	*Coefficient*	*T-ratio*	*Mean of X*
CONSTANT	-0.6715	-2.347**	1.000
FEMALE	-0.8965	-4.658***	0.393
NONIVORIAN	-1.311	-3.164***	0.046
WEST	0.7434	2.602***	0.189
NORTH	1.345	2.174**	0.028
URBAN	0.8380	3.211***	0.621
DADPRIM	0.7305	2.989***	0.376
MUMPRIM	1.144	1.471	0.139
BOTHPRIM	-2.000	-2.348**	0.115
CONPC1	0.7721E-01	3.860***	28.015
CONPC2	-0.8489E-03	-2.850***	1,604.426
CONPC3	0.2629E-05	2.331**	165,829.972
Log-likelihood	-351.58		

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability

		Predicted	
	TOTAL	*Fail*	*Pass*
Total	667	89	578
Actual			
Fail	194	63	131
Pass	473	26	447

** = significant at 5 percent
*** = significant at 1 percent

Table 3.7 Effect on Probability of Passing the CEPE of the Nought-One Dummy Variables: Full Sample and Own Gender Only Results (percentage)
(Evaluated at means of other explanatory variables; gender specific means in part b)

(a) When estimated for full sample:

	Probability of Passing CEPE
No Change, all variables at means	80
Boy	85
Girl	70
Ivorian	81
Non-Ivorian	53
Western Home	81
Northern Home	88
Urban Home	82
Other Home	66
Neither Parent Primary Schooling	76
Father Some Primary Schooling	87
Mother Some Primary Schooling	91
Both Parents Some Primary Schooling	74

(b) When estimated for single sex samples:

	Probability of Passing CEPE	
	Girls Only	Boys Only
No Change, all variables at means	77	84
Ivorian	78	85
Non-Ivorian	54	54
Western Home	76	85
Northern Home	92	88
Urban Home	79	87
Other Home	63	73
Neither Parent Any Primary Schooling	73	78
Father Some Primary Schooling	86	87
Mother Some Primary Schooling	77	100
Both Parents Some Primary Schooling	72	72

Côte d'Ivoire that girls perform less well than boys in primary school, which are:

i) girls' performance is adversely affected by the low parental demand for their schooling;
ii) there are very few female teachers in Ivorian schools;
iii) girls are disadvantaged by the practice of being taught French in school since they are less likely to talk the language at home than boys. This finding, which emerged from Grisay's survey of French primary school pupils, was explained by Grisay as reflecting the fact that the only adults who can speak French are those who are educated, the majority of whom are men and such men—fathers and uncles—may be more likely to talk to boys than girls;
iv) girls ask less questions in class than boys. Grisay explained this finding by arguing that girls are restrained from actively participating in class by social stereotypes of how a girl should behave; they may fear parents or teachers believing them insolent or gossips.

The likelihood ratio test for whether the logit should be run separately on the subsamples of girls and boys gave a statistic of 10.7, accepting the restriction of a single sample at 10 percent. Thus the model used here for CEPE performance identifies a *pure* gender differential which varies little with the other explanatory variables. The results of the model when estimated separately for each gender are given in Table 3.8. Table 3.7 gives the effects of the explanatory variables on the probability of passing the CEPE evaluated at the means of the other variables.

Household consumption per capita rather than income per capita was selected as an explanatory variable because of the likelihood ratio test statistics given below for retaining one of the following three measures and rejecting the other two (all were entered in cubic forms):

Table 3.8 ·Logit Estimates of Determinants Passing CEPE Conditional Upon Completing Primary School: ages 10-18 years; by sex

	Girls Only		*Boys Only*	
Variable	*Coefficient*	*T-ratio*	*Coefficient*	*T-ratio*
CONSTANT	-2.231	-4.283***	-0.2126	-0.575
NONIVORIAN	-1.139	-1.945**	-1.577	-2.602***
WEST	0.6379	1.233	0.7881	2.204**
NORTH	1.878	1.486	1.049	1.502
URBAN	0.7974	1.681*	0.8920	2.741***
DADPRIM	0.8125	2.272**	0.6010	1.772*
MUMPRIM	-0.2503	0.288	15.50	0.012
BOTHPRIM	-1.102	-1.117	-16.42	-0.012
CONPC1	0.1278	4.013***	0.3840E-01	1.370
CONPC2	-0.1523E-02	-3.331***	-0.3059E-03	-0.662
CONPC3	0.4813E-05	2.832***	0.1021E-05	0.533
Log-likelihood	(Girls)	-145.11	(Boys)	-201.10

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability

		Girls Only Predicted	
	TOTAL	*Fail*	*Pass*
Total	262	82	180
Actual			
Fail	98	56	42
Pass	164	26	138

		Boys only Predicted	
	TOTAL	*Fail*	*Pass*
Total	405	13	392
Actual			
Fail	96	8	88
Pass	309	5	304

** = significant at 5 percent
*** = significant at 1 percent

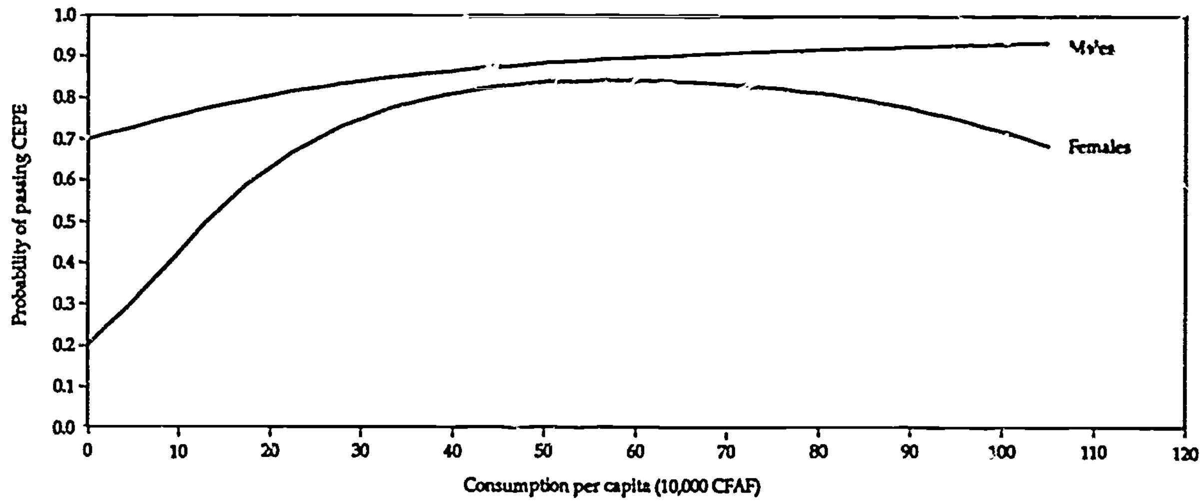
Retaining income per capita only: 21.4 (reject at 5 percent)
Retaining consumption per capita only: 12.4 (not reject at 5 percent)
Retaining food consumption per capita only: 23.5 (reject at 5 percent)

When taken together with the results of similar likelihood ratio tests reported for the other models as discussed in this chapter, these results are part of a pattern; namely that total consumption per capita is the most suitable of the three proposed measures of a households economic status for explaining exam performance whilst income per capita is a better predictor of entry to school (conditional on the individual having the appropriate exam passes in the case of secondary school cycles). This is what one might expect a priori: for income to affect exam performance it must be spent, whether on educational material or just on attaining a comfortable home environment. By contrast, following the human capital theory of schooling, expenditure on education is at least in part an investment and thus a function of savings, the excess of income over consumption.

