SAMFUNNSØKONOMISKE STUDIER

58

INDIVIDUAL LABOUR SUPPLY IN NORWAY

INDIVIDENES TILBUD AV ARBEIDSKRAFT

BY/AV
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STATISTISK SENTRALBYRÅ
CENTRAL BUREAU OF STATISTICS OF NORWAY

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FORORD

Denne publikasjonen inneholder en analyse av sammenhengen mellom individenes atferd på arbeidsmarkedet og deres respektive personlige kjennetegn, så som kjønn, alder, utdanning, ekteskapelig status og barnetall. Arbeidsledighet og daghjemsdekning behandles også. En forsøker å tallfeste hvordan disse faktorene påvirker tilbudet av arbeidskraft. Datamaterialet er i all hovedsak hentet fra arbeidskraftundersøkelsene i 1977, og analysen gjøres ved hjelp av logitmodeller og linær regresjon.

Måling av arbeidskrafttilbudet er imidlertid forbundet med en del prinsipielle og praktiske problemer. Disse problemene blir inngående drøftet i publikasjonen. Den økonometriske modell som anvendes forsøker å ta hensyn til noen av måleproblemene.

Statistisk Sentralbyrå, Oslo, 4. oktober 1984

Arne Øien

PREFACE

This study is an analysis of the relationships between individual persons' labour market behaviour and their respective socio-demographic characteristics, such as sex, age, education, marital status, and parenthood. An attempt is made to estimate the effects of these variables on men's and women's supply of labour. The approach is to apply the logit and linear statistical models to a disaggregate data set prepared from the Norwegian Labour Force Sample Surveys of 1977.

The supply of labour is, however, a concept open to numerous problems of measurement and interpretation. An extensive discussion of these problems is offered, and an econometric model designed to partially circumvent these difficulties is put forward.

Central Bureau of Statistics, Oslo, 4 October 1984

Arne Øien

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The data set used in this study was prepared by Jan Lundsbakken. Jørgen Ouren was responsible for transferring the data to the TROLL computer system, without which this analysis would hardly have been possible. Liv Hansen drew all the diagrams, performed a major part of the manual calculations, and helped prepare the manuscript in innumerable ways. Solveig Wiig did a wonderful typing job. Last, but not least, thanks are due to the author's colleagues in the Socio-demographic Research Unit, in particular to Olaf Foss, John Dagsvik, Olav Ljones, and Per Sevaldson, whose thoughts and ideas pop up at irregular intervals throughout this book. Needless to say, all errors are my own.

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1. INTRODUCTION

Human labour is probably the most important of all factors of production. How much of it is there? What does it depend on? If certain conditions are altered, how much will it change?

With a view to answering the first of these questions - how large is the supply of labour - certain standards have been worked out for the compilation of relevant statistical data. Recommendations have been laid down by the International Labour Office (ILO). In tune with these recommendations the Central Bureau of Statistics of Norway has been conducting Labour Force Sample Surveys (LFSS) every quarter since 1972. Results are published in the Weekly Bulletin of Statistics and also in the yearly publication NOS Labour Market Statistics.

The statistics provide estimates of the number of persons employed and unemployed and the amount of hours worked by those having a job. Only work for pay or profit is counted as employment. Housework and other services that are not being paid for fall outside the scope of the labour market statistics and of the analysis presented in this publication.

However, even if we limit our attention to market work only, it is not immediately clear how one should go about to measure the amount of labour available. The main reason for this is that the assessment of unemployment - i.e. of how many persons are willing to work but unable to find a job - is a rather intricate matter.

Therefore, in chapter 2 of this book, we address the question of how to define and measure the supply of labour. The principles and definitions underlying current labour market statistics are examined, and certain weaknesses inherent in this nomenclature are pointed out. Two essentially different methodologies designed to remedy these weaknesses are discussed.

Having established certain principles as to how the supply of labour may be measured, we ask, in chapter 3: On what factors does it depend? This is, of course, not a question that can be answered by data alone. A theory is required. As our point of departure, we choose the well-known neoclassical theory of utility maximization. An econometric model of individual labour supply is then developed, and a method to estimate the reduced form parameters of this model is proposed.

The strength of such a multivariate model is that its parameters are interpretable as partial effects. That is, the model enables us to separate out the impact of each independent variable. Such information is essential whenever one wants to predict how the supply of labour will respond to a given change in a particular explanatory variable.

The present labour supply study is, of course, not the first one to apply multivariate statistical techniques to a disaggregate data set. Others have paved the ground. Suffice it to mention the seminal work by Bowen and Finegan (1969), who studied individual labour force participation by means of linear regression analysis, and that of Ljones (1979), who was the first to apply such a method to Norwegian data.

The novelty of the present analysis can be summarized in four points. (i) Its scope is virtually all persons[1] between 16 and 74 years of age resident in Norway, irrespective of sex or marital status. (ii) Labour force participation is analyzed using a binary logit model. Thus, we avoid the logical and practical difficulties inherent in the application of linear regression models to binary response variables. (iii) Separate probability models are estimated for labour force participation and hours worked. The overall effect on labour supply (measured in hours) of a given independent variable is decomposed into one part originating from a change in participation rates and another part being due to changes in the average length of the working week. (iv) The split between full-time, part-time, and no working is studied by means of a trinomial logit model.

Perhaps the most important weakness of the present analysis is that wage data have not been available for the respondents in question. The estimation of structural relationships is therefore ruled out. Using, however, a supplemantary data set taken from the Norwegian Fertility Survey of 1977, we are still able to shed a small bit of light upon the role played by the wage variable.

Chapters 4 and 5 of this book are devoted to the empirical study of women's and men's labour supply, respectively. The relationship between labour force participation and unemployment is described. The partial effects on labour supply of age, marital status, level of education, motherhood, and the capacity of day-care centres are all examined. In

chapter 6 the main results are summarized and the respective patterns of labour market behaviour are compared between men and women.

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[1] Two categories are, however, left out: Persons stating "school or university attendance" or "compulsory military or civil service" as their principal activity during the survey week. The labour supply generated by these persons is small.

For the sake of completeness, certain supplementary results etc. have been gathered in the appendices. Appendix 1 contains detailed variable definitions and summary sample statistics. In appendix 2 we report the results of certain tests for interaction between some of the independent variables as they influence labour force participation. Appendix 3 is a collection of supplementary tables providing, i.a., estimated contrasts between any pair of possible values assumed by a given qualitative independent variable. Contrasts are given for the effect on labour force participation as well as for the effect on full-time as against part-time work.

Readers only interested in the main results and conclusions may limit their reading to sections 4.1, 5.1, and 6. Theory and method are explained briefly in section 3.1.

2. MEASURING THE SUPPLY OF LABOUR

2.1. Motivation: what is the problem?

Measuring the supply of labour, individually or on an aggregate basis, is not as straightforward as it may seem. However, the problem is usually swept under the rug. Let us spell it out, briefly.

The great bulk of the applied economic literature treats aggregate labour supply as a concept synonymous to the labour force. The latter concept is (roughly speaking) defined as the sum of those having a job (the employed) and those actively seeking one (the unemployed).

This way of measuring labour supply leaves out the so-called hidden unemployed. Thus, in all probability the overall level of labour supply is consistently underestimated. More serious, however, is the fact that even variations in labour supply (temporal or cross-sectional) tend to be underrated. This is so because the labour force has been shown to vary systematically with the tightness of the labour market: the higher the jobless rate, the smaller the labour force. One explanation for this is the so-called discouraged worker hypothesis, saying that during periods (or in areas) of slack labour market, many people refrain from job-seeking because they believe their search will be in vain. Now, the fact that these people withdraw from or fail to enter the labour force can hardly be interpreted as a shift in labour supply. Rather it must be seen as a reflection of insufficient demand. If, nevertheless, labour supply is defined synonymously with the labour force, we seem to be faced with a supply function that shifts (automatically) whenever demand does. It goes without saying that the identification of supply and demand within such a model is difficult, if not altogether nonsensical. At the very least, the adoption of such a conceptual framework has an important bearing on the political and economic interpretation of whatever results are derived from aneconometric model.

The difficulties involved in constructing a meaningful labour supply measure are enhanced by the fact that labour is no ordinary commodity.

Labour is a service, and as such invisible unless traded. Often the excess supply of labour (if any) fails to manifest itself physically, as would be the case of goods piled up in an inventory.

The labour market is subject to strong institutional constraints. Wages and working hours are often fixed by convention, law, or collective bargaining agreements.

Labour is a source of income. Thus, in the theory of consumer behaviour, the wage rate has the double role of determining the budget constraint as well as the "price of leisure". This means that the sign of the wage effect on labour supply remains a priori uncertain. Some individuals may have a so-called backward-bending labour supply curve.

Labour is heterogeneous. Put simply, it matters to the employer whom (s)he buys labour from. It also matters to the employee whom (s)he is working for. Jobs are different in more than one respect. Labour may be short in some market segments while in excess supply elsewhere. When choosing among different jobs, the (prospective) worker takes into consideration a large number of different job characteristics other than wage. The wage rate may not even be the most important.

In other words, labour supply is a <u>function</u>, and a fairly complex one. For each and every <u>individual</u> the number of potential labour supply determinants is quite large. When we say "the supply of labour", what do we really mean: supply at what wage, under what conditions, in which segment? The number of possible arguments in the <u>aggregate</u> labour supply function is almost endless. Including all of them is impossible. We must choose-but how?

These are, in short, the intriguing questions to which this chapter is addressed. While not aiming to solve once and for all the problems of measuring labour supply, we believe that a clarification of the conceptual framework in this area is desirable, and that such a clarification is essential to the understanding and interpretation of the empirical results to follow.

Thus, in sections 2.3 and 2.4 we discuss two essentially different approaches to the construction of a reasonably satisfactory labour supply concept. Somewhat misleadingly, we have termed them the <u>survey</u> approach and the <u>econometric</u> approach.

The former method emphasizes the use of a diversified conceptual framework so as to adequately describe the labour market barriers and opportunities facing individuals in widely different life situations. This method presupposes data of a highly disaggregate and detailed nature, obtainable only through personal interviews (whence the term "survey approach"). To a large extent the data used are based on <u>stated preferences</u>. The interpretation of such data, in particular their use for purposes of prediction, raises important and interesting methodological questions. Drawing heavily on the works by Farm (1977) and Foss (1980), we shall see that for our purpose, the definitions recommended by the ILO, or those adopted in the Norwegian Labour Force Sample Surveys (LFSS), are not really precise enough.

The "econometric approach" on the other hand, is characterized by the use of aggregate measures of labour market tightness as indicators of hidden unemployment or labour underutilization. These data are typically of a revealed preference nature; they pertain, in principle, to factual events and behaviour. Even the econometric approach may, however, very well be based on survey data (as in our case). Moreover, the interpretional problems of the two respective methods are not unrelated. The conceptual difficulties encountered in the survey approach have an important bearing even on econometric analyses.

Before we turn to discussing the theoretical questions involved, we shall, as a service to the reader, have a look at the principles and definitions underlying current labour market statistics. Readers familiar these matters may want to skip section 2.2.

2.2 The nomenclature of current labour market statistics

The ILO recommendations

International recommendations for the compilation of labour market statistics have been laid down by the ILO (International Labour Office 1976).

According to the ILO recommendations, employment is defined as follows:

> "Definition of employment
>
> 6. (1) Persons in employment consist of all persons above a specified age in the following categories:
>
> (a) at work; persons who performed some work for pay or profit during a specified brief period, either one week or one day;
>
> (b) with a job but not at work; persons who, having already worked in their present job, were temporarily absent during the specified period because of illness or injury, industrial dispute, vacation or other leave of absence, absence without leave, or temporary disorganisation of work due to such reasons as bad weather or mechanical breakdown.
>
> (2) Employers and workers on own account should be included among the employed and may be classified as "at work" or "not at work" on the same basis as other employed persons.

(3) Unpaid family workers currently assisting in the operation of a business or farm are considered as employed if they worked for at least one-third of the normal working time during the specified period.

(4) The following categories of persons are not considered as employed:

(a) workers who during the specified period were on temporary or indefinite lay-off without pay;

(b) persons without jobs or businesses or farms who had arranged to start a new job or business or farm at a date subsequent to the period of reference;

(c) unpaid members of the family who worked for less than one-third of the normal working time during the specified period in a family business or farm." (ILO 1976: 28 f.)

As for unemployment, the definition goes:

"Definition of unemployment

7. (1) Persons in unemployment consist of all persons above a specified age who, on the specified day or for a specified week, were in the following categories:

(a) workers available for employment whose contract of employment had been terminated or temporarily suspended and who were without a job and seeking work for pay or profit;

(b) persons who were available for work (except for minor illness) during the specified period and were seeking work for pay or profit, who were never previously employed or whose most recent status was other than that of employee (i.e. former employers, etc.), or who had been in retirement;

(c) persons without a job and currently available for work who had made arrangements to start a new job at a date subsequent to the specified period;

(d) persons on temporary or indefinite lay-off without pay.

(2) The following categories of persons are not considered to be unemployed:

(a) persons intending to establish their own business or farm, but who had not yet arranged to do so, who were not seeking work for pay or profit;

(b) former unpaid family workers not at work and not seeking work for pay or profit." (ibid.)

Finally, we have:

"Definition of labour force

4. The civilian labour force consists of all civilians who fulfil the requirements for inclusion among the employed or the unemployed, as defined in paragraphs 6 and 7 (....).

5. The total labour force is the sum of the civilian labour force and the armed forces." (ibid.)

Clearly, the ILO recommendations leave considerable room for interpretation and judgement. In some cases, it is possible to think of verbal reformulations that would immediately sharpen the definitions adopted. Suffice it to mention such terms as "a specified age", "available for work", "some work", or "seeking work". Moreover, even if these concepts were made more precise, there would be several ways to implement them, or to make them operational. Interviewing a random sample of the population would be one such method. Even so, it has been demonstrated that the outcome of sample surveys can be quite sensitive to even minor changes in the phrasing and order of the interview questions, or to other characteristic aspects of the interview situation. This is particularly so if the questions asked are of a subjective nature, i.e. if the respondent is asked to pass a judgement rather than describe a factual event (Foss 1980: 102f. and 207f.; Hoem and Ljones 1974: 43f.).

The Norwegian Labour Force Sample Surveys (LFSS).

The conceptual framework of the Norwegian Labour Force Sample Surveys is based on the ILO recommendations.

The reference period is *one week*. Four surveys are taken each year in February, May, August, and November. The age limits are *16 to 74 years* inclusive, as measured on December 31 of the survey year.

The ILO concept "some work" has been specified as "at least one hour of work last week". This is the general cut-off point for being classified as an "employed person at work". "One-third of the normal working time" has been set to 10 hours per week. Below this limit unpaid family workers are not classified as employed. The ILO formulation "available for work, ...without a job, and seeking work for pay or profit" has been translated into "looking for work by registration at the Employment and Seamen's Office, advertising, responding to advertisement, contacting employers, etc.". Thus, the rather subjective terms "seeking work" and "available for work" have been made operational by specifying some objective *activity* criterion that the respondents will have to fulfil in order to be classified as unemployed. It is not sufficient to "want" a job, the respondent is required to have taken certain *actions directed at the job market.*

When must this action have been taken? According to the Norwegian LFSS, the job-search must have taken place some time during the 2-month period prior to the interview, and be directed at such jobs as the respondent "would have been able to take" during the reference week.

As of 1977, however, the 2-month limit was not explicitly spelt out in the questionnaire, but only in the interviewer's guide. This probably has the effect that certain respondents believe the search must also have occurred during the reference week, so that the number of unemployed is somewhat underestimated.

Moreover, as of 1977 certain (non-employed) categories were not asked about their job-seeking activity; these include persons who state "ill", "disabled", or "military conscript" as their principal activity during the reference week. Hence these persons could not possibly have been classified as "unemployed".

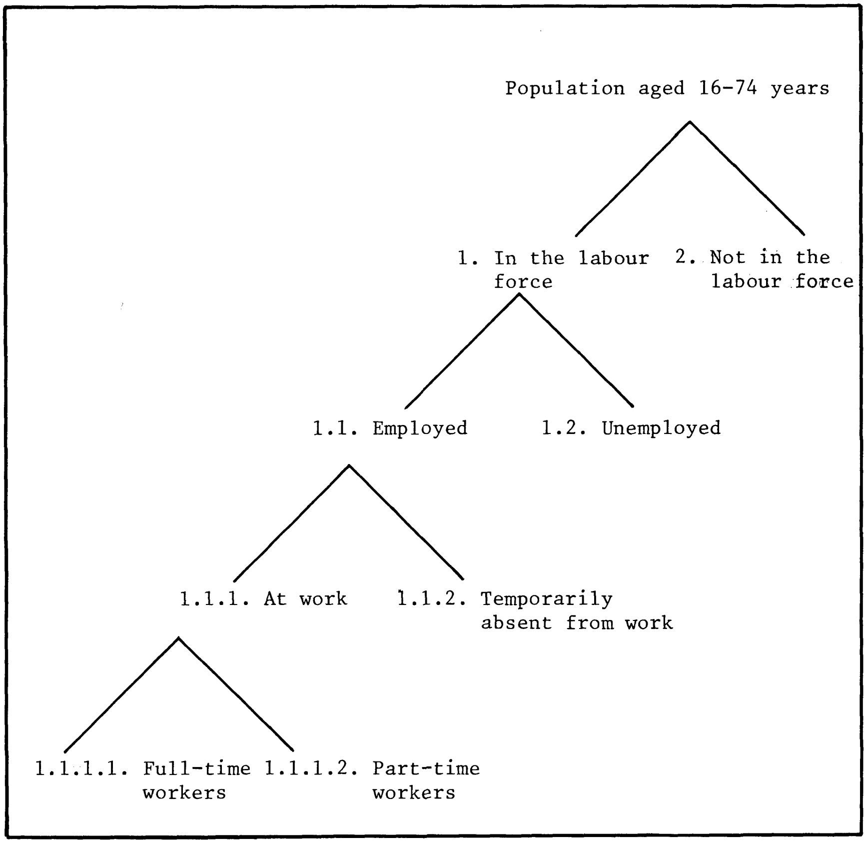
The definition of the total Norwegian labour force is slightly different from the ILO recommendation, in that conscripts are not included. So are, however, officers and other salaried military personnel. Thus, the Norwegian labour force as defined by the LFSS does not exactly correspond to the ILO-defined "civilian labour force", nor to the ILO-defined "total labour force". It is something in between.

The structure of the LFSS conceptual framework is summarized in figure 2.1. The population aged 16-74 is divided into two main groups: those in the labour force (1) and those not in the labour force (2). The labour force consists of employed (1.1) and unemployed (1.2) persons. Unemployed persons are those who did not work as much as one hour for pay or profit during the reference week, nor were temporarily absent from such a job, but who had taken certain actions to obtain a job for the specified week, who had been laid off without pay, or who were waiting to report to a new job at a date later than the reference week. Employed persons comprise persons at work (1.1.1) and persons temporarily absent from work (1.1.2). Persons at work are those who performed at least one hour's work for pay or profit during the reference week. The temporarily absent include those who during the entire survey week were absent from their job, owing to such reasons as illness, labour conflict, vacation, leave, absenteeism, bad weather, or mechanical failure. Note that the following categories are not counted as temporarily absent: (i) persons absent during part of the survey week (at work), and (ii) persons laid off without pay (unemployed).

Finally, in this study employed persons at work are subdivided into two groups: full-time workers and part-time workers. The cut-off point has been set to 30 hours of work during the reference week.

The advantages and disadvantages of the LFSS conceptual framework are discussed further in section 2.3.

Figure 2.1. Structure of the LFSS conceptual framework



Registered (ESO) unemployment

The most commonly used measure of unemployment in Norway is not the LFSS figure. Statistics are also gathered by the Employment and Seamen's Offices (ESO). These statistics are compiled on the last working day of every month and cover persons who have recently registered (or renewed their registration) as unemployed at the local ESO. Counted as unemployed are persons who are "without any job for pay or profit", but who "seek work for pay or profit" and are "available for" and "able to take" the kind of work sought. Students in search of a temporary holiday job are, however, not counted as unemployed by the ESO (Foss 1980: 54f.)

With a little lenience, even this measure can be seen as an adaptation of the ILO recommendation, provided we take "seeking work" to mean "seeking work through registration at the ESO". Other job-search strategies are not counted. Hence, this way of defining unemployment is narrower than the LFSS concept. Traditionally, the latter statistic has exceeded the ESO figure by some 50 to 100 per cent (Foss 1980: 58).

The incentives for registration at the ESO are two-fold: (i) to obtain a job, and/or (ii) to become eligible for unemployment benefits through the National Insurance Scheme. This scheme

> "..... provides benefits in the form of a daily cash allowance during unemployment, grants to cover costs of moving to and establishing a home and a new place of work, and retraining grants. Further, the scheme affords assistance towards obtaining a livelihood, grants in connection with vocational rehabilitation and wage subsidies for unemployed persons who are given municipal work." (National Insurance Institution 1978: 20f.)

As a general rule,

> "The daily allowance is Nkr 15 per day plus either 1/1000 of earned income in the last completed calendar year, or 1/1000 of the average earned income for the last three completed calendar years, if this amounts to more" (ibid.)

Registration at the ESO is a necessary condition for obtaining the daily cash allowance. It is, however, not a sufficient one:

> "To become eligible for a daily cash allowance the insured person must have been in public or private employment in Norway or on board a Norwegian ship in foreign trade, and in such employment he must have had an earned income amounting to not less than 75 per cent of the basic amount, i.e. Nkr 8 850 as from 1 January 1976 and Nkr 9 075 as from 1 May 1976, either in the completed calendar year or on an average in the three last completed calendar years. (.....)
> To be eligible for the daily cash allowance, the unemployed person must be able to work and willing to accept any work which the employment office finds suitable for him. He must also keep in regular touch with the employment office as long as he is receiving the daily allowance." (ibid.)

Also, to receive unemployment benefits the persons must be unemployed "through no fault of their own". Persons without a previous income or who quit their last job voluntarily receive no cach allowance. For persons in this category who also do not expect the ESO to be able to find them a job, the economic incentive to register at the ESO is practically nil. This is probably why the propensity to register is particularly low among housewives and youths.

As compared with the LFSS figure, the statistics on registered unemployment have one important advantage: they are void of sampling error and available at the municipal level. This is precisely why they are of interest to the analysis presented in this publication.

2.3. The survey approach

If we want to find out about people's supply of labour, it may not be a bad idea to ask them.

We then face the problem of exactly how to phrase the question(s). The questionnaire adopted by the Norwegian LFSS in 1977 may be viewed as one practical solution to this problem.

Based on the information collected by means of this questionnaire, the respondents may be classified as either "employed", "unemployed", or "not in the labour force".

The criteria for classifying a person as employed are relatively objective and straightforward. If, however, we accept the idea that the total supply of labour includes not only those presently having a job, but also a number of individuals who for various reasons do not have one (the "unemployed"), we must decide how the latter group is to be delimited.

As demonstrated, i.a., by Foss (1980), the unemployment concept derivable from the LFSS questionnaire relies on some rather arbitrary methodological choices. Many of these choices can be seen as relating to the definition of labour market availability, cf. the ILO recommendation. What does it mean to be "available for employment"? Another set of conventions relates to the delimitation of relevant job-seeking activities. What does it mean to "seek work"?

These questions are discussed at some length by Farm (1977:55ff.). He observes that availability for work is not a property of the individual per se. Rather, it expresses a relationship between the individual and the existing set of jobs. A given individual may be available for certain jobs, while unavailable for others, due to such factors as physical disabilities, family constraints, inadequate skills, or the attitudes of employers. This means that availability depends on the composition of labour demand, notably on the content and requirements of each individual job, as defined by the employer. Availability can be altered either by training the individuals, by changing their environment, or by modifying the jobs. To define certain individuals as "unavailable for work", and thus as "not in the labour force", amounts to tacitly assuming that the job content cannot be changed. The mismatch between job requirements and individual skills is concealed, and hence also the amount of labour underutilization.

Simular difficulties apply to the job-seeking criterion embedded in the ILO definition. This criterion, too, depends on labour demand characteristics, as illustrated by the fact that the rate of unemployment observed within a given community is sometimes seen to rise as a result of announced, job-generating business expansions. It seems that this phenomenon can only be explained by reference to a latent supply of labour, which, however, remains invisible as long as there are no vacant jobs for which to apply.

The following extended definition of unemployment has been put forward by Farm (1977:59, our translation):

Definition 2.1: An individual is said to be unemployed on a certain occasion if (s)he (i) is without a job, (ii) is capable of performing a job, and (iii) wants a job.

The criteria (ii) and (iii) may in turn, be defined as below (adapted from Farm (1977)).

Definition 2.2: An individual is said to be capable of performing a job if there exists at least one job on the actual labour market which (s)he would be able to perform, in the sense that (s)he fulfils the physical and mental requirements set by the job, including those requirements which pertain to previous education and training.

Definition 2.3: An individual is said to want a job if there exists at least one job on the actual labour market which the individual would accept to hold, given the wage, time schedule and all other conditions specified in the job contract.

Note that definitions 2.2 and 2.3 refer to the existence of any job, vacant or not, within the "actual labour market", which should be thought of as extending throughout the entire nation.

With these qualifications, definition 2.1 is seen to constitute a relatively wide definition of unemployment. In fact, it is too wide, in the sense that no link is established between the individual's capabilities on the one hand and his/her wishes on the other hand. Persons with totally unrealistic job expectations risk being classified as unemployed although unable to perform any of those jobs which they want, and unwilling to do any of those jobs for which they qualify.

To exclude such persons form the unemployment concept, definition 2.1 may be modified as follows (Farm 1977:64):

<u>Definition 2.4</u>: An individual is said to be unemployed on a certain occasion if (i) (s)he is without a job, and (ii) there exists at least one job on the actual labour market which (s)he (a) is capable of performing <u>and</u> (b) would accept to hold, in the senses specified in definitions 2.2 and 2.3, respectively.

In other words, to be counted as unemployed, the individual must be willing to accept some job for which (s)he would also qualify.

This convention still leaves us with one more loose end: when does a person "qualify" for a given job? In principle, one might consider the possibility of establishing a set of objective personal attributes ($\underset{\sim}{x}_i$, say) measurable for every individual i, and a criterion function $f_j(\underset{\sim}{x})$ for every job j, such that individual i is said to qualify for job j if and only if $f_j(\underset{\sim}{x}_i)$ exceeds some prespecified value c.

A conceptually much simpler approach, however, would be to sharpen definition 2.2 by adding

<u>Definition 2.5:</u> An individual is said to fulfil the physical and mental requirements set by a given job if, being the only applicant for the job, (s)he would have been hired by the employer.

Taken together, definitions 2.2, 2.3, 2.4 and 2.5 form a fairly wide and yet precise definition of unemployment. When this unemployment measure is added to the employment figure, a labour supply measure is obtained that should be no more than marginally affected by changes in the firms' demand for labour.

There is, of course, a catch: This new labour supply measure is not operational. To quantify unemployment according to the above definition, every member of the population (or at least every LFSS respondent), would have to answer two hypothetical questions, viz. (i) "is there a job that you would be able to get if there were no other applicants?" and (ii) "if yes, would you take it?"

A large number of respondents would not possess the information necessary to answer such questions. And out of those who did give an answer, an unknowable fraction would be wrong.

To construct operational measures of labour supply, the hypothetical and/or unanswerable questions implicit in definitions 2.2 through 2.5 must somehow be translated into more practicable ones.

Thus, persons classified as unemployed according to the Norwegian LFSS are those who were without a job, but who were on temporary lay-off or who had been "looking for work by registration at the ESO, advertising, responding to advertisement, contacting employers etc." It may be argued that the job-seeking activity shown by these persons is proof of their willingness and ability to work. In other words, all active job-seekers may be assumed to fulfil the conditions specified in definition 2.4.

The converse is, however, not true. A number of individuals who want a job and who are also capable of working do not seek work, for the simple reason that they believe their search would be in vain. These persons have been termed "discouraged workers".

Since 1967 the US. Current Population Survey (CPS) has included questions purporting to shed light on the size and characteristics of this group[1]. For all persons not in the labour force the question

(1) "Does the respondent want a regular job now, either full-time or part-time?"

is asked. If the answer is "yes" or "maybe", a second question is asked:

(2) "What are the reasons the respondent is not looking for work?"

Those persons who are reported as not looking for work for one or more of the following five reasons are classified as discouraged workers:

1. Believes no work available in line of work or area.
2. Couldn't find any work.
3. Lacks necessary schooling, training, skills or experience.
4. Employers think [the person is] too young or too old.
5. Other personal handicap in finding [a] job.

Reasons 1 and 2 may be termed "job-market reasons", relating to the subject's perception of the local job market, while reasons 3, 4, and 5 may be referred to as "personal reasons", since they express the subject's view of his or her own personal problems in finding work. Again, however, it must be observed that the split between these two sets of reasons is arbitrary in the sense that it probably depends on the size and composition of labour demand. Whether or not such handicaps cited

[1] The below discussion is based on Finegan (1978).

as reasons 3, 4, or 5 would constitute an impediment to employment depends on the hiring standards of employers.

In 1977, the number of discouraged workers thus classified in the United States amounted to 1.010 million, as compared to an unemployment toll of 6.855 millions and a total civilian labour force of 97.401 millions (Finegan 1978:9).

The total number of persons not in the labour force who were reported to "want a regular job now" was, however, no less than 5.942 millions. The reasons given for not seeking work were, other than "discouragement", school attendance (1.534 million), ill health or disability (0.753 million), home responsibilities (1.253 million), and other (1.392 million).

Thus, the number of persons "wanting a job" was almost as large as the number of unemployed. However, only one sixth of these were officially classified as "discouraged".

Even the Norwegian LFSS contain a few questions that throw light on this issue, although the information gathered is less detailed than in the American CPS. In the LFSS, most[1] persons not in the labour force are asked (our translation):

(3) "Would you say that you have a need for or would wish to have a job?"

Those who answer "yes", or "yes, under certain circumstances" are termed "potential job-seekers". Respondents falling in this category are asked the following question:

(4) "Was the principal reason for your not seeking work that no suitable jobs were available? By suitable jobs we are referring to such aspects as type of work, working hours, and travel time."

Those who answer "yes" to this question constitute the closest Norwegian analogue to what is termed "discouraged workers" in the US Current Population Survey.

The estimated numbers of "potential job-seekers" and "discouraged workers" in Norway 1977 are shown in table 2.1. Persons stating "suitable work not available" as their principal reason for not being in the labour force are more than twice as numerous as those classified as unemployed. If these persons were to be included in the labour force, the latter would increase by 3.5 per cent. In the US, the number of "discouraged workers" corresponds, in comparison, to only one sixth of the unemployed and only 1 per cent of the labour force.

[1] In 1977, the question was considered inappropriate for conscripts, pensioners and sick or disabled persons.

Table 2.1. Persons 16-74 years, by sex and labour market status. LFSS estimates, 1977. Thousands

	Total	In the labour force			Not in the labour force		
		Total	Employed	Unemployed	Total	Potential job-seekers: Total	Potential job-seekers: Suitable jobs not available
Both sexes ...	2 844	1 851	1 824	27	993	146	64
Males	1 421	1 119	1 108	11	302	24	8
Females	1 423	732	716	16	691	122	56

Sources: NOS A 958 Labour Market Statistics 1977 and Fridstrøm (1981:tables 6.10 and 6.11).

Although the concepts and interwiev techniques used are not strictly comparable between the two countries, it is tempting to hypothesize that the incidence of "discouragement" is higher in Norway than in the US, and that this hypothesis may partly explain the large difference in unemployment rates. Such a hypothesis is by no means implausible in view of the unequal demographic, geographic, social, and economic conditions prevalent in the two countries. Norway is a sparsely populated country, in which relatively large parts of the population are confined to small, regional communities with little mobility between them but with strong internal social networks. In such areas most inhabitants have nearly full information about current job openings and in particular about the fact that there are none, in which case they refrain from job-seeking.

Perhaps the most striking feature of table 2.1 is the apparent difference between male and female labour market behaviour. That women have a lower rate of labour force participation than men is well known. Less well known, however, is the fact, apparent from table 2.1, that 7 out of 8 persons who refrain from job-seeking for lack of suitable job openings, are women. Among persons "wishing to have a job" (i.e. the "potential job-seekers"), 4 out of 5 are women. Male "discouraged workers" are fewer than the "unemployed", while among females the ratio of "discouraged" to "unemployed" is 3.5 to 1.

About 42 per cent of all persons stating "suitable jobs not available" cite "part-time work" as their first requirement for taking a job (Foss 1980:171). The great majority of these persons are women. Thus, the scarcity of part-time jobs may account for part of the difference between male and female "discouragement".

Whether or not to include "discouraged workers" in the unemployment count has been an issue of much debate in the United States. The main argument in favour of such a reclassification is that the "discouraged workers" form part of the nation's potential labour supply. They want to work, are available for work, and some of them would be working or actively seeking work if jobs were more plentiful. The number of "discouraged workers" has been shown to vary systematically over the business cycle, in the same way as the rate of unemployment (Finegan 1978). Hence "discouraged workers" are part of the underutilization of labour not reflected in the unemployment statistics.

Yet such a reclassification was never carried into effect. The arguments against it are several (see Finegan 1978:38-45). Perhaps the most compelling one is that the proposed reclassification would introduce a more subjective element into the unemployment concept and hence also into the labour force statistic.

For all their shortcomings, current unemployment statistics in Norway and in the US do have the advantage of being based on relatively objective criteria. That is, the criteria for being classified as unemployed depend not on the respondent's judgement, but (at least in principle) on his or her actual behaviour[1].

Suppose it were decided to count as unemployed those persons who "have a need for or would wish to have a job", in short, those stating that they "want a job". While a certain fraction of this group would undoubtedly belong to the labour force under certain, more favourable economic conditions, one would never be able to know <u>how big</u> a fraction. This is so because it is not clear under what circumstances the respondent would want a job, whether his/her perception of the job market is at all realistic, or, indeed, what meaning the respondent attaches to such terms as "need", "wish", or "want".

Similar objections apply to the possible inclusion of persons stating "suitable jobs not available" in the unemployment concept.

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[1] This is true only in principle because the person interviewed may be unaware of, lie, or forget about the respondent's activities. Usually only one person is intervived about the conditions of all respondents belonging to the same household.

Again, we do not know what the respondent thinks of as a "suitable job" or, indeed, if such a job is anywhere in existence, cf. definition 2.4 above.

Questions (1) through (4) above are all of an attitudinal nature; they have to do less with the respondents' actual behaviour than with their stated preferences and attitudes. If we were to construct a measure of labour underutilization based on these questions, we would have to solve the problem of how to predict behaviour on the basis of stated attitudes. This problem is by no means trivial, see Foss (1980:207-220) and references cited therein.

It is true that the objective job-seeking behaviour which acts as the crucial criterion for being classified as unemployed does not in itself prove that the respondent is serious in his job-search and has a realistic perception of the job market. However, the probability that this be the case is considered high enough to justify an assumption that active job-seekers do fulfil the conditions specified in definition 2.4.

Now, one might consider the possibility of extending the definition of unemployment by accepting, simply, a wider set of (objective) activities as criteria for "job-seeking" and "availability". Such extensions could take place in several directions. As of today, only activities geared directly towards the job market are considered proof of job-seeking and availability. As was noted above, however, availability is a highly relative concept, in that it depends on the individual's physical and social environment. A particularly important set of labour market barriers is related to the individuals' family situation (childcare responsibilities etc.). Efforts to remove or reduce such barriers, e.g. by applying for entrance at a nearby day-care institution, are not considered job-seeking even if the parent in question may actually have a job waiting for her (or him), and be ready to start working as soon as a suitable day-care arrangement has been found for the child(ren). This convention is a bit paradoxical. It means that highly employable persons that are very close to becoming "available for work" are not included in the labour force, while job-seekers with very small chances of finding an employer willing to hire them are in the labour force. Says Foss (1980:33):

> "At a given point of time persons in the labour force as well as persons outside the labour force may be thought of as varying in the "strength" of their labour force attachment and their attachment to other "spheres of activity" - labour market attachment being conceived of as a

> a continuum rather than a dichotomy. Even though we can easily mention several relevant indicators of labour force attachment, there is no easy way of determining the relative strength of a person's total attachment or the ordinal location of different groups of persons on a conceptual continuum representing the degree of attachment to the labour market. Labour force attachment must be considered as a multi-dimensional phenomenon and it is not possible or relevant to collapse these dimensions into a single measure. Established measures of "unemployment" do not, for example, necessarily guarantee that the persons that are thus classified, constitute the group of non-employed persons with the objectively strongest labour force attachment. If we regard other "attachment indicators" than those specified in the ILO recommendations, we may find that groups outside the labour force (according to the ILO definitions) in several ways can be said to show an attachment to the labour market which is as strong as (and in some ways even stronger than) that of persons counted as ["unemployed"] by established metods."

Yet, if another set of activities were to be added to the list of approved job-seeking criteria, it would be hard to distinguish between those cases in which a given activity (not directed at the job market) had an increased "availability" purpose and those motivated by entirely different concerns. Again, obtaining such information would be difficult without resorting to hypothetical or attitudinal questions.

To sum up, the conceptual framework and conventions underlying current labour market statistics are such that the estimated labour force is unlikely to capture all those individuals who, according to certain ideal standards of measurement, should be considered part of the country's labour supply. At most points of time, the measured labour force will be smaller than the hypothetical full employment labour force.

All efforts to extend the usual unemployment concept so as to comprise a larger portion of the assumed "true" supply of labour seem, however, to be inflicted with serious methodological shortcomings. In particular, it appears that a more comprehensive labour supply concept cannot be constructed without departing even further from the principle that the statistical measures be based on objective criteria.

One final remark is in order here. This section has focused on methods to determine the number of individuals supplying a smaller or higher amount of labour. However, since this amount may vary drastically between different members of the labour force, or over time, a more accurate labour supply measure would be expressed in hours rather than in persons. The total number of hours worked during the survey week is stated by each LFSS respondent. Whether or not this figure is interpretable as the respondent's "true" supply of labour is, however, subject to debate in much the same way as the labour force concept. On the one hand it is clear that many individuals are not free to choose the exact length of their own working week. On the other hand questions about the respondent's desired amount of working hours are necessarily of a hypothetical nature. There seems to be no way to solve this dilemma in a fully satisfactory manner; rather it should be made part of the qualifications attached to any labour supply measure based on hours of work data.

