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ESSAYS ON DISCRIMINATION, WELFARE AND LABOR SUPPLY

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Abstract

This thesis consists of five papers in applied labor economics. The first two papers are related to wage discrimination between males and females, whereas the next two papers are related to labor supply and welfare participation, while the last paper analysis early retirement in Sweden.

"Endogenous Schooling and the Distribution of the Gender Wage Gap" studies how the unexplained wage gap is affected by treating years of education in a traditional wage equation as endogenous. The result shows that the estimated wage differentials are substantial higher using OLS than using IV and panel data. We find that the gender wage gap differs substantially across different values of work experience and that the unexplained gender wage gap has increased over time.

"Occupational Gender Composition and Wages in Sweden" analyzes the relationship between wages and occupational gender segregation in Sweden controlling for non-random selection into an occupation. The result shows that the unexplained gender wage gap is largest in female dominated occupations and smallest in male dominated occupations. Females' experience earnings profile is steeper for women in male dominated occupations. Ignoring occupational segregation produces considerable higher estimates of the unexplained part of the gender wage gap.

"Household Labor Supply and Welfare Participation in Sweden" studies the joint effects of the tax and benefit systems on household labor supply. The estimates from the structural model yielded small wage and income elasticities. A tax simulation showed that reducing the progressivity in the Swedish tax system may have considerable welfare effects. The effect on working hours from the reform was quite small, while the evaluation of a change in the welfare system showed that the stigma-effect had a substantial impact.

"Labor Supply and Welfare Participation of Single Mothers in Sweden" analyzes the effects of changes in income taxes, cost of childcare and social assistance on labor supply for single mothers households. The results show that there is a positive and significant stigma-effect associated with welfare participation. Fixed costs of working is an important factor in a single mothers' decision to enter the labor market. We find a negative covariance between social assistance and labor supply, which implies self-selection into welfare. Welfare effects from the childcare reform are quite similar across all income deciles, even if predicted increases in hours of work are substantial for the poorest single mother households.

"Early Retirement in Sweden" studies the determinants of early retirement from the Swedish labor market for both males and females. The result shows that there is heterogeneity in the underlying preference structure and that the probability of a complete early withdrawal from the labor market increases with age. Blue collar workers have lower probability to take part time pension and full early age retirement than workers from other occupational schemes. Finally, we find that economic incitements affect the decision of Swedish workers to leave the labor market.

Keywords: *Gender Wage Gap, Endogenous Schooling, Panel Data, Occupational Segregation, Labor Supply, Welfare Participation, Unobserved Heterogeneity, Tax Simulation, Early Retirement, Occupational Pension*.

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Contents

Endogenous Schooling and the Distribution of the Gender Wage Gap*

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Abstract

Previous studies on the gender wage gap have relied on OLS when estimating the wage equations. However, a number of recent studies, devoted to estimating the return to education, have shown that OLS may produce biased estimates for a number of reasons. As a consequence, previous results regarding the gender wage gap may also be biased. In this paper, we first estimate wage equations using instrumental variables procedures and panel data and then investigate the distribution of the gender wage gap. The results indicate that OLS may seriously overestimate the unexplained gender wage gap.

Key words: Return to Education, Endogenous Schooling, Panel Data, Unobserved Heterogeneity, Distribution of Wage Gap.

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1 Introduction

Estimating the return to education and the gender wage gap has been the target for labor economists for decades. In most empirical work, the wage functions are derived from the Mincer equation (Mincer (1974)), relating the logarithm of earnings (or wages) linearly to years of education. For some time, labor economists have argued that estimating this relationship by ordinary least squares (OLS) produces biased estimates for at least three reasons. First, the OLS estimates may suffer from "ability bias". Specifically, if schooling decisions are positively correlated with ability and ability only enters the wage equation through an additive error term, then it can be shown that OLS yields estimates that are upward biased.[1] Second, if the measure of educational attainment (usually years of education) is measured with error, this will produce a negative bias in the OLS estimates of the return to education (see Griliches (1977)). Third, schooling attainment can be thought of as a function of ability and subjective discount rates. If these two arguments are correlated (presumably negative), this will also bias the OLS estimates (see Card (1995) and Belzil and Hansen (2000)).