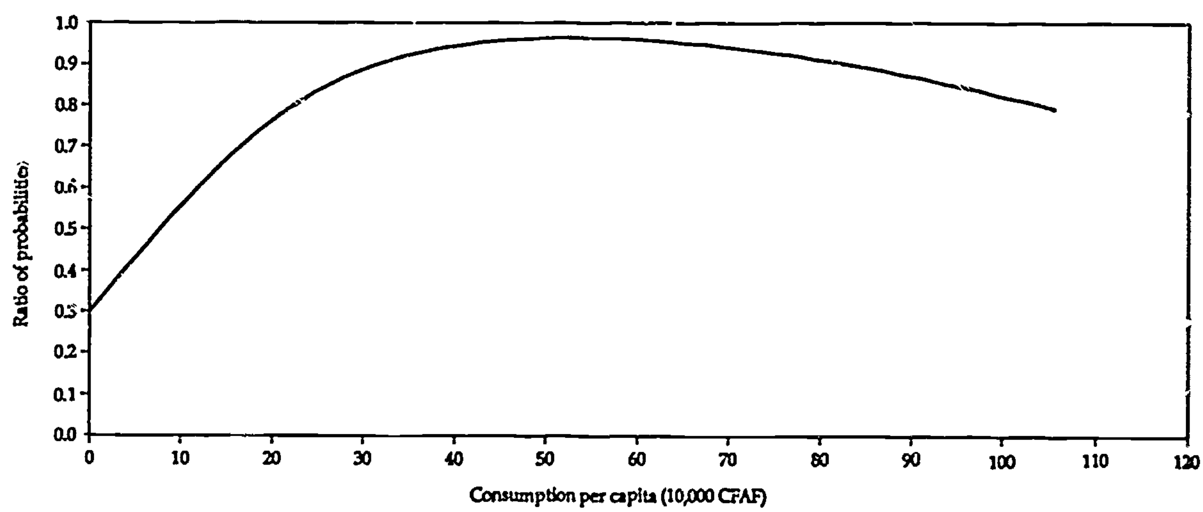
Total consumption per capita affects CEPE performance in a similar way to that in which income per capita affected primary school enrollment. This can be seen in Figure 3.3, which plots consumption per capita against the corresponding probabilities, calculated

Figure 3.3 Consumption and CEPE Performance*

a) Probability of Passing the Exam: by Gender



b) Ratio of Probabilities of Passing CEPE: Girls to Boys



*Probabilities evaluated at the means of the other regressors.

from Table 3.8, for girls and boys of passing the CEPE when evaluated at the means of the other explanatory variables. Specifically, there is a powerful but diminishing positive effect for most of the range of values of consumption per capita of interest, but once again for relatively well-to-do families the effects on girls are perverse. These effects are perverse only for girls from households that consume around 600,000 CFA Francs per capita per annum. Consumption has more significant and sizable effects on girl's exam performance than it does on boys; so much so that Figure 3.3 reveals that the relative gender differential in the probability of passing the exam predicted at the means of the other variables falls with consumption per capita until the level at which the possibly spurious non-monotonic consumption effects for girls begin. Thus at low-to-medium levels of consumption, girls perform worse than boys whilst at medium-to-fairly high levels they almost equal them.

If either parent has received some schooling, a child's chances in the CEPE are improved, although the coefficient for maternal education is insignificant

Table 3.9 Multinomial Logit Estimates of Determinants of Enrolling in Lower Secondary School Conditional Upon Passing the CEPE: ages 10-18 years

Variable	*Public School* Coefficient	T-ratio	*Private School* Coefficient	T-ratio	Mean of X
CONSTANT	-28.933	-3.380***	-20.823	-2.157**	1.000
FEMALE	0.9539E-01	0.270	0.8408	2.074**	0.347
AGE1	3.332	3.026***	1.741	1.401	15.903
AGE2	-0.1047E-01	-2.895***	-0.5023E-01	-1.229	256.038
NONIVORIAN	-1.164	-1.396	-0.4932	-0.512	0.030
NORTH	-1.377	-2.099**	-1.489	-1.332	0.032
ABIDJAN	-0.4764	-1.114	0.87852	1.949*	0.228
MUMPRIM	1.251	1.893*	1.180	1.669	0.150
DADPRIM	0.5663	1.449	0.6564	0.148	0.425
ESTCEPE1	2.051	1.604	4.062	2.505**	0.754
INCPC1	0.1061	2.898***	0.1210	2.732***	46.505
INCPC2	-0.1027E-03	-2.222**	-0.1340E-02	-2.478**	2,920.443
INCPC3	0.2713E-05	1.703*	0.3841E-05	2.143**	261,647.950
Log-Likelihood	-370.29				

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability
Lower Secondary Schooling

	TOTAL	*Predicted* None	Public	Private
Total	473	25	407	41
None	76	15	55	6
Actual				
Public	302	8	277	17
Private	95	2	72	18

* = significant at 10 percent
** = significant at 5 percent
*** = significant at 1 percent

despite being larger than that for paternal education and possibly due to its collinearity with the interaction term for whether both parents have received some schooling. As might be expected, this interaction term is negative, but it is perverse that its effects are such as to outweigh either of the other two parental education dummies individually and indeed collectively.

The children of non-Ivorians are less likely to pass the CEPE if they complete primary school, as are children from rural areas, with the exception of those from the North. This exception may reflect a strong sample selection effect: since primary school enrollment is, as we have seen in the previous section, significantly lower in the North, only the very able receive schooling. This suggests that policies targeted at increasing enrollment in the North which are desirable on equity grounds, and from Table 3.4b potentially especially beneficial to females, need not lower *academic standards.*

Entry to Lower Secondary School Conditional on Passing the CEPE

A multinomial logit with three outcomes—no secondary schooling at all, some at a public school and some at a private school—was estimated over all those aged 10-18 who had passed the CEPE and were known not to be still at primary school.

The results of the logit are reported in Tables 3 9 and 3.10 and confirm the initial impression from the descriptive statistics in Table 3.1 that girls suffer from failure to pass the CEPE rather than low demand for lower secondary schooling conditional upon passing the exam. Indeed, the female dummy variable is positive in both its appearance and significantly so in its effect on the probability of going to a private secondary school. However, this is not to say that low demand by parents for the schooling of girls as opposed to boys is unimportant since, as has been shown, low

Table 3.10 Effect on Probability of Going to Lower Secondary School of the Nought-One Dummy Variables: Full Sample and Own Gender Only Results (percentage)
(Evaluated at means of other explanatory variables; gender specific means in part b)

a) When estimated for full sample:

	Probability of Going to:		
	State School	*Private School*	*No School*
No Change: all variables evaluated at means	75	18	7
Girl	68	27	
Boy	78	15	
Non-Ivorian	57	27	16
Ivorian	75	18	7
Northern Home	68	11	21
Abidjan Home	55	38	7
Other Home	80	14	6
Mother Some Primary	78	19	2
Mother No Primary	74	18	8
Father Some Primary	80	15	5
Father No Primary	71	21	8