2.4. The econometric approach

In several countries it has been observed that as the economy is picking up speed after passing the bottom of the business cycle, employment grows by a larger number than the reduction in unemployment. Conversely, when economic activity is slowing down, the rise in unemployment is far smaller than the decrease in employment. In other words, the labour force expands and contracts depending on the level of economic activity. The labour force is bigger at full employment than during a recession. A similar relationship is detectable in cross-sectional data sets as well. The difference between the full employment labour force and the labour force at any particular level of economic activity is often referred to as "hidden unemployment".

The existence of such "hidden unemployment", of a magnitude not invariant under shifts in labour demand, is, on the face of it, a nuisance to anyone proposing to construct autonomous and operational supply and demand relations. On the other hand, if it were possible to model the mechanism by which the labour force responds to labour demand variations, this responsiveness might possibly be turned into an asset.

Let p denote the overall rate of labour force participation (within some population subgroup), let $\underset{\sim}{e}$ denote a vector of variables describing labour market tightness, and let $\underset{\sim}{x}$ denote a vector of additional explanatory variables. Postulate

$$(2.1) \quad p = f(\underset{\sim}{x},\underset{\sim}{e}) + u ,$$

i.e. p is a function of $\underset{\sim}{x}$, $\underset{\sim}{e}$, and a random disturbance term u with zero expectation. If a set of observations $\{p_i, \underset{\sim}{x}_i, \underset{\sim}{e}_i; \quad i=1,2,\ldots,n\}$ exists on p, $\underset{\sim}{x}$ and $\underset{\sim}{e}$, and a suitable parametric functional form is assumed for f, the relation (2.1) can be calibrated using standard statistical estimation techniques. Denote by

$$(2.2) \qquad \hat{p}_i = \hat{f}(\underset{\sim}{x}_i, \underset{\sim}{e}_i)$$

the predicted value of $E(p_i)$ obtained by plugging the i'th set of variable values into the estimated relation.

Now, denote by $\underset{\sim}{e}^0$ the value of $\underset{\sim}{e}$ corresponding to full employment. The corresponding rate of LFP can be inferred as

$$(2.3) \quad \hat{p}_i^0 = \hat{f}(\underset{\sim}{x}_i, \underset{\sim}{e}^0).$$

p_i^0 is interpretable as the hypothetical rate of LFP at sample point i, given full employment. The difference $\hat{p}_i^0 - \hat{p}_i$ between the hypothetical and the observed rate of LFP expresses "hidden unemployment".

Assuming that a reasonably big sample has been used to estimate f, $\hat{p}_i^0$ depends only marginally on $\underset{\sim}{e}_i$, the tightness of the labour market at sample point i. Apparently, $\hat{p}_i^0$ is precisely what we want: a labour supply measure invariant under shifts in labour demand.

Whether or not this invariance is more than apparent depends, however, on the precise form and content of relation (2.1): what variables form part of $\underset{\sim}{x}$ and $\underset{\sim}{e}$, how are they measured, and how is $\underset{\sim}{e}^0$ determined? The measure $\hat{p}_i^0$ is, in short, subject to all the usual pitfalls of econometric misspecification.

Within the framework of the traditonal market economic model, the supply of labour corresponds to the locus of equilibria under shifts in labour demand, and vice versa. Thus, for supply and demand to be identifiable each function must contain at least one argument (the "shift parameter(s)") not included in the other. The precise meaning of both concepts boils down to a question of what arguments are included in either function. But this is, in fact, a question subject to the analyst's own discretionary choice. Depending on the exact purpose of his analysis and on his theoretical and empirical insight, intuition, or prejudice, he may choose to specify his model in a lot of different ways. The definition of supply and demand will, therefore, ultimately have to rest on some rather arbitrary methodological choices.

This is not to say that just about any specification will do. A prime concern in all econometric work is to build relations that are as autonomous (stable) as possible (Haavelmo 1944). Indeed, this entire discussion departed from the idea that the size of the labour force is not autonomous with respect to certain types of structural change. What we have in mind here are such phenomena as changes in international commodity prices, exchange rates, production technology, or governments expenditure. These are factors which we, in line with established economic theory and usage, prefer to think of as affecting the aggregate <u>demand</u> for labour, through their relation to the firms' total volume of production.

Now, assume that a theory has been built specifying the arguments entering the supply and demand functions, respectively. Let us, however, make the rather realistic assumption that the analyst is unable to measure all the variables playing a role in his theory. Returning to relation (2.1), we may think of $\underset{\sim}{x}$ as that subset of labour supply determinants for which data are available in a given study.

Then, if (any element of) $\underset{\sim}{e}$ is correlated with some relevant explanatory variables not included in $\underset{\sim}{x}$, there is an omitted variable bias. Such a bias may be hard to avoid if, e.g., one is confined to work with a data set not including wage. Under tight labour market conditions, wages are usually bid up. Thus, the increase in LFP observed under such circumstances is explicable in at least two ways. Our interest is focused on the hypothesis (i) that, even if all the arguments of the labour supply function are kept constant, LFP increases with positive shifts in labour demand because more people expect being able to find a job. To the extent, however, that labour supply is not completely inelastic, an equally plausible explanation could be (ii) that LFP increases because jobs become better paid. In other words, all the arguments of the labour supply function[1] are <u>not</u> constant, and (part of) the observerved LFP increase reflects merely a movement along the wage axis.

[1] Although the choice of arguments to enter the labour supply function is discretionary, few economists would disagree that for most purposes the wage rate is among the more obvious explanatory factors.

In the US, the cyclical variations in the labour force have been shown to be far greater than the simultaneous changes in "discouraged workers", as estimated by the Current Population Survey. Finegan (1978:37) observes:

> "... many persons who decide to enter the labor market when unemployment is low are not reported as discouraged workers when unemployment is high, and many of the persons reported as discouraged when unemployment is high do not take jobs or look for work when unemployment is low. That is why the concepts of hidden unemployment and discouraged workers, which appear to be so similar, are actually so different.
>
> Consequently, the count of discouraged workers in a year of high unemployment, while interesting and important in its own right, tells us nothing about how many persons can be expected to enter the labor force when unemployment falls."

In Norway, the time series relationship between labour force, unemployment and "discouraged workers" have never been properly analyzed. Survey data on "discouraged workers" were, until the late 1970s, not even published, due to the difficulties of interpretation inherent in such attitudinal data.

The disagreement between the CPS "discouraged workers" statistics and the econometrically derived "hidden unemployment" estimates may be thought to lend support to the wage level explanation of labour force fluctuations: Although many persons do not "want a job <u>now</u>" and hence are not classified as discouraged, they do, in fact, want one when the market wages have been bid up sufficiently.

Other explanations, more compatible with the "discouragement" hypothesis, are, however, also possible. Some would see the differences between the "discouraged workers" and the "hidden unemployment" figures simply as another illustration of the difficulty of predicting behaviour from attitude measurements. A number of interview technical reasons may be cited in partial explanation of the low "discouraged workers" statistics: many respondents may refuse to admit that they "want a job" but have failed in obtaining one; the fact that only one person is usually interviewed about the conditions of all household members may seriously distort the statements; persons choosing to go to school because they cannot get a job are not classified as "discouraged"; etc. (see Finegan 1978).

The discouraged worker mechanism has a counteracting force known as "added workers". When unemployment is high, some households choose to send more members into the labour force as a safeguard against reduced family income when primary breadwinners lose their jobs. The observed response of the labour force to variations in labour market tightness is a net result of the discouraged and added worker effects (and possibly others).

Empirical analyses from a large number of countries are almost unanimous in indicating that the discouraged worker effect outweighs the added worker effect. Whether aggregate or disaggregate, cross-sectional or temporal, most studies indicate a highly significant, positive relationship between the size of the labour force and the tightness of the labour market. As a proxy for the latter the observed rate of unemployment is the most commonly used measure, generally yielding a negative coefficient value (Bowen and Finegan 1969, Flaim 1973, Mincer 1966 and 1973, Finegan 1978, Egle 1979, Salais 1971 and 1977, Eymard-Duvernay and Salais 1975, Vannebo 1977). Other measures of labour market tightness have, however, also been tried. In examining married women's labour force participation, Ljones (1979) uses a female job opportunity index calculated by applying nationwide, sector-specific female-male employment ratios to the industry mix of the local labour market. Female LFP is found to vary positively with this job opportunity index. A similar index is used by Nakamura et al. (1979) and by Nakamura and Nakamura (1981), who, however, also include the unemployment rate. Both variables come out significant with the expected sign.

3. AN ECONOMETRIC MODEL OF INDIVIDUAL LABOUR MARKET BEHAVIOUR

3.1. Overview

In the terminology of modern economic labour market theory, each individual (i, say) is said to have a reservation wage rate (w_{0i}, say) given as the lowest wage rate at which she would accept to work. If the market wage rate (w_i, say) exceeds the reservation wage rate, the individual supplies labour to the market. In this case, the quantity of labour supplied is referred to as the notional hours of work (t_i, say), determined so as to yield a utility-maximizing combination of purchasing power and leisure. If the reservation wage rate is higher than the market wage rate, the person's labour supply is zero. In this case, the notional hours of work are reduced to a purely hypothetical construct defined as the amount of work the individual would choose to do if institutional constraints compelled her to take a job.

In mathematical notation,

$$(3.6) \quad t_i = \underset{\sim}{\theta}'\underset{\sim}{z}_i + v_i ,$$

$$(3.11) \quad \bar{t}_i = \begin{cases} t_i & \text{if } w_i > w_{0i} \\ 0 & \text{if } w_i \leq w_{0i} \end{cases} .$$

Here, $\underset{\sim}{z}_i$ is a vector of explanatory variables characterizing individual i and/or her social and economic environment, v_i is a zero expectation random disturbance term, $\underset{\sim}{\theta}$ a vector of unknown parameters, and $\bar{t}_i$ is the supply of labour generated by individual i.

Note that the notional hours may be positive even if the supply of labour is not. This occurs if, e.g., there are certain (fixed) time or money costs of labour market entry, such as travel expenses, childcare arrangements, etc.

The notional hours are, however, observable only when $w_i > w_{0i}$ (implying $\bar{t}_i > 0$). We are therefore confined to estimate relation (3.6) using a sample of working individuals only. This gives rise to a so-called selectivity bias affecting the least squares estimator of $\underset{\sim}{\theta}$. In other words, a standard linear regression analysis is unable to provide consistent estimates of the parameters of the theoretical notional hours equation. What we do get is an estimate ($\hat{\underset{\sim}{\alpha}}$, say) measuring how each independent variable affects the supply of labour, given that the latter is positive. Under certain (weak) assumptions, $\hat{\underset{\sim}{\alpha}}$ is unbiased ($E(\hat{\underset{\sim}{\alpha}}) = \underset{\sim}{\alpha}$).

Denote by p_i^0 the probability of a positive labour supply, i.e.

(3.27) $p_i^0 = P(w_i > w_{0i})$.

The expected supply of labour generated by person i is

$$E(\bar{t}_i) = E[t_i \mid w_i > w_{0i}] \cdot P(w_i > w_{0i}) + 0 \cdot P(w_i \leq w_{0i}) \tag{3.31}$$
$$= \underset{\sim}{\alpha}' \underset{\sim}{z}_i \cdot p_i^0.$$

To evaluate this expression we must find an estimate of p_i^0. Drawing on recent advances in the theory of job market rationing and discouraged workers, and making a few simplifying but convenient assumptions, we derive the following logistic model of individual LFP (p_i):

$$p_i \approx P[\log w_i > \log w_{0i} - \varepsilon n_i - e_i] \tag{3.21}$$
$$= [1 + \exp(-\underset{\sim}{\beta}' \underset{\sim}{x}_i)]^{-1} \tag{3.24}$$

where

$$\underset{\sim}{x}_i = (\underset{\sim}{z}_i' \; n_i)'. \tag{3.26}$$

Here, n_i is the local rate of unemployment, $\underset{\sim}{\beta}$ and ε are unknown parameters, e_i is an error term and $\underset{\sim}{z}_i$ is the vector of individual or contextual characteristics introduced already in equation (3.6).

In other words, labour force participation depends on the same factors as do the notional hours of work, although quite possibly with different coefficients. In addition, LFP is assumed to vary with the tightness of the labour market, as described by the jobless rate n_i. According to this theory, the probability (p_i^0) of a non-zero labour supply is given simply as the LFP probability evaluated at zero unemployment:

$$p_i^0 = p_i \Big|_{n_i = 0}.$$

A consistent and efficient estimate of β is found by applying maximum likelihood techniques to the logit model (3.24). Plugging this estimate back into the model, and setting $n_i = 0$, the probabilities p_i^0 and hence also the expected supply of labour hours (3.31) can be consistently estimated.

The variables assumed to form part of the explanatory vector $\tilde{z}_i$ are the following: age, marital status, education, the number and age of children, and the relative capacity of local kindergartens. Separate models are estimated for men and women. Our data set is taken from the Norwegian Labour Force Sample Surveys of 1977 and consists of 9 276 female and 8 922 male respondents.

A further account of the theoretical model and of its practical empirical implementation is given in sections 3.2 to 3.6 below.

3.2. Theoretical basis: a utility model with fixed costs of market work

In specifying an operational model of individual labour supply, we shall build on recent advances in the field of labour market econometrics (see, e.g. Smith 1981). These developments rely, in turn, on the neo-classical theory of utility maximization.

A fairly general framework for analyzing individual labour market behaviour has been put forward by Cogan (1981). The basic ideas underlying Cogan's model are illustrated by figure 3.1.

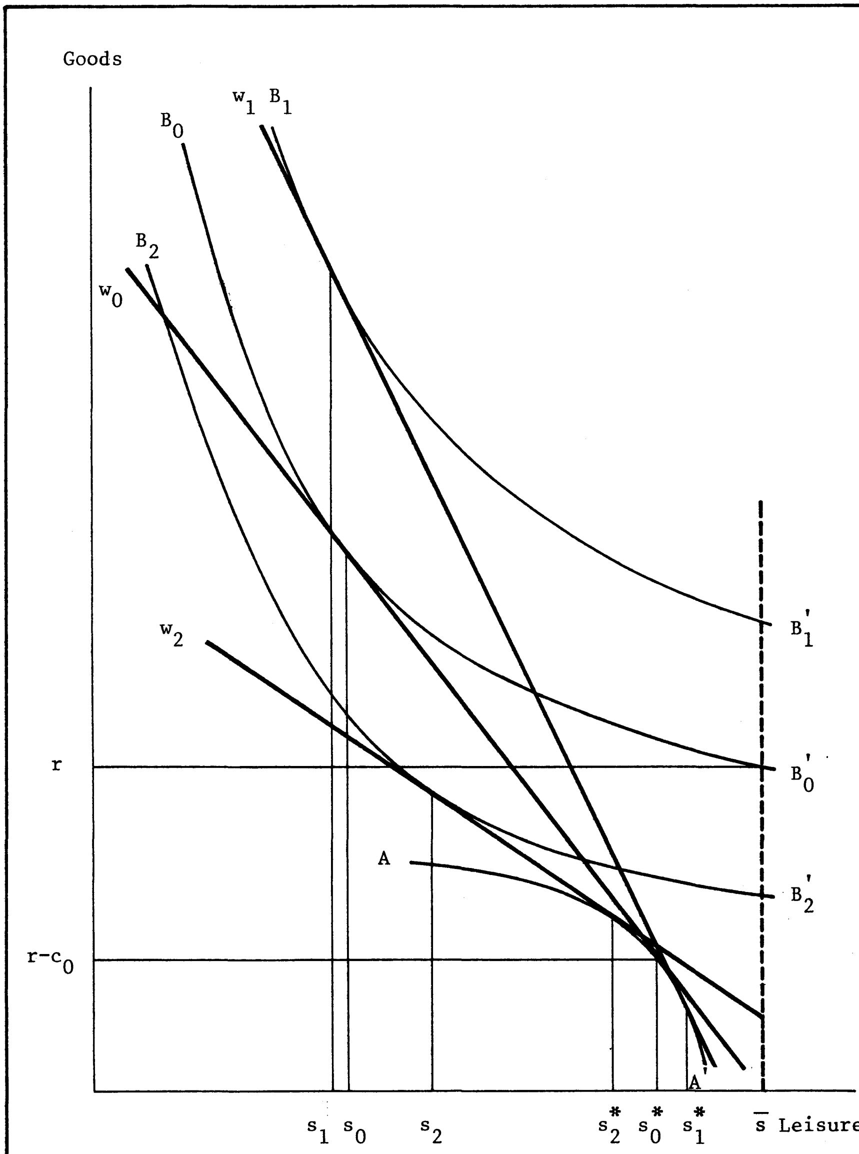
B_0B_0', B_1B_1', and B_2B_2' represent the indifference map of a given individual choosing between leisure (s)[1] and a composite commodity (a). The individual has an unearned income equal to r, originating from financial assets spouse's work for pay or profit, or other sources. Thus, without working the individual is able to reach a utility level given by the curve B_0B_0'.

If the individual elects to work in the market, certain (fixed) time and money costs of labour market entry are incurred. In the most general formulation, there is substitutability between time and money costs of work, as given by the entry cost opportunity curve AA'. Fixed costs of work may include such items as clothing, transport, and childcare expenses, while time costs may include travel time and time spent bringing children to and from their day-care institution.

The fact that residential housing rents decrease with the distance to the city centre has traditionally been interpreted as evidence that households do choose between time and money costs of work.

[1] "Leisure" in this study means "time not devoted to market work". In other words, "leisure" includes household work, childcare etc.

Figure 3.1. Labour supply with fixed time and money costs of work



The lowest wage at which the individual will supply labour is given by w_0, the slope of the tangent line between the entry cost opportunity curve and the indifference curve passing through the point $(\bar{s},r)$. ($\bar{s}$ is the maximal amount of leisure.)

w_0 is called the reservation wage rate. The amount of labour supplied at this wage rate is often referred to as reservation hours. In our diagram, the individual's reservation hours are given by $s_0^* - s_0 \equiv t_0$ (say), obtained by substracting $\bar{s}-s_0^*$, the time cost of labour market entry, from $\bar{s} - s_0$, the total amount of leisure foregone in order to work t_0 hours.

Now, assume that the market wage rate for our sample individual is $w_2 < w_0$. At this (low) wage rate, the individual will choose to incur more time costs and less money costs of work, i.e. $\bar{s} - s_2^*$ hours will be spent going to and from work etc. The best attainable utility level given that the individual works at a wage rate w_2 is represented by the curve B_2B_2'. The corresponding number of working hours at this wage rate, $s_2^* - s_2 \equiv t_2$, is referred to as the notional hours of work.

However, since the utility level B_2B_2' is inferior to B_0B_0', the individual will choose not to work at all. Thus, in this case the notional hours of work are inobservable; they consititute a purely hypothetical concept.

If, on the other hand, the market wage rate is $w_1 > w_0$, the individual is able to reach a higher level of utility by working than not. Here the supply of labour equals the notional hours of work, given by $s_1^* - s_1 \equiv t_1$.

Notional hours t depend on the market wage rate w, on the fixed time and money costs $\bar{s} - s^*$ and c, on unearned income r, on a vector of observable exogeneous variables $\underset{\sim}{z}_1$ determining the utility structure, and on a random disturbance term u_1 reflecting unobserved factors entering the utility function. Assuming a linear relationship, we may write, for individual i,

$$(3.1) \quad t_i = \gamma_1 \log w_i + \gamma_2 (\bar{s}-s_i^*) + \gamma_3(r_i-c_i) + \underset{\sim}{\chi}'\underset{\sim}{z}_{1i} + u_{1i}.$$

Here and in the sequel, Greek letters denote unknown parameters.

The optimal trade-off between time and money costs of work depends on the market wage rate, on a set of exogenous variables $\underset{\sim}{z}_2$ reflecting differences in participation cost constraints among individuals, and on a set of disturbance terms u_2 and u_3:

$$(3.2) \quad \bar{s}-s_i^* = \delta_1 \log w_i + \underset{\sim}{\delta}'\underset{\sim}{z}_{2i} + u_{2i},$$

and

$$(3.3) \quad c_i = \zeta_1 \log w_i + \underset{\sim}{\zeta}'\underset{\sim}{z}_{2i} + u_{3i}.$$

The market wage rate is, in turn, determined by

$$(3.4) \quad \log w_i = \underset{\sim}{\eta}'\underset{\sim}{z}_{3i} + u_{4i}.$$

Substituting (3.2) and (3.3) into (3.1) we obtain

$$(3.5) \quad t_i = (\gamma_1+\gamma_2\delta_1-\gamma_3\zeta_1)\log w_i + \gamma_3 r_i + \underset{\sim}{\chi}'\underset{\sim}{z}_{1i} + (\gamma_2\underset{\sim}{\delta}-\gamma_3\underset{\sim}{\zeta})'\underset{\sim}{z}_{2i}$$
$$+ u_{1i} + \gamma_2 u_{2i} - \gamma_3 u_{3i}.$$

The reduced form equation for t_i is obtained by substituting (3.4) into (3.5):

$$(3.6) \quad t_i = \underset{\sim}{\theta}'\underset{\sim}{z}_i + v_i,$$

where we have defined

$$(3.7) \quad \underset{\sim}{\theta}' = (\underset{\sim}{\chi}' \quad (\gamma_2\underset{\sim}{\delta}-\gamma_3\underset{\sim}{\zeta})' \quad (\gamma_1+\gamma_2\delta_1-\gamma_3\zeta_1)\underset{\sim}{\eta}' \quad \gamma_3),$$

(3.8) $\underset{\sim}{z}_i = (\underset{\sim}{z}'_{1i} \quad \underset{\sim}{z}'_{2i} \quad \underset{\sim}{z}'_{3i} \quad r_i)'$,

and

(3.9) $v_i = u_{1i} + \gamma_2 u_{2i} - \gamma_3 u_{3i} + (\gamma_1\gamma_2\delta_1 - \gamma_3\zeta_1)u_{4i}$.

Let w_{0i} denote the reservation wage rate. Note that the reservation hours, t_{0i}, are given by plugging $w_i = w_{0i}$ into the right hand side of (3.5). However, equation (3.4) does not apply to w_{0i}. Thus, to solve for t_{0i} and w_{0i} we need one more relation, expressing the fact that the slope of the reservation wage line is determined by the relative positions of the entry cost opportunity curve and the indifference curve passing through $(\bar{s},r)$:

(3.10) $$\log w_{0i} = \underset{\sim}{\xi}'_1 \underset{\sim}{z}_{1i} + \underset{\sim}{\xi}'_2 \underset{\sim}{z}_{2i} + \xi_3 r_i + u_{5i}$$

$$= \underset{\sim}{\xi}' \underset{\sim}{z}_i + u_{5i},$$

where $\underset{\sim}{\xi}' = (\underset{\sim}{\xi}'_1 \ \underset{\sim}{\xi}'_2 \ \underset{\sim}{0}' \ \xi_3)$, cf. (3.8).

We are now ready to define the <u>labour supply function</u> $\bar{t}_i$:

(3.11) $$\bar{t}_i = \begin{cases} t_i & \text{if } w_i > w_{0i} \\ 0 & \text{if } w_i \le w_{0i} \end{cases}.$$

That is, if the market wage rate exceeds the reservation wage rate, the individual is said to supply an amount of labour identical to the notional hours of work. When the opposite condition holds, the supply of labour is zero.

3.3. Accounting for job market rationing and costs of search

Assume that the job market is rationed so that individual i is not necessarily able to obtain a job at her prevailing market wage rate w_i. If the probability of obtaining a job is perceived by the individual as sufficiently low, she may not find it worthwhile to enter the labour force, although she would have done so had the number of job openings appeared more promising. We have previously referred to such individuals as "discouraged workers".

A theory linking the discouraged worker hypothesis to the general theory of individual labour supply has been put forward recently by Eaton and Quandt (1983) and, independently, by Dagsvik (1984). The below discussion is based partly on their works.

Let ρ_i denote the probability that individual i will obtain a job, given that she enters the labour force. Assume that ρ_i is known to the individual. (Alternatively, we may think of ρ_i as the individual's subjectively perceived probability of getting a job.) The probability that individual i will belong to the labour force can be written as

$$(3.12) \quad p_i = P[\rho_i \cdot \sup_{t_i} U(r_i-c_i-m_i+w_it_i,t_i) + (1-\rho_i)U(r_i-m_i,0) > U(r_i, 0)]$$

where $U(r,t)$ is the utility of consumption r and leisure $\bar{s}-t$, while m_i denotes the cost of job search.[1] Labour force participation occurs if the expected utility of job search exceeds the utility level attained when no work is done.

[1] For simplicity of exposition only money costs of search are considered. The argument is easily extended so as to handle also time costs of search, however without the conclusions being affected.

The indirect utility on the left hand side of the inequality sign is, obviously, an increasing function of w_i. Thus, there exists a wage rate $\tilde{w}_{0i}$ (say) such that the bracketed expression holds as an equality:

(3.13) $\rho_i \cdot \sup_{t_i} U(r_i - c_i - m_i + \tilde{w}_{0i} t_i,\ t_i) + (1-\rho_i) U(r_i - m_i,\ 0) = U(r_i,\ 0).$

In other words, the LFP probability can be written

(3.14) $p_i = P[\log w_i > \log \tilde{w}_{0i}].$

We may suitably refer to $\tilde{w}_{0i}$ as the "effective reservation wage rate in the presence of job market rationing and costs of search". It is a function of m_i and ρ_i:

(3.15) $\tilde{w}_{0i} = \tilde{w}_{0i}(m_i,\ \rho_i).$

Recall that the "theoretical" reservation wage rate (w_{0i}), as defined by (3.10), fulfils the condition

(3.16) $\sup_{t_i} U(r_i - c_i + w_{0i} t_i,\ t_i) = U(r_i,\ 0).$

Comparing (3.13) and (3.16), we note that when $\rho_i = 1$, $\tilde{w}_{0i}$ differs from w_{0i} only in that there is an extra fixed cost of labour market entry to be compensated. Thus, the value $\tilde{w}_{0i}(m_i, 1)$ is equal to the slope of the reservation wage line which results from shifting the entry cost opportunity curve upwards by an amount m_i. Indeed, if we assume that the search cost is included in the fixed costs of work considered by the individual, we may write

$$(3.17)\quad w_{0i} = \tilde{w}_{0i}(m_i,1).$$

Denote by n_i the local rate of unemployment. It may not be unreasonable to assume that

$$(3.18)\quad 1-\rho_i = \nu n_i$$

i.e. the (perceived) probability of not obtaining a job is proportional (identical if $\nu=1$) to the prevailing jobless rate. Also, assume that the cost of job search for different individuals is randomly distributed around a mean value $\bar{m}$:

$$(3.19)\quad m_i = \bar{m} + u_{6i},$$

u_{6i} being a zero expectation error term.

Now, expanding $\log \tilde{w}_{0i}$ around $\rho_i=1$ and its derivative around $m_i = 0$, we may, in view of (3.17), rewrite (3.14) as

$$(3.20)\quad p_i \approx P\left[\log w_i > \log \tilde{w}_{0i}(m_i,1) + (\rho_i - 1)\frac{\partial \log \tilde{w}_{0i}(m_i,1)}{\partial \rho_i}\right]$$

$$\approx P\left[\log w_i > \log w_{0i} - \nu n_i \cdot \left[\frac{\partial \log \tilde{w}_{0i}(0,1)}{\partial \rho_i} + m_i \frac{\partial^2 \log \tilde{w}_{0i}(0,1)}{\partial \rho_i \partial m_i}\right]\right].$$

Note that the first derivative inside the last pair of brackets vanishes since the hiring probability ρ_i does not matter when search is free of costs.

(To see this, plug $m_i = 0$ into (3.13) and compare with (3.16).) Thus

$$(3.21)\quad p_i \approx P[\log w_i > \log\ w_{0i} - \varepsilon n_i - e_i]$$

$$= P[u_{5i} - u_{4i} - e_i < \underset{\sim}{\eta}'\underset{\sim}{z}_{3i} - \underset{\sim}{\xi}'\underset{\sim}{z}_i + \varepsilon n_i],$$

where we have defined

$$(3.22)\quad \varepsilon = \nu\bar{m}\cdot\frac{\partial^2 \log \tilde{w}_{0i}(0,1)}{\partial\rho_i\partial m_i}$$

and

$$(3.23)\quad e_i = \nu n_i u_{6i}\cdot\frac{\partial^2 \log \tilde{w}_{0i}(0,1)}{\partial\rho_i\partial m_i}\cdot$$

Assuming that $u_{5i} - u_{4i} - e_i$ is logistically distributed with, say, variance $\sigma^2\pi^2/3$, one can write

$$(3.24)\quad p_i = [1 + \exp(-\underset{\sim}{\beta}'\underset{\sim}{x}_i)]^{-1}$$

where

$$(3.25)\quad \underset{\sim}{\beta}' = -(\underset{\sim}{\xi}_1'\ \underset{\sim}{\xi}_2'\ \underset{\sim}{\eta}'\ \xi_3\ -\varepsilon\ )/\sigma$$

and

$$(3.26)\quad \underset{\sim}{x}_i = (\underset{\sim}{z}_i'\ n_i)'.$$

Denote by p_i^0 the LFP probability evaluated at n_i=0. Note (from (3.21) and (3.23)) that

$$(3.27)\quad p_i^0 = P(w_i > w_{0i}),$$

i.e. p_i^0 coincides with the probability of a non-zero labour supply, as defined by (3.11).

3.4. Empirical specification: an imperfect but operational measure of labour supply

In order to estimate the above model, the variables entering the model must be specified in a way compatible with available data sets. For the present study, the set of candidate explanatory variables has, by and large, been constrained to those collected through the Norwegian Labour Force Sample Surveys, with a couple of additions.

Lacking data on market wages (w_i), reservation wages (w_{0i}), non-wage income (r_i), and fixed costs of work ($\bar{s}-s_i^*$ and c_i), the best we can do is to estimate the reduced form (3.6). To the extent that the same variables form part of $\underset{\sim}{z}_{1i}$, $\underset{\sim}{z}_{2i}$, and $\underset{\sim}{z}_{3i}$, only the combined effect can be identified.

The following variables are assumed to enter $\underset{\sim}{z}_i$: age (AGE), old-age pension eligibility (RETAGE), marital status (MAR), level of education (EDUC), number of children (NUMCH, for women), age of youngest child (AGECH, for women), and the relative capacity of local kindergartens (KINDG, for married women with young children).

Age and education may be assumed to form part of the variable set $\underset{\sim}{z}_{1i}$ influencing the indifference structure, as well as of $\underset{\sim}{z}_{3i}$, which determines the market wage rate. Child status obviously affects the location of the entry cost opportunity curve (through $\underset{\sim}{z}_{2i}$), as well as the indifference structure. The availability of local day-care institutions is used as a proxy for time and money costs of labour market entry among married women with small children. Non-wage income (r_i) is, unfortunately, not available in our data set. As proxies for r_i, marital status and old-age pension eligibility are used.

The jobless rate applicable to individual i (n_i) is set equal to the local rate of registered unemployment (UNEMP). Since in our data set several regions are found to exhibit zero jobless rates, UNEMP=0 is used as a (natural) reference point, representing, by convention, the "full employment" level of unemployment. It should be understood that this (zero) level of unemployment may not be attainable as a national average, due to friction in the labour market.

The conventions and definitions underlying all the variables used are stated in greater detail in chapter 4 and in appendix 1. The advantages and disadvantages of each variable and the possible misspecification biases occurring are discussed at some length in chapter 4.

Since the notional hours of work (t_i) are not observed unless $\bar{t}_i > 0$, we are confined to estimate relation (3.6) using a sample consisting of working individuals only. As a result of this, standard least squares estimators of Θ are subject to a so-called selectivity bias (see, e.g., Heckman 1981). Taking the conditional expectation of t_i, given $w_i > w_{0i}$, we have (cf. (3.6)):

$$(3.28) \quad E[t_i | w_i > w_{0i}] = \underset{\sim}{\Theta}' \underset{\sim}{z}_i + \omega_i$$

where we have defined

$$(3.29) \quad \omega_i = E[v_i | u_{5i} - u_{4i} < \underset{\sim}{\eta}' \underset{\sim}{z}_{3i} - \underset{\sim}{\xi}' \underset{\sim}{z}_i].$$

In the formulation (3.28), the selectivity bias is seen to originate from the omission of an independent variable whose expected value ω_i equals the conditional expectation of the disturbance term given the sample selection rule. In general, this conditional expectation does not vanish, (i) because the composite error term v_i is a function, inter alia, of u_{4i} (cf. (3.9)), and (ii) because all the error terms $u_{1i}, u_{2i}, \ldots, u_{5i}$ may very well be correlated.

From equation (3.29) it is evident that ω_i is, in fact, a function of $\underset{\sim}{z}_i$. Assume that we can write

(3.30) $\omega_i = \underset{\sim}{\gamma}'\underset{\sim}{z}_i,$

i.e., ω_i is a linear function of $\underset{\sim}{z}_i$. Obviously, when single-equation estimation techniques are applied to (3.6) (for $\bar{t}_i > 0$), the result is an unbiased estimator not of $\underset{\sim}{\theta}$, but of the - in fact - more interesting parameter $\underset{\sim}{\alpha} = \underset{\sim}{\theta} + \underset{\sim}{\gamma}$. $\underset{\sim}{\alpha}$ has an interpretation as the vector of differential effects on the supply of labour, given that the latter is positive.

Thus, under the assumption (3.30), selectivity represents no big cause for alarm, since the expected supply of labour from individual i remains calculable as

$$(3.31)\quad E(\bar{t}_i) = E[t_i \mid w_i > w_{0i}]\cdot P(w_i > w_{0i}) + 0\cdot P(w_i \leqq w_{0i})$$

$$= (\underset{\sim}{\theta}'\underset{\sim}{z}_i + \omega_i)\cdot p_i^0$$

$$= \underset{\sim}{\alpha}'\underset{\sim}{z}_i\cdot p_i^0.$$

All we need, in addition to $\hat{\underset{\sim}{\alpha}}$, is a consistent estimator of p_i^0. But this can be obtained by applying maximum likelihood (ML) estimation techniques to the logit model (3.24), and then plugging the ML estimates and $n_i = 0$ back into the model (McFadden 1974).

By summing our estimate of $E(\bar{t}_i)$ over all members of a given population subgroup, the amount of labour hours generated by that subpopulation under "full employment" conditions can be evaluated.

3.5. A note on statistical methodology

The merit of Cogan's modelling framework lies in its relative generality. In allowing for the existence of fixed time and money costs of work, the possibility that these costs may be zero is by no means ruled out. It is interesting to see what happens in this special case.

If there are no fixed costs of work, the entry cost opportunity curve AA' degenerates into the single point $(\bar{s},r)$ (cf. figure 3.1.). The (theoretical) reservation wage rate coincides with the slope of the indifference curve passing through $(\bar{s}, r)$, reservation hours equal zero, and the supply of labour is positive if and only if notional hours are. In others words,

$$(3.32)\quad p_i^0 = P[w_i > w_{0i}]$$

$$= P[t_i > 0]$$

$$= P[\underset{\sim}{\theta}'\underset{\sim}{z}_i + v_i > 0].$$

Thus, the implication of no fixed costs of work is that the statistical model has a very restrictive structure, in that the same linear combination factors $\underset{\sim}{\theta}'\underset{\sim}{z}_i$ determine both the participation probability and the notional hours of work. That is, not only are the independent variables identical, so are their coefficients!

Assume further that the job market is not rationed, and that v_i is normally distributed. In this case we are left with the familiar tobit model (Tobin 1958).

In other words, the statistical procedure adopted in this study differs from the tobit model in being less restrictive on three accounts: (i) the possibility of fixed costs of work is allowed for, implying that the coefficients and variable sets may differ between the hours equation and the participation equation, (ii) job market rationing is not ruled out, and (iii) no assumption is made concerning the shape of the error distribution in the hours of work equation.

In an empirical study of married women's labour supply, Cogan (1981) compares the two models and concludes that the constrained model (with no fixed costs) results in significantly lower explanatory power and seriously biased estimates. Similar results were obtained in an early phase of this research, when the tobit model was, indeed, tried out as an alternative to the unconstrained two-step linear/logit procedure[1].

[1] Reproducing these results is outside the scope of this publication. They are, however, available from the author upon request.

The normality assumption implicit in the tobit model is not entirely innocuous either. Arabmazar and Schmidt (1982) have shown that the tobit estimators may be quite sensitive to departures from normality. Our experiments with the tobit model resulted in residual distributions that were quite visibly non-normal, especially for male respondents[1].

The two-step linear/logit technique used in this study allows us to compute the effects of all exogenous variables (i) on labour force participation, (ii) on the conditionally expected number of hours supplied, given a positive supply, and (iii) on the unconditionally expected number of hours supplied, calculated as the product of (i) and (ii).

The mean hours of work may increase as a consequence of either a large number of persons working a little more, or a smaller group of workers increasing their working week from, say, 20 to 40 hours. It is therefore of some interest to see how the exogenous factors affect (iv) the split between part-time and full-time work.

If the disturbance terms of the hours equation have a known distribution, as is assumed in the tobit model, the proportion of workers in each hours interval is calculable. Rather than assuming a particular error distribution, however, we have estimated the split between full-time, part-time, and no work by applying a trinomial logit model[2] to the complete data set.

From this model the conditional probabilities of full-time and part-time work, given a positive amount of work, can be calculated. By multiplying these probabilities by the full employment LFP probability we

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[1] See note 1, previous page.

[2] The trinomial logit model looks like this:

$$(3.33)\quad p_{ij} = \exp(\underset{\sim}{\beta}_j'\underset{\sim}{x}_i)/\sum_{k=1}^{3}\exp(\underset{\sim}{\beta}_k'\underset{\sim}{x}_i),$$

where p_{ij} is the probability that individual i chooses alternative j, $\underset{\sim}{x}_i$ is the vector of explanatory variables, and $\underset{\sim}{\beta}_k$ (k=1,2,3) are alternative-specific parameter vectors. Without loss of generality we may put $\underset{\sim}{\beta}_3=\underset{\sim}{0}$. It is easily shown that

$$(3.34)\quad p_{i1}/\,(p_{i1}+p_{i2}) = [1+\exp(-(\underset{\sim}{\beta}_1-\underset{\sim}{\beta}_2)'\underset{\sim}{x}_i)]^{-1},$$

i.e., the conditional probability of choosing alternative 1 (full-time), given that 1 <u>or</u> 2 (part-time) is chosen, is simply a binomial logit function with parameter vector $\underset{\sim}{\beta}_1-\underset{\sim}{\beta}_2$ (cf. equation (3.24)).

are able to split the supply of labour between "full-time supply" and "part-time supply". The interpretation of these last statistics is, however, subject to certain difficulties due to the possible existence of institutional constraints in the labour market. These problems were alluded to already at the end of section 2.3 and are discussed at further length in section 4.2.