Despite the recognitions of the inability of OLS to obtain unbiased and consistent estimates in wage regressions, it has been the principal estimator in studies regarding the gender wage gap. This is surprising. Arguably, the bias in OLS will not only affect the wage parameters, but also the wage gap itself. The strong reliance on OLS in the previous literature motivates us to study the effect of treating schooling as an endogenous variable on the unexplained gender wage gap.[2]

Consequently, our first problem is how to obtain unbiased or consistent estimates of the wage regressions.[3] One way to reduce the problem of ability bias is to include proxy variables for labor market ability in the wage equation. Typical variables that have served as proxies for ability in previous literature include scores from IQ-tests (e.g. Blackburn and Neumark (1993) and Griliches (1977)) and family background information (e.g. Blackburn and Neumark (1995) and Lam and Schoeni (1993)). A common result from these studies is that when proxy variables are included in the wage equation, the estimated return to education fall. However, even if these procedures may reduce the ability bias, we still face problems with bias due to measurement error and subjective discounting. To overcome these hurdles we may use instrumental variable techniques instead. The problem then reduces to finding suitable instruments. One approach is to use exogenous influences that affect schooling attainment. For

[1] However, as argued by Griliches (1977), individuals with high ability may have higher opportunity costs associated with schooling. If this is the case, schooling decisions are negatively correlated with ability and OLS estimates are *downward* biased.

[2] It should be noted that other regressors, such as work experience, might also be endogenous.

[3] Card (1999) provides an excellent survey of methods and results in the literature on the return to schooling.

instance, information on changes in the minimum school leaving age has been used (see Harmon and Walker (1995)). Card (1993) used information on proximity to colleges as an instrument, with the motivation that the cost of attending college is lower if there is a college close to where the student resides. The results from these two studies indicate that OLS substantially underestimate the true return to education. Another branch of the literature has relied on using data on twins or siblings, combined with a first-difference estimator, as a way of reducing the problem with ability bias. For example, Ashenfelter and Kreuger (1994) used data on identical twins and reported first-difference estimates of the marginal return to schooling in the order of 12-16 percent. These estimates were substantially higher than the corresponding OLS estimates.

When panel data is available, we do not have to rely on the cross-sectional approaches outlined above. First, we could assume that ability is an individual specific, time-invariant effect and estimate the parameters with a fixed-effect estimator. This would produce unbiased estimates of the return to education and is the approach taken by Angrist and Newey (1991). They used data from the NLSY and report fixed-effects estimates which are roughly twice as large as OLS estimates. However, the fixed-effects estimates can only be identified from individuals that return to school.[4] Second, since it is difficult to find suitable instruments, we could adopt a procedure proposed by Hausman and Taylor (1981) (hereafter HT). They derived a two-stage least-squares random effects estimator, in which identification is achieved through the use of individual means over time of the included time-varying regressors in the primary equation. The main advantage of this procedure is that it *does not require any exclusion restrictions* in order to identify the parameters, as opposed to cross-sectional procedures.

Given a set of consistent estimates of the parameters in the wage equation, we are able to assess the proportion of the observed gender wage gap that is attributable to unobserved factors (such as ability, preferences and discrimination). We adopt a methodology suggested by Jenkins (1994) which is based on the distribution of wage differences. This approach contrasts the standard practice to decompose the *average* wage gap into explained and unexplained parts (e.g. Oaxaca and Ransom (1994), Blinder (1973) and Oaxaca (1973)). As Jenkins argues *"the measurement and analysis of earnings discrimination should take into account the complete distribution of discrimination experienced."* That is, focusing on differences in averages does not provide a complete picture of the unexplained gender wage gap and this measure may be consistent with very different distributions of wage gaps. To exemplify, suppose that the unexplained

[4] Angrist and Newey reported that about 20 percent of the continuously employed men included in their sample experienced some increases in schooling between 1983 and 1987. This number seems quite high and is most likely due to the very young sample they used (aged 18-26 in 1983). According to our sample, very few individuals return to acquire more schooling (4.6 percent of the women and 2.4 percent of the men). In such a case, the estimated return using fixed-effects would reflect the marginal return of these people, and would not necessarily coincide with the average, marginal return in the sample.

part of the average wage gap is found to be 10%. From this measure we are not able to infer whether all women in the sample faces a wage gap of 10%, or if half the women experience a gap of 20% and the other half experience no gap at all. Clearly, relying only on differences in averages can be quite misleading.