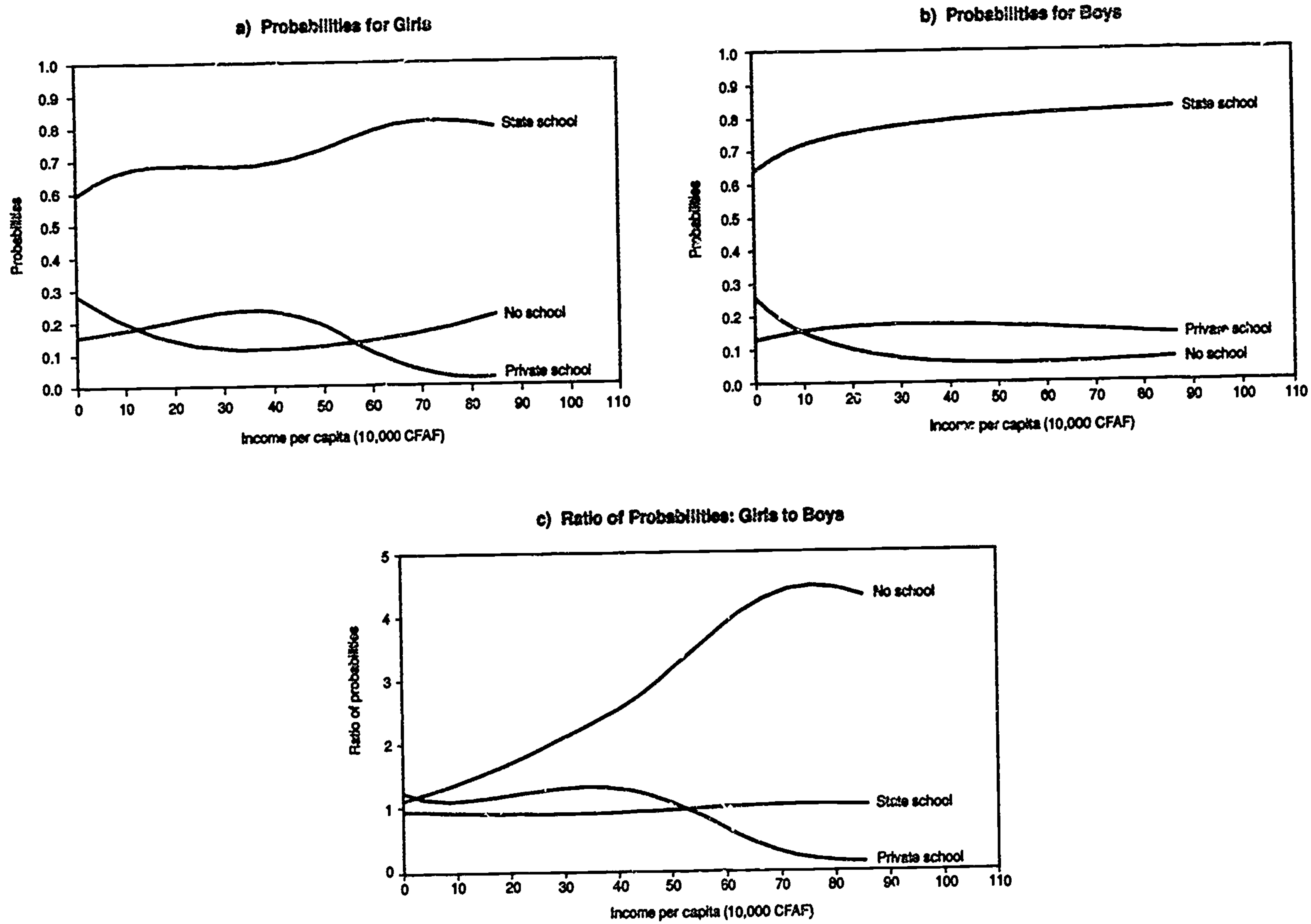
b) When estimated for single sex samples:

	Girls Only			*Boys Only*		
	State	*Private*	*No School*	*State*	*Private*	*No School*
No Change	69	21	11	79	17	5
Non-Ivorian	23	27	50	81	13	6
Ivorian	70	20	10	78	17	5
Northern Home	50	0	50	69	18	13
Abidjan home	49	45	6	53	41	7
Other Home	68	21	11	83	13	4
Mother some Primary	74	22	3	81	17	2
Mother no Primary	66	20	14	78	17	6
Father some Primary	73	17	10	83	14	3
Father no Primary	62	26	11	75	18	7

demand is responsible for less girls entering primary school and thus being able to sit the CEPE and, as has been suggested by Grisay, low demand may affect motivation and thus exam performance.

The reason why the female dummy is significantly positive in its effect on the probability of going to private school is hard to determine. There are two possible classes of explanations: arguments to do with rationing, suggesting that girls are more likely to go to private schools because they cannot get into state schools, and arguments about demand, positing that the demand for private schooling is greater for girls than for boys. The former could be stated thus: girls are more likely to go to private schools than boys because even if they pass the CEPE their marks are generally lower than those of boys and good CEPE marks rather than mere passes are often required for admission to state schools. The second argument could be valid if, for example, private schools are more likely to be single sex than state schools and if parents are more concerned that their daughters go to single sex schools than that their sons do. However, most private and public secondary schools in the Côte d'Ivoire are mixed. Various conflicting considerations can be put forward for and against these two hypotheses. One is that the probability of girls going to private schools falls at high incomes. This is shown in Figure 3.4a, which plots the effect of income on the probability of girls receiving lower secondary schooling, evaluated at the means of the other regressors and using the results of the model when estimated solely on the subsample of girls as reported in Table 3.11. This finding runs counter to the idea that girls are more likely to go to private schools for demand reasons since one would expect the demand for relatively expensive private schooling to rise with income. Furthermore, the finding is consistent with the hypothesis that girls are more likely to be rationed out of the state schools even if they pass the CEPE because, as the previous section showed, it is at low incomes (or low levels of consumption at least) that girls' performance is inferior to that of boys. Against this argument, one would expect the variable for predicted exam performance to control for rationing effects. However, the

Figure 3.4 Income and Lower Secondary School Choice*



* Probabilities evaluated at the means of other variables.

Table 3.11 Multinomial Logit Estimates of Determinants of Enrolling in Lower Secondary School Conditional On Passing the CEPE: ages 10-18 years; by sex

	Girls		*Boys*	
Variable	*Coefficient*	*T-ratio*	*Coefficient*	*T-ratio*
Public School				
CONSTANT	9.747	0.433	-39.380	-3.719***
AGE1	-1.810	-0.618	4.677	3.480***
AGE2	0.6596E-01	0.695	-0.1497	-3.398***
NONIVORIAN	-2.789	-2.023**	-0.2133	-0.158
NORTH	-1.838	-1.178	-1.312	-1.741*
ABIDJAN	0.2855	0.376	-0.9499	-1.708*
MUMPRIM	1.558	1.379	1.299	1.440
DADPRIM	0.2714	0.409	0.9526	1.856*
ESTCEPE2	1.638	0.980	1.926	0.980
INCPC1	0.1109	1.806*	0.1311	2.3[illegible]3**
INCPC2	-0.1336E-02	-1.411	-0.1207E-02	-1.849*
INCPC3	0.4581E-05	1.111	0.3001E-05	1.450
Private School				
CONSTANT	29.018	1.155	-36.887	-2.711***
AGE1	-4.506	-1.383	3.801	2.198**
AGEY2	0.1535	1.453	-0.1168	-2.070**
NONIVORIAN	-1.342	-1.013	-0.4921	-0.299
NORTH	-16.370	-0.012	-0.7745	-0.659
ABIDJAN	1.355	1.725*	0.6733	1.143
MUMPRIM	1.618	1.362	1.167	1.182
DADPRIM	-0.3355	-0.449	0.6242	1.058
ESTCEPE2	3.125	1.487	3.024	1.198
INCPC1	0.3033E-01	0.274	0.1550	2.371**
INCPC2	0.1227E-02	0.459	-0.1537E-02	-2.089**
INCPC3	-0.1804E-04	-0.912	0.4119E-05	1.823*
Log-likelihood	(Girls Only)	-129.68	(Boys Only)	-230.21

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability

Girls Only			*Predicted Lower Secondary Schooling*	
	TOTAL	*None*	*Public*	*Private*
Total	164	14	125	25
None	26	8	17	1
Actual				
Public	95	5	79	11
Private	43	1	29	13

Boys Only			*Predicted Lower Secondary Schooling*	
	TOTAL	*None*	*Public*	*Private*
Total	309	20	273	18
None	50	10	36	4
Actual				
Public	207	7	194	6
Private	52	3	43	6

* = significant at 10 percent
** = significant at 5 percent
*** = significant at 1 percent

variable is far from a perfect predictor of exam performance since it is generated without using any direct measure of intelligence or primary school quality. A final consideration is that if the rationing explanation was correct, it would imply a negative coefficient on the female dummy variable in the government equation rather than a positive one in the private equation. Consequently, there is an unresolved question over whether the fact that girls are less likely to enrol in state schools conditional on having passed the CEPE is d' to their not being eligible to enrol or because then parents prefer to school them privately.