The statistical procedure adopted in this study may thus be summarized as follows:

(i) The parameters $\underset{\sim}{\beta}$ of the LFP equation are estimated by applying maximum likelihood techniques to the binary logit model (3.24).

(ii) Full employment LFP probabilities p_i^0 (3.32) are estimated by setting unemployment equal to zero and plugging the $\underset{\sim}{\beta}$ estimate back into (3.24).

(iii) The parameters $\underset{\sim}{\theta}$ of the notional hours function are estimated by applying ordinary least squares to equation (3.6). The estimates are inconsistent due to selectivity bias. The conditonal supply of labour $\bar{t}_i$, given $\bar{t}_i>0$ is, however, consistently estimated.

(iv) The unconditionally expected supply of labour (3.31) is calculated as the product of the full employment LFP probability and the conditionally expected supply measured in hours.

(v) Parameters governing the choice between full-time and part-time work are estimated by means of a trinomial logit model applied to the complete data (i.e., including those not working).

(vi) The supply of labour measured in persons is divided into full-time and part-time supply by multiplying p_i^0 by the conditional full-time, resp. part-time probabilities derivable from (v).

Separate parameter sets are estimated for the male and female part of the sample.

3.6. The data set

The data set used in this study consists of a little less than half the entire LFSS sample for 1977. Students and conscripts are excluded, and no respondent is included more than once, although most of them were interviewed twice during the year. (The LFSS sample is a rotating panel.) Our data set contains 9 176 women and 8 922 men.

The variables used in this study are listed in table 3.1, which is largely self-explanatory. Further details on sample selection and variable nomenclature are given in appendix 1.

Table 3.1. List of dependent and independent variables used in this study

Variable name	Definition
Dependent variables:	
LFP	Dummy for labour force participation
FTW	" " full-time work (30+ hours/week)
PTW	" " part-time work (1-29 hours/week)
T	Hours worked last week
Independent variables:	
CONSTANT	1
AGE	1977 minus year of birth
RETAGE	Dummy for passed retirement age (67 + years)
EDUC1	" " 0-9 years of schooling (level 1)
EDUC2	" " 10 " " " (" 2)
EDUC3	" " 11-12 " " " (" 3)
EDUC4	" " 13-14 " " " (" 4)
EDUC5	" " 15+ " " " (" 5)
MAR1	" " unmarried
MAR2	" " married
MAR3	" " previously married
NUMCH0	" " no child under 16 years
NUMCH1	" " 1 " " " "
NUMCH2	" " 2 children under 16 years
NUMCH3	" " 3+ " " " "
AGECH1	" " child(ren), youngest child 0-2 years
AGECH2	" " " " " 3-6 "
AGECH3	" " " " " 7-15 "
UNEMP	Local rate of unemployment (per cent of labour force)
KINDG	Number of children in day-care institutions per child less than 7 years old resident in the municipality

4. FEMALE LABOUR SUPPLY

4.1. Introduction and overview

This chapter is concerned with the empirical ananlysis of female labour supply. The principal findings may be summarized as follows.

The relationship between female LFP and unemployment is negative and highly significant (figure 4.3). Thus, it appears that the so-called discouraged worker effect outweighs the added worker effect. When (or where) the labour market is slack, a number of persons that would otherwise have attempted to get a job, give up or fail to try. This phenomenon has an important bearing on the interpretation of unemployment statistics. It implies the existence of a large hidden unemployment almost proportional to and possibly much bigger than the observed part.

The effects of marital status and education are also highly significant. By and large, female LFP increases significantly with each level of education, but the pattern of variation is strikingly different between married, unmarried, and previously married women (figure 4.4). Among married women, there is little difference in LFP from the lowest (0-9 years) to the medium (10-12 years) level of education. Typically the difference may amount to about 8 percentage points, depending on age and child status. From the medium to the university level of education (13+ years), there is, however, a difference in the area of 25 percentage points.

Among unmarried women this pattern is, in a sense, reversed. Here, the big difference is from the lowest to the medium level of education, typically 25 percentage points. From the medium to the university level of education the LFP differentials are much smaller (around 6 percentage points).

Previously married women are something in between. Among these, the corresponding LFP differentials between the lowest, medium and highest level of education are typically 17 and 12 percentage points, respectively.

This also means that when one controls for education, the effect of marital status on LFP is very small at the lowest and highest levels of education, but substantial (20-25 percentage points) at the medium level. (Note that the number and age of children are already controlled for.)

In an attempt to shed light upon the factors behind the peculiar pattern of variation observed for married women, a set of regressions were run on another data set taken from the Norwegian Fertility Survey of 1977. Here, wage data are available for both husband and wife. It is found that the wage rates of married men increase systematically with their wives' level of education (figure 4.6). The same is, of course, true of the wives' own wage rates.

Consider the differences between married women with 0-9, 10, and 11-12 years of schooling, respectively. Their estimated average wage rates are, in that order, 25.66, 27.36, and 31.02 kroner per hour (as of 1977). It seems possible, however, that the impetus derived from these increments in the women's own wage rate is offset by similar differentials in the wage rates of their husbands, enabling families with higher levels of education to "afford more leisure". In such a case, the wives will usually be the ones to spend more time at home, since women at the low and medium levels of education have significantly lower wage rates than their respective husbands. Women with 13-14 years of schooling are, on the other hand, almost on a par with them, while those having completed 15 or more years of schooling enjoy, on the average, considerably higher wages than their husbands.

In the light of these wage differentials, the pattern of LFP rates observed for married women with unequal levels of education becomes less inexplicable. This is, of course, not to say that other, non- economic explanations are ruled out.

While female LFP rates vary greatly with education, it is interesting to note that the length of the working week does not. On the average employed women with only compulsory schooling put in only 2-3 hours less per week than do comparable[1] women with a university level education (figure 4.7).

Marital status, by contrast, has a generally stronger impact on the length of the working week than on labour force participation. Once they have a job, unmarried women put in about 4-5 hours more per week than do their married sisters. LFP rates are, as already metioned, not significantly different across marital groups, except for women at the medium level education.

Children seem to constitute the most important barrier to female labour force participation. If LFP among childless women with a certain age, marital status and education is set at 55 per cent, comparable women

[1] That is, given the same age, marital status, and child status.

having one child under 2 would be down to 20 per cent. With two children, of which the younger is under 2, the rate drops to 13 per cent. As the children grow up, however, their mothers' LFP climbs back to 20 per cent (younger child aged 3-6), and later to about 41 per cent (both children aged 7-15) (figure 4.9).

Thus, the age of the youngest child means more than their number. Also, the effect of children is drastically smaller among women with a university level education (figure 4.13).

Job-holding mothers have a 5 to 10 hours shorter working week than comparable married women without children (figure 4.10). The length of the working week does not, however, vary a lot with the age of the youngest child (except when there are 3 or more children).

For married women with pre-school children, labour force participation appears to increase with the availability of institutionalized day-care services (kindergartens). The effect is rather small, though. If, e.g., the relative capacity of kindergartens rises from, say 11.7 per cent (the national average) to 33.9 per cent (the Oslo level), the LFP rate of married women with pre-school children can be expected to grow by an estimated 5 percentage points (figure 4.13).

When controlling for marital status, education and children, women in their 30's seem to have the highest LFP of all age groups. Female LFP drops sharply after the age of 50 (figure 4.14). This is, at least, the case in our cross-section of women surveyed in 1977. However, it need not be true of each particular cohort.

The length of the working week is not nearly as age-dependent as LFP is. Married women under 40 may have an average working week of about 33 hours (given 11-12 years' education and no children under 16). The corresponding figure for 65-year-olds is no smaller than 26 hours per week. Taking account, however, of the large differences in LFP, the per capita labour supply of 65-year-old women is seen to be less than one third of that generated by childless women under 40 (figure 4.15). Again, cohort effects may be playing an important role.

At 67, all Norwegians are entitled to old-age pension from the National Insurance Institution. The effect of old-age pension eligibility has been estimated at 13 percentage points. That is, if the age limit were lowered by one year, the LFP of 66-year-old women would drop from 29 to 16 per cent (as of 1977).

Further details on the effects of each particular explanatory variable are presented and discussed below. In section 4.2, the "raw" results of each round of estimation are reported. Parameter estimates are given for (i) the binomial logit model of labour force participation, (ii) the trinomial logit model of full-time, part-time, and no work, and (iii) the linear regression model of hours worked. Certain tests for interaction in the binomial logit model are described and discussed. The possible biases due to institutional rigidities affecting the individuals' choice of working time are pointed out and discussed.

Sections 4.3 to 4.7 all rely on the results presented in section 4.2. However, each section is concerned with only one or two explanatory variables, that are dealt with in detail. Section 4.3 concerns the relationship between female labour force participation and unemployment. Education and marriage are the subject of section 4.4, the number and age of children are treated in section 4.5, kindergartens in section 4.6, and the effect of age in section 4.7.

4.2. General empirical results and method

Labour force participation

Results from the binomial logit analysis of female labour force participation are summarized in table 4.1.

The set of explanatory variables in the model includes age, level of education, marital status, the number and age of children, the local rate of unemployment, and the relative capacity of local day-care institutions.

The column $\hat{\beta}_g$ of table 4.1 contains maximum likelihood estimates of the logit model parameters. Standard errors are given in the column $\hat{\sigma}_g$. The parameters β_g have an interpretation as log-odds ratios. The (additive) increment in the LFP probability corresponding to a given log-odds ratio will depend on the "initial" value of the probability. The column $\Delta\hat{p}$ presents one set of such increments, namely those resulting when we take the sample mean ($\bar{p}$ = 0.5455) as our initial value.

As shown by the asterisks, most parameters are different from zero at the 1 per cent level of significance, when tested against a two-sided alternative. All but three parameters are significant at the 10 per cent level[1].

[1] For testing against one-sided alternatives, one and two asterisks correspond to 5 and 0.5 levels of significance, respectively.

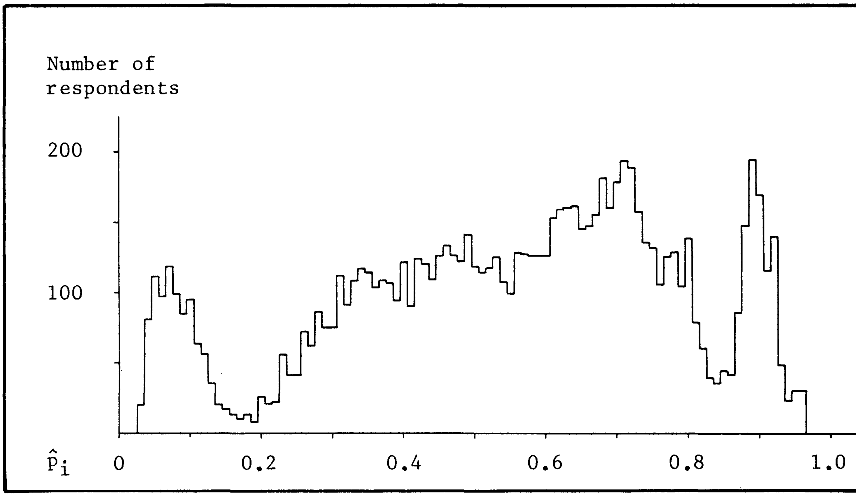
Table 4.1. Estimation results from the binomial logit model of female labour force participation. Women 16-74 years not in school. 1977

g	Independent variable x_{ig}	Parameter estimate $\hat{\beta}_g$	Standard error $\hat{\sigma}_g$	Additive effect at sample mean $\Delta\hat{p}$
1	CONSTANT	-0.9002**	0.2898	.
2	AGE/100	13.9569**	1.5192	
3	$(AGE/100)^2$	-21.3279**	1.7707	-0.014[1]
4	RETAGE	-0.8463**	0.1515	-0.206
5	EDUC2	1.1397**	0.1443	0.244
6	EDUC3	1.1867**	0.1570	0.252
7	EDUC4	1.3934**	0.1098	0.283
8	EDUC5	1.8622**	0.2066	0.340
9	MAR2	-0.1622	0.1139	-0.040
10	MAR3	0.0938	0.1394	0.023
11	MAR2 • (EDUC2 + EDUC3)	-0.8327**	0.1517	-0.203
12	MAR3 • (EDUC2 + EDUC3)	-0.4486*	0.2060	-0.112
13	NUMCH1	-0.0505	0.0844	-0.013
14	NUMCH2	-0.5605**	0.0916	-0.139
15	NUMCH3	-0.7915**	0.1096	-0.193
16	AGECH1	-1.4887**	0.1132	-0.332
17	AGECH2	-0.9598**	0.1025	-0.231
18	UNEMP	-0.1134**	0.0385	-0.028
19	KINDG • MAR2.(AGECH1 + AGECH2)	0.9444*	0.5170	0.210

log L = - 5204.18 n = 9276

[1] Calculated effect of one extra year at AGE = 45.145 (sample mean).
* Significantly different from 0 at the 10 per cent level against a two-sided alternative.
** Significantly different from 0 at the 1 per cent level against a two-sided alternative.

Figure 4.1. Histogram of estimated LFP probabilities. Female sample



By plugging the parameter estimates back into the logit model (3.24), it is possible to compute an estimated LFP probability for each and every respondent in the sample. The histogram of such probability estimates is shown in figure 4.1. The great bulk of the sample is seen to be scattered rather evenly between $\hat{p} = 0.2$ and $\hat{p} = 0.8$. In addition, there is a certain congestion of respondents in the area $0.04<\hat{p}<0.1$ and around $\hat{p} = 0.9$. The variation between different members of the sample is quite pronounced, the minimal and maximal values of $\hat{p}$ being 0.025 and 0.964, respectively.

Tests for interaction

Prior to the estimation of the model presented in table 4.1, certain tests for interaction were performed, cf. appendix 2. These tests were made using a full sample but a reduced set of independent variables of interest. (Certain computer technical constraints made it cumbersome to draw a random subsample of our data set.) From a methodological viewpoint this procedure is, of course, far from ideal, in that the parameter set of the final model is influenced by the same disturbance terms as its estimators. Strictly speaking the sampling distribution of such estimators is not known.

The first test concerns the possible interaction between marital status and the number and age of children. One can think of several a priori reasons that could make the impact of children larger for married than for unmarried women. A single parent is subject to a reduced tax rate ("class 2"), thus increasing her effective (after tax) market wage rate as compared to a married person. Secondly, single parents are usually given priority among applicants for kindergarten attendance, the market for which is strictly rationed. Thus fixed costs of market work may be smaller for single mothers. Thirdly, the amount of unearned income probably varies systematically with marital status, although one cannot conclude from a priori reasoning how a given difference should affect the impact of children on labour force participation. Besides, the magnitude of the effective differences in unearned income is open to debate. According to the income statistics, single mothers have considerably lower incomes than married couples with children, also per household member (NOS B 94 Income Statistics 1976: tables 50 and 101). However, the statistics do not tell us what fraction of married couples' income is not earned by the wife. Also the income statistics fail to cover certain tax-exempt contributions to unwed mothers payable under the National Insurance Scheme.[1]

At any rate, the hypothesis of zero interaction is not rejected (table A2.1 of appendix 2). More precisely, we have not been able to reject the hypothesis that the effect of children, <u>as measured by the log-odds ratios</u>, is equal between married and unmarried or previously married women. Now, this does <u>not</u> mean that whether or not the woman has

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[1] According to the scheme of family allowances, a single parent receives benefits for one child more than the real number of supported children under 16. Also, unwed mothers who are not cohabiting with the child's father are entitled to so-called maternity grants, assistance benefits, educational allowances, and transitional allowance. The <u>maternity grant</u> is a lump-sum allowance payable at the time of the delivery. It amounts to 37 per cent of the "base amount" of the National Insurance Scheme (Nkr 13 100 in 1977). <u>Assistance benefits</u> cover childcare expenses for women at work or during education (20 per cent of "base amount" annually, or higher when need is documented). <u>Educational allowances</u> cover schooling expenses. <u>The transitional allowance</u> is given to unwed mothers who are unable to support themselves owing to the care of children. The annual amount granted equals the "base amount" (National Insurance Institution 1978).

a husband, the expected rate of LFP will drop by the same number of percentage points. Since the general level of LFP is higher among unmarried women (i.e., their "initial" value is closer to 1), the expected drop in LFP due to children remains smaller than for married women. This non-linear feature of the logit model may be one reason why interaction terms turn out insignificant in our case.

The second test for interaction focuses on age and marital status. Here, too, the null hypothesis of zero interaction is not rejected, implying that in all marital groups, the same second-degree polynomial can be used to model the effect of age on LFP (table A2.2).

Thirdly, a couple of tests were carried out to unravel the simultaneous effects of education and marriage. The log-odds ratios for medium vs. low level education were found to differ significantly between the three marital groups. Consequently the corresponding interaction terms were included in the final model. Similar interaction terms involving high level education were, however, not found statistically significant (table A2.3).

In an attempt to curb the cost of estimation, the set of preliminary tests for interaction was, alas, limited to the above three. Later it became apparent that certain other interaction terms should also have been included in the model. Jensen (1981b:13) observes that the effect of children on female LFP is largest for women at the medium level of education. To examine this hypothesis a reduced logit model including interaction terms between education and children was estimated. These interaction terms were found significant (see table A2.4). The same applies to the interaction between the number of children and the age of the youngest child (table A2.5), and to the interaction between unemployment and level of education. The outcome of these tests, while not incorporated in the final model, is used to qualify the results of the latter (sections 4.4 and 4.5).

Part-time or full-time work

The choice between full-time and part-time work has been analyzed using a logit model with three alternative outcomes: (1) full-time work (at least 30 hours during the survey week), (2) part-time work (1-29 hours), or (3) no work (0 hours).

The third category includes not only persons outside the labour force, but also the unemployed and the temporarily absent from work.

Table 4.2. Estimation results from the trinomial logit model of female part-time and full-time work. Women 16-74 years not in school. 1977

	Independent variable	Full-time vs. no work		Part-time vs. no work		Full-time vs. part-time work	
		Parameter estimate	Standard error	Parameter estimate	Standard error	Parameter estimate	Standard error
g	x_{ig}	$\hat{\beta}_{1g}$	$\hat{\sigma}_{1g}$	$\hat{\beta}_{2g}$	$\hat{\sigma}_{2g}$	$\hat{\beta}_{1g}-\hat{\beta}_{2g}$	$\hat{\sigma}_{12g}$
1	CONSTANT ..	-1.4823**	0.2972	-3.7546**	0.3711	2.2723**	0.3908
2	AGE/100 ...	15.0830**	1.6886	15.2725**	1.9341	-0.1895	2.0937
3	$(AGE/100)^2$	-23.1299**	2.0096	-18.8723**	2.2258	-4.2576*	2.4524
4	RETAGE	-0.9542**	0.2153	-0.8978**	0.1964	-0.0564	0.2705
5	EDUC2	0.6173**	0.0635	0.2057**	0.0644	0.4116**	0.0738
6	EDUC3	0.7248**	0.0988	0.2974**	0.1081	0.4274**	0.1162
7	EDUC4	1.4135**	0.1160	0.9662**	0.1172	0.4473**	0.1180
8	EDUC5	1.8381**	0.1938	1.2481**	0.2012	0.5900**	0.1783
9	MAR2	-0.7805**	0.0959	0.2401*	0.1239	-1.0206**	0.1244
10	MAR3	-0.1531	0.1211	0.2049	0.1492	-0.3580*	0.1535
11	NUMCH1	-0.3770**	0.0945	0.3897**	0.0966	-0.7666**	0.1014
12	NUMCH2	-0.9958**	0.1068	0.2844**	0.1052	-1.2802**	0.1149
13	NUMCH3	-1.2857**	0.1362	0.1471	0.1250	-1.4328**	0.1468
14	AGECH1	-1.4261**	0.1219	-1.2046**	0.1210	-0.2215	0.1427
15	AGECH2	-0.9358**	0.1106	-0.6002**	0.1012	-0.3356**	0.1198

log L = -8163.3 n = 9276

Legend: See table 4.1.

Due to the costs of computation, the trinomial model had to be even more parsimonious in variables than the binomial version considered above. Interaction terms were neither included nor tested for.

Estimation results are reported in table 4.2. The columns $\hat{\beta}_{1g}$ and $\hat{\beta}_{2g}$ contain log-odds ratios for full-time vs. no work and part-time vs. no work, respectively. The log-odds ratios for full-time vs. part-time work follow as the differences $\hat{\beta}_{1g}-\hat{\beta}_{2g}$. Almost all parameters are significantly different from zero.

In wishing to estimate, respectively, the supply of full-time and part-time labour we face a rather difficult identification problem. Individuals are not free to choose the length of their working week. In the survey of part-time work conducted in 1978, 29 per cent of all full-time working women stated that they wanted a shorter working week. Among married women this percentage was 38 (see table 4.3). Out of those working part-time, 5 per cent wanted fewer and 13 per cent wanted more working hours. (In the part-time survey the cut-off point between part-time and full-time was set between 34 and 35 hours.)

Table 4.3. Employed persons wanting shorter or longer working week, by sex and/or marital status. Sample estimates. 1978

	Part-time workers[1]			Full-time workers[2]	
	Total	Want shorter week	Want longer week	Total	Want shorter week
	1 000 persons				
Men	59	(2)	11	1 080	220
Women	380	20	51	369	107
Married women	299	17	36	210	79
	Per cent				
Men	100	(3)	19	100	20
Women	100	5	13	100	29
Married women	100	6	12	100	38

[1] 1-34 hours. [2] 35 hours and over.
Source: Ellingsæter (1979:61).

In the same survey, persons working part-time were asked how they would choose between housework and full-time market work (assuming part-time jobs were not available). 46 per cent stated that they would choose housework, as against 23 per cent who would convert to full-time (Ellingsæter 1979:62). Excluding the "don't know" and "no answer" categories, the figures become 67 and 33 per cent, respectively.

In a special addition to the Labour Force Sample Survey of the 4th quarter 1976, persons outside the labour force were questioned about their relation to the job market. About 27 per cent of these, or a number corresponding to 243 000 persons, stated that they "have a need for or might wish to have a job". 10 out of these 27 per cent (i.e., 85 000 persons) reported "lack [of] suitable job opportunities" ask their principal reason for not looking for work (Foss 1980:126-127). When asked what kind of work would fit the respondent's idea of a suitable job, 31 per cent of those "needing" or "wishing" a job stated "part-time work" or "flexible working hours" as their first requirement. Among housewives this percentage was 48 (Kommunal- og arbeidsdepartementet 1980:87).

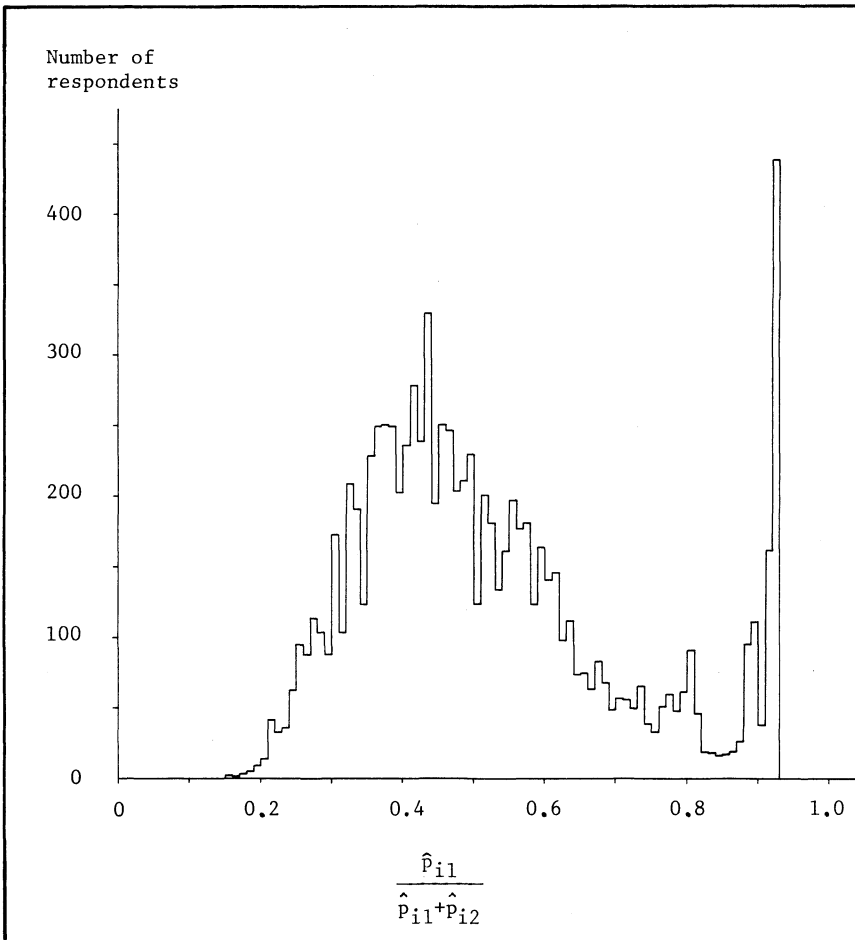
With due regard to the difficulties associated with interpreting attitudinal survey questions (cf. Foss 1980:102-107), one can hardly escape the conclusion that the observed number of hours worked fails to reflect the "true" supply of labour. This is so because the choice of working time is subject to strong institutional constraints (legislation, standardized working contracts, collective agreements etc.). It is likely that a large number of workers would choose a shorter or longer working week if the labour market were more flexible in this respect. In addition, such an increased flexibility would bring new persons into the labour force. The rate of part-time employment would probably increase, at the expense of both full-time workers and non-employed persons.

We have found no efficient way to circumvent this problem. In the binomial analysis of labour force participation, the unemployment variable represents an attempt, however insufficient, to control for variations in labour demand. For the analysis of part-time vs. full-time employment a similarly relevant measure of aggregate disequilibrium can hardly be found.

Now, given that we focus our interest on <u>differentials</u> rather than on the <u>general level</u> of labour supply, the identification problem may not be too serious. It may be argued that a general shortage of part-time jobs should not affect women with a particular age, education, marital status, or family size any more than other women. Put more precisely, it is conceivable that out of the estimates reported in table 4.2, only the constant terms are biased. If that is the case, all log-odds ratios of interest to us are unaffected.

Conditional probabilities of full-time work, given some work, are calculable in terms of the contrasts $\hat{\beta}_{1g} - \hat{\beta}_{2g}$ only (formula (3.34)). The histogram of such conditional probabilities for the entire sample (predicted) is shown in figure 4.2. The sample mean is 0.521. The histogram is, however, bimodal and skew. A majority of the respondents have conditional full-time probabilities that are smaller than the sample mean. On the other hand the highest peak of the histogram occurs around 0.90.

Figure 4.2. Histogram of estimated conditional full-time probabilities. Female sample



When seen together, figures 4.1 and 4.2 might be taken to indicate that the female supply of labour consists, roughly speaking, of three rather different categories. There is one relatively small group of women who participate in the labour force to a very limited extent (3-15 per cent). Then there is a large and very heterogenous groups with LFP probabilities ranging from 20 to 80 per cent. Finally, there is a relatively small group of highly job-oriented women with LFP rates around 90 per cent. Apparently, these women also work full-time to an overwhelming extent. Or aren't these the same women? In general, who are the different groups, and why does their supply of labour differ so drastically? These are some of the questions addressed in section 4.4-4.8 below.

Hours of work

A linear regression analysis was carried out to study variations in the length of the working week among employed women at work. Results are reported in table 4.4.

The model contains the same personal attribute variables as were included in the binomial logit model. More interaction terms are, however, included.

Each parameter α_g is interpretable as the partial effect of variable g on the conditionally expected length of the working week, given at least 1 hour's work. Only respondents fulfilling this condition were used in the regression sample.

The fit is poor (R^2=0.125). Only a small fraction of the variation in hours worked is explained by the variables used. Most interaction terms are statistically insignificant by the t-test. Most main effects are, however, significant, at least at the 10 per cent level.

Because of the market imperfections discussed above, the parameters are not readily interpretable as differentials in labour supply. The reservations made in relation to the trinomial logit model results apply almost verbatim, cf. also the remarks on pages 87-88 below.

Table 4.4. Estimation results from the linear regression model of hours worked by employed women at work. Dependent variable: T

g	Independent variable x_{ig}	Parameter estimate $\hat{\alpha}_g$	Standard error $\hat{\sigma}_g$
1	CONSTANT	32.224**	2.118
2	AGE	0.304**	0.118
3	$(AGE)^2/100$	-0.534**	0.141
4	RETAGE	-1.057	1.647
5	EDUC2	1.954*	0.822
6	EDUC3	1.807	1.068
7	EDUC4	3.018*	1.328
8	EDUC5	2.452	1.996
9	MAR2	-4.829**	0.915
10	MAR3	-1.497	0.879
11	NUMCH1	-2.760*	1.308
12	NUMCH2	-7.478**	1.923
13	NUMCH3	-6.000*	2.991
14	AGECH2•NUMCH1	-0.243	1.206
15	AGECH2•NUMCH2	-2.044*	1.099
16	AGECH2•NUMCH3	-3.901**	1.487
17	AGECH1•NUMCH1	-0.176	1.265
18	AGECH1•NUMCH2	-1.109	1.353
19	AGECH1•NUMCH3	-4.357*	1.902
20	MAR2•EDUC2	0.131	0.968
21	MAR2•EDUC3	0.295	1.351
22	MAR2•EDUC4	-0.211	1.554
23	MAR2•EDUC5	0.877	2.312
24	MAR2•NUMCH1	-2.067	1.359
25	MAR2•NUMCH2	0.646	1.961
26	MAR2•NUMCH3	-0.378	3.019

$R^2 = 0.1251$ $\hat{\sigma} = 12.378$

Legend: See table 4.1.

4.3. Labour market tightness

The unemployment parameter is significantly smaller than zero at the 1 per cent level (table 4.1). Thus, when unemployment expands, the labour force contracts. The discouraged worker effect seems to dominate the added worker effect.

Assume that the rate of unemployment adequately describes the lack of local job opportunities, so that when no one is unemployed there will be no added or discouraged workers, either. Under this assumption, a fairly natural measure of labour supply would be the hypothetical size of the labour force under zero unemployment. This is the approach adopted here, although the weaknesses are not hard to spot.

First of all, the basic assumption is an oversimplification. The absence of suitable job opportunities for a person with given characteristics does not necessarily bear any relationship to the aggregate rate of unemployment.[1] Labour is heterogeneous. The mismatch between supply and demand as regards the length of the working week is just one example. Similar friction exists when it comes to education, experience, geographic location, or perhaps even sex.

The measure of unemployment used is subject to rather arbitrary variations. Since LFSS unemployment rates are inestimable at the regional level (too small sample), our measure is based on registered unemployment. The size of the regional labour force is not known, either, so that in calculating the jobless rate we have had to use a standardized figure in the denominator. This standardized labour force is calculated by applying nationwide LFP rates, broken down by age and sex, to the resident population in each region. Registered unemployment, LFP rates, and population figures are taken from the official labour market and population statistics for 1977[2]. Each region is defined so as to coincide with one of the primary projection regions (PPR) that are used by the Central Bureau of Statistics in their population projections (Sørensen 1975). There are 96 such PPRs. On the average, each PPR comprises about 4 ½ municipalities. The PPRs have been constructed so as to minimize commuting between them. Thus, each PPR is probably not too far from constituting a regional labour market.

[1] Indeed, the interaction testing results reported in table A2.7 bear testimony to this. These results show that unemployment does not reduce the LFP of women with a university level education. Since the model does not take account of this there is an upward bias in the unemployment parameter estimate, i.e. as far as women at the low and medium levels of education are concerned the effect of unemployment is underestimated in absolute terms. See also section 4.4 for a discussion of this. [2] Registered unemployment, being available on a monthly basis, is measured at the time of the LFSS interview. For nationwide LFP rates and population statistics we use, however, annual data.

Even so, it is likely that some PPRs cover too large an area to function as one market. On the other hand some are too small or too close. Whatever regional labour market subdivision were used, commuting between "markets" could never be eliminated.

The weaknesses inherent in the statistics on registered unemployment were discussed in sections 2.2 and 2.3. Usually this figure remains lower than the LFSS rate of unemployed, which corresponds more closely to ILO recommendations. As seen from the econometric viewpoint, perhaps the most important drawback has to do with the fact that the registered number of unemployed is a main political target variable. In areas experiencing high jobless rates, the government will often prescribe certain extraordinary employment measures so as to bring the figure down. This practice, while socially and politically understandable, tends to undermine the autonomy of our relation. In some cases the jobless rate is kept artificially low, compared to the "real" slackness or excess supply in the local labour market.

We use one aggregate local rate of unemployment for both men and women. The idea is that both groups compete on the same labour market, or at least that sex per se does not matter to an employer. This assumption may be a bit far-fetched. Anyway, the propensity to register as unemployed is much lower among women.[1] One reason for this is probably that more women, having recently entered the labour force, are not entitled to unemployment allowance, cf. section 2.2.

Last, but not least, there is an omitted variable bias.[2] Although a number of personal attribute variables are included in the model, one can never completely control for differences in attitudes and tastes. Persons resident in relatively less developed regions (with high unemployment rates) may, e.g., be thought of as generally less oriented towards the labour market than those living in urban or highly industrialized districts - some of which have gone to town precisely to seek work. Also, it is possible that wages are higher in areas with little unemployment. This means that the (net) discouraged worker effect is probably overestimated (in absolute value) by our model.

---

[1] In 1977, the ratio of registered to LFSS unemployment was about 0.9 for men and 0.4 for women (NOS A 958 Labour Market Statistics 1977, tables 3 and 61). [2] Maximum likelihood estimates are generally not unbiased. They are, however, consistent and asymptotically unbiased provided the model is correctly specified. By "bias" we always mean asymptotic bias.

When setting UNEMP = 0 for all respondents in the sample and recalculating the LFP probabilities, the overall female sample mean of LFP increases from 0.5455 to 0.5640. By our convention, then, the latter figure represents the rate of female labour supply in 1977 measured in persons (not in school).

The sample mean of UNEMP is 0.8544. According to our model, reducing this figure to zero is consistent with a 3.4 per cent increase in the female labour force.

Figure 4.3 describes the estimated relatonship between female LFP and unemployment (the dashed line)[1]. In addition a (solid) line representing female employment has been drawn, calculated as the LFP rate multiplied by 1 minus the rate of unemployment[2]. The distance between the two curves represents observed female unemployment. The gap between LFP and the dotted line on top represents, in principle, hidden female unemployment, defined as the difference between the actual labour force and the predicted labour force at zero unemployment.

Unless the discouraged worker effect is grossly overestimated by our model, the rate of hidden unemployment appears to be several times higher than the observed (LFSS or registered) rate of unemployment. This is true for all non-zero levels of unemployment. Previous studies based on attitudinal survey data have come up with similar results (Foss 1980; cf. also Fridstrøm 1981: 84-90).

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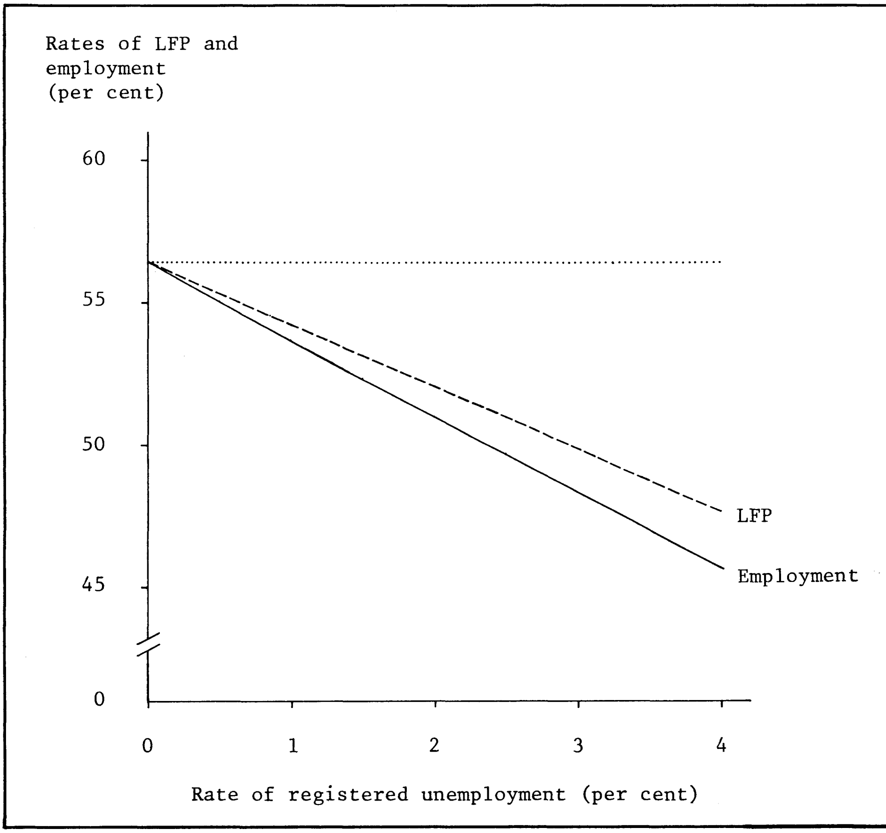
[1] The curve has been computed as

$$f(\text{UNEMP}) = \frac{1}{n} \sum_{i=1}^{n} \left[1 + \exp\left(- \sum_{g \neq 18} x_{ig}\hat{\beta}_g - \text{UNEMP} \cdot \hat{\beta}_{18}\right)\right]^{-1}$$

i.e. recalculating the LFP probability of every respondent under varying rates of unemployment, and taking the mean of all these predicted LFP probabilities (cf. table 4.1). A uniform rate of unemployment in all regions is assumed.