Another attractive feature of the distributional approach is the ability to examine the sources of variations in the unexplained gender wage gap. For instance, we can use this framework to investigate if the gap differs both across different individual characteristics (such as educational attainment) and over time.

The outline of the paper is as follows. In the following section, the empirical framework is discussed. We briefly outline how we estimate the wage equations and describe different measures of the unexplained wage gap. Section 3 provides a description of the data. Section 4 presents the results, with focus on the size and composition of the unexplained wage gap. The paper ends with conclusions and discussion in Section 5.

2 Empirical framework

In this section, we will first briefly describe how we obtain the estimates of a reduced form wage equation, accounting for (potentially) endogenous explanatory variables. After this, we will describe the various measures of the unexplained wage gap that have been used in the literature, both the traditional approach based on a decomposition of differences in average wages, and a newer approach which is based on the whole distribution of wage differences.

2.1 Estimating wage equations with endogenous regressors

The empirical relationship to be estimated in this paper follows a traditional Mincer (1974) equation that relates the logarithm of hourly wages (LNW_{it}) to a vector of time-invariant characteristics (Z_i) as well as to a vector of time-varying characteristics (X_{it}). Formally, the equations to be estimated are:

$$LNW_{it} = \alpha^w + \gamma^{w\prime} Z_i + \beta^{w\prime} X_{it} + \mu_i^w + \varepsilon_{it}^w \qquad \forall i \in W_t \tag{1a}$$

and

$$LNW_{it} = \alpha^m + \gamma^{m\prime} Z_i + \beta^{m\prime} X_{it} + \mu_i^m + \varepsilon_{it}^m \qquad \forall i \in M_t \tag{1b}$$

where W_t denotes the set of women in the sample at time t and M_t denotes the set of men. Further, $i = 1, ..., N_t^w$ for women, $i = 1, ..., N_t^m$ for men, and $t = 1, ..., T_i$. Sub index i denotes individuals. The last two terms of the above relationships concern the error structure of the model. Unobserved, time-invariant individual effects (such as, ability and motivation) are assumed to be captured in μ_i, while ε_{it} is the remainder disturbance, reflecting effects of unobservable variables that vary both across individuals and over time. It is assumed that: (1) the sequence $\{\varepsilon_{it}\}$ consists of normal i.i.d. random variables with mean zero and variance σ_ε^2; (2) ε_{it} and μ_i are mutually independent; and finally, (3) μ_i is normally distributed with mean zero and a homoscedastic variance σ_μ^2. As argued in the introduction, it is reasonable to assume that the included human capital variables (especially years of education but also work experience) are correlated with unobserved ability (μ_i).[5]

When applying instrumental variables (IV) techniques in studies using cross-sectional data, a problem is to find exogenous variables that can be used as instruments, and at the same time, be legitimately excluded from the wage equation. If panel data information is available, individual means over time of all included time-varying regressors can serve as valid instruments (see HT (1981)).[6] In this paper, we use the HT approach, modified to account for the fact that we have an unbalanced panel. Identification is obtained using the within-person means of both endogenous and exogenous time-varying variables. Specifically, the instruments are (i) each time-varying endogenous variable, expressed as deviations from within-person means, (ii) each time-varying exogenous variable, expressed both as deviations from within-person means and as within-person means, and (iii) each time-invariant exogenous variable. In our sample, the time-varying variables are: actual work experience, area of residence, marital status, and time dummy variables.

To implement the HT method, we must make assumptions about which regressors are correlated with the person-specific error component (μ_i). We assume the exogenous variables include the following time-invariant regressors: education of both parents, father's occupation while growing up, the work status of the mother while growing up, and the immigrant status of the parents. Education as well as all the time-varying variables are potentially correlated with the person-specific error component (μ_i).

If the regression models above (1a and 1b) are correctly specified, and if our assumptions regarding the instruments are valid, then the estimates obtained using HT are consistent. In this case, they are also more efficient than fixed-effect (within-person) estimates as HT uses cross-person as well as within-person variation in the sample. Moreover, effects of time-invariant regressors are identified in HT but not in fixed-effects framework. We will use a Hausman-test, based on the difference between the HT and the fixed-effects estimates, to evaluate the validity of HT.

[5] It is assumed throughout this paper that the stochastic term ϵ_{it} is uncorrelated with all included covariates.

[6] When the panel is balanced, there exists even more efficient instruments (see Breusch, Mizon and Schmidt (1989) and Amemiya and MaCurdy (1986)).