The likelihood ratio test statistic for whether the model should be estimated separately for the two sexes is 21.74 which is insignificant at 10 percent with 24 degrees of freedom. This implies that overall the interactions between the effects of gender and other explanatory variables are not so large as to reject the gender unspecific model. (For the purposes of the test the gender unspecific ESTCEPE1 was used rather than ESTCEPE2). However, as recorded in Table 3.11, several variables do exhibit differential effects by gender, one of which is income.

Likelihood ratio tests were performed to see whether household income, consumption or food consumption per capita should be included in the model. All were entered in cubic forms. The results were:

Retaining income per capita only. 11.7 (not reject at 5 percent)
Retaining consumption per capita only. 20.2 (reject at 5 percent)
Retaining food consumption per capita only: 24.9 (reject at 5 percent)

Following these results, only the income per capita cubic terms were entered. It can be seen that at low incomes (under 20,000 CFA Francs per capita per annum) there are powerful income effects increasing the probability of going to lower secondary school. However, as plotted in Figure 3.4c, relative gender differentials fall with income and once more the effect of income becomes perverse at high levels for girls. Furthermore, as has been noted above, the probability of girls going to a private school falls for incomes above the mean.

Another variable of interest is the probability of an individual passing the CEPE. This is entered as a proxy for how well individuals performed in the CEPE. If the private sector caters for those who have not performed well enough in the exams to be included in the ration of state school places one would expect the probability of passi g the CEPE to have a strong positive effect on the probability of going to a state school and a weaker—and maybe even negative—effect on the probability of going to a private school. However, although positive in all appearances, the variable is only significant in its effect on the probability of going to a private lower secondary school. Indeed, using the esults of Table 3.9 and evaluating at the mean probabilities, a 10 percent rise in the expected probability of passing the exam increases the likelihood of successful CEPE candidates going to private lower secondary school by 38 percent whilst the likelihood of their going to state lower secondary school rises by only 0.5 percent. The reason why improved exam performance increases the probability of going to a private lower secondary school are the same as those suggested in the overview to account for the failure of the private sector to cater for those CEPE failures rationed out of the state school system. That is to say, private schools may ration entry by the exam to try to maintain a reputation for quality, or, alternatively, it may be that parents believe it worthwhile spending large sums of money on private school fees only if their child is a particularly promising student.

The dummy variable for some maternal education dummy in Table 3.9 has both larger coefficients and higher t-ratios than that for paternal education As can be seen from Table 3.10b, its effects for girls are more marked than those for boys: at the means of the other regressors, having a mother who has received some primary schooling cuts the conditional probability of a not being enrolled in lower secondary school by 11 percentage points for girls but only 4 points for boys By contrast paternal education is more sizable and significant as a determinant of boys' rather than girls' enrollment.

Evaluated at the means of the other variables, living in Abidjan more than doubles ones' chances of going to a private school. This could be interpreted as proxying higher rates of return to secondary schooling due to greater access to the wage employment sector, but the insignificance of a dummy variable for living in a urba. ea in preliminary estimates seems to contradict this interpretation. More likely, the variable simply reflects the greater availability of private schools in Abidjan (private schools would tend to set up in areas of greatest population density ceteris paribus since distance from the school is likely to reduce demand). This effect is significant only for girls.

Despite their superior performance in the CEPE, primary school leavers from the North are much less likely to enrol: an effect which could reflect lo returns to education in this less developed region alternatively, lower availability or attractiveness (d e to quality or distance from home) of secondary schools; that the variable has a more negative and significant coefficient in the state school equation inclines one to the latter explanation. Girls are at a par-

Table 3.12 Logit Estimates of Determinants of Passing the BEPC Exam Conditional Upon Completing Lower Secondary School: age 30 and under

Variable	Coefficient	T-ratio	Mean of X
CONSTANT	0.6436	3.096***	1.000
FEMALE	-0.4820	-1.771*	0.295
MUMCEPE	-1.301	-1.325	0.033
MUMBEPC	1.756	1.202	0.020
DADCEPE	0.7696	1.649*	0.270
DADBEPC	-0.5781	-1.064	0.153
CONPC1	0.1204E-01	2.569**	40.775
Log-likelihood	-199.25		

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability

	Predicted		
	TOTAL	Fail	Pass
Total	359	1	358
Actual			
Fail	94	1	93
Pass	265	0	264

* = significant at 10 percent
** = significant at 5 percent
*** = significant at 1 percent

ticular disadvantage if they come from the North: Table 3.10 shows that—evaluated at the means of other regressors—girls from the North have a 50 percent chance of not enrolling in lower secondary school if they pass the CEPE whereas boys have only a 13 percent chance. This may just be consistent with the evidence across time and across countries of gender differentials being highest in areas of low educational development; but could also reflect a reluctance of the predominantly Muslim area to send their girls to secular state schools. Being a non-Ivorian is also a major disadvantage to girls—increasing the probability of their not going to lower secondary school conditional on passing the CEPE fivefold when evaluated at the mean of the other regressors—whereas it has a positive although insignificant effect on boys.

3.5 Entry to Upper Cycle Secondary School

Access to the upper cycle of secondary school was analyzed using the same framework as that for access to the lower cycle. That is to say passing the leaving exam for the previous cycle of schooling, in this case the BEPC, was assumed to be a prerequisite for entry to the next grade and a logit for performance in this exam was estimated for those within a certain age range who had completed the last grade of the previous cycle. The subsample of successful BEPC candidates was then used to estimate the determinants of the choice of whether to enter the upper secondary school and if so whether to go into the private or the public sector. This last choice was modelled using a multinomial logit.

The age range chosen for this analysis was rather liberal with an upper age limit of 30. This was partly due the relatively late ages at which Ivorians enter upper secondary school—those observed in the first grade of the upper secondary cycle were aged between 18 and 24. Attempts were made to estimate the models on a sample of those under 25 but this yielded very poor results, with few explanatory variables significant at the 10 percent level. This was probably due in part to the very small size of the subsample that resulted. However, the doubling of the sample size by extending the upper age limit to 30 renders suspect some of the results that follow due to problems of measurement error and of endogeneity mentioned earlier. In particular, it is likely that those who have completed lower secondary school and are aged around 30 will be contributing a substantial amount to household income—and thus affecting household consumption. The effect of this on the behavior of household income as a determinant of entry choice is unclear although it is likely that those who were enrolled in upper secondary school will eventually earn more than those who were not (after all this may be why they chose to enter), it is also clear that in the

Table 3.13 Multinomial Logit Estimates of Determinants of Enrolling in Upper Secondary School Conditional Upon Passing the BEPC: age 30 and under

	Parameter Estimates				
	Public School		*Private School*		
Variable	*Coefficient*	*T-ratio*	*Coefficient*	*T-ratio*	*Mean of X*
CONSTANT	4.650	0.566	-16.987	-1.314	1.000
FEMALE	-0.7821	-2.192**	-0.6650	-1.247	0.279
AGE1	-0.1173	-0.174	1.736	1.576	24.109
AGE2	-0.1984E-02	-0.[illegible]5	-0.4318E-01	-1.851*	595.098
ABIDJAN	-0.6580	-1.977**	-0.3477	-0.732	0.426
MUMPRIM	0.4332	0.503	1.783	1.673*	0.075
DADSEC	-1.259	-2.338**	-1.895	-2.196**	0.147
DADWAGED	1.066	2.310**	0.7011	1.071	0.313
INCPC1	0.1844E-01	2.877***	0.6193E-02	0.602	38.772
Log-likelihood	-211.83				

Frequencies of actual and predicted outcomes
Predicted outcome has maximum probability

		Predicted Upper Secondary Schooling			
		TOTAL	*None*	*Public*	*Private*
Total		265	50	214	1
	None	73	31	42	0
Actual	Public	159	15	144	0
	Private	33	4	28	1

* = significant at 10 percent
** = significant at 5 percent
*** = significant at 1 percent

short run attendance in school will prevent individuals earning a sizable income and even after leaving school may leave them lower on the seniority ladder than their less educated peers.