[2] Assuming that the male and female rates of unemployment coincide, and disregarding the difference between LFSS and ESO unemployment.

Figure 4.3. The estimated relationship between female LFP and regional unemployment



4.4. Education and marriage

Effect on labour force participation

It is well known that female labour force participation increases with the level of completed education (Ljones 1979, Vannebo 1977). This is so even when certain other labour supply determinants (age, marital status, presence of children) are controlled for, cf. table 4.1.

A common economic explanation for this is that high education means higher productivity and hence a higher obtainable wage rate. However, other job factors than the wage rate can also be expected to vary with the amount of schooling, in the sense that the kind of jobs offered to the better educated is generally seen as more interesting, attractive, or rewarding. A third, and perhaps even more important explanation, lies in the hypothesis that persons with different types of education have fundamentally different preferences. In micro-economic jargon: their indifference maps differ. It is conceivable that (i) such differences might arise as a <u>result</u> of the education process. Perhaps more interesting, however, is (ii) the selection mechanism inducing basically more career-minded people to seek higher formal education, with the very purpose of obtaining a (better) job. In this case, the education variable is not completely exogenous. Although the act of going to school precedes the act of taking a job, the <u>decision</u> to become a student may very well be conditioned by the individual's plans for a future career. Or, at the very least, the two variables must be seen as simultaneous and determined jointly by some third, "attitudinal" or other background variable.

In the context of prediction, the distinction between these two interpretations - (i) or (ii) - becomes important. In the first case, measures to educate more people will (in the long run) add to the labour force. In the second case, however, such measures are not likely to have much impact.

With these qualifications in mind, let us have a closer look at the parameter estimates in table 4.1. Six dummy variables are used to describe the respondents' level of education, two of which are, however, terms of interaction between eduation and marital status. The reference category is "9 or fewer years of schooling" (EDUC1 = 1, cf. table 3.1), so that all six parameters express log-odds ratios with respect to this lowest level of education[1].

Preliminary tests for interaction suggested that the effect of education was not the same for all marital categories, or vice versa (see appendix 2). Married women with 10-12 years of schooling were found to have comparatively low participation rates. The same was true, although to a lesser extent, of previously married women at the same level of education.

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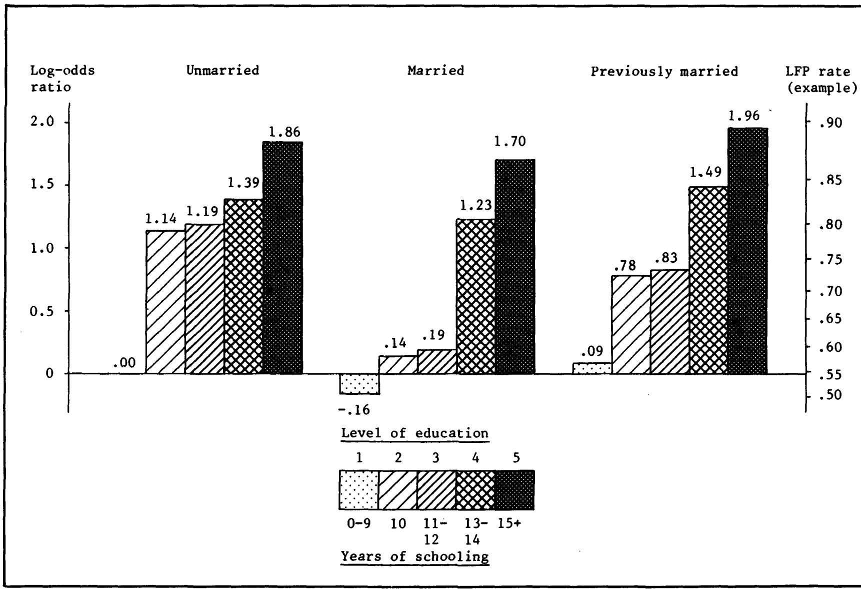
[1] The choice of reference category is arbitrary. Contrasts between the remaining four educational categories are, in fact, just as interesting. The complete set of estimated log-odds ratios between different levels of education is shown in table A3.1 of appendix 3. Standard errors are given for all contrasts.

The effect of education and marriage on women's labour force participation is illustrated by figure 4.4. On the left vertical axis, we measure log-odds ratios with respect to a reference group - consisting of unmarried women with level 1 education (0-9 years). The right vertical axis measures the corresponding LFP rates, assuming that the reference group has a participation rate equal to the overall sample mean - (p = 0.5455)[1].

Figure 4.4. The effect of education and marriage on female LFP



[1] This last qualification is essential. The right-hand scale of figure 4.4 would have been different had we chosen another reference point (e.g. p = 0.8596 as in figure 5.10).

There are statistically significant differences between most levels of education, except between levels 2 and 3 (see table A3.1). Also, for unmarried women, the differences between levels 2, 3, and 4 are insignificant.

Marital status is, by and large, statistically significant only for women with 10-12 years of schooling (table A3.2). Among these, married women have substantially lower LFP than the previously married, who in turn work less than the unmarried. Note that, according to the procedures adopted in the 1977 labour force sample surveys, women cohabiting without marriage are, in principle, counted as married. Also, recall that the presence of children is controlled for in our model.

Among unmarried women, the large "step" in LFP is seen to occur from the lowest (1) to the medium (2-3) level of education. For married women, the big difference is between the medium and the university (4-5) level. Previously married women exhibit a more regular pattern, with approximately the same difference between the low and the medium level as between the medium and the university level.

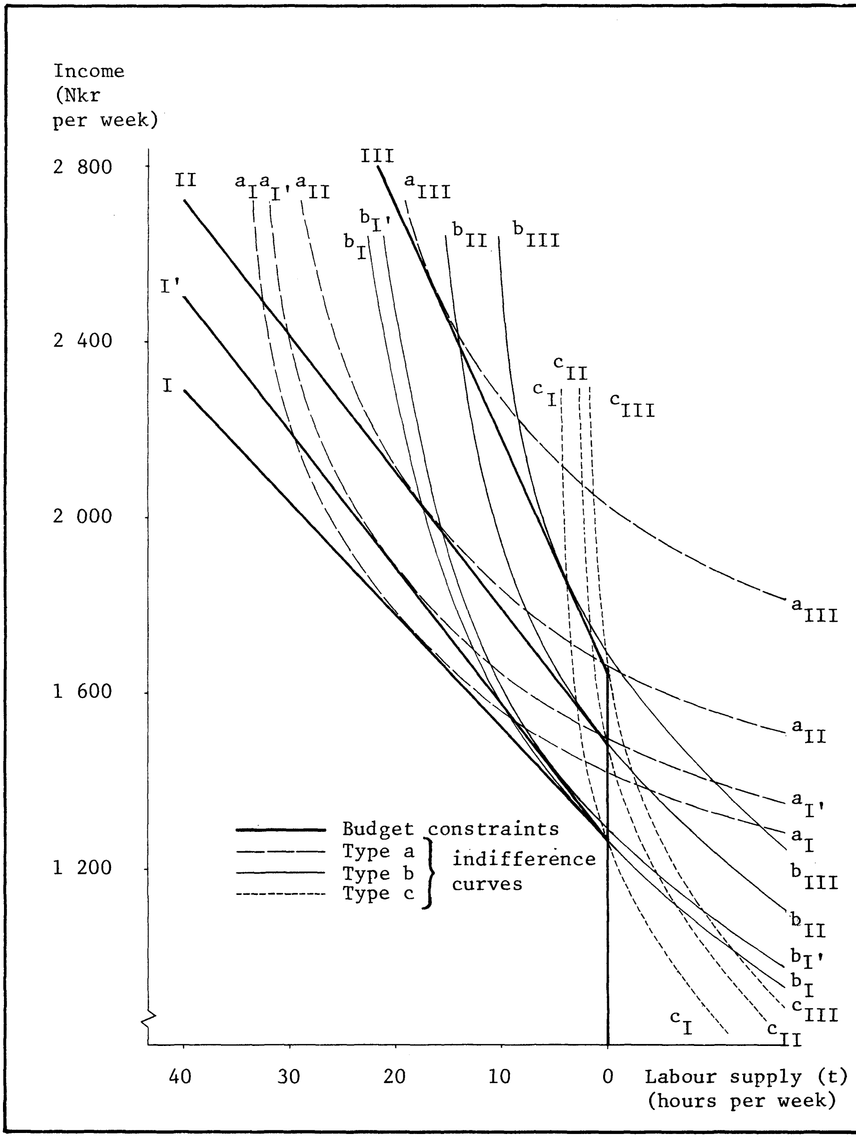
Why is it that the picture is so different between the three marital groups? The same pattern was found and discussed by Ljones (1979: 68), based on census data for 1970. Thus it seems that we are faced with a fairly stable structural phenomenon.

Based on micro-economic theory one would expect married women to work less than unmarried or previously married. The addition of a husband's income to the total family budget acts, as far as the wife's behaviour is concerned, as a pure income effect, reducing the supply of labour unless leisure is an "inferior good". Our findings for the medium level education group are in line with this hypothesis.

More surprising is the finding that this mechanism does not seem to play any large role among those with the lowest level of education, nor among the very best educated. We are also puzzled at the observation that among married women the difference between levels of education 1 and 2 is, relatively speaking, very small.

The explanation is probably to be sought in a combination of wage differentials, "attitudinal" factors, and social selection mechanisms. It has been argued that the unusually low participation rate among unmarried women with little education might be due to an overrepresentation, within this group, of persons chronically ill or disabled (Ljones 1979: 62ff.).

Figure 4.5. Married women's supply of labour



As for the comparatively low LFP among married women at the medium level of education, there might be a trade-off between income and substitution effects at work. Compared to the lowest educated, women at the medium level have a higher obtainable own wage, yielding presumably a positive substitution effect and a negative income effect. Now, in all probability their husbands also have on the average a higher income, thus strengthening the negative income effect on the wife's labour supply. Ljones (1979: 136ff.) makes the interesting observation that when the wife's own education is controlled for (together with age, children etc.), female LFP decreases, by and large, with the husband's level of education. Thus, Ljones includes the educational characteristics of both spouses in the model. In his model the wife's LFP is found to vary more regularly with her level of education than in our case. Theeuwes (1981) finds a negative relationship between the husband's wage rate and the wife's participation.

It is instructive in this respect to consider a diagram adapted from Ben-Porath (1973). Figure 4.5 shows the possible indifference maps of three different (types of) individuals, say a, b, and c.[1] Assume, for the sake of the argument, (i) that every individual falls into one of these three categories (i.e. all individuals of a given "type" have identical preferences), and that (ii) the proportion falling into each category is constant across educational levels. Denote these proportions by p_a, p_b, and p_c, respectively.

Assume that line I defines the budget constraint for women with a given (low) level of education. The slope corresponds to the women's market wage rate, while the intercept with the vertical line t=0 equals the husband's wage (plus other non-labour income).

Among women with this wage structure the LFP rate will equal p_a.[2] Individuals of type a supply a certain amount of labour. Individuals of type b and c represent corner solutions and remain outside the labour force.

Ben-Porath considers an increase in the women's wage rate as represented by the rotated budget line I'. Note that at this wage rate group b is drawn into the labour force, jumping from the indifference curve b_I-b_I to the curve b_I'-b_I'. Ben-Porath makes the pertinent observation that

[1] For expository reasons the diagram assumes no fixed costs of labour market entry. [2] We assume all the way through this argument that there is no discouraged worker effect.

"As the wage rate increases, LFP ... can only increase, because those who are not in the labor force (such as b and c ...) can be drawn into it (and there is some wage rate at which everybody will be), and nobody who is in the labor force will leave. For those who are out of the labor force there is no income effect in a wage rise, and only substitution works, and those who are in may reduce their hours because of an income effect, but never to zero. The elasticity of LFP ... with respect to [the wage rate] must be positive. Its size depends on the distribution of [individuals between] the groups [a, b, and c, i.e.] according to "tastes"." (Ben-Porath 1973: 702).

In our graphical example, the LFP rate increases from p_a to p_a+p_b, while the supply of labour as measured in hours may be only marginally affected.

Next consider an extension of Ben-Porath's argument to cover the case where, simultaneously with the wage increase, non-labour income also grows. Let, e.g., the line II represent the budget constraint for women at a level of education higher than I. Line II runs parallel to I', i.e. the female wage rate is as assumed in case I'. Yet given the higher level of non-labour income, group b is seen to remain outside the labour force (curve b_{II}-b_{II}). Thus, between levels I and II there is no difference in total participation.

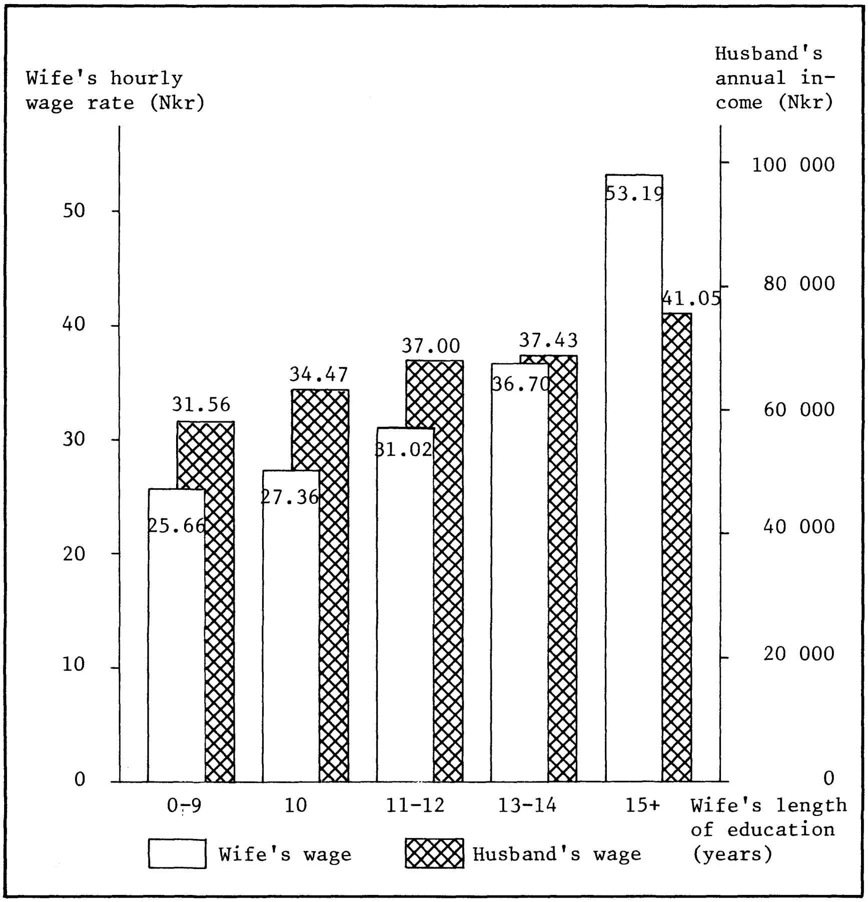
Finally, consider an even higher level of education, characterized by the budget constraint III. Here, individuals of type b eventually become economically active, and LFP rises to p_a+p_b. The positive substitution effect more than offsets the negative income effect of the concomitant increase in non-labour revenue.

To examine this hypothesis empirically, a few linear regressions were run using data from the Norwegian fertility survey of 1977. Our sample consists of married women between 18 and 44 years old. The full results are reported in tables A3.3 and A3.4 of appendix 3. For the sake of comparability with the results reported in table 4.1 and figure 4.4, the same set of personal attribute variables was used as in our binomial logit analysis.

Our interest is focused on the partial relationship between the woman's level of education and the wage rates of both spouses. Our findings in this matter are summarized in figure 4.6. The white columns represent married women's average hourly wage rate[1], as a function of the

[1] Two regressions were run, one for respondents in the labour force (a) and one for those not in the labour force (b). The latter group, having no wage rate, stated their expected obtainable wage. In figure 4.5 the two regressions have been weighted together according to the size of either sample. The wage rates are those predicted by the regressions for a 30-34 years old married woman without children.

Figure 4.6. Estimated wage rates of husband and wife, by wife's level of education. 1977



women's level of education. The shaded columns picture the husband's annual income[1],[2].

It is evident from the diagram that the wife's level of education is correlated not only with her own wage rate, but also with that of her husband. The latter relationship must be due mainly to a selection process. Møglestue (1975:49ff.) has shown that there is a strong tendency for husband and wife to have similar levels of education.

Not surprisingly, the husband's wage rate varies not nearly as much as that of the wife when the couples are grouped according to her level of education. However, between levels 1, 2, and 3 the increase in the husband's income is seen to be about the same size as the increase in the wife's own wage rate. Thus, in these cases the total income effect due to the husband may very well offset the substitution effect of the wife's wage rise. Between levels 3 and 4, however, the income effect due to the husband is negligible, while the wife's wage rate increases considerably. From level 4 to level 5 the husband's income is seen to rise about 10 per cent, but the wife's wage rate is increased by a full 45 per cent, so that here, too, the substitution effect is more likely to prevail.

Indeed, at this point we may reveal to the reader that the budget constraints of figure 4.5 have been drawn in a way corresponding exactly to the wage rates shown in figure 4.6. The lines I, II, and III represent the budget constraints of an average married woman with level of education 1, 3, and 5, respectively. Weekly non-labour income has been set equal to 40 times the husband's wage rate.

In reality non-labour income consists of more components than the husband's wage. In a recent study by Aamodt (1982) it is documented that the choice of education in Norway is highly dependent on social background, notably on parental income (op.cit., tables 7.9-7.10). Thus it may be assumed that women with higher education are also generally wealthier. The same applies to their husbands.

Yet the husband's wage probably remains the most important item

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[1] In comparing annual income data to hourly wage rates, a conversion factor of 1 840 hours per year was used (46 working weeks, 40 hours per week). [2] Even for the husband's income, two separate regressions were run (table A3.4). One includes the number and age of children, the other does not. The first regression ensures comparability with the results obtained by the logit model of female LFP in that the set of attribute variables is practically identical. This regression is the basis for drawing figure 4.5. The second regression was run in order to test for a possible simultaneity bias in the education parameters (fatherhood is probably endogenous with respect to education and income). However, the contrasts between the education parameters do not change drastically - all parameters are reduced by some 12-13 per cent.

of female non-labour income. The data on wife's and husband's wage rates seem, altogether, to provide a fairly plausible economic explanation of the variation in LFP among married women with varying educational background.

Now, the assumption (ii) that groups of individuals with unequal levels of education be identically distributed as between different preference structures is obviously far-fetched. Realistically, there are few reasons to believe that women in different educational groups respond to economic incentives in the same way. As already mentioned education is not completely exogenous in relation to labour supply.

All university level fields of study in Norway can be seen as highly specialized and clearly vocational training programs. Women undertaking such studies are generally job-oriented. Their predisposition is to work in the market rather than in the household. This predisposition is strong enough to carry through whether or not the woman has a (wealthy) husband, yielding high LFP rates in all marital groups.

Also, it is generally assumed that academic studies tend to enhance certain "liberal" values, encouraging e.g. women to take an equal part with men in market work. Thus, any "predisposition" present is only apt to get strengthened through the course of a university program.

A large part of the medium level education is, on the other hand, non-vocational in character (viz. the "examen artium"). Among married women with this kind of background, there seems to be a certain percentage who prefer to take on the more traditional roles of women (children, housework). They can afford this because they have husbands with a fairly good income.

Referring again to figure 4.5, this amounts to saying that in the higher educational categories (III) the proportion of type a individuals will be large, while at the low level (I), type c will be comparatively frequent. Part of the estimated LFP differentials is due to such heterogeneity between educational categories.

Similar objections, although perhaps to a lesser extent, apply to the use of marital status as an independent variable. All three variables - labour supply, education and marital status - are choice variables that could be more or less simultaneously determined. Behind this simultaneous process lurk such elusive factors as social and genetic background, attitudes, and "tastes".

Yet another source of bias arises from the fact that we are estimating a reduced form equation. The parameters of such an equation may reflect both supply and demand variations. In principle, variations in labour demand is "controlled for" through the unemployment variable. But obviously this crude measure does not adequately describe the different demand situations faced by individuals with unequal levels of education. One might suspect that a slack labour market affects individuals with a weak educational background most. To test this hypothesis a reduced logit model with interaction between unemployment and level of education was estimated (table A2.7). The results are appalling. Between the low and the medium levels of education the interaction term is statistically insignificant. At both levels the effect of unemployment on female LFP is, as expected, significantly negative. For women with a university level education, however, the "effect" of unemployment appears to be highly positive!

This finding is a bit hard to explain. Surely there must be some kind of spurious correlation at work. Registered unemployment is, e.g., negatively correlated with the degree of urbanization[1]. Using data from the Norwegian population census of 1970, Birkeland (1977: 33) found an interesting pattern of cross-variation between female LFP, level of education, and geographic region. She showed that among women at the low and medium levels of education, LFP was, by and large, highest in the Oslo area and lower in most other districts. At the university level of education the regional pattern of variation was, however, rather the opposite. Here, the rate in Oslo was no higher than the national average (82 per cent for married women at level 5), while the highest LFP rates were found in the three northernmost counties (89-90 per cent). As of 1977 the registered rate of unemployment in these counties was more than twice the national average.

About the reasons behind this regional pattern of variation one can only speculate. Probably it must have something to do with the fact, already mentioned in section 4.3, that the population mix of a given region is not exogenous; it is influenced by the characteristics of the regional labour market. Highly educated women do not migrate to rural or less developed regions unless they are able to get a job there. Yet, why their rate of LFP should be <u>higher</u> in the rural districts than in the

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[1] From the male sample, the following regression equation was estimated (standard errors in parentheses):

$$UNEMP_i = \underset{(0.016)}{1.225^{**}} - \underset{(0.022)}{0.526^{**}} \cdot URB_i + u_i$$

n=8922 R=-0.247

Here, URB_i denotes the proportion of inhabitants in respondent i's municipality who live in an urban area. R is the sample correlation coefficient.

cities we are unable to explain. It may be of some relevance to note that in the male sample, the interaction between unemployment and education was found totally insignificant (results not reproduced here). Thus we are faced with an exclusively female phenomenon.

What are the consequences of this in terms of (asymptotic) bias? The interaction test does not give reason to believe that the parameter estimates for EDUC2 or EDUC3 be biased. However, the estimates concerning EDUC4 and EDUC5 are. The interaction term between either one of these two variables and UNEMP may be regarded as an omitted variable positively correlated with EDUC4, resp. EDUC5. The result of this is an upward bias in the estimation of both parameters[1]. It is unlikely that this bias be very large, though[2].

To sum up, the parameter estimates concerning level of education reflect a number of different effects: (i) Women seeking higher education are generally more career-minded. (ii) As a <u>result</u> of formal training they also <u>become</u> more job-oriented. (iii) The market for skilled or highly qualified labour is comparatively tight. (iv) The market wage increases with the individual's level of education. (v) Women with higher education also tend to have husbands with generally higher education and income. (vi) Women with higher education are generally wealthier.

Items (i) through (iv) tend to increase the observed LFP differentials between women with unequal educational background. Items (v) and (vi) work in the opposite direction.

Which of them belong in a supply equation? Ultimately this is a matter of discretionary choice, cf. the discussion in section 2.4. In our opinion, factor (iii) is the only one which clearly does not belong in a supply function. However, the bias caused by this factor was found

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[1] Results concerning the bias caused by omitted variables in the multinominal logit model are derived in a recent article by Lee (1982). Strictly speaking these results have only been shown to hold when the omitted variable is either dichotomous or normally distributed. We feel confident, however, that similar results must hold even in our case if only by analogy to the linear regression case. [2] Assuming that the correct model is one in which unemployment does not affect the LFP of women with at least 13 years of schooling, the biases of $\hat{\beta}_7$ and $\hat{\beta}_8$ should be approximately equal to the unemployment parameter times the sample mean of unemployment, i.e. to $0.1134 \times 0.8544 = 0.0969$ (cf. table 4.1).

to be small. As for the other five items, whether or not we want to model them as supply determinants will depend on our time perspective as well as on precisely what economic experiment we have in mind. Suppose, for instance, that we want to measure the effect of educating more people. Then definitely factors (ii) and (iv) are relevant. But unless the added schooling capacity be reserved for women, one can hardly avoid the effect (v). In the very long run even (i) and (vi) will have some bearing. Furthermore, one must consider that an increased schooling incidence is likely to affect such variables as marriage and child-bearing. Thus, in our model there is an additional effect channelled through the dummies for marital status and children.

Effect on hours of work

Estimates from the linear regression model of hours worked were given in table 4.4. We note that the differences between the five levels of education are not very large, and barely statistically significant. Married women work, however, significantly shorter weeks than other women.

The model includes first-order interaction terms between marriage and education, none of which are, however, significant.

The effect of education and marriage on the average length of the working week is illustrated by figure 4.7. Four representative examples have been worked out. Each example is described by five columns, one for every level of education. Each column conveys three pieces of information. The total length of the column equals the mean number of hours worked by women at work. The white part of the column denotes the average reckoned over all women (not in school). The difference between these two figures results from the incidence of non-participation, the effect of which is represented by the shaded area.

The lengths of the white columns have been calculated by multiplying the hours estimates for women at work by their corresponding LFP rates (at zero unemployment). Thus, the white columns have an interpretation as the average supply of labour, measured in hours, within each population subgroup[1].

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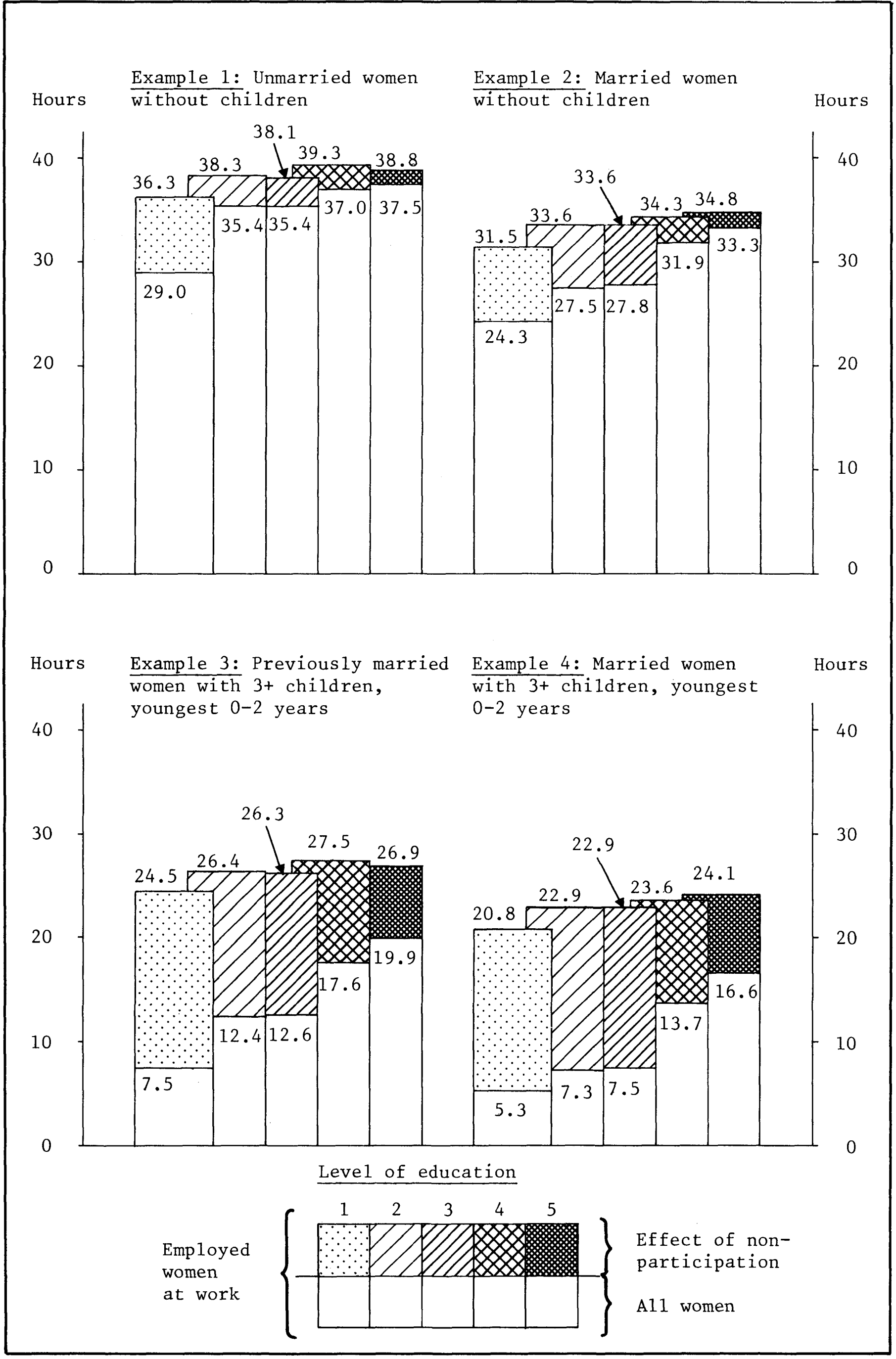
[1] Note that the supply of working hours, as defined here, includes a potential supply from all members of the labour force, not only from "employed persons at work". The unemployed and the temporarily absent are counted, too, by the same logic as adopted later when calculating full-time and part-time rates (see page 89). One might, of course, argue that workers absent from their job do not supply labour, but this ultimately remains a matter of semantics. Anyway, our main interest is focused on the differentials in labour supply between various groups, rather than on its level.

Figure 4.7. Estimated weekly hours of work, by level of education. Mean values calculated for 35-year-old women in different socio-demographic groups



Married women in their mid-30s and with no children typically[1] have a working week of 32 to 35 hours. Their unmarried sisters put in about 4 hours more, on the average. As for the previously married, they seem to work on the average 3 or 4 hours more per week than the married.

The variation across levels of education seems to follow very much the same pattern in all marital groups and regardless of children in the family. Only women with the lowest level of education can be singled out as working shorter weeks than the others.

These figures apply when looking at the working population only, i.e. when conditioning on labour supply, the dependent variable of interest. To avoid a biased impression one should examine mean values calculated not only over those actually working, but over all members of the relevant (sub)population. Because LFP rates vary greatly between socio-demographic groups, these figures show much larger differences. Whereas childless, unmarried women at the highest level of education supply an average of 37 hours of work per week, those with the biggest families (and small children) supply less than 6 hours to the market. The pattern disclosed in figure 4.4 is recognizable also in figure 4.7, in that for married women, the big difference in labour supply goes between medium and high level education, while for unmarried women the only large gap is found between the low and the medium level.

It is interesting to note that most of the variation in labour supply across levels of education is due to differences in LFP rates. The length of the working week is very similar for all levels of education, while LFP rates vary greatly. This is precisely what we would expect on the basis of figure 4.5 above. In the determination of hours worked there is both a substitution and an income effect of wage, while LFP depends only on the substitution effect. In addition there is a non-labour income effect on both hours and LFP.

If we examine the variation across marital groups, the emerging pattern is a bit different. Here, LFP rates do not vary as much (except at the medium level of education), while the difference in the working week may amount to 4 or 5 hours on the average.

It is conceivable that this pattern of variation be influenced not only by "pure" supply factors, but also by institutional inflexibilities in the choice of working hours, or even by demand variations. This

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[1] One might ask how "typical" this is, given the poor explanatory power of the model. The figures stated are expected values for the groups in question. The spread around each value is, however, large.

would be the case if institutional or demand factors were correlated with the variables of the model, so that women with different attributes were affected to different degrees by the omitted variables. For instance, it seems reasonable to assume that the strength of women's bargaining position vis-à-vis their employer may be positively related to their level of education. University educated women stand a better chance to work the desired number of hours per week. If this is the case, the bias due to institutional rigidites (as measured in absolute value) will be negatively related to the length of education. If part-time work is what many women desire, the estimated number of hours supplied by women at work may be too high, especially for women at the low and medium levels of education.

On the other hand, LFP rates are also affected, and in the opposite direction. The fact that many women are unable to obtain a job with the desired number of working hours may keep some of them from entering the labour force. The sign of the overall bias is therefore hard to tell.

As for marital status, it seems unlikely that there be any sizable bias present. The marital status of a worker should be irrelevant to the employer as well as in relation to any institutional constraint. It is possible, though, that a more flexible labour market would enhance further the "spread" between women in different marital groups, simply kecause a larger fraction of married women are presently overemployed (see table 4.3). Again, there are two effects at work, running in opposite directions. Increased flexibility is also liable to lure more married women into the labour force.

Effect on part-time versus full-time work

Women at the lowest level of education (0-9 years) have a significantly stronger "preference" for part-time work (as opposed to full-time) than women in all other educational categories. The differences between the remaining four categories are, however, statistically insignificant (see table A3.5 of appendix 3).

The estimated log-odds ratio between levels 2 (10 years) and 1 is 0.4116. At the sample mean (52 per cent conditional full-time probability), this translates into a 10 percentage points arithmetic difference. That is, if 52 per cent of all women with level 1 education "prefer" a full-time to a part-time job, at level 2 the figure would be about 62 per cent, when comparing women with the same age and family situation (marriage, children). At level 5 we would have an estimated 66 per cent "preferring" full-time to part-time employment, although the difference between levels 2 and 5 could be due to sampling error.

The difference in the length of the working week between employed women with, respectively, 0-9 and 10 years of schooling was found to be about 2 hours (figure 4.7). This two-hour difference comes about because the part-time rate is 10 per cent higher among (employed) women with the lowest level of education.

As for marital status, all contrasts are significant at the 5 per cent level. That is, married women have a stronger preference for part-time work than do previously married women, who in turn prefer part-time to a larger extent than unmarried women do (table A3.6 of appendix 3).

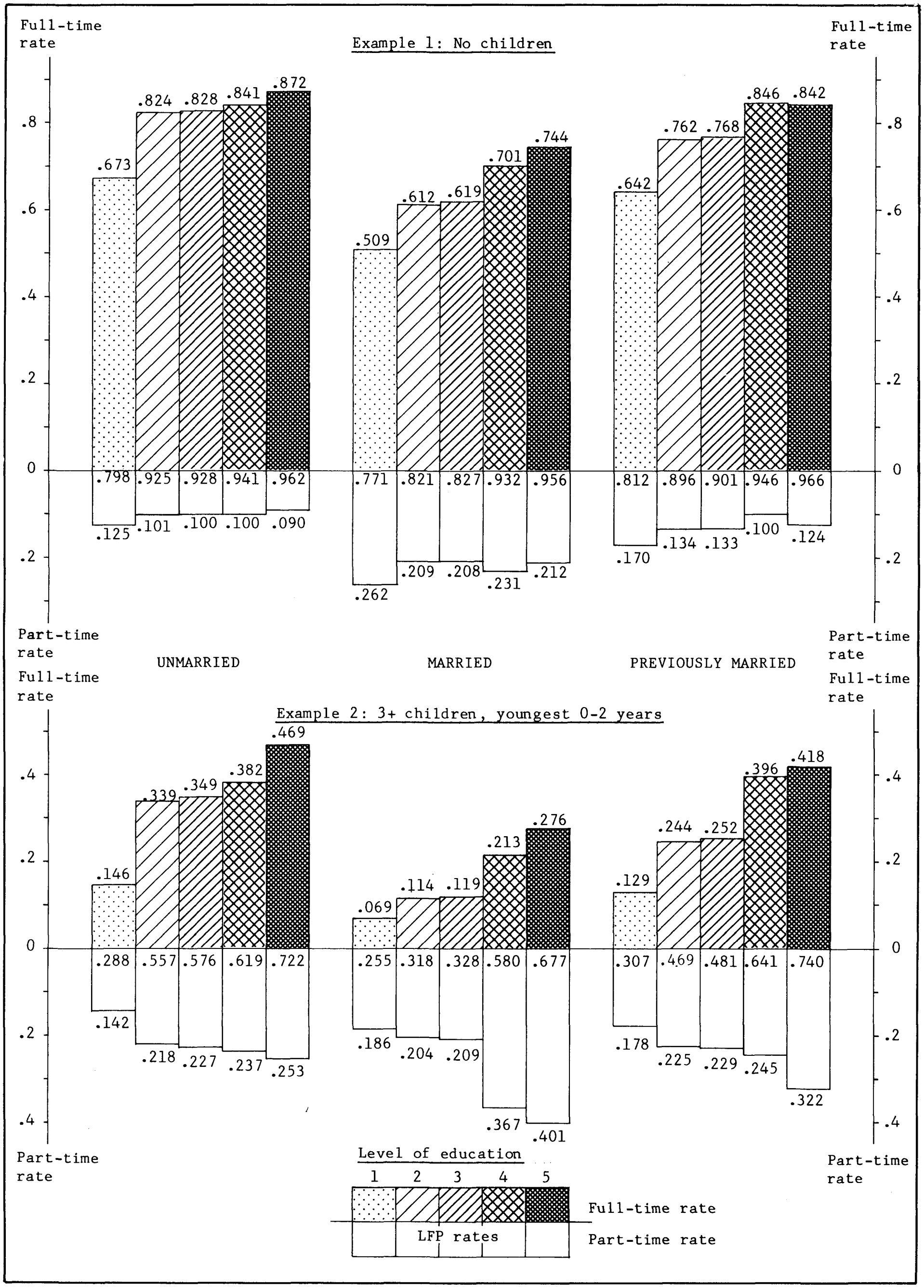
The log-odds ratio between married and unmarried is -1.0206. This corresponds to a 24 percentage points difference as evaluated at the sample mean. That is, if unmarried women prefer full-time over part-time at a rate of 52 per cent, the rate among married women with comparable attributes would be 28 per cent. This compares to an about 5 hours reduction of the working week, as seen from table 4.4 and figure 4.7.

So far we have been studying the conditional probabilities of full-time and part-time work, given at least one hour's employment. What can we say about unconditional probabilities, i.e. about full-time and part-time labour supply as per cent of the entire population 16-74 years? The answer to this cannot easily be read off table 4.2. Even if the full-time rate increases in relation to part-time (i.e., $\beta_{1g} - \beta_{2g} > 0$), it does not necessarily mean that part-time becomes less common. Both may grow at the expense of "no work".

Because of the non-linear structure of the logit model, the result depends on all other variables in the model. Two examples have been computed. By multiplying the conditional probabilities deducible from the trinomial logit model with the corresponding full employment LFP probabilities of the binomial model we derive measures of full-time and part-time labour supply, respectively. The assumption inherent in this procedure is that labour force participants who did not work during the survey week (i.e., were unemployed or temporarily absent) supply part-time and full-time labour in the same proportion as other labour force participants with equivalent attributes.[1]

[1] Confer note 1, page 85.

Figure 4.8. Estimated full-time, part-time, and total LFP probabilities, by level of education and marital status. 35-year-old women with different family types



The results are shown in figure 4.8. Example 1 concerns women at the age of 35, without children, i.e. a fairly "labour force prone" group. Example 2 belongs rather at the other extreme: women aged 35, with at least three children, of which the youngest is 0-2 years old.[1]

The shaded area of each column represents the full-time labour supply rate, while the white part shows the part-time rate. Taken together the two parts equal the LFP rate at zero unemployment.

One striking feature appears from all the diagrams: the same (groups of) persons who have a low LFP rate also have a comparatively strong tendency to hold part-time jobs. This applies in particular to married women and also to women with only compulsory education (level 1).

Women without children have, generally speaking, high LFP rates (example 1), and the great majority work full-time. For this highly job-oriented group, education does not have much impact on LFP, except that women at level 1 seem to fall a bit short compared to the others. Within each marital group, part-time rates do not vary a lot, while full-time rates are strongly influenced by education. In fact, part-time and full-time rates seem to vary in the opposite direction as functions of education. Full-time rates increase while part-time rates are lowered through higher education.

Among women around 35 years and without children, LFP is not very dependent on marital status. Married women come out with almost as high LFP rates as the other two groups (except at levels of education 2 and 3). A larger proportion work, however, part-time.

Turning to the less job-oriented group (example 2), the pattern is a bit different. For women with a large family, part-time and full-time rates change in the same (positive) direction when moving to higher levels of education. Education is seen to have a strong impact on labour supply, no matter what measure we use. Among married women at level 1 of education, only (an estimated) 7 per cent hold full-time jobs, while about 19 per cent work part-time. At the highest level of education, the corresponding figures are 28 and 40 per cent, respectively.

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[1] Having chosen 35-year-old women as our example obviously does not mean that the diagram is irrelevant for other ages. 30- or 40-year-old women will have very similar labour supply rates. Even for 20, 50 and 70-year-olds, labour supply rates will vary across marital status, levels of education, and (if applicable) the number and age of children, in a manner consistent with the pattern shown in figure 4.8.

Again, marital status seems to affect the choice between part-time and full-time, but not the total LFP rate (with the exception, again, of women with medium level education). Now, there are not a lot of unmarried women having three or more children (neither in the population nor in the sample), so that in example 2 the variation across marital groups should be interpreted with caution.

The results of our posterior test for interaction between level of education and presence of children suggest that the variation across educational groups in example 2 may be somewhat underestimated (tables A2.5 - A2.6). In general, the LFP probabilities of university educated women with pre-school children appear to be biased downward in our model. The same is true of women with low and medium education and no pre-school children. For other groups the bias is positive. This essentially means that in example 2, the LFP differentials between various levels of education are even larger than indicated by figure 4.8. In example 1, however, the true differentials are actually overstated by the diagram.

4.5. Children in the family

Effect on labour force participation

The impact of child-bearing on women's labour market behaviour, or vice versa, has long been a subject of study. Recent works based on Norwegian data include Ljones (1979), Jensen (1981 a and b), and Bjøru (1981). All have found a strong covariation between female labour supply and the number and age of children.

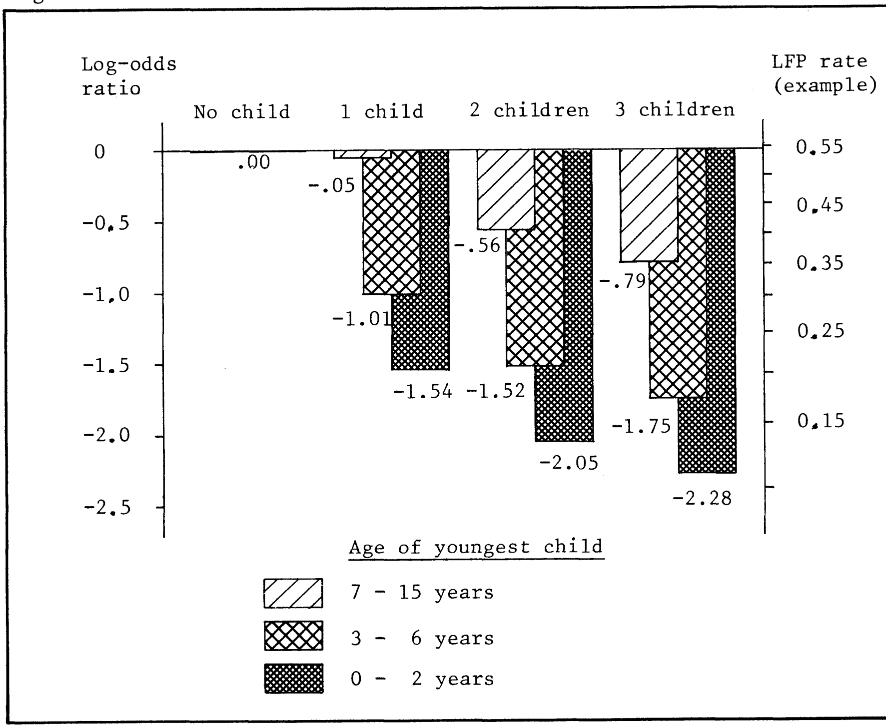
Our results are in line with these general findings. Log-odds ratios for participation versus non-participation are shown graphically in figure 4.9. Women without children are used as the reference group. The full set of contrasts is given in table A3.6 of appendix 3. Almost all contrasts are significant.

As is obvious from figure 4.9, female LFP decreases drastically with the number of children. Even more important, however, is the youngest child's age.

One child aged 7-15 hardly makes any difference on LFP (log-odds ratio insignificant). Two children betwen 7 and 15 yield, however, a log-odds ratio of -0.56, lowering the LFP rate from (for instance) 55 to 41 per cent[1]. With three children the LFP rate drops to 35 per cent.

[1] Taking the sample mean as our "initial" value, i.e. setting the LFP rate of the reference group (women without children) to 0.5455.

Figure 4.9. The effect of children on female LFP



One child less than 7 years old does make a difference, however: the LFP rate drops (in our example) from 55 to 31 per cent when the child is 3-6 years old, and all the way to 20 per cent if the child is even younger. For women with three or more children, of which the youngest is less than 3 years old, the corresponding LFP rate would be down to 11 per cent.

The age of the children seems to have a stronger impact than their number. A similar result was found by Ljones (1979:134).

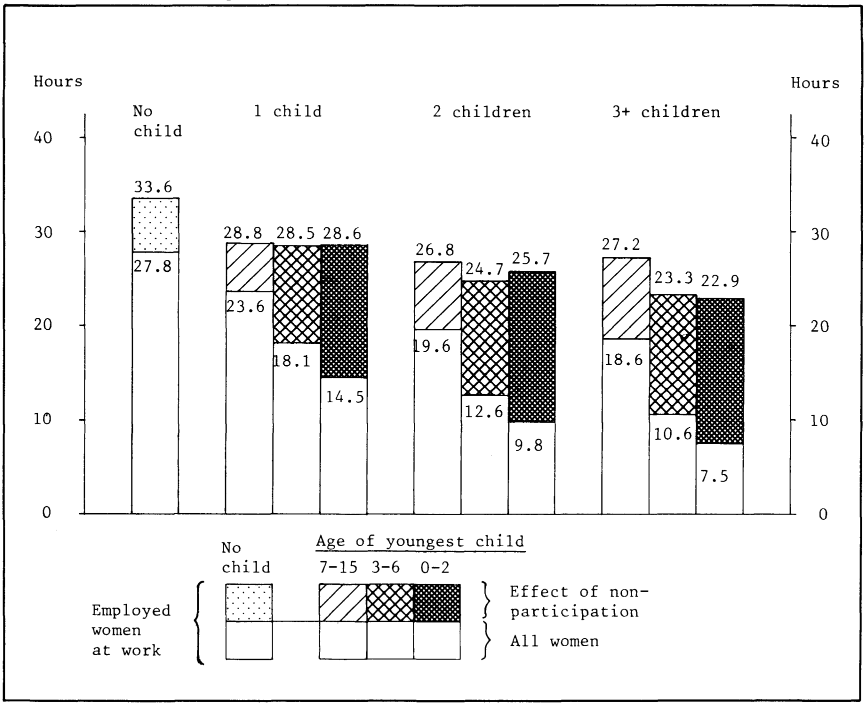
When interpreting figure 4.9, one should, however, keep in mind that the model includes a variable, called KINDG, measuring the availability of institutionalized day-care services (see the next section). The sample mean of KINDG is 0.1175. In figure 4.9, this variable has been set to zero. On the average, therefore, log-odds ratios for children aged 0-2 or 3-6 are 0.1175 x 0.9444 = 0.111 higher than shown in the diagram, thus reducing the gap between age 7-15 and age 0-6. Still the gap remains large.

Results from our posterior test for interaction between the number of children and the age of the youngest child are reported in table A2.4. Essentially, the test suggests that our main model overestimates the rate of LFP among women with one child aged 7-15 years, while underestimating the LFP of women with one pre-school child. For women with 2 or more children the bias appears to be negligible. In other words, for women with one child the differences according to the child's age are smaller than shown in figure 4.9.

Effect on hours of work

Figure 4.10 summarizes our linear regression results concerning the impact of children on their mothers' working week. The format is similar to that of figure 4.7 above.

Figure 4.10. Estimated weekly hours of work, by the number and age of children. Mean values calculated for 35-year-old married women with 11-12 years of schooling



Looking first at the working population only, we notice that married women with one child tend to work on the average 5 hours shorter weeks than those without children. The difference is significant at the 1 per cent level[1].

For women having only one child, the child's age does not seem to matter. Among women having more than one child, however, there is a (statistically significant) tendency to put in fewer hours as long as the youngest child is under school age. Put another way, the number of children (above one) seems to matter only if they are small[2]. Sticking to the group of women with at least one child under 7, we derive an average working week of 28½ hours with one child, 25 hours with two children, and 23 hours among women with three or more children under 16.

When we look at the white parts of the columns the barriers against market work experienced by women with small children stand out more clearly. According to the diagram, the supply of labour from married women at the medium level education is almost cut in half through the arrival of the first child (from 28 to 14½ hours). With three or more children, of which one is under 3, the supply of market work is only about one forth (7½ hours) of the amount offered by "comparable" women without children.

Effect on part-time versus full-time work

How do women with and without children divide their labour supply between part-time and full-time jobs? Log-odds ratios and probabilities in answer to this question appear in figures 4.11 and 4.12.

Mothers with children under 16 favour part-time over full-time work much more often than other women, and more so the more children they have (figure 4.11). The difference between two and three or more children is, however, not statistically significant (see table A3.8).

Using, again, the sample mean as our point of reference, the fraction of women "preferring" full-time over part-time jobs can be expected to drop (e.g.) from 52 to 28 per cent as a result of the first child (aged 0-2). One child aged 7-15 corresponds, in comparison, to

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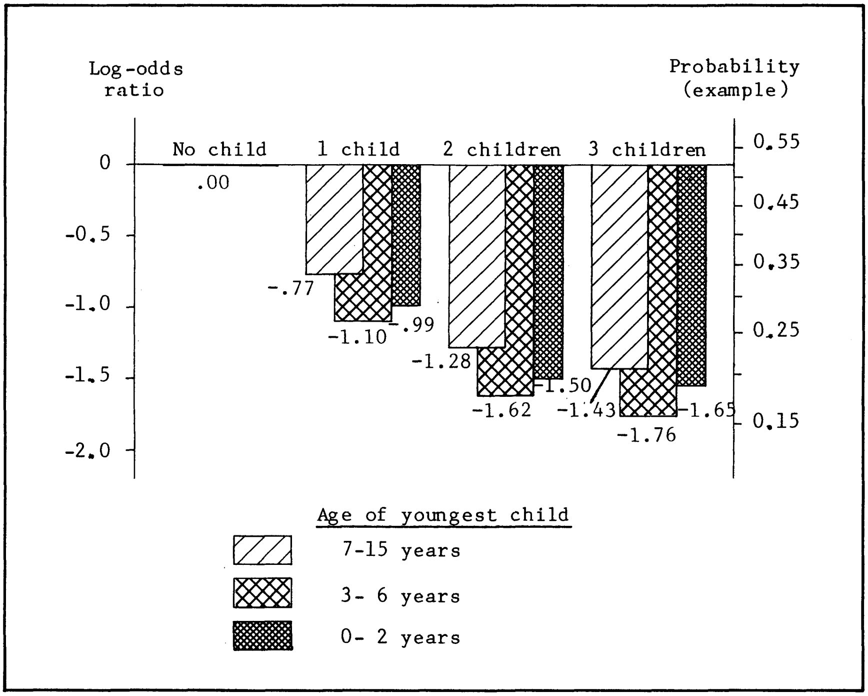
[1] From table 4.4: $\hat{\alpha}_{10}+\hat{\alpha}_{23} = -2.760-2.067 = -4.827$. From the covariance matrix (not reproduced here):
$\text{var}(\hat{\alpha}_{10}+\hat{\alpha}_{23}) = \text{var}\ \hat{\alpha}_{10}+\text{var}\ \hat{\alpha}_{23}+2\ \text{cov}(\hat{\alpha}_{10},\hat{\alpha}_{23}) = 3.9127$.
This yields a t-ratio of $-4.827/\sqrt{3.9127} = -2.441$, and a significant probability of 0.7 per cent against a one-sided alternative. [2] The model includes a full set of interaction terms between the number of children and the age of the youngest child. Thus, the total height of each column in figure 4.10 is determined independently.

about 34 per cent full-time. Two children of which the youngest is 0-2 years, yields a full-time proportion of about 20 per cent (figure 4.11).

The difference between age 7-15 and age 3-6 is statistically significant, while the contrast between age 3-6 and age 0-2 is not (table A3.8). Yet the apparent tendency for employed women with very young children to work at least as much full-time as those with somewhat older children is puzzling. Is the reason for this that women withdraw from the labour force altogether when having their babies, only to reenter the labour force for a part-time job after a few years?

The answer, as apparent from figure 4.12, seems to be yes. However, the pattern comes out rather differently between groups with different levels of education.

Figure 4.11. The effect of children on women's "choice" between full-time and part-time work



Women at the lowest level of education definitely do withdraw from the labour force when having their first child. The total LFP rate drops (for instance) from 77 to 42 per cent (example 1, one child aged 0-2). The full-time rate drops from 51 to 18 per cent, while the part-time rate shrinks by about 2 percentage points. Thus, although some women may change from full-time to part-time employment, enough part-time working women withdraw from the labour force that the net effect on the part-time rate is negative[1].

As the child reaches the age of 3-6, typically[2] about 13 per cent of the women reenter the labour force, although (as a net figure) 9 out of these 13 per cent take part-time jobs. As the child reaches school age, the LFP rate again becomes almost as high as for women without children. However, the split between full-time and part-time is different: an estimated 40 per cent of women with one child aged 7-15 work part-time, as against 26 per cent for those without children.

The effect of having more than one child is to further reduce both the part-time and the full-time rate, although the full-time rate decreases more. As a general rule, the part-time employment rate is higher among mothers than among other women with low education although their total LFP rate is much lower.

Examples 2 and 3 concern married women with medium and high level education, respectively. At the medium level the pattern looks very similar to example 1 (level 1), although employment rates (especially full-time) are generally higher. In this group, the first child seems to bring about a net increase in the part-time rate, while further children will reduce it.

Turning to the highest level of education the picture is different. Here the first child produces only a moderate fall in LFP, typically from 96 to 82 per cent for married women. The full-time rate, however, shrinks by no less than 28 percentage points, from (e.g.) 74 ½ to 46 ½ per cent. As a net result, half of these 28 per cent leave the labour force, the other half change to part-time work. With more children, the part-time rate continues to grow, while the full-time rate

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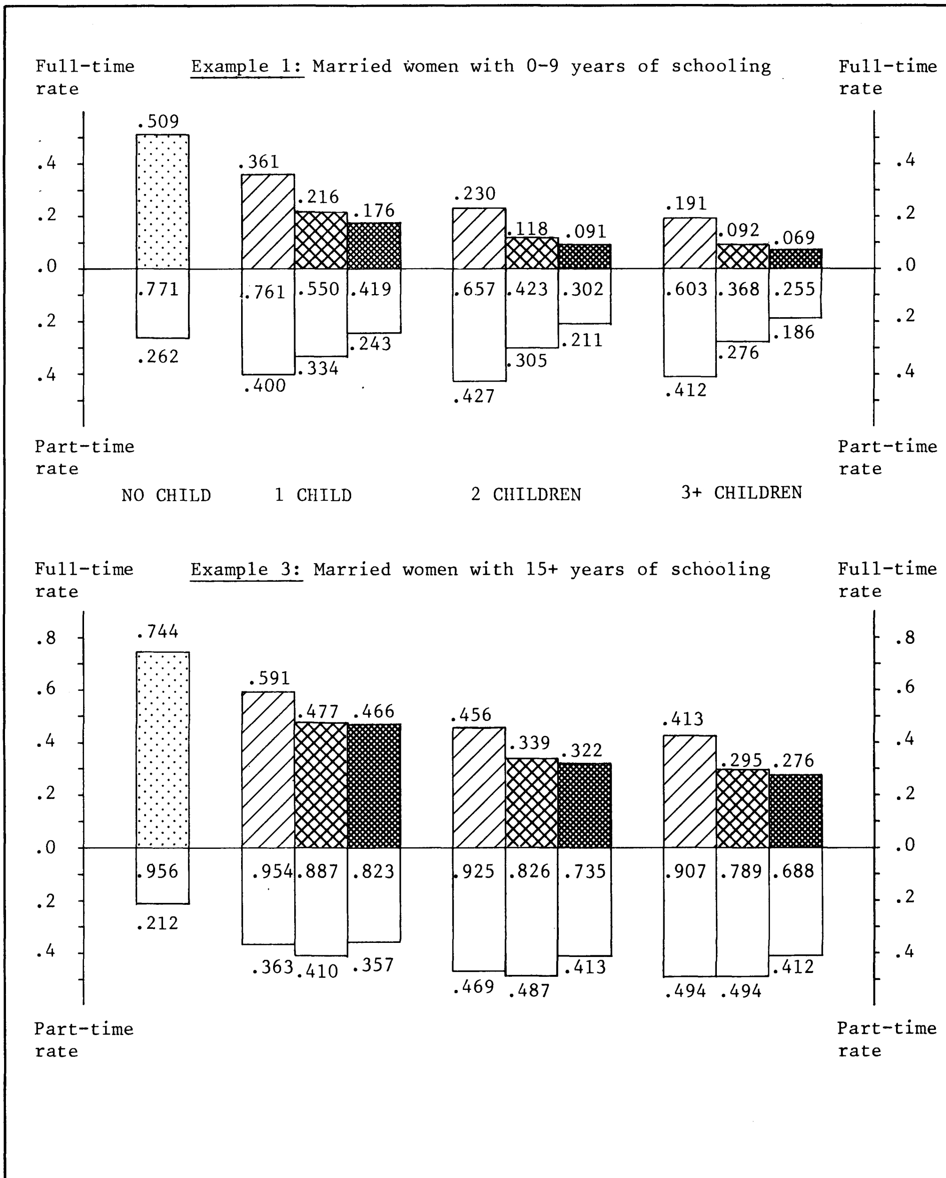
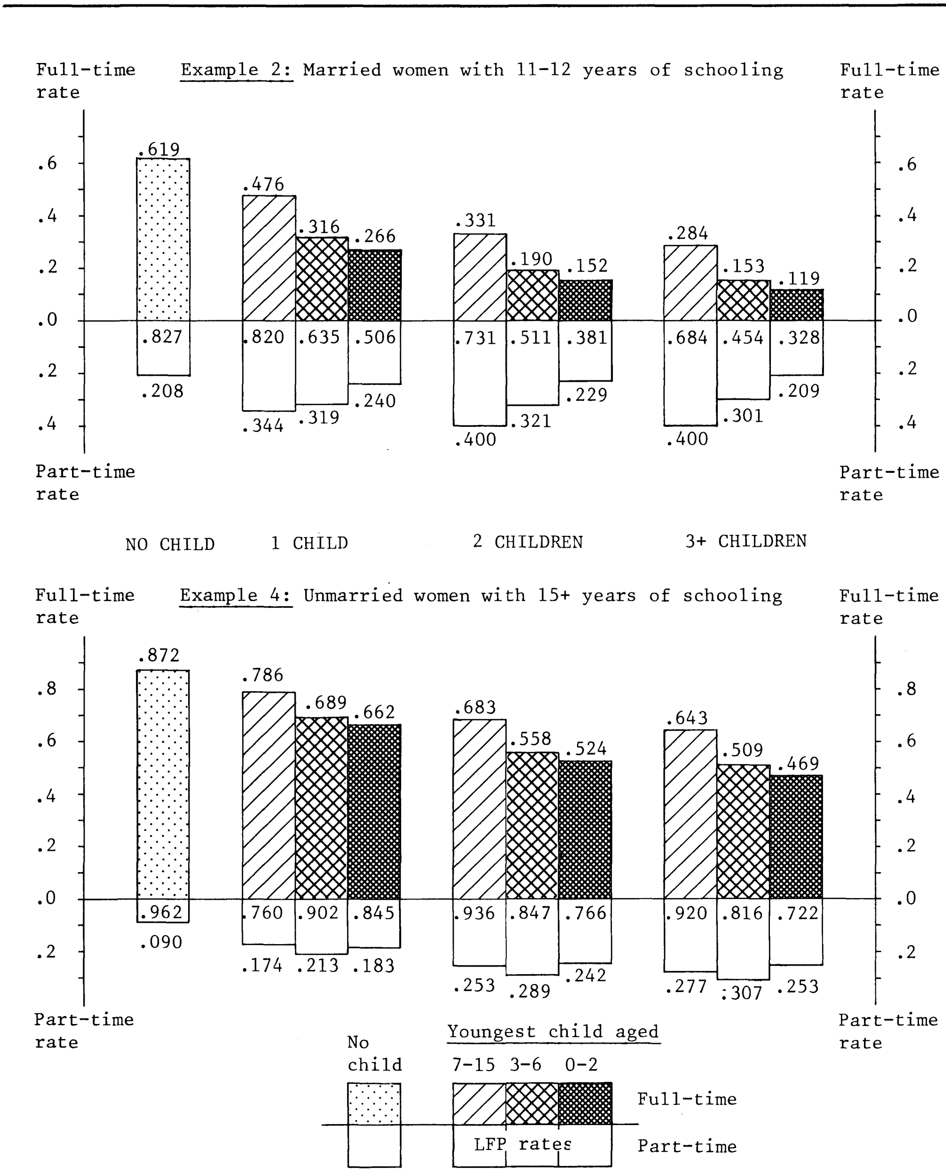
[1] The figure stated are averages for women with children in the three-year age interval 0-2. During the first year or so after childbirth the LFP rate will probably be lower than indicated by our figures, i.e. more women withdraw from the labour force, but some return before their children reach three years of age. [2] Recall that these estimates are examples, valid for 35-year-old married women. For other socio-demographic groups the figures will be similar, but not identical, cf. note 1 on page 91.

Figure 4.12. Estimated full-time, part-time, and total LFP probabilities, by the number education



and age of children. 35-year-old women with different types of marital status and



shrinks by another 14 percentage units for the second child and about 5 percentage units for the third child. The age of the youngest child clearly affects full-time rates, but hardly part-time rates[1].

Finally, a forth example is included in order to facilitate comparison between marital groups. Example 4 is like example 3, except that the former applies to unmarried women. The pattern of variation across the number and age of children is seen to be very much the same as for the married. This relationship appears to be "real", in that the interaction between children and marital status was tested for but found insignificant (appendix 2). Perhaps one explanation for this lies in the neat structure of the logit model. We note that, even if the log-odds ratios are the same in the two cases, the impact of children on labour supply rates as measured in percentage points is smaller for unmarried than for married women. This is so because the LFP and full-time rates of unmarried women are generally closer to 1.

A note on cross-sectional versus cohort data

Now, is it really true that women go in and out of the labour force as children are born and grow up, or are we comparing different groups of women with different patterns of behaviour throughout their lifetime? Our data set is a cross-section of women during a given year. The respondents belong to different cohorts (generations). During the 1970s the cross-sectional age profile of female LFP has risen year by year, so that virtually all birth cohorts have been subject to rising LFP rates (Fridstrøm 1981:31, Bjøru 1981). No temporary lapses in female LFP due to child bearing are detectable.

Yet, the fact that such lapses are not discernible in aggregate data does not rule out their existence. Women belonging to the same cohort will have their children at different times: some in their teens,

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[1] It might be argued that since interaction terms between education and children have not been included in the model, we are not really in a position to compare the impact of children between separate educational categories. To some extent this is true. The patterns materializing in figure 4.12 are a product of the model structure as well as of the data. Here, however, the results shown in tables A2.5 and A2.6 come in handy. As noted at the end of section 4.3, the lesson from the interaction test is that LFP rates are underestimated for women with pre-school children and a university level education, as well as for women without pre-school children and holding a low or medium level education. For the remaining groups the bias is, by and large, positive. But this implies that the effect of children is even more dependent on the woman's education than what appears from figure 4.12. At the highest level of education the effect of children (esp. of the age of the youngest) is quite small, while the converse is true at the low and medium levels.

some in their twenties, and some in their thirties. When adding together all these individual life histories the aggregate time profile will be rather smooth. Temporary withdrawals from the labour force are unlikely to show, unless there is a strong tendency for the members of a given birth cohort to have their children simultaneously.

The difference between women with and without small children, as illustrated by figure 4.12, is so large that it is bound to reflect temporal variations in individual women's labour supply. Taking account of any "cohort effects" present could only marginally modify the picture. It is conceivable, though, that by the time this study is published (1984) all labour supply rates estimated in our model will have increased by several percentage points. In that case, young childless women as of 1977 who have (say) their first child within 7 years will experience a somewhat smaller drop in LFP than what follows from figure 4.12. On the other hand the increase in labour supply as the children grow up will be larger than indicated in the diagram.

The impact of children on female LFP is probably diminishing over time. During the 1970s, LFP rates rose particularly fast among married women with children (Bjøru 1981:83f., Moen 1981:94). In other words, if our model were reestimated using data from 1984, the suspicion is that child status would come out with somewhat smaller coefficient values than for 1977.

We revert to the question of cohort effects in section 4.7, when discussing how to interpret the age variable.

4.6. Kindergartens

Our binomial logit model of labour force participation includes a variable measuring the relative capacity of kindergartens within the respondent's district of residence[1]. The parameter comes out as significantly larger than zero at the 5 per cent level. It has the expected

[1] To be specific T, KINDG is defined as the proportion of children under 7 years of age resident in the municipality who attend in-door, private or public day-care institutions. The variable is made to apply, however, only to married women having children in the relevant age group (0-6). For other respondents it has been set to zero. (Since unwed or single mothers are generally given priority in whatever day-care centres exist, the variable is considered less relevant for these.) Note that informal day-care or nursing arrangements between relatives, neighbours etc. are not included, even if they are being paid for. Thus, the day-care capacity, as defined here, can be regarded as a politically determined variable, at least in principle.

sign. However the effect is moderate. As the relative capacity of day-care institutions increases by, say, 10 percentage points, the log-odds of labour force participation for married women with child(ren) under 7 goes up by 0.9444 x 0.1 = 0.0944.

Even so, the effect of building new day-care institutions is most probably overestimated by our model. There is an omitted variable bias. Certain labour supply determinants not included in the model are likely to be correlated with the day-care variable. Suffice it to mention (i) women's own attitude towards market work, (ii) the attitudes prevalent in the local social environment, and (iii) the industrial structure and level of economic development in the respondent's district of residence. All three factors vary along the urban-rural dimension, as does also very clearly the day-care variable. As a general rule, the relative day-care capacity is highest in the larger towns.

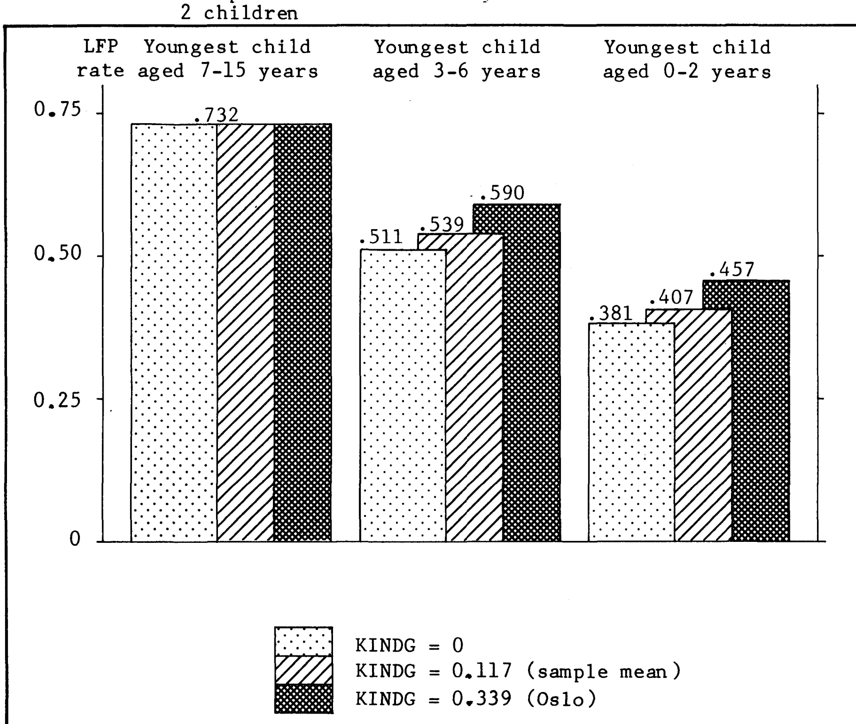
As for (ii), the link between this factor and the day-care variable is more direct than it may seem, in that kindergartens are a municipal responsibility. To the extent that the opinion of the voters are reflected in the governing bodies of the municipality, attitudes regarding women's social role and the division of labour between household and market will have a bearing on the amount of municipal funds allotted to day-care centres. These same attitudes will, of course, also affect women's perceived barriers against market work, i.e. influence their labour supply.

A large part of the day-care institutions are only short-time, i.e. offering their clients less than a full working day's stay. In 1978, this was the case for almost 60 per cent of all children attending day-care institutions (Weekly Bulletin of Statistics No. 50, 1978). In fact, 30 per cent of the children could spend less than 16 hours per week in the day-care centre. This is hardly suficient to bring both parents into the labour force. In other words, extending the opening hours of existing institutions might have as big an effect on female LFP as the opening of new ones. The estimates presented here rely in a sense on the assumption of a constant split between short-time and full-time institutions.

With these qualifications in mind, let us have a closer look at the estimated effect of day-care centres. Again, an example has been worked out, concerning 35-year-old married women with 12 years of schooling, and two children. Figure 4.13 shows the expected LFP rates of such women under three alternative assumptions about the day-care variable. To facilitate comparison, women without children in the pre-school age are also shown in the diagram.

When no day-care institutions exist (which is the case for 14 per cent of the respondents), the estimated LFP rate of women with youngest child 3-6 years is 51 per cent. With an average level of day-care capacity, the LFP rate climbs to 54 per cent, while the Oslo level of kindergarten density corresponds to 59 per cent LFP. The comparable rate of women with only school-age children is 73 per cent. In other words, even with a relatively well built-out kindergarten network, women with small children fall behind.

Figure 4.13. The effect of kindergartens on female LFP. Estimated LFP probabilities for 35-year-old married women with 2 children



What would happen if, for instance, the kindergarten capacity in all municipalities were doubled? By plugging this assumption into the model and recalculating predicted LFP rates, we have been able to evaluate the effect on the labour force. The LFP rate of married women with pre-school children can be expected to grow by about 5 per cent, or by 2.3 percentage points, from 46.8 to 49.1. Since, however, married women with pre-school children constitute only 20 per cent of the working-age female population, the effect on the total female labour force is a mere 0.8 per cent.

A more drastic assumption would be to set KINDG = 1, i.e. assume that there be enough kindergartens for all pre-school children. As a result of this, the average of all married women with pre-school children would rise to an estimated 67 per cent and be almost on a par with those having only school-age children. The total female labour force expands by 7 per cent.

Beware, however, that in these hypothetical examples we have been extrapolating far beyound the variable range[1] occurring in the sample, so that these results are even more uncertain than those shown in the diagram.

As the capacity of kindergartens increases by, say, 100 children, how many women enter the labour force? Hard to say, but a tentative answer could be as follows. According to the Family Statistics (NOS A 951), in 1977 there were about 420 000 children under 7 in Norway, shared among approximately 300 000 families. Doubling the capacity means providing space for an additional 0.117 x 420 000 = 49 000 children[2]. Thereby, the LFP rate of the affected women was found to increase by 2.3 percentage units, i.e. the labour force expands by some 0.023 x 300 000 = 6 900 women. Hence, for every 100 new admissions to the kindergartens, an estimated 100 x 6 900/49 000 = 14 women will enter the labour force.

This figure is only indicative. In reality the result will depend on a number of factors not included in our model, of which the price of day care services is one. As argued by Torp (1981), however, the price of kindergarten services is not likely to affect female LFP as long as the market is rationed and there is excess demand for such services. Only the total quantity supplied counts. In this context, interpreting KINDG as a proxy for (part of) the fixed costs of market work does not seem too far-fetched.

[1] $KINDG_{max}$ = 0.372 in sample. [2] 0.117 is the sample mean of KINDG.

4.7. Age

Effect on labour force participation

A 2nd degree polynomial and a dummy variable for those eligible for old-age pension under the National Insurance Scheme are used to model the relationship between female LFP and age, measured in years. LFP increases as a function of age up until the age of 33, when it starts to fall. A sudden fall in labour force participation in detectable at the age of 67, when all Norwegians are entitled to old-age pension from the National Insurance Institution.

The age variables, in conjunction with the educational characteristics to be discussed below, probably act as reasonable proxies for the opportunity cost of remaining outside the labour force.

The parameter β_4 has an interpretation as the effect of lowering the general age of retirement. If our underlying assumptions hold, i.e., "true" model etc., lowering the age limit from 67 to 66 will reduce the log-odds for 66-year-olds by approximately $-\hat{\beta}_4 = 0.8463$. By recalculating predicted LFP probabilities for all 66-year-old women in the sample, their LFP is found to drop from 29 to 16 per cent. The total female labour force is reduced by 0.44 per cent.

Now, these results should be interpreted with caution. Had, e.g., a 3rd degree polynomial in age been used, allowing LFP to fall faster over the upper age interval, the effect of RETAGE would probably have come out smaller. There is no a priori reason why a second degree polynomial should be considered more "correct", although we might add that for the male sample the addition of a third degree term did not come out as statistically significant. As for the female sample, no similar experiment was made.

When used for prediction of future events, the age variable is altogether a treacherous one. When, as in our case, only cross-sectional data are used, one cannot distinguish between the cohort effect and the "pure age effect". The latter can be thought of as the changes in the dependent variable that occur over time when a particular birth cohort of individuals pass through the life cycle. The former, on the other hand, expresses the fact that people born during different time periods may show different economic or social behaviour through their entire life cycle, i.e. at every given age.

In fact, female labour force participation in Norway is a case in point. As mentioned in section 4.5, participation rates increased markedly during the 1970s. When no other variable than age is controlled for, this increase appears to be for the greater part a cohort effect: younger female generations are altogether more inclined to work for pay in the market (Fridstrøm 1981: 31).

However, during the 1970s a number of socio-economic factors did change simultaneously. Some of these factors are described precisely by such independent variables as are used in our analysis: Fertility fell drastically (Moen 1981), so did nuptiality rates (Brunborg 1979), while divorce rates increased (ibid:). The average level of education also rose, and the total capacity of day-care institutions for children increased several times. Thus, when education, marital status, presence of children, and the capacity of kindergartens are used as explanatory variables, one should be able to capture part of those changes in female labour market behaviour which occurred through the 1970s.

All the parameter estimates shown in table 4.1 have the "right" sign, in the sense that they imply a rising overall labour force participation rate for women when the explanatory variables for the "average" woman develop as they actually did over the past decade. Still, it remains an open question whether the variables included in the model are sufficient to explain <u>all</u> the increase in female LFP during the 1970s. Most probably, they are not. There may still be "cohort effects" not accounted for, due, e.g., to such elusive factors as changes in "tastes" or "attitudes". To examine whether the estimated relation is in fact time-stable, a simulation experiment would have to be done, applying the parameters of our model to (a random sample of) the population in certain selected past years, and comparing the result with the <u>observed</u> size of the female labour force. Such a simulation has, however, been beyond the scope of this study.

The relationship between age and female LFP is pictured in figure 4.14. Each curve applies to a particular socio-demograhic group, as defined by the variables (other than age) which are used in our model (cf. table 4.1).

The uppermost curve (a) shows the predicted probability that an unmarried woman without children and with the highest level of education belongs to the labour force, given a zero local rate of unemployment. We note that for this group of women, LFP is higher than 90 per cent for all age groups up to 55 years.

Curve (b) applies to married women without children and with 10 years of schooling, still assuming a zero local rate of registered unemployment. Their LFP rate is substantially lower than that of group (a) at every given age. Women belonging to group (b), unlike group (a), seldom work after reaching the retirement age (less than 10 per cent).

Curve (c) differs from curve (b) only in that the local rate of unemployment has been set to 4 per cent, close to the sample maximum.

Curve (d) applies to married women with a maximal amount of labour market barriers: only compulsory schooling, 3 or more children, of which the youngest is 0-2 years, 4 per cent unemployment, and no kindergartens.

Recall that these curves represent cross-sections of women not in school in 1977. They do not depict the labour force participation of an "average" woman through her life-cycle. Most women will see their socio-demographic attributes change over time, i.e., they will "jump" from one curve to another. Notice, e.g., that curve (d) does not extend beyond the age of 45, since it is rather unlikely that women much above this age would have 2-year-old children.