Passing the BEPC

The results of the logit in Table 3.12 for the determinants of success in the BEPC are interesting in that the results reinforce the conclusions drawn from the analysis of CEPE success as far as gender and consumption are concerned.

In particular, the most significant variable in the model is the gender dummy. The coefficient on the female variable implies that at the mean of the other variables, females have only a 44 percent chance of passing the exam whereas males have a 63 percent chance. The likelihood ratio test statistic for restricting the model to being run on the whole sample, as opposed to be separately estimated for each sex, was 1.66 and thus the restriction to a single sample was not rejected at 5 percent significance with 6 degrees of freedom. However, the coefficients on the household consumption per capita did differ between the sexes, although both were significant. They were:

Subsample	Coefficient	T-statistic	Mean of CONPC1
Males	0.9525E-02	1.633	36.870
Females	0.1510E-01	1.945	50.094

Thus if they are still in the school system at the end of the lower secondary cycle, girls on average are likely to come from better off households than boys. Furthermore, whether their households are indeed well off is more important in determining whether they pass the BEPC for girls than it is for boys.

As with the determinants of CEPE success, consumption per capita was the preferred measure of a household's economic status with the likelihood ratio tests for whether income, consumption or food consumption should be entered as explanatory variables giving the following results:

Retaining income per capita only 4.8 (reject at 10 percent)
Retaining consumption per capita only. 0.2 (not reject at 5 percent)

Table 3.14 Effect on Probability of Going to Upper Secondary School of the Nought-One Dummy Variables: Full Sample Results (percentage)
(Evaluated at means of other explanatory variables)

	Probability of Going to:		
When estimated for full sample:	*State School*	*Private School*	*No School*
No Change	62	17	21
Girl	53	15	32
Boy	65	17	18
Abidjan Home	55	17	28
Other Home	67	16	17
Father some Secondary	46	7	47
Father no Primary	63	19	18
Mother some Primary	46	43	11
Mother no Primary	63	15	22
Father in Employment	73	15	12
Father not in Employment	56	17	27

Retaining food consumption per capita only: 5.0 (reject at 10 percent)

Further likelihood ratio tests did not reject the hypothesis of consumption per capita entering linearly.

As with the determinants of the model for CEPE success, the effects of parental education are perversely non-monotonic.

Entry to Upper Secondary School Conditional on Passing the BEPC.

The multinomial logit estimated for those successful BEPC candidates is presented in its final form in Table 3.13. Unlike the multinomial logit for entry to lower secondary school conditional on having passed the CEPE, the equation for upper secondary school entry conditional on obtaining the BEPC shows that girls are at a disadvantage in entering both state and private sectors, although only significantly so as regards the former sector. The percentage effects of the explanatory variables that were entered into the final form of the model are shown in Table 3.14, evaluated at the means of the other regressors. Using this measure, and thus controlling for other observed characteristics such as the higher household income of female successful BEPC candidates, being female makes one almost twice as likely not to be enrolled in upper secondary school. The difference in these findings about gender differentials in access to the two cycles of lower secondary school may be accounted by one or more of the following hypotheses:

1) The age range for the sample used for entry to upper secondary school is more extended than that for entry to lower secondary school and thus will reflect enrollment decisions made longer ago. This may explain the different findings since gender differentials in secondary school enrollment have declined over time as total enrollment has risen.

2) The period of 18-24 when people enter upper secondary school is very likely to coincide with the time many young women get married and stay at home to look after young children, something which will prevent them attending school.

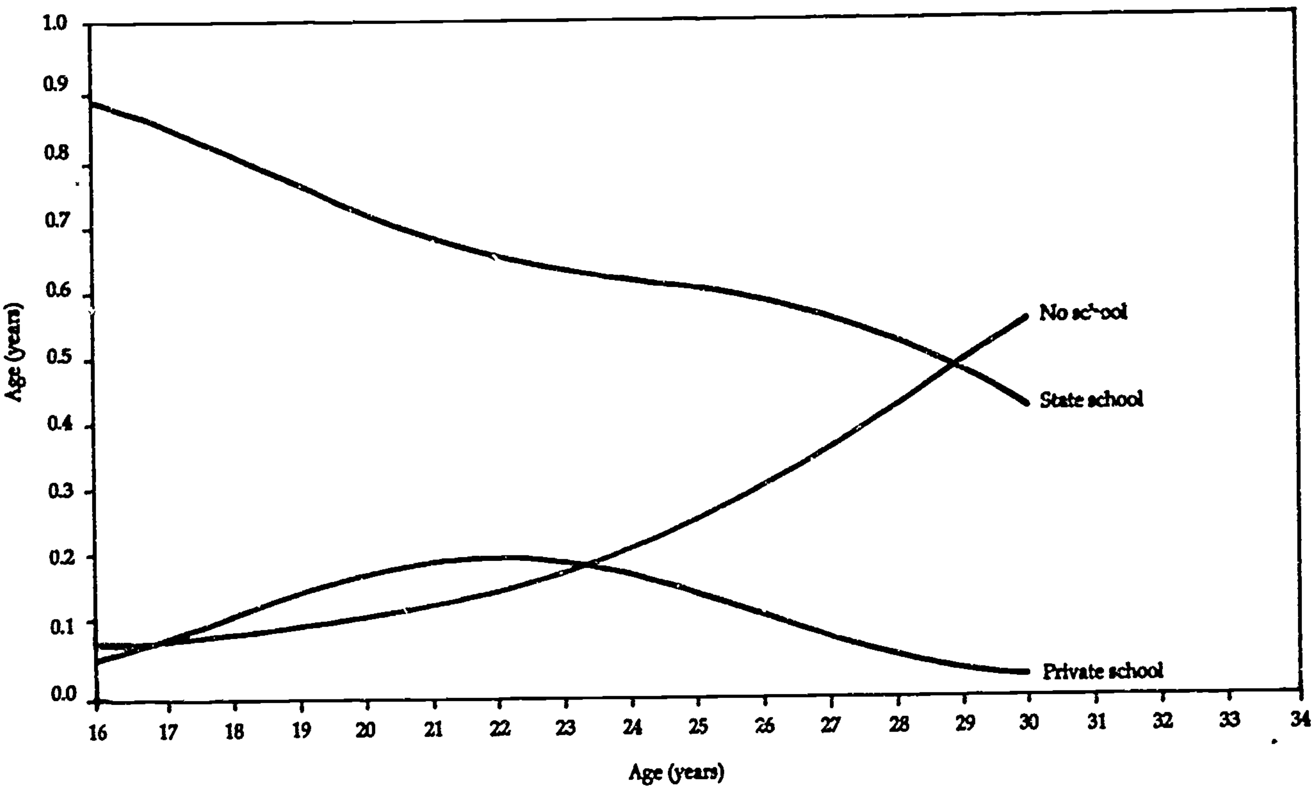
3) Upper secondary school may be valued more in terms of the economic advantages it may bring than are lower levels of schooling: in particular, its costs—both direct and opportunity costs—are likely to be higher, but so is its association with remunerative wage employment. Assessing schooling in these terms may be to the disadvantage of girls if there is discrimination or other reasons why they are less likely to participate in the wage sector.

An attempt was made to estimate the model separately for girls and for boys, but the iterative program for estimation failed when it yielded parameter values implying taking a logarithm of a non-positive value. Attempts to investigate gender differentials by use of interaction terms did not yield any significant results, possibly due to the small numbers involved.

The results of the likelihood ratio tests for which measure of household income or consumption per capita should be used gave the following results:

Retaining income per capita only: 6.4 (not reject at 5 percent)
Retaining consumption per capita only: 12.9 (reject at 5 percent)
Retaining food consumption per capita only: 13 [illegible]
(reject at 5 percent)

Figure 3.5 Age and Upper Secondary School Choice*



*Probabilities evaluated at the means of other regressors.