Figure 4.14. Female LFP as a function of age. Selected socio-demographic groups

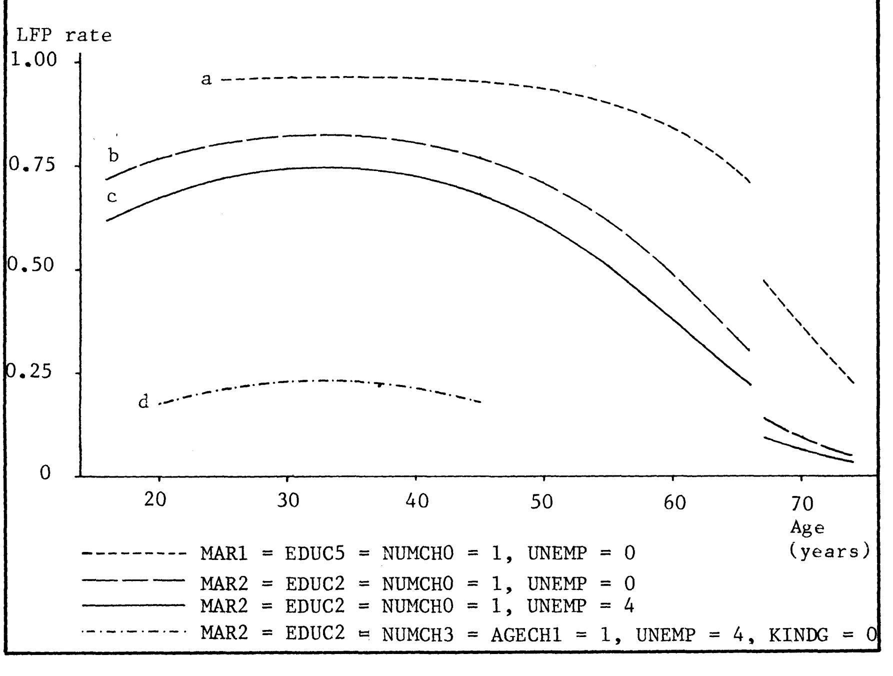
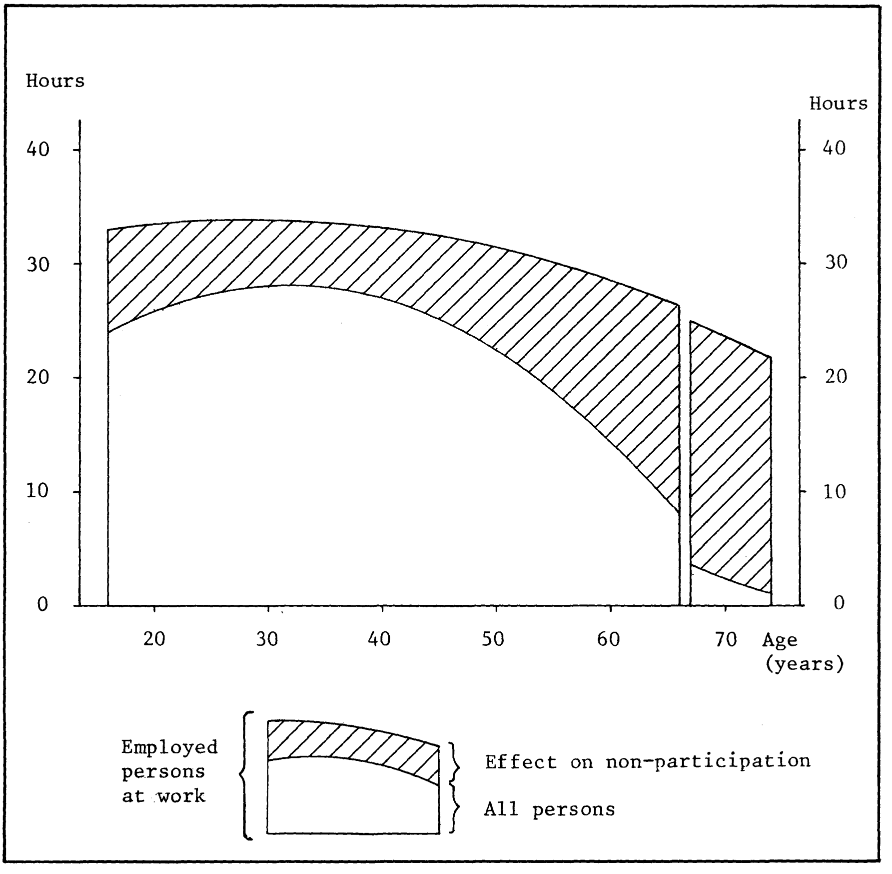


Figure 4.15. Estimated weekly hours of work, as a function of age. Mean values calculated for married women with 11-12 years of schooling and no children



Effect on hours of work

Age has a statistically significant, although relatively modest effect on the length of the working week among employed women. This is evident from figure 4.15 (upper curve). The RETAGE variable is, however, insignificant. It appears that women reaching the age of 67 withdraw from the labour force rather than reduce their working week.

The lower curve represents the "supply" of working hours calculated as an average for all women at a given age. This curve shows more variation than the upper one, since LFP rates vary markedly with age.

Note that marital status, education, and children are held constant. Thus, the diagram describes the partial effect of age *per se*, be it a cohort effect or a "pure" age effect.

Effect on part-time versus full-time work

Figure 4.16 shows the conditional probability of full-time work, i.e. the fraction of women "preferring" full-time to part-time. In a cross-sectional sample, this fraction appears to decrease rather steadily with age. It is unaffected by old-age pension eligibility.

Figure 4.17 shows unconditional full-time and part-time supply rates as functions of age. The total width of the shaded area equals the LFP rate. Our example concerns married women with 11-12 years education and no children. Zero unemployment is assumed.

The full-time supply rate is highest among women between 25 and 30 years old. Part-time, on the other hand, is most common between 50 and 55. The total LFP rate remains high up into the 50s. However, as age increases, a marked transition seems to take place from full-time to part-time work.

Again, the cohorts may behave differently. It is likely that a major part of the variation in figure 4.17 is attributable to cohort differences.

Both the full-time and the part-time supply rates are seen to drop an extra 6 to 8 percentage points as the retirement age is reached.

Figure 4.16. The female conditional full-time probability, as a function of age. Selected socio-demographic groups

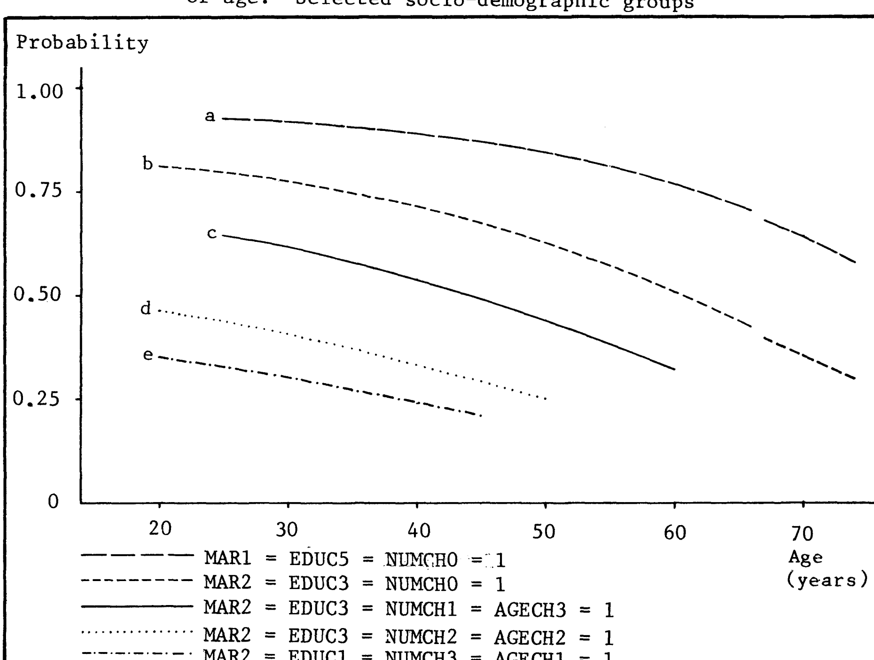
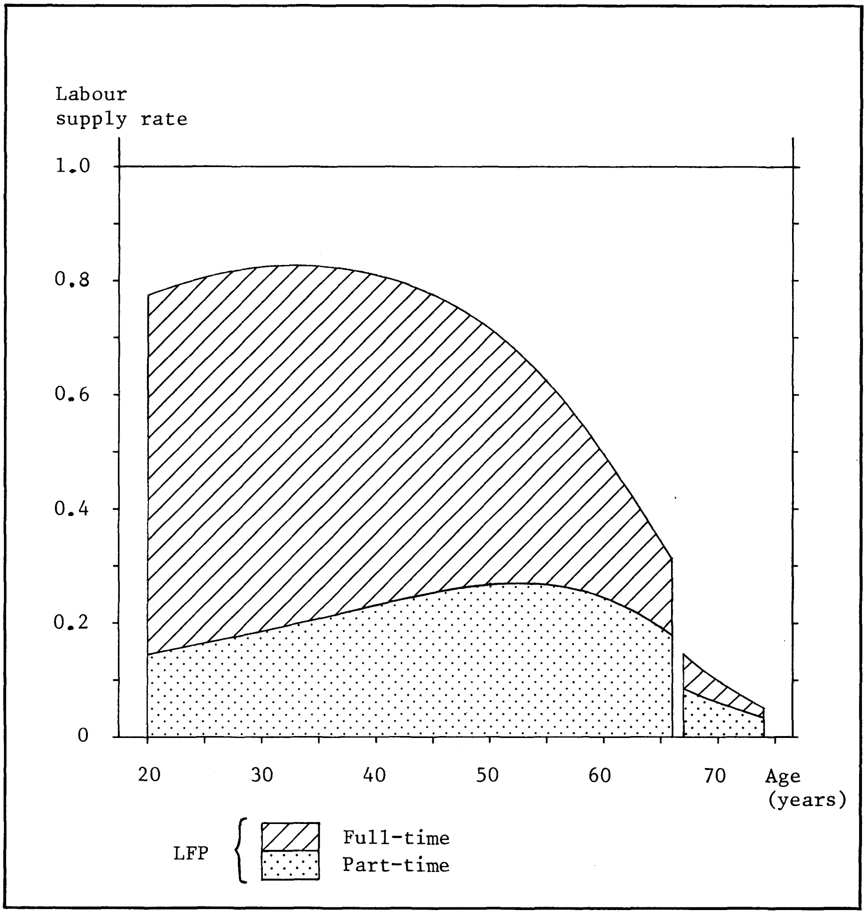


Figure 4.17. Estimated full-time, part-time, and total LFP probabilities, as functions of age. Married women with 11-12 years of schooling and no children



5. MALE LABOUR SUPPLY

5.1. Introduction and overview

Male labour force participation, like female, is negatively related to the rate of unemployment (figure 5.3). As estimated by our model, a one percentage point increase in the jobless rate corresponds to an about 3 percentage points decrease in LFP, for men as well as for women. There is, however, reason to suspect that the model overestimates the discouraged worker effect.

Marital status has a statistically significant "effect" on male LFP. A comparatively large proportion of unmarried men in the middle age intervals do not participate in the labour force. This phenomenon is particularly noticeable among unmarried men with only compulsory schooling, of which 24 per cent remain outside the labour force at the age of 50. Married and previously married men, by contrast, exhibit LFP rates close to 100 per cent up until this age (figure 5.8).

Men's working week typically exceeds 40 hours, no matter their marital status (figure 5.5). No fewer than 1 out of 20 male respondents report working weeks exceeding 60 hours, and 1 out of 5 put in more than 45 hours per week.

Male LFP starts to drop around the age of 55. At 66, the rate is down to 66 per cent. The estimated effect of old-age pension eligibility is 16 percentage points, i.e. a little bigger than for women. As of 1977, close to 50 per cent of all 67-year-old men were still in the labour force, although many of them with a reduced working week (figure 5.7).

The effect of education is also statistically significant, although small as measured in per cent (2-3 points) (figure 5.10). The exception is, again, unmarried men with no more than 9 years education. These fall at least 10 percentage points behind. The work week is about equally long among men at all levels of education (figure 5.11).

Further details are given in sections 5.2 to 5.6.

5.2. General empirical results

Labour force participation

The results of our binomial logit analysis of male labour force participation are given in table 5.1, the format of which is exactly like table 4.1.

Explanatory variables include age, education, marital status, and unemployment. The effects of age and education are assumed to differ between unmarried men and the rest of the male population. This assumption is based on certain preliminary tests for interaction (reported in appendix 2).

All but two parameters are significantly different from zero at the 1 per cent level. The "direct" effects of MAR2 and MAR3 are, however, found insignificant. This does not mean that marriage makes no difference. The effect of marital status is captured by the interaction terms.

The estimated LFP probabilities of the male sample are distributed as shown in figure 5.1. More than half the sample have estimated LFP probabilities above 0.95. The sample mean is $\hat{p}$ = 0.8596. The histogram is, however, extremely skewed. About one half per cent of the men have estimated LFP probabilities of less than 0.10.

Our preliminary testing results suggest that there is significant interaction between all three variables age, marital status, and education. The interaction between age and education was, however, found insignificant except for educational level 4 (table A2.10). This latter interaction term was, however, not built into the model since no plausible economic interpretation could be given.[1]

[1] The model with all 1st order interaction terms between education and (the) age (polynomial) suggests that for educational level 4, the curvature of the age profile is sharper than for other levels of education. However, the difference in LFP probability is appreciable only for men above the age of 65 (see figure A2.1).

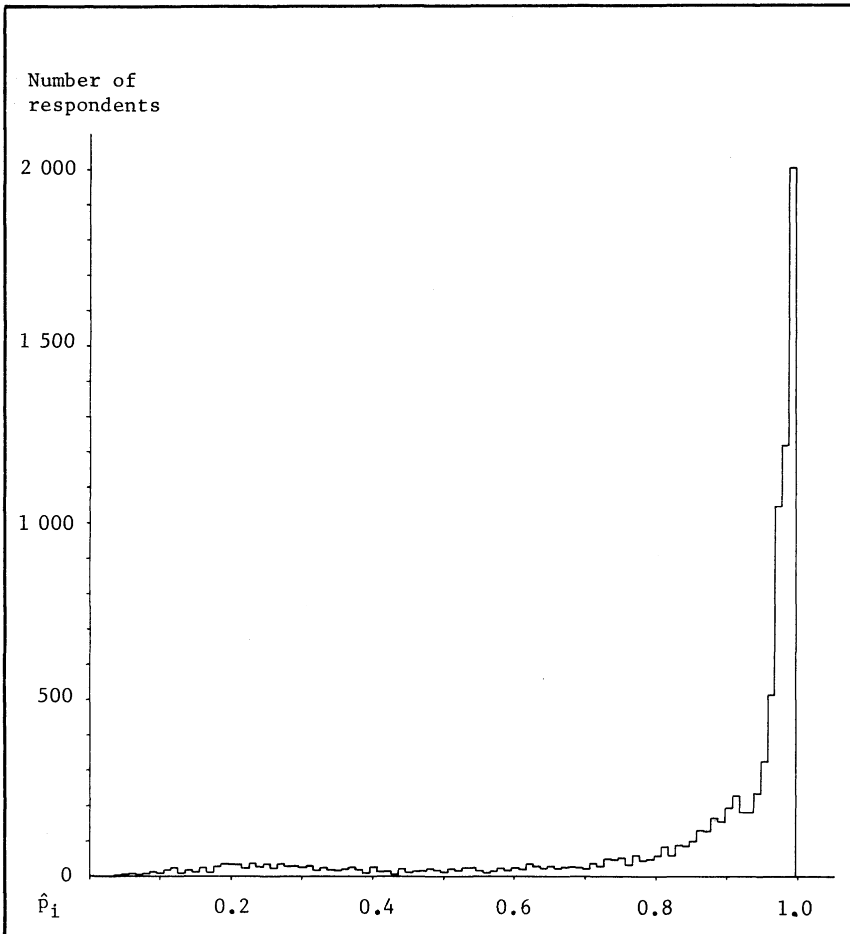
Table 5.1. Estimation results from the binomial logit model of male labour force participation. Men 16-74 years not in school, excl. conscripts. 1977

g	Independent variable x_{ig}	Parameter estimate $\hat{\beta}_{1g}$	Standard error $\hat{\sigma}_{1g}$	Additive effect at sample mean $\hat{\Delta} p$
1	CONSTANT	3.6412**	0.5685	.
2	AGE/100	12.7211**	4.4783	-0.012[1]
3	(AGE/100)^2	-24.4349**	4.3693	
4	RETAGE	-0.7240**	0.1549	-0.112
5	EDUC2	0.5437**	0.1130	0.054
6	EDUC3	0.5930**	0.1217	0.058
7	EDUC4	1.1044**	0.2536	0.089
8	EDUC5	1.8116**	0.2669	0.115
9	MAR2	-0.7620	1.1905	-0.119
10	MAR3	-1.5704	1.1960	-0.300
11	(AGE/100) • MAR1	-13.9184**	4.9095	0.005[1]
12	(AGE/100)^2 • MAR1	20.1243**	4.8819	
13	EDUC1 • MAR1	-0.8368**	0.2009	-0.133
14	UNEMP	-0.2416**	0.0567	-0.032

log L = -2 178.65 n = 8 922

[1] Calculated effect of one extra year at AGE = 45.137 (sample mean)
Legend: See table 4.1.

Figure 5.1. Histogram of estimated LFP probabilities. Male sample



As for the interaction terms involving marital status, it turns out that status "unmarried" is the one that differs (significantly) fro the others. This applies to both sets of interaction terms, i.e. with age as well as education. Thus, a few product terms involving the dumm variable MAR1 were sufficient to model this relationship.

The partial effects of all variables are discussed in sections 5.3-5.6 below.

Part-time or full-time work

Table 5.2 presents the results of our trinomial logit analysis of full-time, part-time, and no work. The format is like table 4.2.

The trinomial model incorporates the same personal attribute variables as our binomial model of male LFP (including the interaction terms). Each variable has a significant effect on the log-odds between full-time and no work. All variables except age have a significant effect on the log-odds between part-time and no work. (For unmarried men even the age effect is significant.) However, most variables affect full-time and part-time in much the same way. Thus, the log-odds between full-time and part-time are largely unaffected by other variables than age.

The histogram of conditional full-time probabilities is shown in figure 5.2. This probability is not lower than 0.33 for any member of the male sample. For 80 per cent of the respondents, it exceeds 0.90.

The sample mean is 0.9127.

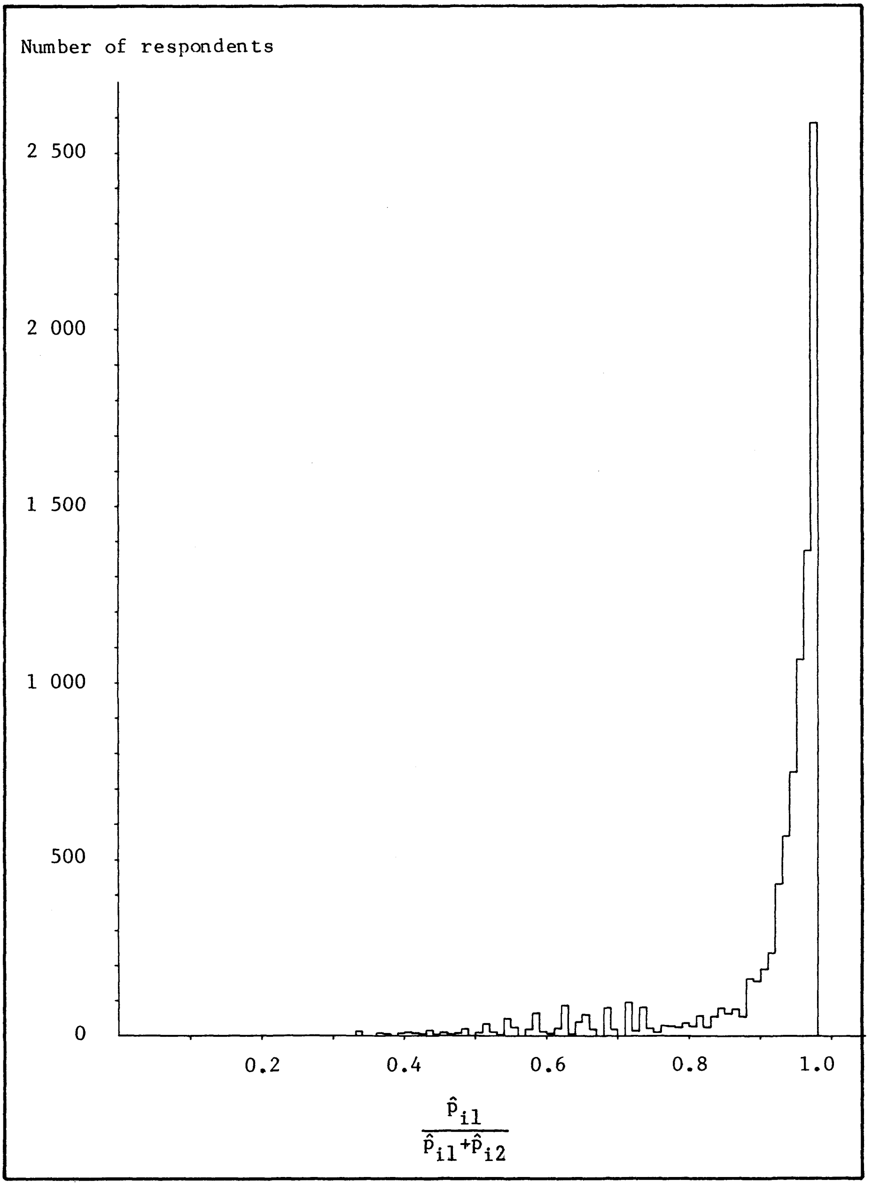
20 per cent of all men working more than 34 hours per week state that they would prefer a shorter working week (table 4.3). Thus, if workers could choose more freely how many hours to work, the fraction working part-time might have been somewhat higher, cf. the discussion in section 4.2.

Table 5.2. Estimation results from the trinomial logit model of male part-time and full-time work. Men 16-74 years not in school, excl. conscripts. 1977

		Full-time vs. no work		Part-time vs. no work		Full-time vs. part-time work	
g	Independent variable x_{ig}	Parameter estimate $\hat{\beta}_{1g}$	Standard error $\hat{\sigma}_{1g}$	Parameter estimate $\hat{\beta}_{2g}$	Standard error $\hat{\sigma}_{2g}$	Parameter estimate $\hat{\beta}_{1g}-\hat{\beta}_{2g}$	Standard error $\hat{\sigma}_{12g}$
1	CONSTANT	0.4794	0.4340	0.8941	0.7608	-0.4147	0.7192
2	AGE/100	21.9252**	2.5469	4.4636	4.3718	17.4616**	3.9881
3	$(AGE/100)^2$ ...	-29.4124**	2.7027	-6.9564	4.6372	-22.4560**	4.2761
4	RETAGE	-0.8428**	0.1443	-0.1288	0.2428	-0.7140**	0.2423
5	EDUC2	0.5527**	0.0898	0.2922*	0.1615	0.2605*	0.1523
6	EDUC3	0.5067**	0.0927	0.3252*	0.1664	0.1815	0.1557
7	EDUC4	0.7901**	0.1652	0.7296**	0.2685	0.0605	0.2357
8	EDUC5	1.2453**	0.1795	1.6963**	0.2344	-0.4510*	0.1878
9	MAR2	-2.1535**	0.6701	-2.7038*	1.1393	0.5503	1.0513
10	MAR3	-3.1379**	0.6811	-2.9444*	1.1515	-0.1935	1.0677
11	(AGE/100)•MAR1	-15.4436**	3.2575	-17.1766**	5.5456	1.7330	5.2521
12	$(AGE/100)^2$•MAR1	18.2172**	3.5943	18.9710**	5.9729	-0.7538	5.7839
13	EDUC1• MAR1 ..	-0.3072*	0.1461	0.3916	0.2608	-0.6988**	0.2416
log L = -4907.08		n = 8922					

Legend: See table 4.1.

Figure 5.2. Histogram of estimated conditional full-time probabilities. Male sample



Hours of work

As with the female sample, a linear regression analysis was carried out to examine the variations in hours worked among employed men at work (table 5.3).

The model incorporates age, education, and marital status. The variation according to age is modelled by means of a 3rd degree polynomial. A set of 1st order interaction terms between status "unmarried" and the age polynomial is included. In addition, a set of 2nd order interaction terms has been included, calculated as the first set times the EDUC1 variable. All interaction terms are statistically significant. So are all the direct effects, except for EDUC5. Yet the model explains no more than 4.5 per cent of the total (squared) variation in hours worked (R^2 =0.0453).

An alternative model was also tried, including no 2nd order interaction terms but a full set of 1st order terms between marital status and education. The latter interaction terms were, however, found insignificant.

The frequency distribution of hours worked by the male sample shows, not unexpectedly, that there is a large number of respondents stating "40 hours". We are hard put to determine to what extent this reflects a strictly enforced regime of standardized working contracts, or merely the fact that respondents tend to state round numbers. (This latter tendency is noticeable over the entire range 0-99 hours.) At any rate, it seems implausible that 45 per cent of all working men would choose to work exactly 40 hours per week, unless there are strong institutional quantity constraints in the labour market. Thus, the "true" supply of labour from employed men is probably more diversified than suggested by the observed frequency distribution of hours worked.

Table 5.3. Estimation results from the linear regression model of hours worked by employed men at work. Dependent variable: T

g	Independent variable x_{ig}	Parameter estimate $\hat{\alpha}_g$	Standard error $\hat{\sigma}_g$
1	CONSTANT	100.387**	18.318
2	$(AGE)^{-1} \cdot 100$	5.125*	2.399
3	AGE	1.654**	0.455
4	$(AGE)^2/100$	-1.538**	0.358
5	RETAGE	-4.100**	1.078
6	EDUC2	1.222**	0.389
7	EDUC3	0.905*	0.395
8	EDUC4	-1.039*	0.571
9	EDUC5	0.151	0.546
10	MAR2	-111.230**	24.705
11	MAR3	-113.304**	24.706
12	$[(AGE)^{-1} \cdot 100] \cdot$MAR1	-12.289**	2.936
13	AGE•MAR1	-3.195**	0.677
14	$[(AGE)^2/100] \cdot$MAR1	2.837**	0.593
15	$[(AGE)^{-1} \cdot 100] \cdot$MAR1•EDUC1	-0.922*	0.486
16	AGE•MAR1•EDUC1	0.315*	0.149
17	$[(AGE)^2/100] \cdot$EDUC1	-0.547*	0.243
	$R^2 = 0.0453$ $\hat{\sigma} = 11.202$		

Even so, it is interesting to observe that a full 21 per cent of all working men put in more than 45 hours per week, 13 per cent work 50 hours and more, and 5 ½ per cent are above 60 hours. There are respondents reporting more than 90 hours of work per week (0.3 per cent).

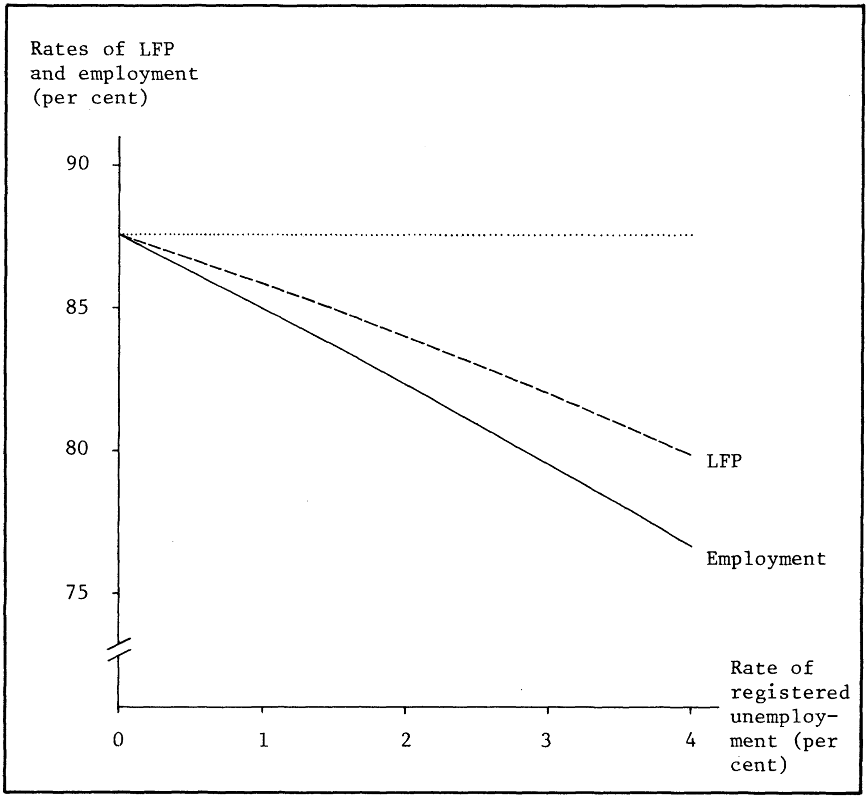
5.3. Labour market tightness

When the rate of unemployment increases by 1 percentage point, the log-odds of male LFP drops by an estimated 0.2416. At the sample mean, this corresponds to a 3.2 percentage points reduction in male LFP. This is our estimate of the (net) male discouraged worker effect. For women the corresponding figure was 2.8 percentage points (table 4.1, column $\Delta\hat{p}$). Thus, it seems that as the labour market slackens, about as many men and women leave (or fail to enter) the labour force. But since the female labour force is smaller, the relative impact of unemployment is larger for women.

As mentioned in section 4.3, however, the discouraged worker effect is probably overestimated in our model.

Figur 5.3 describes the relationship between unemployment, male LFP, and male employment, as estimated by our model. Under zero unemployment the male labour force would have been an estimated 1.8 per cent larger than observed in 1977. As for women, hidden unemployment appears to dominate open unemployment, although to a somewhat lesser extent.

Figure 5.3. The estimated relationship between male LFP and regional unemployment



5.4. Marriage

Effect on labour force participation

Even for men, the rates of LFP differ significantly between different marital groups. The pattern is, however, contrary to what we find for women. Married and previously married men are more economically active than unmarried. The difference between married and previously married is insignificant when controlling for age and education.

In our model there is interaction between marital status and the two other variables. This means that the log-odds ratios for marital status will be different depending on both age and education. An attempt to summarize the information on the effect of marital status has been made in figure 5.4.

The thick lines represent estimated log-odds ratios at different ages. The shaded area around each line represents a set of 59 conditional 10 per cent confidence intervals (one for each age between 16 and 74). Note that this is not the same as a simultaneous confidence region for the entire curve. At a given age we are "90 per cent confident" that the log-odds ratio lies somewhere between the upper and the lower dotted line. We can, however, not be "90 per cent confident" that this is true for all age levels simultaneously.

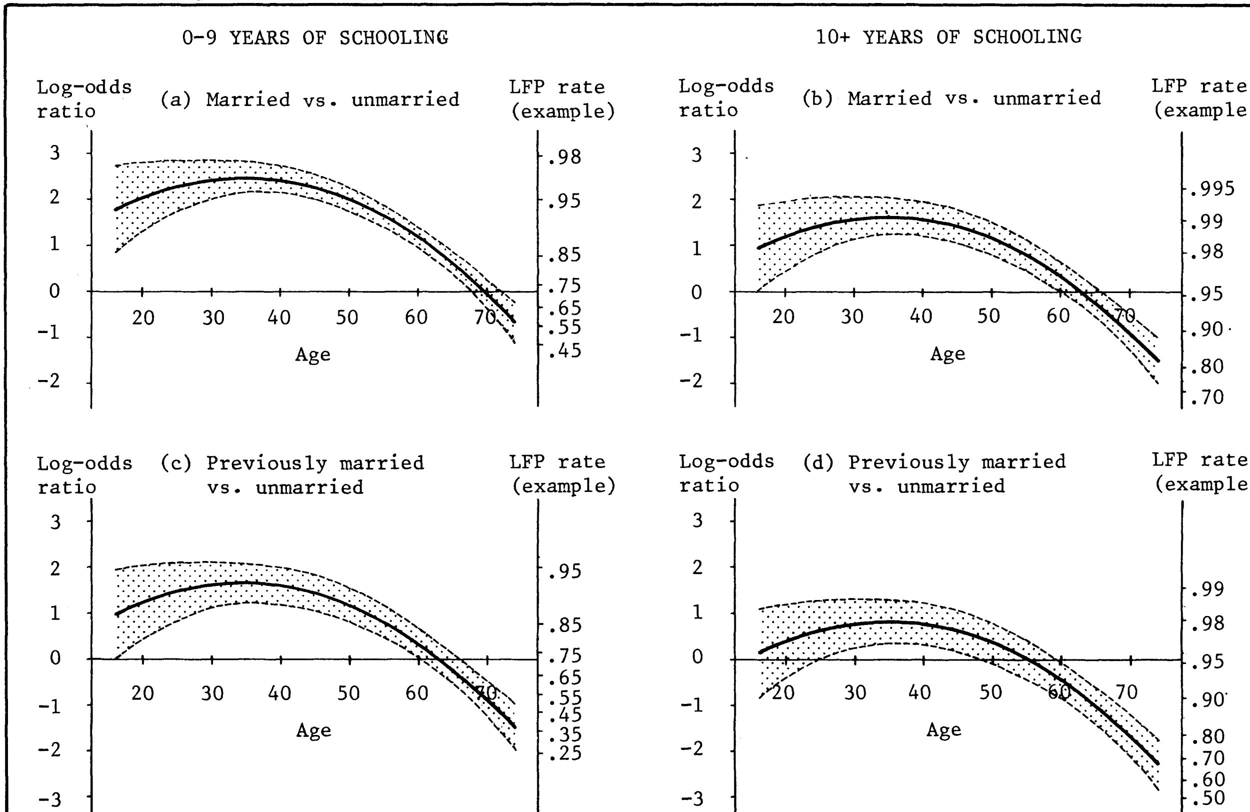
The right-hand scale of each diagram shows the LFP rates of married, resp. previously married men corresponding to the estimated log-odds ratios. The points of reference are the average LFP rates of unmarried men at educational level 1 (0-9 years, p=0.721), resp. levels 2-5 (10+ years, p=0.953).

At the lowest level of education, the difference between married and unmarried men is significant at all ages under 67 (curve(a)). Married men have substantially higher LFP rates than unmarried. Previously married men are also more economically active than unmarried. However, this difference is statistically significant only up to the age of 60.

Among men with more than the compulsory amount of education (levels 2-5), the differentials between marital groups are generally much smaller. At the age of 40, typically 99 per cent of all married men belong to the labour force, as against 95 per cent for the unmarried. The difference between previously married and unmarried men is hardly significant.

According to the model, there is a tendency for old, unmarried men to have a comparatively high rate of LFP. (Log-odds ratios are significantly below zero in the highest age intervals.)

Figure 5.4. The effect of marital status on male LFP. Bands of 90 per cent confidence intervals for log-odds ratios, conditional on age and education



How are these differences to be explained?

Labour force participation is, comparatively speaking, particularly low for middle-aged, unmarried men at the lowest level of education. One hypothesis could be that within this group, there is an over-representation of disabled and less resourceful men. This hypothesis is strengthened by the fact that a similar tendency was found in the female sample. The highly significant interaction term between MAR1 and EDUC1 seems inexplicable unless such a selection mechanism exists.

Even at other levels of education, however, there is a significant difference between married and unmarried men. Married men participate more in the labour force. Apparently, marriage has diametrically opposite effects on male and female LFP.

According to the traditional sex role pattern in western industrialized nations, men are strongly committed to the task of supporting not only themselves, but also their families. The larger their family, the higher income they need. Within this normative framework, there are few legitimate reasons for a married man to remain outside the labour force. Possibly the role expectations confronting unmarried men are somehow less commanding.

On the female side the role expectations are very different. Up until recently, married women were mainly expected to remain at home. At the very least it was (and is) considered fully legitimate for married women to dedicate themselves to their family, especially if they have children. Married women's labour force participation used to be the exception rather than the rule (cf. Foss 1980: 43-52; Ljones 1979: 34-49).

To the extent that these traditions were still in effect in 1977, the observed pattern of variation across sex and marital status is not too surprising.

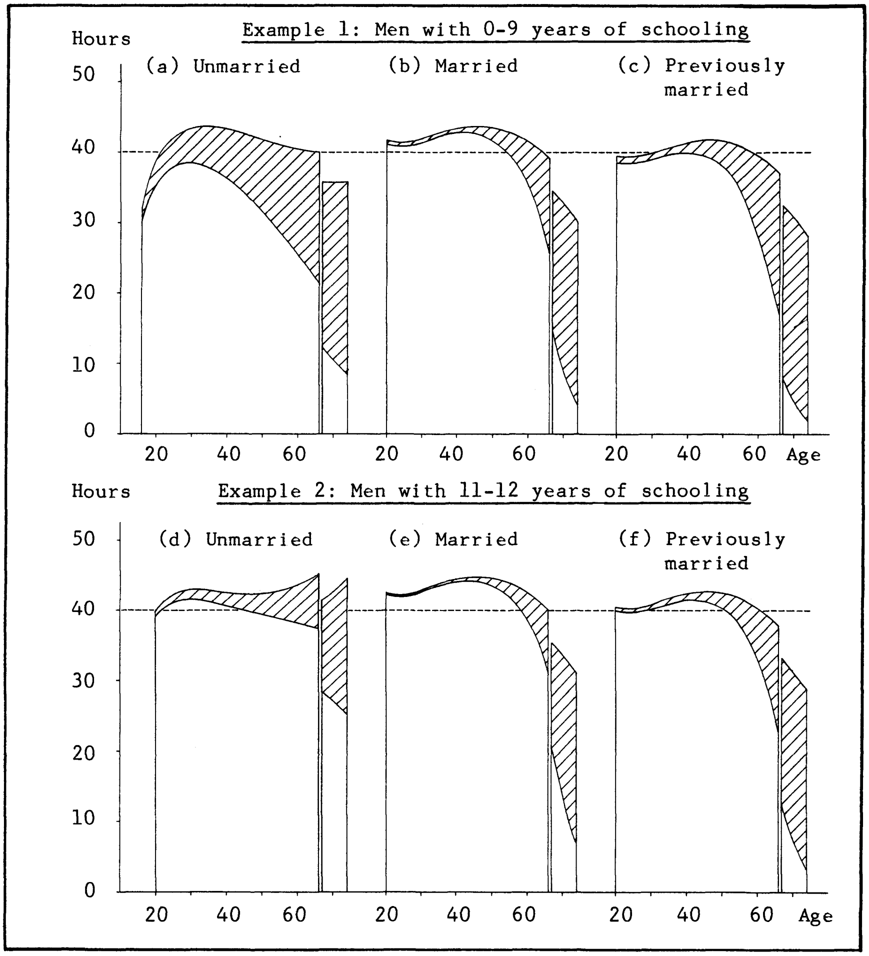
Effect on hours of work

How does marital status affect the length of men's working week? Our model results in this respect are illustrated by figure 5.5[1].