These test results unambiguously find that income per capita is a preferable explanatory variable to the other two measures. Per capita income has a significant positive effect in the state schooling equation. Evaluating at the mean probability, increasing household per capita income by 100,000 CFA Francs per annum increases the probability of a successful BEPC candidate of enroling in a state upper secondary school by 4 percent, although the probability of going private falls by 1 percent.

Having a mother with some education or a father in employment both significantly increase the conditional probability of upper secondary enrollment. However, the only significant paternal education and regional dummy variables have perverse effects, possibly due to sample selection problems.

Finally, it should be noted that a likelihood ratio test for rejecting the quadratic age terms in Table 3.13, which were not individually significant, rejected the hypothesis that its joint effects were insignificant at 5 percent. As Figure 3.5 shows, the effects implied by the quadratic at the mean of the other explanatory variables are very marked. In particular, the probability of not having gone to upper secondary school is over 60 percent for 30 year olds but falls to less than 10 percent for 18 year olds. This improvement in educational provision is largely in the form of increased state school places and there is no evidence of falling state school provision in the recent years of structural adjustment as there was with primary schooling.

4. Policy Implications

Recall that the basic phenomena we are analyzing are that women are far less likely than men to participate in the labor market, and that girls are far less likely to be educated than are boys. We have shown that the *benign* explanation for this in terms of female specialization in child rearing does not hold. The gender bias is a problem which reflects either differentially low aspirations on the part of females or discrimination on the part of employers and possibly parents. What can public policy achieve to alleviate this problem? We consider possible interventions in the labor market and in the education system in turn.

In developed countries public interventions in the labor market have focused upon equal pay legislation and upon the encouragement of provisions to make child rearing compatible with employment. Neither of these interventions appear to be particularly appropriate for the Côte d'Ivoire. Once in the labor market, women earn equal pay to that of men, controlling for their characteristics. Women's participation in the labor market, though low, does not appear to be explained by child rearing obligations. Evidently the extended family functions well as a system of child care. Nevertheless, women are less likely than men with similar characteristics to enter the labor market and one interpretation of this is that employers discriminate against women. By disaggregating the labor market as between the public and private sectors and, within the latter as between the unionized and non-unionized components, we were able to trace the phenomenon of differential recruitment to the private unionized sector. Within the public sector there was no evidence of gender discrimination and the private, non-unionized sector is not statistically distinguishable from the public sector. Hence, public sector recruitment practices, which might have been thought to be an obvious point for policy intervention, appear to be unbiased. One hypothesis which might account for the source of the bias being in the private unionized sector is if the unions were male dominated and functioned so as to exclude women. However, a different research approach would be needed to investigate this further. Were the hypothesis substantiated, it would suggest that an appropriate government intervention would be to influence the gender bias in unions.

Taking recruitment practices as given, the other policy intervention which would increase the participation of women is the expansion of education. Gender differences in participation narrow as education is increased. This leads us to a discussion of policies aimed at increasing female access to education.

We have distinguished between three points of access to the educational system, the entry to primary schooling, the entry to public secondary schooling and, failing that, the entry to private secondary schooling. Of these, the gender bias in the first is the most severe. Girls are 91 percent more likely not to go to primary school than boys. Indeed, the predominant reason why girls are less likely to go to secondary school is that many fewer of them have been to primary school (or have dropped out before completion). Having completed primary schooling, they are 37 percent less likely than boys to proceed to secondary school. Only at the level of upper secondary schooling does the differential become reasonably narrow. Having completed lower secondary schooling girls are 14 percent more likely than boys not to proceed to upper secondary school. Our analysis of the determinants of primary school enrolment found gender to be the most significant of all the variables investigated. Nor does this bias against girls appear to be alleviated by household income. Increased household income increases the likelihood that children will be sent to school but it actually widens the gender gap. At no income level does the chance of a girl being sent to school even reach the chance of a boy from the poorest household. Against this, parental education and urbanization do narrow the absolute gender gap. However it appears that left to the *automatic* development processes, the diffusion of parental education, rising incomes, and urbanization, it will be a long time before the gender gap is eliminated. The analysis suggests that, left only to income growth, around a quarter of girls will receive no education even when all boys attend school. The policy implication of this is surely that it will be girls who benefit differentially from the compulsory introduction of universal primary education by 1987, and that without such government intervention a substantial group of the female population would have remained uneducated under even the most favorable development scenarios.

The problems of access to secondary education are quite different. We have established that, conditional upon completed primary education girls are less likely to attend secondary school because they perform markedly less well in the CEPE examination which rations entry. In our analysis of the determinants of exam performance gender was the most significant variable. Note that this is quite a surprising result since in many developed countries girls tend to do better than boys at this stage in the educational system. But now, quite unlike the gender gap in access to primary schooling, the gap in examination performance narrows dramatically as per capita consumption increases. Almost the entire problem of girls' under performance is accounted for by the very wide gap between girls and boys from the poorest households. It is tempting to interpret this as indicating that when the household is driven in extremis to make agonising priorities a more generalized preference is given to boys over girls. Note, however, that Deaton and Benjamin (1988) did not find any tendency on average for households to favor boys over girls. This problem is amenable to policy in three distinct ways. First, the problem is evidently closely bound up with that of poverty. All those policies which raise the living standards of the poorest households will tend to reduce the problem of girls' under performance in the CEPE Second, that under performance is occurring in an *educational production function* in which the government controls both the inputs of teaching resources and the measurement of performance It can therefore re-target teaching resources towards girls from poor households. and/or revise the examination in such a way as differentially to increase their pass rate Finally, girls' under performance in the CEPE matters in particular because it is used by the government as the rationing device for a scarce public resource, namely secondary school places. The government faces a problem only because its current rationing device differentially deselects girls from poor households The government, could, therefore, overcome the problem by supplementing the examination with other criteria so that the net effect was gender neutral.

Interventions in the labor market and in access to education can be expected to be mutually reinforcing Improving the chances of women gaining employment in the private, unionized sector will enhance the incentives for parents to invest in their daughters' education and provide girls with role models for motivation in study. Improving the access of girls to schooling narrows the gender differential in participation in wage employment.

Appendix

Table A.1 Parameters of Reduced Form Equation (Whole Sample)

	(1)	(2)
CONSTANT	-11.69120	-6.22790
	(-13.392)**	(-14.190)**
AGEY	0.42579	0.22480
	(9.231)**	(9.495)**
AGE2	-0.00572	-0.00304
	(-9.687)**	(-9.988)**
CEPE	0.64041	0.30752
	(2.218)**	(1.979)**
BEPC	0.08707	0.03259
	(0.237)	(0.191)
HIGHER	-0.48487	-0.24675
	(-0.830)	(-0.754)
FGOVT	0.44049	0.23356
	(1.851)*	(1.763)*
HEAD	1.70700	0.98793
	(8.419)**	(8.769)**
MARRIED	0.46622	0.23963
	(2.295)**	(2.158)**
MGOVT	0.39658	0.[illegible]3765
	(0.542)	(0.641)
CHILD	0.06834	0.02988
	(0.331)	(0.264)
ABIDJ	0.67901	0.36805
	(4.810)**	(4.795)**
SEX	0.97189	0.55040
	(5.353)**	(5.568)**
GRSP	0.12533	0.07782
	(2.796)**	(3.240)**
GRS	0.09144	0.05097
	(1.372)	(1.352)
GRS2	0.24211	0.14424
	(1.523)	(1.606)
YRU	0.24186	0.11552
	(1.571)	(1.504)
NAT	0.53316	0.22260
	(2.870)**	(2.239)**
Log-Likelihood	-729.92939	-733.06000

* = significant at 10 percent
** = significant at 5 percent

	Predicted			
	0	1	0	1
Actual	1,726	83	1,728	81
	214	301	213	302

Table A.2 Parameters of Reduced Form Equation (Male)

	(1)	(2)
CONSTANT	-10.15780 (-9.792)**	-5.67350 (-10.329)**
AGEY	0.42308 (7.532)**	0.23715 (7.870)**
AGE2	-0.00588 (-8.357)**	-0.00331 (-8.823)**
CEPE	0.57045 (1.628)	0.25674 (1.311)
BEPC	-0.28691 (-0.790)	-0.17332 (-0.826)
HIGHER	-0.58614 (-0.903)	-0.32356 (-0.866)
HEAD	1.62906 (5.876)**	0.94046 (5.811)**
MARRIED	0.18391 (0.583)	0.11951 (0.664)
FGOVT	0.15163 (0.490)	0.05684 (0.319)
MGOVT	-0.17228 (-0.171)	-0.00818 (-0.013)
CHILD	0.53041 (1.866)*	0.31498 (1.877)*
ABIDJ	0.71333 (4.199)**	0.38580 (4.008)**
GRSP	0.06638 (1.235)	0.05013 (1.666)*
GRS	-0.00096 (-0.012)	0.00048 (0.010)
GRS2	0.40200 (2.095)**	0.23450 (2.150)**
YRU	0.08683 (0.537)	0.05097 (0.595)
NAT	0.73122 (3.439)**	0.36196 (3.023)**
Log-Likelihood	-478.26001	-480.79999

* = significant at 10 percent
** = significant at 5 percent

	Predicted			
	0	1	0	1
Actual	600	79	600	79
	128	277	129	276

Table A.3 Parameters of Reduced Form Equation (Female)

	(1)	(2)
CONSTANT	-14.78420 (-8.143)**	-6.84330 (-8.296)**
AGEY	0.51968 (5.129)**	0.23577 (4.880)**
AGE2	-0.00637 (-4.594)**	-0.00294 (-4.428)**
CEPE	0.64107 (1.218)	0.35329 (1.330)
BEPC	0.59176 (1.709)*	0.48784 (1.609)
HIGHER	-0.22075 (-0.162)	-0.24365 (-0.312)
HEAD	1.41168 (3.134)**	0.75911 (3.064)**
MARRIED	0.48545 (1.455)	0.23353 (1.325)
FGOVT	0.69776 (1.905)*	0.47190 (2.458)**
MGOVT	0.28729 (0.272)	0.07158 (0.129)
CHILD	-0.44091 (-1.394)	-0.24986 (-1.533)
ABIDJ	0.54193 (1.987)**	0.37189 (2.699)**
GRSP	0.30573 (3.265)**	0.13605 (3.055)**
GRS	0.75394 (2.173)**	0.12112 (1.882)*
GRS2	-0.00885 (-0.031)	0.01932 (0.113)
YRU	0.72031 (1.359)	0.43882 (1.458)
NAT	0.69479 (1.445)	0.17813 (0.850)
Log-Likelihood	-220.00999	-224.02000

* = significant at 10 percent
** = significant at 5 percent

	Predicted			
	0	1	0	1
Actual	1,117	13	1,115	15
	62	48	65	45

Table A.4 One Equation Smith Model

	Coefficient	*t-Statistic*	*Mean*	*Standard*
CONSTANT	3.6881	17.956	—	—
AGEY	0.0301	5.993	34.76[illegible]	9.0099
BEPC	0.1226	1.138	0.2563	0.4370
HIGHER	0.3325	1.659	0.1223	0.3280
TECH	0.2563	2.682	0.3612	0.4808
YRP	0.0243	1.028	4.6738	2.4083
YRS	0.0993	3.372	2.1495	1.8835
YRS2	0.1283	2.445	0.6990	1.1953
YRU	0.1133	3.057	0.4757	1.4882
YRST	0.0564	2.000	1.1592	1.5209
ABIDJ	0.0225	0.337	0.615	0.4869
YEARS	0.0300	2.585	8.6890	8.0243
YEARS2	-0.0005	-1.416	139.7626	232.0397
PUB	0.1250	0.480	0.5107	0.5004
PUB*UNION	0.0118	0.114	0.3903	0.4883
PRIV*UNION	0.3607	3.933	0.2194	0.4143
PUB*SEX	0.1201	1.187	0.3981	0.4900
PRIV*SEX	-0.0586	-0.540	0.3883	0.4878
PUB*NAT	-0.0344	-0.170	0.4874	0.5003
PRIV*NAT	0.1153	1.142	0.3515	0.4779
PUB*CEPE	0.3571	2.001	0.4291	0.4954
PRIV*CEPE	0.2884	1.896	0.2757	0.4473
R-Squared	0.6292	—	—	—

— = not applicable

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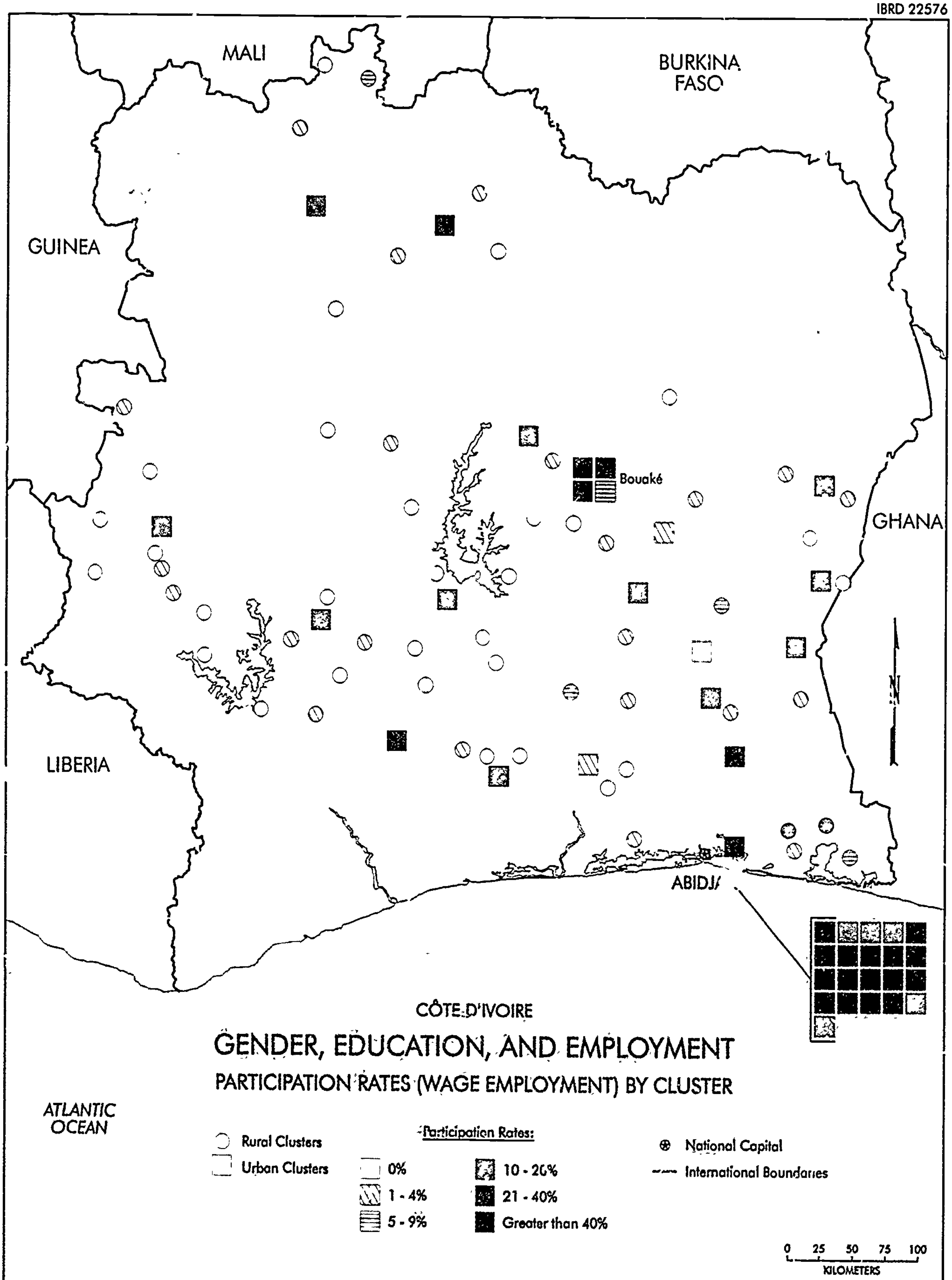