The average length of the working week among those who work (upper curve) does not vary very much with marital status. Yet there is a tendency for young unmarried men to work comparatively short weeks, while older unmarried men appear to put in more hours than do married and previously married men at similar ages.

[1] This diagram has the same format as the corresponding diagram for women (figure 4.15). However, since in the male model there is interaction between age, education and marital status, the age profile is different depending on the latter two variables.

Figure 5.5. Estimated weekly hours of work, as a function of age and marital status. Mean values calculated for men with two types of education



Legend: See figure 4.15.

The lower curves of figure 5.5 represent the male supply of working hours as an average over employed and non-employed persons alike. Due mainly to their low rate of LFP, unmarried men at the lowest level of education are seen to supply considerably fewer hours of labour than their married or previously married brothers.

Effect on part-time versus full-time work

Figure 5.6 shows, by analogy to figure 5.4, the effect of marital status on men's "preference" for full-time as against part-time work. Log-odds ratios for full-time vs. part-time are measured on the left-hand scale and conditional full-time probabilities on the right-hand scale. The reference probability equals the mean estimated conditional full-time rate for unmarried men at the respective educational levels (0.846 for level 1 and 0.946 for levels 2-5).

Among men at the lowest level of education, the difference between married and unmarried is statistically significant, at least at the 10 per cent level. As measured by the fraction of employed men working full-time, the difference between the two marital groups typically may amount to around 7 per cent, however depending on age.

At the higher levels of education there is a tendency for unmarried men to have a stronger preference for full-time work than the previously married. Between married and unmarried men the difference is insignificant.

In figure 5.7 the information on LFP and part-time/full-time rates has been put together.

Up to the age of 60, married men work most. Their total LFP rates are high, while part-time rates are low. Until the age of 50, almost 100 per cent of all married men are in the labour force. Only between 2 and 4 per cent work part-time. This is true at both the low and medium levels of education.

Unmarried men have lower LFP rates and higher part-time rates. At the medium level of education the difference in LFP is moderate, in the lower age brackets almost non-existent. At the age of 50, unmarried men with 11-12 years education have an estimated LFP rate of 93 per cent, as against 98 per cent among married, and 95 per cent among previously married men.

Among men with only compulsory schooling, the differences between marital groups are more striking. An estimated 24 per cent of all unmarried men stay outside the labour force at the age of 50, as against 4 and 9 per cent among married and previously married men, respectively. Full-time rates differ even more across marital groups than LFP rates do.

Figure 5.6. The effect of marital status on men's "choice" between full-time and part-time work. Bands of 90 per cent confidence intervals for log-odds ratios between full-time and part-time, conditional on age and education

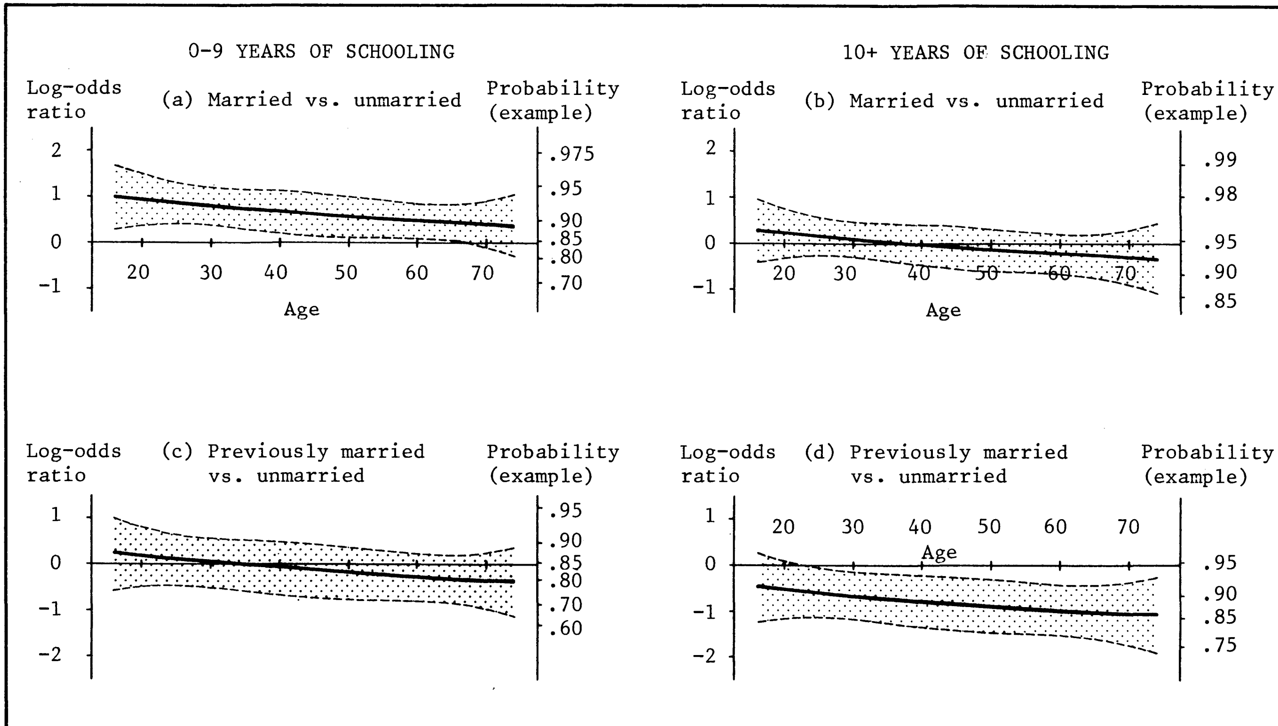
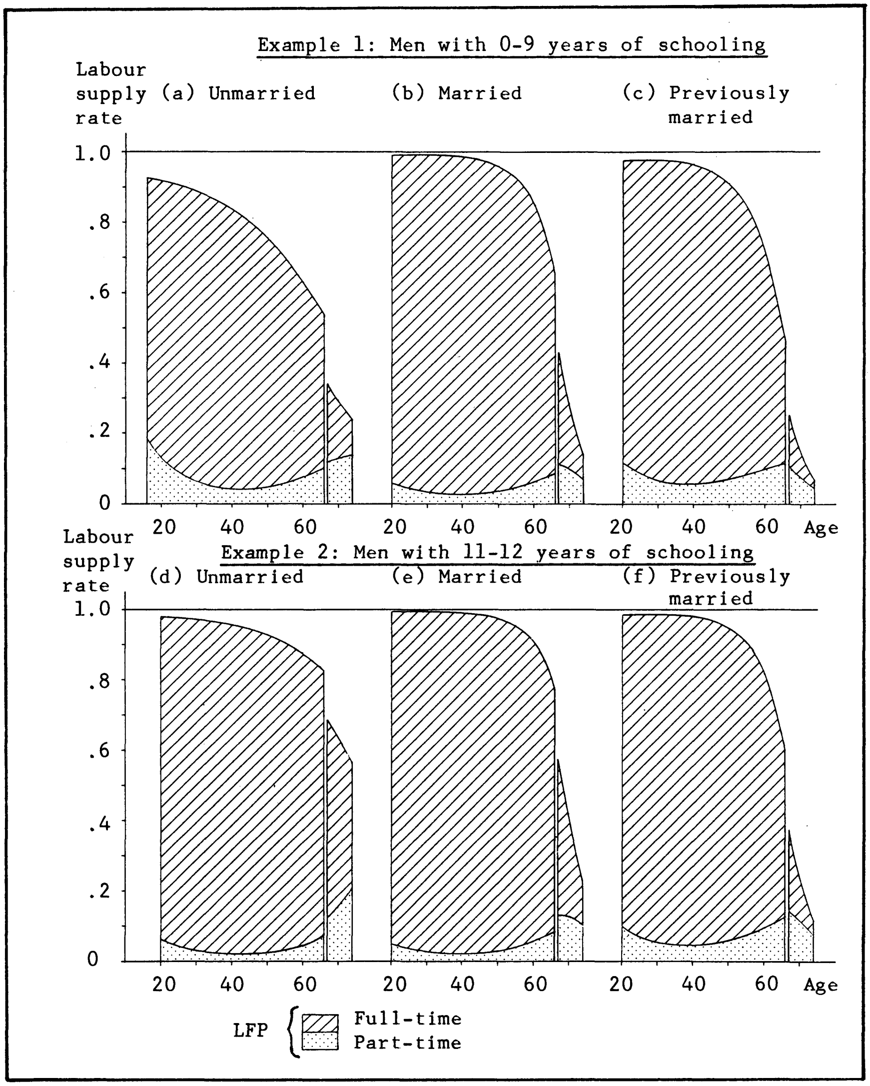


Figure 5.7. Estimated full-time, part-time, and total LFP probabilities, by age and marital status. Men with two types of education



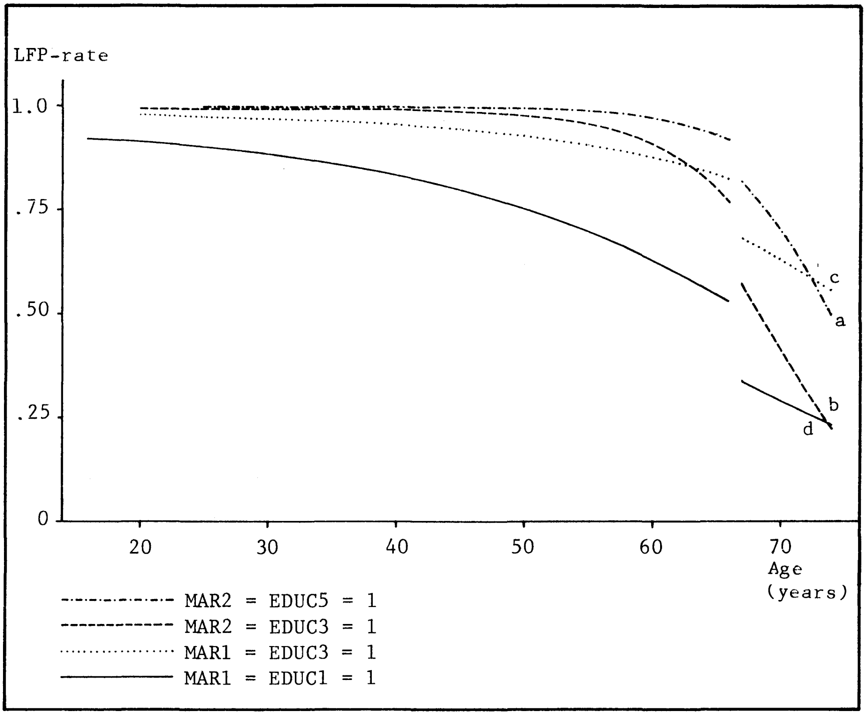
5.5. Age

Effect on labour force participation

Figure 5.8 illustrates how, within groups of men with different types of education and marital status, estimated LFP rates vary with age. The uppermost curve (a) applies to married men with the highest (5th) level of education. Curve (b) represents married men at the 3rd level of education. Curve (c) differs from (b) only with respect to marital status; (c) represents unmarried men. So does curve (d), although here the length of education is 9 years or less. All curves have been calculated on the assumption of zero unemployment.

Among married and previously married younger men, LFP rates are generally close to 100 per cent, but start to drop sharply between 50 and 60 years of age. As noted in the previous section the age profile for unmarried men is very different.

Figure 5.8. Male LFP as a function of age. Selected socio-demographic groups



There is a marked drop in male LFP rates at the age of 67. Recalculating LFP rates for all 66-year-old men, under the assumption that they, too, be eligible for old-age pension, yields an estimated 16 percentage points reduction in their LFP, from 66 to 50 per cent. The total male labour force shrinks by about 0.3 per cent.

Although male LFP is generally very high, it is interesting to note that no fewer than one out of three men withdraw from the labour force before they reach the general retirement age. On the other hand almost half of all men continue working for some time after reaching this age.

Effect on hours of work

To examine the effect of age on the length of the working week, refer back to figure 5.5.

Among employed men under 67, the number of hours varies little with age. Typically the mean working week exceeds 40 hours. The only significant exception concerns unmarried men in their teens.

A certain drop is detectable, though, in the highest age bracket (67-74). The estimated effect of old-age pension eligibility is -4.1 hours. Even 70-year-olds work, however, if they work, an average of at least 30 hours per week.

In general, age seems to affect the male supply of labour not through reductions in the working week, but by lowering the rate of LFP. It is, however, conceivable that this picture be influenced by the existing quantity constraints in the labour market and not only by differences in the "true" supply of labour.

Effect on part-time versus full-time work

The effect of age on men's "choice" between full-time and part-time work is illustrated in figure 5.9. This effect is seen to be very much the same regardless of marital status and education. Full-time rates are highest in the middle age intervals. Up to the age of 65 the great majority of men "prefer" full-time jobs.

There is an interesting tendency for male part-time rates to increase towards the end of the men's working life (figure 5.6). In some cases[1] the part-time rate appears to rise abruptly as the pension age is

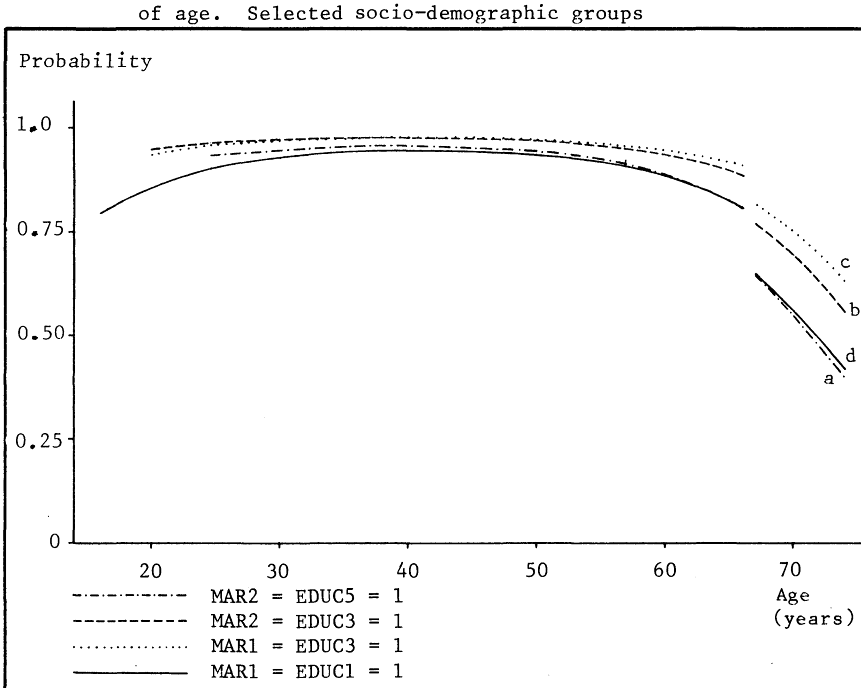
[1] Since the model includes no interaction terms between the RETAGE variable and marital status or education, we are not in a position to compare the effect of old-age pension between different socio-demographic groups. The effects appearing in the diagrams are estimated on the basis of pooled information from the entire sample.

reached. This observation may seem to have an obvious economic explanation:

> "Persons between the ages of 67 and 70 can take out 1/4, 1/2, 3/4 or the whole of their old-age pension. However, the pension plus any earned income must not total more than 80 per cent of previous earned income. Special regulations cover persons who previously had a lower earned income. The full retirement pension is always given after the age of 70." (National Insurance Institution 1979 : 16.)

Previous studies indicate that in the male population, cohort effects on LFP were almost negligible during the 1970s. The male cohort curves coincide approximately with the cross-sectional age profile in any given year (Fridstrøm 1981 : 32). Thus, in the male case, the observed variation by age is for the most part interpretable as a "pure" effect of ageing.

Figure 5.9. The male conditional full-time probability, as a function of age. Selected socio-demographic groups



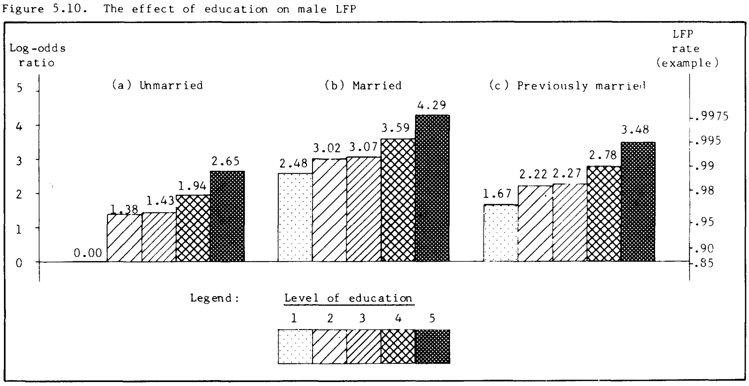
5.6. Education

Effect on labour force participation

Estimated log-odds ratios between different levels of education are shown in figure 5.10[1]. The full set of contrasts, with standard errors, is given in table A3.9. All contrasts are significant, except between educational levels 2 and 3.

Education is seen to have a clear effect on male LFP in terms of log-odds ratios. Since, however, men in general have very high LFP rates, the differences are not very large when measured in percentage points. Again, there is one important exception: unmarried men with low education typically fall at least 10 percentage points behind.

Figure 5.10. The effect of education on male LFP



[1] The diagram shows log-odds ratios calculated at the age of 35. Since in the model education interacts with marital status, which in turn interacts with age, the relative position of the three marital groups would be different at another age. However, the differences between educational levels are constant within each marital category.

Effect on hours of work

Our linear regression results on the effect of education on hours worked are summarized in figure 5.11.

Among married and previously married men, those with a medium level education work the longest week. The differences with respect to levels 1 and 4-5 are small, but statistically significant, at least at the 10 per cent level.

There is a curious tendency for men with 13-14 years education (level 4) to work somewhat shorter weeks (on the average) than those at levels 1 or 5. These differences are just barely significant at the 10 per cent level.

Among unmarried men, those with the lowest education work the longest week. However their LFP, and hence their overall supply, is low. The total supply of hours from unmarried men with more than the compulsory amount of schooling is comparable to the supply generated by previously married men. On the average, both groups supply about 2 hours less per week than married men with a similar level of education.

Figure 5.11. Estimated weekly hours of work, by level of education. Mean values calculated for 35-year-old men with different types of marital status

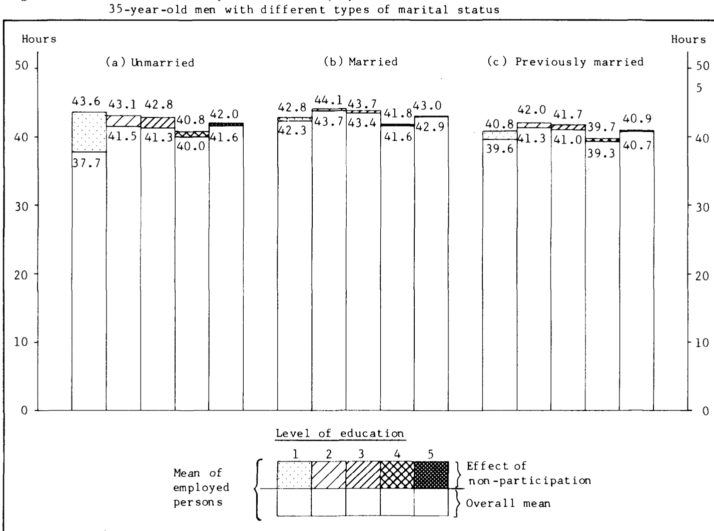
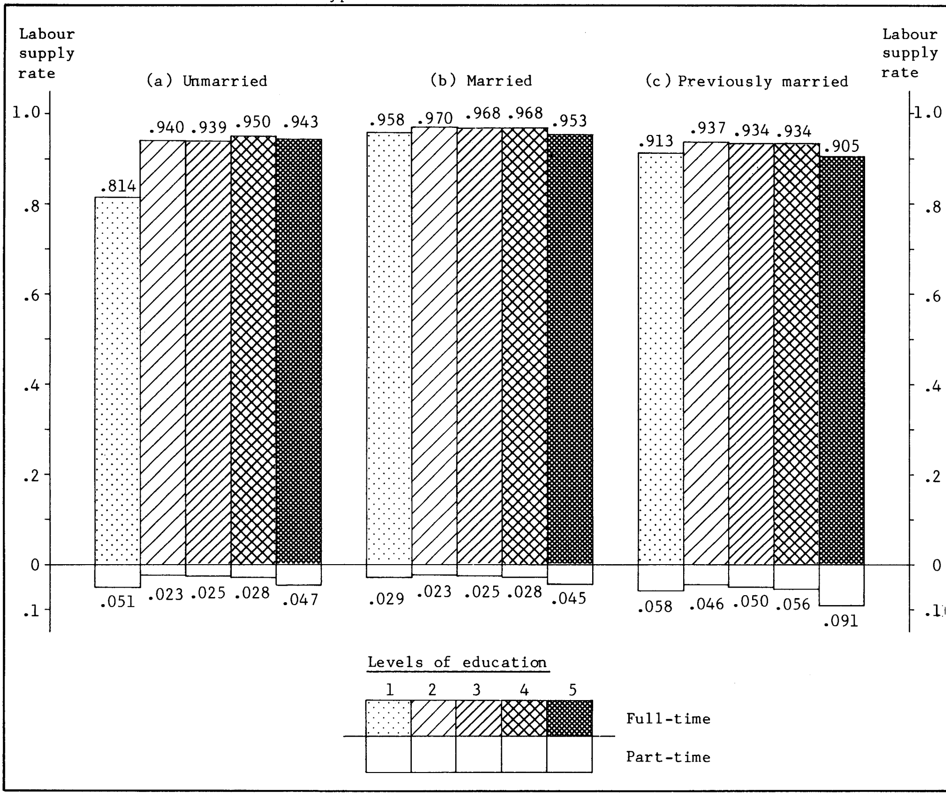


Figure 5.12. Estimated full-time and part-time probabilities, by level of education. 35-year-old men with different types of marital status



Effect on part-time versus full-time work

The effect of education on men's "preference" between part-time and full-time work is significant only in certain cases. The relevant log-odds ratios are given i table A3.10 of appendix 3. Figure 5.12 depicts male part-time and full-time supply rates by marital status and education.

Among unmarried men, the propensity to work part-time is significantly higher at the lowest and highest levels of education than at levels 2 to 4. Other contrasts are insignificant. Among married and previously married men, only those with the highest level of education are significantly more inclined to work part-time. In general the differences are small in terms of both log-odds and probabilities.

Interestingly, with respect to the labour supply of men at the top two levels of education, figures 5.11 and 5.12 are in apparent contradiction. Although the part-time rates are higher for men at the very highest level (15+ years), these men have a longer average working week than those at the second highest level (13-14 years). This probably means that although a comparatively large fraction of the best educated men work part-time, there is another contingent who work unusually long weeks. It may seem as if men at the highest level of education are less uniform in their relation to the labour market than other males.

We are, however, not in a position to tell whether the higher part-time rate among the best educated men reflects a difference in preferences or in opportunities.

In the first case, we may be faced with an example of the "backward-bending supply curve", since wages are strongly correlated with education. However, an equally plausible explanation could be that men with a higher education are generally less influenced by the traditional sex role pattern. They take a more equal part with their wives in housework, and spend more time with their children (Central Bureau of Statistics 1977: table 20, and NOS B 378 The Time Budget Survey 1980-81: tables 8 and 9). Also, university educated men tend to have wives with a similar education (Møglestue 1975:49). Such wives are almost on a par with their husbands in regard to labour force participation.

It may, however, also be the case that university graduates are simply in a better bargaining position vis-à-vis their employers. This could be all the more important as white-collar jobs may lend themselves more easily to part-time employment.

6. SUMMARY AND CONCLUSIONS

In Norway as of 1977, the single most important barrier to female labour force participation was the presence of young children. On the average the advent of the first child reduces the mother's supply of labour by approximately 50 per cent over the subsequent two to three years. A large number of women leave the labour force, and those who remain in, reduce the length of their working week by an average of 5 hours. As the child reaches school age, most mothers have returned to the labour force although they continue to put in about 5 hours less than do comparable women without children.

If there are more than one child in the family, the mother's supply of labour is reduced further. However, the age of the youngest child matters more than the number. This is in spite of the fact that the age of the youngest child has a comparatively small effect on the length of the working week. Almost all of the reduction in the mothers' labour supply comes through reduced participation. The number of children under 16, on the other hand, affects both labour force participation and the length of the mothers' working week, albeit to a lesser extent.

The presence of children in the family is generally thought to alter women's optimal trade-off between non-market time and money, whereby a reduced supply of labour to the market results. In addition, there is probably a fixed time and money cost of labour market entry to be compensated, related to the need for external childcare arrangements.

Thus, the availability of nearby kindergartens is found to have a modest, but statistically significant impact on the labour force participation of women with pre-school children.

The effect of children on female labour supply, although large on the average, is quite variable between women at different levels of education. Few university educated women withdraw from the labour force as a result of child-bearing, while this is all the more common among women with only compulsory education.

In general, women's level of education has a large bearing on the size of their labour supply. Women with a university level education supply on the average 9-12 hours more per week than those with only compulsory schooling. Most of this difference is due to variations in labour force participation rather than in hours worked by those employed.

Men at all levels of education work more than women. Male labour supply is, generally speaking, only marginally affected by the length of education. A curious exception occurs for unmarried men with only compulsory schooling. These have significantly lower participation rates than the rest of the male population. Possibly there is a social selection mechanism at work, producing an overrepresentation of less resourceful men among the unmarried as well as among those lacking higher education. This hypothesis is to some extent corroborated by the fact that even among unmarried women, the labour supply is, comparatively speaking, unusually low among those at the lowest level of education.

In general, however, marital status works in opposite directions between men and women. Married women supply less labour than do the previously married, who in turn supply less than the unmarried. Among men, the converse is true, although here the differences are small.

For some reason, marital status has a particularly big impact on women at the medium level of education (10-12 years). Married women in this educational category have remarkably low LFP rates. Several possible explanations exist. In this study, it has been demonstrated that the wage levels of husband and wife have an interesting pattern of variation between wives with different levels of education. This pattern of variation may be expected to generate a set of income and substitution effects on female labour supply which happens to be fully consistent with the observed LFP differentials.

At the medium level of education, marital status affects female labour supply both through the rate of participation and through the length of the working week. At the lowest and highest levels of education, however, female participation rates vary very little with marital status, while the preference for part-time work is significantly more pronounced among married women.

Age is the single most important determinant of male labour supply, in the sense that almost all groups of men supply an average of about 40 hours per week up until the age of 60, when male labour supply starts to drop sharply. Becoming eligible for old-age pension (at 67) has the general effect of reducing men's labour supply by an estimated 10 hours per week. Quite a few men are, however, economically active into their seventies. Among women this is very uncommon.

When controlling for marital status, education, and children, hours worked by women in the labour force vary relatively little with age. However, participation rates do, being highest in the age bracket 30-40, and quite low above the age of 60.

Now, it is impossible to tell, on the basis of our data set only, to what extent the observed age profile of labour supply reflects the ("pure") effect of ageing or simply the differing life styles of different generations.

From a methodological point of view it is of some interest to note that while certain exogenous factors seem to affect only labour force participation, others work through their association with the length of the working week. Within the traditional utility-maximizing framework, this amounts to saying that there are certain fixed costs of market work. A model intended to explain both LFP and hours worked should not be tied down by the constraint that both response variables be governed by the same linear combination of factors.

A second argument in favour of an unconstrained model is the existence of job market rationing, as evidenced by today's (1984) high unemployment figures. If the labour market is rationed, one might expect that certain parts of the actual labour supply fail to manifest themselves in the employment or unemployment statistics, because some people give up job-seeking, or fail to try. Using the local rate of unemployment as an explanatory factor of labour force participation, we are able to estimate the relationship between LFP and labour market tightness. The strength of this relationship, although probably overestimated by our model, is such as to indicate the existence of a large hidden unemployment.

Defining and measuring the supply of labour is, generally speaking, open to numerous pitfalls and difficulties. Apart from the fact that some people refrain from job-seeking for lack of suitable opportunities, many members of the labour force would probably also have chosen to put in a different amount of labour, had the market not been subject to a number of institutional rigidities. These conceptual problems are only partially circumvented through the method adopted in this study. The estimates derived might still be sensitive to certain kinds of institutional change.

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Appendix 1

Description of the data set

A1.1. Scope of sample

The data set used in this study is a subset of the LFSS sample for 1977.

The Norwegian LFSS are conducted quarterly. Each quarter about 6 000 households are drawn at random, according to a two-stage sampling scheme (see Thomsen 1977). In each selected household all persons between 16 and 74 years of age are asked to be interviewed. The total number of respondents thus amounts to approximately 10 000 per quarter.

The LFSS sample is a rotating panel. Each household participates four times over a 6-quarter period, with three quarters elapsing between the second and the third interview (i.e., interviews are taken in, say, quarters 1, 2, 5, and 6). Thus, about 7 out of 8 interviews taken in 1977 concern respondents that participate twice during that year. In our data set, however, no respondent appears more than once. To be specific, our sample has been made to consist (i) of persons born on days 1 to 15 and interviewed in the 1st or 3rd quarter, and (ii) of persons born on days 16 to 31 and interviewed in the 2nd or 4th quarter. In this way we are left with just about half the entire LFSS sample from 1977.

In the LFSS, respondents working less than 21 hours for pay or profit during the survey week, without being classified as temporarily absent from work, are asked about their principal activity during the survey week. Persons stating "school or university attendance" or "compulsory military or civil service" as their principal activity have been excluded from our data set. This fact should be kept in mind particularly when interpreting the effect of the age variable.

After the exclusion of students and conscripts, our data set consists of 9 176 women and 8 922 men.

A1.2. Variable definitions

The variables used in this study are listed in table 3.1, which is largely self-explanatory. A few clarifications are, however, in order.

The AGE variable is measured in whole years as of December 31st, 1977.

Educational categories are based on the Norwegian Standard Classification of Education (Central Bureau of Statistics 1973). The lowest level (EDUC1) comprises persons with only compulsory schooling, or less[1].

[1] Questions about educational attainment are sensitive. The sample contains about 2 per cent missing values on this variable. Rather than removing these respondents from the sample, we have assumed that they belong in the lowest educational group (EDUC1). This assumption seems consistent with the experience gathered by the LFSS interviewers (Ellingsæter 1981: 17-18). There seems to be consensus among the interviewers that the reason why some people are reluctant to answer is their embarrassment about their own (low) level of education.

"Married" in this study means, in principle, "married or cohabiting without marriage". Since, however, as of 1977 this was not explicitly spelt out in the LFSS questionnaire, but only in the interviewer's guide, there is a risk that some respondents living in cohabitation without marriage may have been classified as "unmarried". "Previously married" includes widow(er)s as well as separated and divorced persons.

The Norwegian LFSS provide information on the number and age of children under 16 who belong to the household. The questionnaire does not, however, enable us to unravel the family relations between the respective household members. Hence, prior to this study a rather elaborate recoding procedure was embarked on to single out the (probable) mother of the children present in a given household. Age and marital status were the main criteria used to choose between different female household members. In most households there is only one married woman, and usally only one has the appropriate age compared to the youngest child present. The principles of the recoding procedure are explained in further detail in the next section.

Thus, "child(ren)" in this study means, in principle, "child(ren) born (or adopted) by the respondent". For the male sample the information on children in the household has not been exploited.

The variables UNEMP and KINDG are the only ones taken from sources other than the LFSS queationnaire. The local rate of unemployment is calculated on the basis of statistics provided by the Directorate of Labour and refers to unemployment registered at the local Employment and Seamen's Offices (see section 4.3 for detailed definitions). The local rate of unemployment has been linked to each respondent's record using the code for his/her resident municipality, which is available from the LFSS data file.

In the same way, the kindergarten coverage in the respondent's residence area has been added to the variable list. The kindergarten statistics are taken from the seres is "Nye distriktstall", published by the Central Bureau of Statistics. Kindergarten statistics are compiled as of the 15th of December every year. For respondents interviewed in the 1st or 2nd quarter of 1977, the KINDG variable refers to the situation as of December 1976, while respondents interviewed in the 3rd or 4th quarter have been assigned KINDG values applicable in December 1977. For further details, see section 4.6.

In table A1.1, certain summary statistical measures describing the data set are presented (e.g., sample means of all variables etc.).

Table A1.1. Summary sample statistics

Variable name	Female sample Mean	Standard deviation	Min. value	Max. value	Male sample Mean	Standard deviation	Min. value	Max. value
LFP	.5455	.4979	0	1	.8596	.3473	0	1
FTW	.2821	.4500	0	1	.7480	.4341	0	1
PTW	.2096	.4068	0	1	.0488	.2152	0	1
T	14.25	17.46	0	99	32.38	20.48	0	99
AGE	45.14	15.99	16	74	45.14	15.88	16	74
RETAGE	.1073	.3094	0	1	.1054	.3075	0	1
EDUC1	.4475	.4972	0	1	.3793	.4852	0	1
EDUC2	.3679	.4822	0	1	.2555	.4361	0	1
EDUC3	.0921	.2889	0	1	.2291	.4201	0	1
EDUC4	.0685	.2525	0	1	.0652	.2468	0	1
EDUC5	.0240	.1531	0	1	.0708	.2564	0	1
MAR1	.1598	.3663	0	1	.2135	.4096	0	1
MAR2	.7136	.4521	0	1	.7447	.4360	0	1
MAR3	.1267	.3324	0	1	.0418	.2001	0	1
NUMCH0	.5942	.4910	0	1	..	..	..	..
NUMCH1	.1573	.3640	0	1	..	..	..	..
NUMCH2	.1602	.3667	0	1	..	..	..	..
NUMCH3	.0883	.2835	0	1	..	..	..	..
AGECH1	.0980	.2972	0	1	..	..	..	..
AGECH2	.1137	.3173	0	1	..	..	..	..
AGECH3	.1940	.3954	0	1	..	..	..	..
UNEMP	.8544	.6238	0	4.4	.8710	.6437	0	4.4
KINDG	.1175	.1003	0	.372	..	..	..	..

A1.3. On the definition of the child status variables

As mentioned above, the LFSS questionnaire does not provide information about the number and age of children born (or adopted) by the respondent. What we do get is (i) the number of children under 16 belonging to the household, and (ii) the age of the youngest of these children.

This means, e.g., that in a household consisting of mother, father, a daughter of 18, and a son of 14, all three family members above 16 will receive identical codes as far as child status is concerned. The children present in a given respondent's household may be her offspring, her adopted children, her (great-) grandchildren, her siblings, or completely unrelated.

To sort out the mother of the children present in a given household the following procedure was used:

(i) In households with one, and only one, woman of the appropriate age (defined as not less than 16 and not more than 45 years older than the youngest child), she is designated as the mother of all children present.

(ii) In households with more than one woman of the appropriate age, married women are "preferred" to previously married women, who in turn are "preferred" to unmarried women.

In the great majority of households, the above rules uniquely determines one of the household members as the mother. However,

(iii) in all households containing more than one woman (irrespective of marital status) with the appropriate age, the record of every household member was inspected manually. A final decision about the probable family relationship between the household members was made based on information about the respondents' sex, age and marital status, the number of children present, and the age of the youngest child.

In some cases it became clear that, although several respondents had been coded using the same household number, they considered themselves as belonging to two different households, having stated different answers to the questions on child status. In the great majority of those cases where both households included children, the smaller one was found to consist simply of one unmarried woman under 25 and one child under 7. In these cases two women within the "double" household were pointed out as having children.

In cases where all respondents under a given household number had stated equal answers to the child status question, it was impossible to designate more than one woman as the mother of all children present, although this may have produced some erroneous results.

In general, it is clear that the decision reached in each particular case is only our best guess as to the true identity of the children's mother.

As a crude check on the validity of this recoding procedure, the obtained child status codes were cross-tabulated against marital status and compared to the official family statistics for 1977. The data are given in tables A1.2 and A1.3.

Table A1.2. Sample cross-tabulation between marital status and child status. Women not in school with children aged 0-15. Per cent[1]

Number of children 0-15 years and age of youngest child	All women with children (1)	Of which: Married (2)	Of which: Unmarried or previously married (3)
Total	100.0	91.3	8.7
1 child 0-15 years	38.8	33.6	5.2
Aged 0-2 "	8.8	7.8	1.0
" 3-6 "	7.0	5.8	1.1
" 7-15 "	23.0	20.0	3.1
2 children 0-15 years	39.5	37.0	2.5
Youngest 0-2 "	9.6	9.2	0.5
" 3-6 "	12.1	11.5	0.6
" 7-15 "	17.8	16.3	1.4
3+ children 0-15 years	21.9	20.7	1.0
Youngest 0-2 "	5.7	5.5	0.3
" 3-6 "	9.0	8.8	0.2
" 7-15 "	7.1	6.5	0.6

[1] 100 per cent = 3 764 women.

Table A1.3. Families with children aged 0-16, by family type, number of children, and age of youngest child. Per cent[1]. 1977

Number of children 0-16 years and age of youngest child	All families with children (1)	Of which: Married couples (2)	Of which: Single mother or father (3)
Total	100.0	88.5	11.5
1 child 0-16 years	40.4	33.0	7.4
Aged 0-2 "	11.2	9.1	2.1
" 3-6 "	7.0	5.3	1.8
" 7-16 "	22.1	18.6	3.5
2 children 0-16 years	38.5	35.7	2.9
Youngest 0-2 "	10.1	9.6	0.4
" 3-6 "	11.9	11.1	0.8
" 7-16 "	16.6	14.9	1.7
3 children 0-16 years	21.1	19.9	1.2
Youngest 0-2 "	5.4	5.2	0.2
" 3-6 "	8.3	7.9	0.4
" 7-16 "	7.5	6.8	0.6

[1] 100 per cent = 559 699 families.