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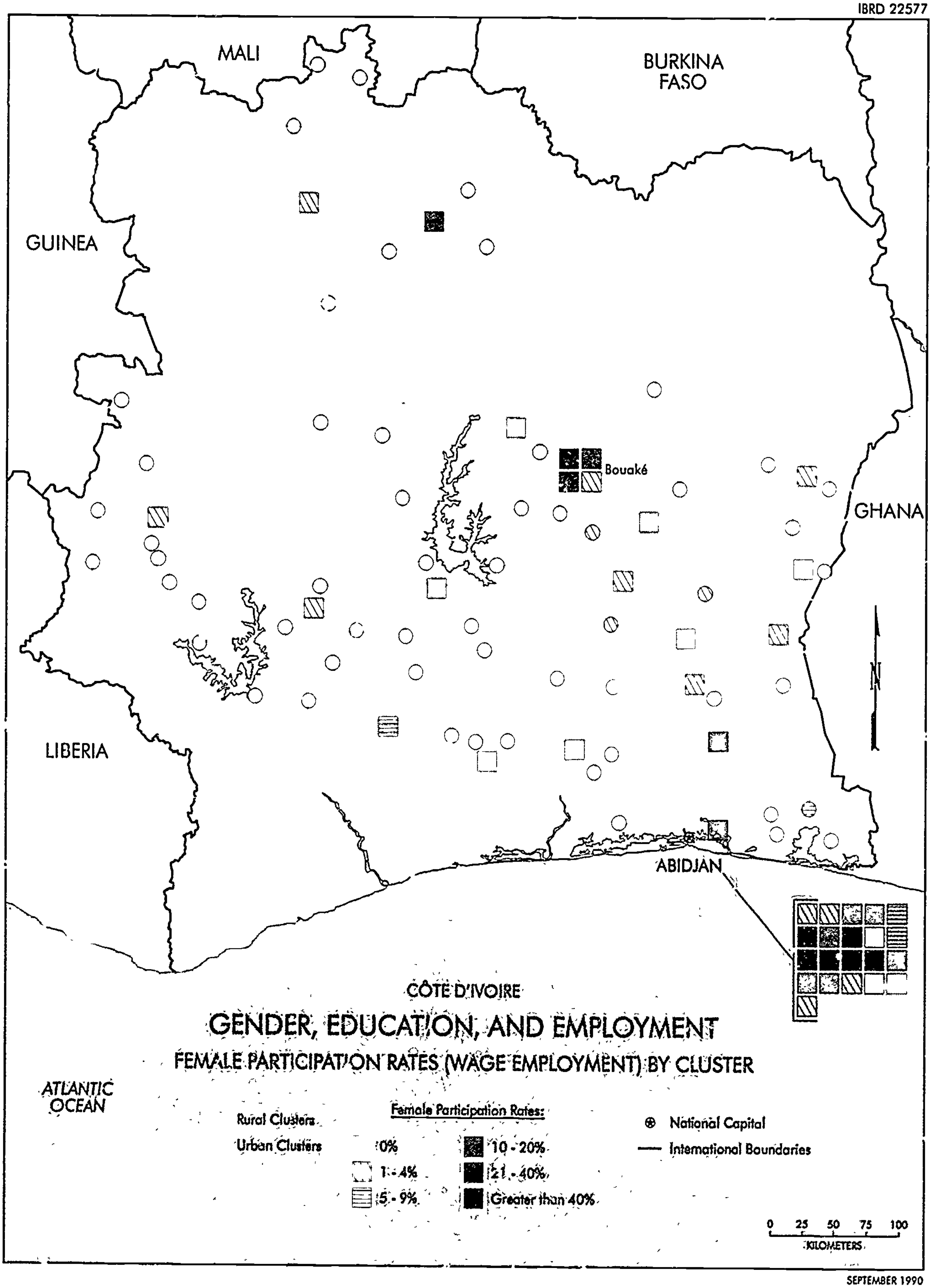
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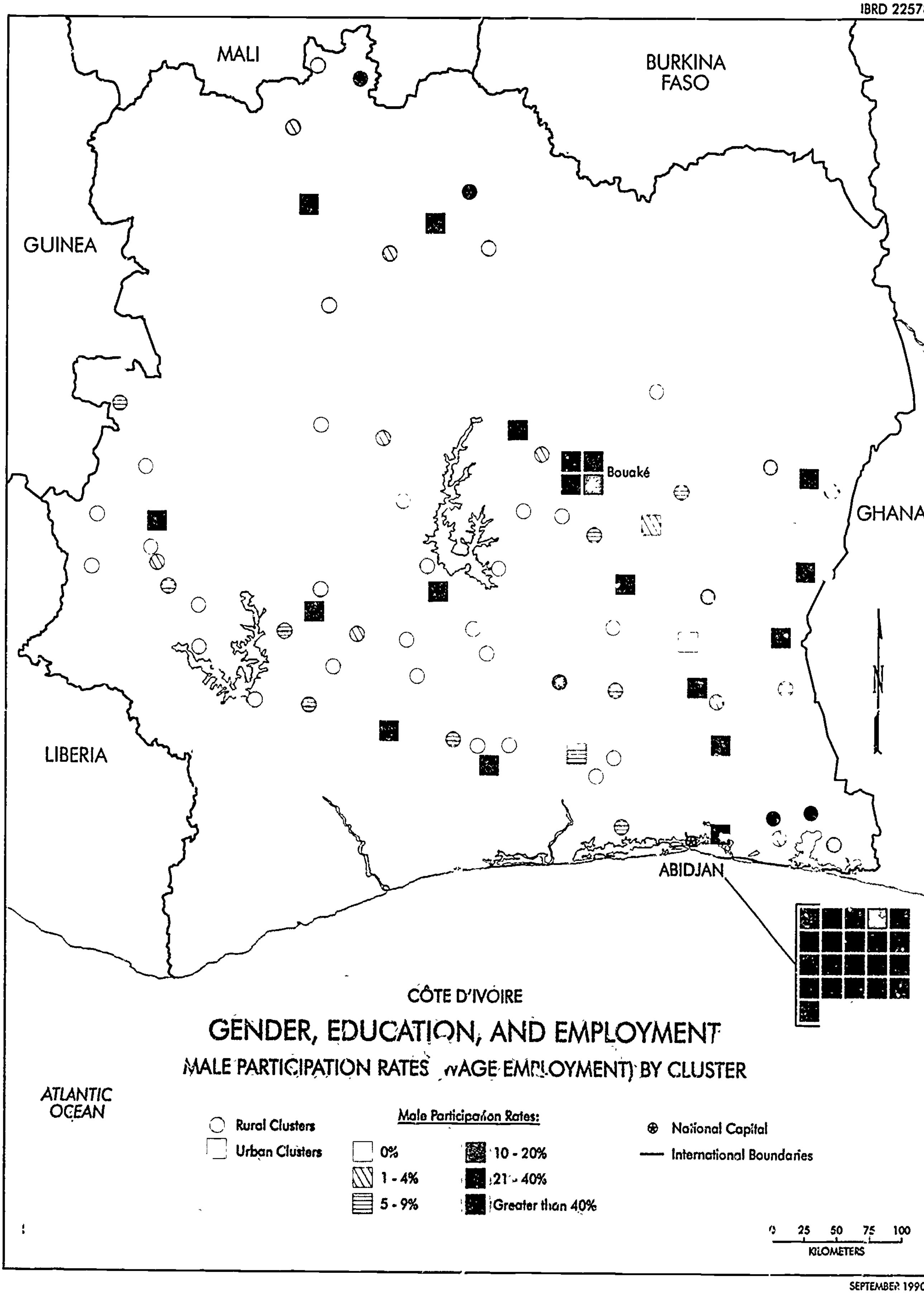
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