Source: NOS A 951 Family Statistics 1977, table 14.

The sample cross-tabulation cannot be expected to coincide exactly with the corresponding family statistics, for six reasons: (i) The families include a few single fathers with custody of children. (ii) In the family statistics, women cohabiting without marriage are counted as unmarried. (iii) Women whose principal activity during the survey week was school attendance have been excluded from our sample. (iv) In the family statistics age is measured as completed years as of July 1st 1977. In the LFSS sample age is measured as of December 31st (no matter when the interview was taken). (v) The family statistics cover families with children under 17. In the LFSS sample the age limit is 16. (vi) The LFSS figures are subject to sampling error.

Based on (i) and (ii) one would expect, other things being equal, column (3) of table A1.3 to exceed that of table A1.2. It does.

Condition (iii) probably leads to a slight underrepresentation, in our sample, of women with only one child (assuming that student mothers have on the average fewer children than other mothers). This hypothesis, too, is confirmed by the figures.

Condition (iv) means that the group "youngest child 0-2 years" should be smaller as defined in the LFSS sample than in the family statistics.[1] It is.

It is apparent from the tables that our recoding procedure produces too many mothers with two or more children, while underestimating the number of one-child mothers. This is as suspected. In view of (iii), however, the error appears to affect only a very limited number of respondents.

The split between married and unmarried or previously married mothers in table A1.2 seems entirely plausible in view of the family statistics and qualifications (i) an (ii) above.

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[1] As of midyear 1977, "0-2 years" according to the LFSS means "born between January 1st, 1975 and July 1st, 1977". According to the family statistics the time interval covered is 20 per cent longer: from July 2nd, 1974 to July 1st, 1977.

Appendix 2

Tests for interaction

A2.1 General methodology

There are two standard ways of testing linear hypotheses in the binomial logit model. One method is based on the estimated asymptotic variance-covariance matrix (Σ, say) of the maximum likelihood estimator. The second method is to use the familiar likelihood ratio statistic (McFadden 1974).

Denote by $\underset{\sim}{\beta}$ the (mxl) vector of parameters. Any set of ν simultaneous linear independent hypotheses can be written as

$$\text{(A2.1)} \quad H_0 \; : K\underset{\sim}{\beta} = \underset{\sim}{\kappa},$$

where K denotes a (νxm) constant matrix of rank ν, while $\underset{\sim}{\kappa}$ is a (νxl) constant vector. Under H_0 the statistic

$$\text{(A2.2)} \quad W(H_0 | H_1) = (K\underset{\sim}{\beta}-\underset{\sim}{\kappa})'[K\Sigma K']^{-1}.(K\underset{\sim}{\beta}-\underset{\sim}{\kappa})$$

converges to a chi-square distribution with ν degrees of freedom. Thus, an asymptotic level α test of H_0 versus the alternative

$$\text{(A2.3)} \quad H_1 \; : K\underset{\sim}{\beta} \neq \underset{\sim}{\kappa}$$

is obtained by rejecting H_0 whenever the statistic (A2.2) exceeds the ($1-\alpha$) quantile of the chi-square distribution with ν degrees of freedom (Fridstrøm 1980:113 f.). This test is known as the Wald test.

The test statistic (A2.2) is asymptotically equivalent to the likelihood ratio statistic

$$\text{(A2.4)} \quad -2[\log L(H_0) - \log L(H_1)] \quad ,$$

where $L(H_j)$ denotes the likelihood calculated under $H_j (j=0,1)$. However, the Wald statistic (A2.2) is often less expensive to compute, since it does not require another round of maximum likelihood estimation.

A2.2. Female sample tests

Child status and marital status

Table A2.1 exhibits the results of our preliminary test for interaction between marital status and child status. Under the null hypothesis this interaction is zero, i.e. (cf. mouel H_1)

(A2.5) $H_0 : \beta_{10} = \beta_{11} = \beta_{12} = \beta_{13} = 0$

The null hypothesis can be written alternatively as in equation (A2.1), where, in this case,

(A2.6) $$K = \begin{bmatrix} 0&0&0&0&0&0&0&0&1&0&0&0 \\ 0&0&0&0&0&0&0&0&0&1&0&0 \\ 0&0&0&0&0&0&0&0&0&0&1&0 \\ 0&0&0&0&0&0&0&0&0&0&0&0 \end{bmatrix}$$

and $\underset{\sim}{\kappa}$ is a vector of zeros.

The Wald statistic (A2.2) and the likelihood ratio statistic (A2.4) are both asymptotically chi-square distributed with (in this case) 4 degrees of freedom.

For this first test both statistics were computed and were found to yield practically identical results. The likelihood ratio test yields a chi-square of 2.92, as against 2.96 for the Wald test. The significance probability (p) exceeds 50 per cent, i.e. the null hypothesis is far from being rejected. In other words, the effect of children on female LFP is found not to differ significantly between married women and others.

Table A2.1. Female sample test for interaction between marital status and child status

g	Independent variable x_{ig}	Model H_1 Parameter estimate $\hat{\beta}_g$	Model H_1 Standard error $\hat{\sigma}_g$	Model H_0 Parameter estimate $\hat{\beta}_g$	Model H_0 Standard error $\hat{\sigma}_g$
1	CONSTANT	0.3257	0.2641	0.3132	0.2634
2	AGE/100	-14.0457**	1.4070	-13.8999**	1.3974
3	$(AGE/100)^2$	21.9669**	1.6362	21.8058**	1.6260
4	RETAGE	0.8131**	0.1463	0.8106**	0.1462
5	MAR2	0.4953**	0.0687	0.4719**	0.0640
6	NUMCH2	0.7989**	0.2453	0.5247**	0.0731
7	NUMCH3	0.6073	0.3713	0.7836**	0.0912
8	AGECH1	1.1988**	0.2794	1.2719**	0.0891
9	AGECH2	1.0484**	0.2668	0.8514**	0.0832
10	MAR2•AGECH1	0.0727	0.2889		
11	MAR2•AGECH2	-0.2163	0.2780		
12	MAR2•NUMCH2	-0.2913	0.2538		
13	MAR2•NUMCH3	0.1879	0.3799		
	$\log L(H_j)$	-5382.57		-5384.03	

$-2[\log L(H_0) - \log L(H_1)] = 2.92$, $p = 0.571$

$W(H_0 \mid H_1) = 2.96$, $p = 0.564$

$(H_0: \beta_{10} = \beta_{11} = \beta_{12} = \beta_{13} = 0)$

Age and marital status

Testing results for the interaction between age and marital status are shown in table A2.2. The interaction is insignificant at all levels below 25 per cent. H_0 is not rejected.

Table A2.2. Female sample test for interaction between age and marital status

g	Independent variable x_{ig}	Model H_1: Parameter estimate $\hat{\beta}_g$	Model H_1: Standard error $\hat{\sigma}_g$
1	CONSTANT	0.1177	0.3751
2	AGE/100	-10.8016**	2.0734
3	$(AGE/100)^2$	18.0248**	2.4461
4	RETAGE	0.8962**	0.1561
5	EDUC2+EDUC3	-0.4721**	0.0506
6	EDUC4	-1.4550**	0.1093
7	EDUC5	-1.9246**	0.2066
8	MAR2	1.2716*	0.5109
9	MAR2•AGE/100	-3.7225	2.5404
10	MAR2•$(AGE/100)^2$	3.8058	2.8604
11	NUMCH2	0.5370**	0.0755
12	NUMCH3	0.7801**	0.0940
13	AGECH1	1.4024**	0.1011
14	AGECH2	0.8954**	0.0896
	$\log L(H_1)$	-5227.36	

$W(H_0|H_1) = 2.76$, $p = 0.252$

$(H_0 : \beta_9 = \beta_{10} = 0)$

Marital status and education

As for the possible interaction between marital status and education, two hypotheses were tested. H_1 assumes no interacton between marital status and the dummies for university level education. H_0 assumes, in addition, that the interaction terms involving medium level education are also zero.

As shown in table A2.3, H_1 cannot be rejected. When testing H_0 against H_1, however, H_0 is rejected at the 10 per cent level but not at the 5 per cent level (significance probability = 6.4 per cent).

Table A2.3. Female sample test for interaction between marital status and education

g	Independent variable x_{ig}	Model H_2 Parameter estimate $\hat{\beta}_g$	Model H_2 Standard error $\hat{\sigma}_g$	Model H_1 Parameter estimate $\hat{\beta}_g$	Model H_1 Standard error $\hat{\sigma}_g$
1	CONSTANT	0.9734**	0.2834	0.9648**	0.2833
2	AGE/100	-13.5465**	1.4607	-13.6038**	1.4616
3	(AGE/100)2	20.8153**	1.6730	20.8768**	1.6742
4	RETAGE	0.8670**	0.1482	0.8638**	0.1485
5	EDUC2+EDUC3	-1.1799**	0.1448	-1.1611**	0.1419
6	EDUC4	-1.5961**	0.3025	-1.4323**	0.1092
7	EDUC5	-2.0582**	0.3525	-1.8937**	0.2065
8	MAR2	0.1332	0.1193	0.1545	0.1125
9	MAR3	-0.1342	0.1455	-0.0975	0.1384
10	MAR2(EDUC2+EDUC3)	0.8477**	0.1552	0.8286**	0.1511
11	MAR3(EDUC2+EDUC3)	0.4911*	0.2091	0.4567*	0.2054
12	MAR2(EDUC4+EDUC5)	0.1651	0.3196		
13	MAR3(EDUC4+EDUC5)	0.3797	0.4530		
14	NUMCH2	0.5280**	0.0750	0.5282**	0.0750
15	NUMCH3	0.7684**	0.0933	0.7687**	0.0934
16	AGECH1	1.4145**	0.0949	1.4136**	0.0948
17	AGECH2	0.8951**	0.0864	0.8946**	0.0864
log $L(H_j)$		-5211.02		-5211.3	

$-2[\log L(H_1)-\log L(H_2)] = 0.56, p = 0.756$ $(H_1:\beta_{12}=\beta_{13}=0)$

$W(H_0|H_1)=5.50,\ p=0.064$ $(H_0:\beta_{10}=\beta_{11}=\beta_{12}=\beta_{13}=0)$

Number of children and age of youngest child

When testing for interaction between the number of children and the age of the youngest child, the null hypothesis is rejected at all levels higher than 6 per cent (p = 0.055).

Upon inspecting table A2.4 it becomes clear that, although four interaction dummies are included in the model, one would have been sufficient, since the differences between the four terms are all insignificant. Thus, a model with only one interaction term (given as (AGECH1+AGECH2)•(NUMCH2+NUMCH3)) would have given almost as good a fit as model H_1, and we would be able to reject H_0 at a very low level.

Table A2.4. Female sample test for interaction between the number of children and the age of the youngest child

g	Independent variable x_{ig}	Model H_1 Parameter estimate $\hat{\beta}_g$	Model H_1 Standard error $\hat{\sigma}_g$	Model H_0 Parameter estimate $\hat{\beta}_g$	Model H_0 Standard error $\hat{\sigma}_g$
1	CONSTANT	-0.4898*	0.2720	-0.3763	0.2694
2	AGE/100	14.9277**	1.4833	14.4297**	1.4716
3	$(AGE/100)^2$	-22.9929**	1.7487	-22.5151**	1.7385
4	RETAGE	-0.7736**	0.1492	-0.7555**	0.1492
5	MAR2	-0.4723**	0.0647	-0.4629**	0.0645
6	NUMCH1	-0.2178*	0.0909	-0.0967	0.0825
7	NUMCH2	-0.4795**	0.1029	-0.5842**	0.0890
8	NUMCH3	-0.7504**	0.1429	-0.8482**	0.1065
9	AGECH1	-0.9406**	0.1457	-1.2340**	0.0947
10	AGECH2	-0.5419**	0.1564	-0.8217**	0.0868
11	AGECH1•NUMCH2	-0.4598*	0.1932		
12	AGECH1•NUMCH3	-0.4810*	0.2393		
13	AGECH2•NUMCH2	-0.4381*	0.1978		
14	AGECH2•NUMCH3	-0.3797*	0.2293		
log $L(H_j)$		-5378.73		-5383.35	

$-2[\log L(H_0)-\log L(H_1)] = 9.24$, $p = 0.055$

$(H_0: \beta_{11}=\beta_{12}=\beta_{13}=\beta_{14}=0)$

Education and child status

The interaction between level of education and child status is highly significant ($p<0.001$). This is shown in table A2.5.

Table A2.5. Female sample test for interaction between education and child status

g	Independent variable x_{ig}	Model H_1 Parameter estimate $\hat{\beta}_g$	Model H_1 Standard error $\hat{\sigma}_g$	Model H_0 Parameter estimate $\hat{\beta}_g$	Model H_0 Standard error $\hat{\sigma}_g$
1	CONSTANT	-0.0281	0.2630	-0.0372	0.2595
2	AGE/100	8.8813**	1.3127	9.1436**	1.3056
3	$(AGE/100)^2$	-15.9305**	1.5424	-16.2885**	1.5368
4	RETAGE	-0.8835**	0.1472	-0.8674**	0.1474
5	EDUC2+EDUC3	0.5497**	0.0604	0.4659**	0.0504
6	EDUC4	1.2510**	0.1389	1.4699**	0.1094
7	EDUC5	1.6721**	0.2247	1.9425**	0.2070
8	NUMCH2	-0.3029*	0.1195	-0.5925**	0.0745
9	NUMCH3	-0.5566**	0.1537	-0.8001**	0.0936
10	AGECH1	-1.8751**	0.1690	-1.6174**	0.0905
11	AGECH2	-1.2005**	0.1412	-1.0382**	0.0844
12	NUMCH2•(EDUC2+EDUC3) ..	-0.4420**	0.1489		
13	NUMCH2•(EDUC4+EDUC5) ..	-0.5832*	0.2674		
14	NUMCH3•(EDUC2+EDUC3) ..	-0.3877*	0.1919		
15	NUMCH3•(EDUC4+EDUC5) ..	-0.3847	0.3318		
16	AGECH1•(EDUC2+EDUC3) ..	0.2146	0.1951		
17	AGECH1•(EDUC4+EDUC5) ..	1.1915**	0.3037		
18	AGECH2•(EDUC2+EDUC3) ..	0.1738	0.1744		
19	AGECH2•(EDUC4+EDUC5) ..	0.8690**	0.3164		
log $L(H_j)$		-5239.79		-5254.30	

$-2[\log L(H_0)-\log L(H_1)] = 29.02$, $p = 0.0003$

$(H_0:\beta_{12}=\beta_{13}=\beta_{14}=\beta_{15}=\beta_{16}=\beta_{17}=\beta_{18}=\beta_{19}=0)$

Here, too, we note that several of the interaction parameters are comparable in size. We have $\hat{\beta}_{12} \approx \hat{\beta}_{13}$, $\hat{\beta}_{14} \approx \hat{\beta}_{15}$, $\hat{\beta}_{16} \approx \hat{\beta}_{18}$, and $\hat{\beta}_{17} \approx \hat{\beta}_{19}$. In other words, the number of children affects LFP in the same way for women having either medium or high level education (but differently for those with low education). Also, the effect of education appears to be the same whether the youngest child is 0-2 or 3-6 years old (but differentfor women with only school-age children).

In table A2.6, the log-odds ratios deducible from table A2.5 (model H_1) are compared to those shown in figure 4.9 (which is based on table 4.1). To facilitate the reading of the table, a frame has been drawn around all figures larger than zero.

The model including interaction terms between education and child status is seen to yield higher log-odds ratios for two main groups of women. One consists of women having small children (under 7) and a university level education. The other one consists of women with the opposite set of attributes: low or medium level education but no small children.

For the remaining groups of women the log-odds ratios are, with few exceptions, found to be lower when the interaction terms are including in the model than otherwise.

Consider the figures given for levels 4 and 5 in table A2.6. The difference between those having children younger than 3 and those having only school-age children equals approximately 0.8. This amounts to saying that for women with a university level education, the effect of the youngest child's age is only about half as big as estimated under the main model,cf. figures 4.9 and 4.12 (example 3).

Table A2.6. Differences between the log-odds ratios calculated from table A2.5 (H_1) and those from table 4.1

Age of youngest child	Number of children	Level of education 1	2	3	4	5
..	0	0[1]	.243	.196	-.142	-.190
7-15	1	.051	.293	.246	-.092	-.140
	2	.258	.059	.011	-.468	-.516
	3+	.235	.090	.043	-.292	-.340
3-6	1	-.190	-.323	-.371	.563	.489
	2	.017	-.008	-.056	.160	.113
	3+	-.006	.023	-.024	.336	.288
0-2	1	-.336	.122	.074	.713	.665
	2	-.129	-.113	-.161	.337	.289
	3+	-.152	-.081	-.128	.513	.465

[1] Reference group.

Unemployment and education

The interaction between the level of education and the rate of unemployment is significant at the 1 per cent level ($p = 0.0052$), see table A2.7. However, the interaction term for medium level education (β_9) is insignificant. Thus, all we can say is that the effect of unemployment on LFP is different between women with a university level education and others.

Table A2.7. Female sample test for interaction between level of education and unemployment

g	Independent variable x_{ig}	Model H_1 Parameter estimate $\hat{\beta}_g$	Model H_1 Standard error $\hat{\sigma}_g$	Model H_0 Parameter estimate $\hat{\beta}_g$	Model H_0 Standard error $\hat{\sigma}_g$
1	CONSTANT	-0.3234	0.2238	-0.3198	0.2220
2	AGE/100	5.1991**	1.1023	5.2089**	1.1019
3	$(AGE/100)^2$	-8.3721**	1.2940	-8.3755**	1.2936
4	RETAGE	-1.5530**	0.1414	-1.5474**	0.1412
5	EDUC2+EDUC3	0.4705**	0.0806	0.4114**	0.0483
6	EDUC4	0.7792**	0.1746	1.1703**	0.1050
7	EDUC5	1.2490**	0.2456	1.6511**	0.2006
8	UNEMP	-0.1225*	0.0499	-0.1304**	0.0364
9	UNEMP•(EDUC2+EDUC3) ...	-0.0716	0.0748		
10	UNEMP•(EDUC4+EDUC5) ...	0.5500**	0.2003		
log $L(H_j)$		-5617.14		-5622.39	

$-2[\log L(H_0)-\log L(H_1)] = 10.5$, $p = 0.0052$
$(H_0:\beta_9=\beta_{10}=0)$

A2.3 Male sample tests

Age and marital status

The male sample test for interaction between age and marital status rejects the null hypothesis (of zero interaction) at every reasonable level of significance ($p < 0.00005$). This is shown in table A2.8.

Note, however, that the interaction terms are not significantly different between married and previously married, thus only unmarried men stand out as having a significantly different age profile.

Table A2.8. Male sample test for interaction between age and marital status

g	Independent variable x_{ig}	Model H_1 Parameter estimate $\hat{\beta}_g$	Model H_1 Standard error $\hat{\sigma}_g$	Model H_0 Parameter estimate $\hat{\beta}_g$	Model H_0 Standard error $\hat{\sigma}_g$
1	CONSTANT	3.1712**	0.5582	0.8319*	0.4543
2	AGE/100	-1.6746	2.9661	14.0750**	2.2629
3	$(AGE/100)^2$	-4.1833	3.4142	-24.0070**	2.5116
4	RETAGE	-0.6984**	0.1548	-0.7761**	0.1481
5	EDUC2	0.7510**	0.1022	0.7236**	0.1017
6	EDUC3	0.7777**	0.1125	0.7505**	0.1113
7	EDUC4	1.2515**	0.2471	1.2212**	0.2443
8	EDUC5	1.9607**	0.2646	1.8934**	0.2609
9	MAR2	-0.4256	1.2751	1.0139**	0.1120
10	MAR3	-2.4807	2.4213	0.1971	0.1832
11	(AGE/100)•MAR2	15.2836**	5.1656		
12	(AGE/100)•MAR3	14.4143	9.6064		
13	$(AGE/100)^2$•MAR2	-21.5849**	5.0825		
14	$(AGE/100)^2$•MAR3	-16.9266*	9.1874		
15	UNEMP	-0.2352**	0.0566	-0.2475**	0.0562
	log $L(H_j)$	-2183.61		-2219.99	

$-2[\log L(H_1)-\log(H_2)] = 72.8$, $p = 0.0000$
$(H_0:\beta_{11}=\beta_{12}=\beta_{13}=\beta_{14}=0)$

Marital status and education

Table A2.9 shows estimation results for a binomial logit model of male LFP including interaction terms between marital status and education.

Most single interaction terms are insignificant. However, the likelihood ratio test rejects at the 1 per cent level the simultaneous hypothesis that all interaction terms be zero.

The interaction terms do not differ significantly between them, except maybe for one (β_{12}). Thus, one interaction term would actually have been largely sufficient, capturing the difference between unmarried men at the lowest level of education, and all others.

Table A2.9 Male sample test for interaction between marital status and education

g	Independent variable x_{ig}	Model H_1 Parameter estimate β_g	Standard error g
1	CONSTANT	0.4959	0.4577
2	AGE/100	14.6544**	2.2574
3	$(AGE/100)^2$	-24.4970**	2.5040
4	RETAGE	-0.7785**	0.1478
5	EDUC2	1.3949**	0.2070
6	EDUC3	1.4307**	0.2147
7	EDUC4	1.1026*	0.6051
8	EDUC5	3.0323**	1.1641
9	MAR2	1.2542**	0.1293
10	MAR3	0.4598*	0.2307
11	MAR2•(EDUC2 + EDUC3)	-0.8405**	0.2219
12	MAR2•EDUC4	0.2088	0.6702
13	MAR2•EDUC5	-1.2332	1.1965
14	MAR3•(EDUC2 + EDUC3)	-0.7907*	0.3796
15	MAR3•EDUC4	-0.6155	0.9407
16	MAR3•EDUC5	-1.7130	1.4893
17	UNEMP	-0.2507**	0.0565
	log $L(H_1)$	-2211.06	

$-2[\log L(H_1) - \log L(H_0)] = 17.8$, $p = 0.0068$

(H_0: $\beta_{11} = \beta_{12} = \beta_{13} = \beta_{14} = \beta_{15} = \beta_{16} = \beta_{16} = 0$)[1]

[1] For estimation results under H_0, see table A2.8.

Age and education

The interaction test for age vs. education is shown in table A2.10.

The null hypothesis of zero interaction is rejected. However, only the therms involving EDUC4 are seen to be significant. For men at levels of education 1, 2, 3, or 5, LFP appears to vary with age in practically the same way.

How big is the differnce between, e.g., levels 4 and 5? This is illustrated in figure A2.1. The relationship between age and LFP is quite similar between the two groups up to the age of 65. Above this age, however, men with 13-14 years of schooling seem to be more economically active than others.

Table A2.10. Male sample test for interaction between age and education

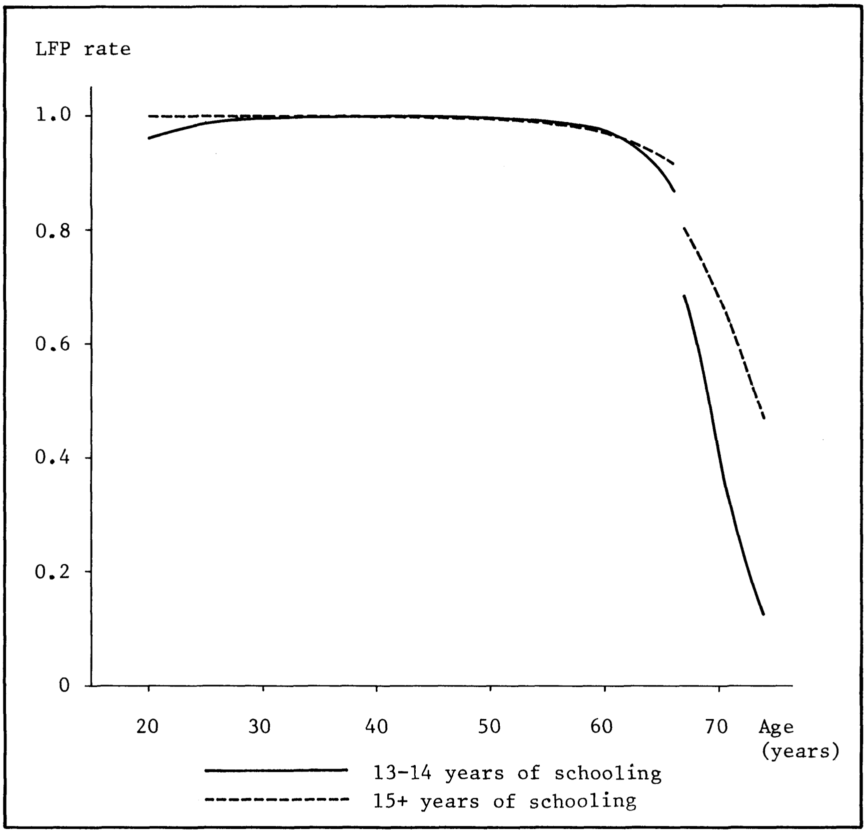
		Model H_1	
g	Independent variable x_{ig}	Parameter estimate $\hat{\beta}_g$	Standard error $\hat{\sigma}_g$
1	CONSTANT	0.6454	0.5088
2	AGE/100	12.9432**	2.4962
3	$(AGE/100)^2$	-21.5608**	2.7528
4	RETAGE	-0.7696**	0.1469
5	EDUC2	1.8039*	0.8461
6	EDUC3	1.8124*	0.8521
7	EDUC4	-6.5072*	2.5899
8	EDUC5	3.9440	4.1479
9	(AGE/100)•(EDUC2+EDUC3)	-0.0361	3.7539
10	(AGE/100)•EDUC4	40.4543**	13.3197
11	(AGE/100•EDUC5	-3.0151	11.2792
12	$(AGE/100)^2$•(EDUC2+EDUC3)	-3.0772	3.8849
13	$(AGE/100)^2$•EDUS4	-43.8928**	14.3595
14	$(AGE/100)^2$•EDUC5	-0.9080	11.4889
15	MAR2	1.0109**	0.1111
16	MAR3	0.2014	0.1825
17	UNEMP	-0.2516**	0.0559
log $L(H_1)$		-2200.70	

$-2[\log L(H_1) - \log L(H_0)] = 38.6$, $p = 0.0000$

$(H_0: \beta_9 = \beta_{10} = \beta_{11} = \beta_{12} = \beta_{13} = \beta_{14} = 0)$[1]

[1] For estimation results under H_0, see table A2.8.

Figure A2.1. The relationship between age and LFP, according to the estimates in table A2.10. Married men with 13-14, resp. 15+ years of schooling



Appendix 3

Table A3.1. The effect of education on female labour force participation. Estimated log-odds ratios of one educational category (h) with respect to another (i). Standard errors in parentheses

Educational category (i)	Educational category (h)			
	10 years	11-12 years	13-14 years	15+ years
			Unmarried	
0-9 years of schooling	1.1397**(0.1443)	1.1867**(0.1570)	1.3934**(0.1098)	1.8622**(0.2066)
10 years of schooling		0.0470 (0.0897)	0.2537 (0.1743)	0.7225* (0.2475)
11-12 years of schooling			0.2067 (0.1846)	0.6755* (0.2547)
13-14 years of schooling				0.4688* (0.2262)
			Married	
0-9 years of schooling	0.3070**(0.0591)	0.3541**(0.0936)	1.3934**(0.1098)	1.8622**(0.2066)
10 years of schooling		0.0470 (0.0897)	1.0863**(0.1120)	1.5551**(0.2073)
11-12 years of schooling			1.0393**(0.1328)	1.5081**(0.2193)
13-14 years of schooling				0.4688* (0.2262)
			Previously married	
0-9 years of schooling	0.6910**(0.1505)	0.7381**(0.1669)	1.3934**(0.1098)	1.8622**(0.2066)
10 years of schooling		0.0470 (0.0897)	0.7023**(0.1822)	1.1711**(0.2529)
11-12 years of schooling			0.6553**(0.1955)	1.1241**(0.2626)
13-14 years schooling				0.4688* (0.2262)

Legend: See table 4.1.

Table A3.2. The effect of marital status on female labour force participation. Estimated log-odds ratios of one marital category (h) with respect to another (i). Standard errors in parentheses

Marital category(i)	Marital category (h)	
	Married	Previously married
	10-12 years of schooling	
Unmarried	-0.9949**(0.1265)	-0.3548* (0.1682)
Married		0.6401**(0.1290)
	0-9 or 13+ years of schooling	
Unmarried	-0.1622 (0.1139)	0.0938 (0.1394)
Married		0.2561* (0.1081)

Legend: See table 4.1.

Table A3.3. Regression results. Married women aged 18-44 years in 1977. Dependent variable: respondent's actual (a) or expected (b) hourly wage (Norwegian kroner)

Independent variable	(a) Respondents in the labour force		(b) Respondents not in the labour force	
	Parameter estimate	Standard error	Parameter estimate	Standard error
CONSTANT	20.562	..	20.635	..
EDUC2	1.915**	0.659	0.856	0.813
EDUC3	4.962**	0.899	7.026**	1.279
EDUC4	9.377**	0.874	17.857**	1.388
EDUC5	23.060**	1.227	45.836**	2.926
Dummy for 20-24 years ...	0.139	3.434	1.956	3.233
" " 25-29 " ...	2.725	3.419	3.353	3.212
" " 30-34 " ...	4.644	3.448	6.857*	3.317
" " 35-39 " ...	6.373*	3.482	5.560	3.421
" " 40-44 " ...	6.519*	3.503	6.268*	3.552
NUMCH1	-0.472	1.043	-0.515	1.811
NUMCH2	-1.525	1.004	-3.271*	1.698
NUMCH3	-3.131**	1.121	-4.693*	1.840
AGECH1	1.968*	0.912	2.718*	1.183
AGECH2	0.751	0.737	2.412*	1.153
R^2	0.258	.	0.565	.
n	1.589	.	389	.
s	10.045	..	6.921	..

Legend: See table 4.1.

Table A3.4. Regression results. Married women aged 18-44 years in 1977. Dependent variable: husband's annual income (Norwegian kroner)

Independent variable[1]	Regression 1		Regression 2	
	Parameter estimate	Standard error	Parameter estimate	Standard error
CONSTANT	41 708	..	42 904	..
EDUC2	5 351**	1 444	4 719**	1 436
EDUC3	10 011**	2 030	8 740**	2 002
EDUC4	10 808**	2 022	9 400**	1 980
EDUC5	17 460**	2 921	15 372**	2 871
Dummy for 20-24 years ...	-488	6 993	991	7 002
" " 25-29 " ...	7 688	6 949	11 501	6 912
" " 30-34 " ...	16 359*	7 029	22 049**	6 914
" " 35-39 " ...	17 228*	7 141	23 747**	6 943
" " 40-44 " ...	18 636**	7 211	25 089**	6 977
NUMCH1	3 516	2 488		
NUMCH2	8 242**	2 401		
NUMCH3	8 039**	2 693		
AGECH1	-2 786	1 876		
AGECH2	-283	1 674		
R^2	0.099	.	0.091	.
n	2 466	.	2 466	.
s	28 017	..	28 097	..

Legend: See table 4.1.

[1] Referring to the attributes of the wife.

Table A3.5. The effect of education on women's "choice" between full-time and part-time work. Estimated log-odds ratios of one educational category (h) with respect to another (i). Standard errors in parentheses.

Educational category (i)	Educational category (h)			
	10 years	11-12 years	13-14 years	15+ years
0-9 years of schooling	0.4116**(0.0738)	0.4274**(0.1162)	0.4473**(0.1180)	0.5900**(0.1783)
10 years of schooling		0.0158 (0.1145)	0.0357 (0.1163)	0.1784 (0.1768)
11-12 years of schooling			0.0198 (0.1467)	0.1625 (0.1980)
13-14 years of schooling				0.1427 (0.1980)

Legend: See table 4.1.

Table A3.6. The effect of marital status on women's "choice" between full-time and part-time work. Estimated log-odds ratios of one marital category (h) with respect to another (i). Standard errors in parentheses.

Marital category (i)	Marital category (h)	
	Married	Previously married
Unmarried	-1.0206**(0.1244)	-0.3580* (0.1535)
Married		0.6626**(0.1086)

Legend: See table 4.1.

Table A3.7. The effect of children on female labour force participation. Estimated log-odds ratios of one family situation (h) with respect to another (i)[1]. Standard errors in parentheses

Family situation (i). Number of children and age of youngest child	Family 1 child aged 7-15	Family 1 child aged 3-6	Family 1 child aged 0-2
0 children	-0.0505	-1.0103**	-1.5391**
	(0.0071)	(0.0136)	(0.0148)
1 child			
Aged 7-15		-0.9598**	-1.4887**
		(0.0105)	(0.0128)
" 3-6			-0.5288**
			(0.0096)
" 0-2			
2 children			
Youngest aged 7-15			
" " 3-6			
" " 0-2			
3+ children			
Youngest aged 7-15			
" " 3-6			

Legend: See table 4.1.

[1] Assuming no day-care institutions available (KINDG = 0).

situation (h)					
2 children, youngest aged			3+ children, youngest aged		
7-15	3-6	0-2	7-15	3-6	0-2
-0.5605**	-1.5203**	-2.0491**	-0.7915**	-1.7513**	-2.2801**
(0.0084)	(0.0130)	(0.0154)	(0.0120)	(0.0141)	(0.0170)
-0.5100**	-1.4698**	-1.9987**	-0.7410**	-1.7008**	-2.2297**
(0.0066)	(0.0153)	(0.0188)	(0.0095)	(0.0157)	(0.0197)
0.4498**	-0.5100**	-1.0389**	0.2188	-0.7410**	-1.2699**
(0.0190)	(0.0066)	(0.0174)	(0.0244)	(0.0095)	(0.0208)
0.9786**	0.0188	-0.5100**	0.7477**	-0.2122	-0.7410**
(0.0201)	(0.0150)	(0.0066)	(0.0251)	(0.0174)	(0.0095)
	-0.9598**	-1.4887**	-0.2310*	-1.1908**	-1.7196**
	(0.0105)	(0.0128)	(0.0088)	(0.0167)	(0.0195)
		-0.5288**	0.7288**	-0.2310*	-0.7598**
		(0.0096)	(0.0218)	(0.0088)	(0.0188)
			1.2577**	0.2979*	-0.2310*
			(0.0236)	(0.0178)	(0.0088)
				-0.9598**	-1.4887**
				(0.0105)	(0.0128)
					-0.5288**
					(0.0096)

Table A3.8. The effect of children on women's "choice" between full-time and part-time work. Estimated log-odds ratios of one family situation (h) with respect to another (i). Standard errors in parentheses

Family situation (i). Number of children and age of youngest child	Family 1 child, aged 7-15	3-6	0-2
0 children	-0.7666** (0.1014)	-1.1022** (0.1379)	-0.9881** (0.1535)
1 child			
Aged 7-15		-0.3356** (0.1198)	-0.2215 (0.1427)
" 3-6			0.1141 (0.1510)
" 0-2			
2 children			
Youngest aged 7-15			
" " 3-6			
" " 0-2			
3+ children			
Youngest aged 7-15			
" " 3-6			

Legend: See table 4.1.

situation (h)					
	2 children, youngest aged			3+ children, youngest aged	
7-15	3-6	0-2	7-15	3-6	0-2
-1.2802**	-1.6158**	-1.5017**	-1.4328**	-1.7684**	-1.6543**
(0.1149)	(0.1395)	(0.1607)	(0.1468)	(0.1567)	(0.1782)
-0.5135**	-0.8491	-0.7350**	-0.6662**	-1.0018**	-0.8876**
(0.1052)	(0.1514)	(0.1754)	(0.1356)	(0.1642)	(0.1888)
-0.1780	-0.5135**	-0.3994*	-0.3306*	-0.6662**	-0.5520**
(0.1670)	(0.1052)	(0.1890)	(0.1962)	(0.1356)	(0.2094)
-0.2921*	-0.6277**	-0.5135**	-0.4447*	-0.7803**	-0.6662**
(0.1790)	(0.1790)	(0.1052)	(0.2045)	(0.1963)	(0.1356)
	-0.3356**	-0.2215	-0.1526	-0.4882**	-0.3741*
	(0.1198)	(0.1427)	(0.1361)	(0.1720)	(0.1909)
		0.1141	0.1830	-0.1526	-0.0385
		(0.1510)	(0.1901)	(0.1361)	(0.2053)
			0.0689	-0.2667	-0.1526
			(0.2033)	(0.2012)	(0.1361)
				-0.3356**	-0.2215
				(0.1198)	(0.1427)
					0.1141
					(0.1510)

Table A3.9. The effect of education on male labour force participation. Estimated log-odds ratios of one educational category (h) with respect to another (i). Standard errors in parentheses

Educational category (i)	Educational category (h)			
	10 years	11-12 years	13-14 years	15+ years
		Unmarried		
0-9 years of schooling	1.3805**(0.1860)	1.4298**(0.1950)	1.9412**(0.3013)	2.6485**(0.3154)
10 years of schooling		0.0493 (0.1332)	0.5607* (0.2605)	1.2679**(0.2736)
11-12 years of schooling			0.5114* (0.2642)	1.2187**(0.2776)
13-14 years of schooling				0.7072* (0.3558)
		Married or previously married		
0-9 years of schooling	0.5437**(0.1130)	0.5930**(0.1217)	1.1044**(0.2536)	1.8116**(0.2669)
10-14 years of schooling		(as for unmarried)		

Legend: See table 4.1.

Table A3.10. The effect of education on men's "choice" between full-time and part-time work. Estimated log-odds ratios of one educational category (h) with respect to another (i). Standard errors in parentheses

Educational category (i)	Educational category (h)							
	10 years		11-12 years		13-14 years		15+ years	
	Unmarried							
0-9 years of schooling	0.9592**	(0.2186)	0.8802**	(0.2257)	0.7593**	(0.2937)	0.2478	(0.2638)
10 years of schooling			-0.0790	(0.1589)	-0.2000	(0.2397)	-0.7114**	(0.1959)
11-12 years of schooling					-0.1210	(0.2416)	-0.6325**	(0.1981)
13-14 years of schooling							-0.5115*	(0.2651)
	Married							
0-9 years of schooling	0.2605*	(0.1523)	0.1815	(0.1557)	0.0605	(0.2357)	-0.4510*	(0.1878)
10-14 years of schooling	(as for unmarried)							

Legend: See table 4.1.

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