As was shown in the introduction, several recent IV studies, based on exogenous sources of variation in educational outcomes, report estimates of the returns to schooling that are substantially higher than the corresponding OLS estimates.[7] As suggested by Card (1995 and 1999), these results may be explained by the existence of heterogeneity in individual returns and by the fact that these studies are based on instruments that influence only the educational decision of individuals with high marginal returns. This conclusion is also consistent with the local average treatment effect (LATE) interpretation of IV (Imbens and Angrist (1994)) according to which IV identifies only the average marginal returns of those who comply with the assignment-to-treatment mechanism implied by the instrument. Thus, the estimate recovered by IV does not necessarily coincide with the average marginal return in the population but rather the average marginal return for the sub-group of the population which is affected by the exogenous variation in educational outcomes. However, in the HT framework, there is no obvious link between the identifying assumptions and educational outcomes. It is therefore possible that the HT estimate of the average marginal return to education is more representative for the population than most of the earlier reported IV studies using cross-sectional data.

2.2 Measuring the Unexplained Gender wage Gap

As was argued in the introduction, there exists (at least) two different ways to measure the unexplained wage gap. The traditional approach uses the regression estimates and average observed characteristics ($\overline{Z}$ and $\overline{X}_t$) to partition the average (log) wage gap, at any time period t, into explained and unexplained parts. Formally, we can write this as:

$$\begin{aligned}\overline{LNW}_t^m - \overline{LNW}_t^w &= (\overline{Z}^m - \overline{Z}^w)\widehat{\gamma}^m + (\overline{X}_t^m - \overline{X}_t^w)\widehat{\beta}^m + \\ &\quad \overline{Z}^w(\widehat{\gamma}^m - \widehat{\gamma}^w) + \overline{X}_t^w(\widehat{\beta}^m - \widehat{\beta}^w) + (\widehat{\alpha}^m - \widehat{\alpha}^w)\end{aligned} \tag{2}$$

where the first two terms represents the part of the observed average wage gap that can be attributed to differences in characteristics, and the last three terms summarizes the unexplained differences in wages. In the above equation we have implicitly assumed the male parameter set as the non-discriminatory (or reference) parameter set. That is, the return to observable characteristics women would face in the absence of unexplained wage differences. While this decomposition is attractive in its simplicity, it is not capable of describing the distribution of the unexplained wage gap.

Instead of relying on differences in average wages, it is more appealing to compare the whole distributions of predicted wages. As a starting point, we need to predict wages for women under two different regimes, one using the set of

[7] Manski and Pepper (2000) obtain lower and upper bound estimates of the return to education using a sample taken from the National Longitudinal Survey of Youth (NSLY). Their results cast doubts on some of the very high returns reported in the literature.

parameters from the regression based on the female sample, and one using the male parameter set. Specifically, we have:

$$\widehat{y}_{it} = \exp(\widehat{\alpha}^{w} + \widehat{\gamma}^{w\prime} Z_i + \widehat{\beta}^{w\prime} X_{it} + \frac{1}{2}(\widehat{\sigma}^2_{\mu w} + \widehat{\sigma}^2_{\varepsilon w})) \qquad \forall i \in W_t \tag{3a}$$

and

$$\widehat{r}_{it} = \exp(\widehat{\alpha}^{m} + \widehat{\gamma}^{m\prime} Z_i + \widehat{\beta}^{m\prime} X_{it} + \frac{1}{2}(\widehat{\sigma}^2_{\mu m} + \widehat{\sigma}^2_{\varepsilon m})) \qquad \forall i \in W_t \tag{3b}$$

where $\widehat{r}_{it}$ can be thought of as the reference wage, that is, the wage woman i would receive in the absence of any unexplained differences.[8] The problem in this framework is how to summarize all the information contained in these two distributions. Jenkins (1994) proposed a number of measures for analyzing the joint distribution of wage differences. The first of these is related to the Gini coefficient, and is based on the area between the Generalized Lorenz Curve (GLC) and the Generalized Concentration Curve (GCC).[9] In the case of no unexplained gap, these two curves will coincide and the larger the area between the curves is, the larger is the unexplained wage gap.[10] Formally, this index (C_t) is defined as:

$$C_t = \left(1 + \frac{1}{2N_t^w}\right)\left(\frac{\overline{r}_t - \overline{y}_t}{\overline{y}_t}\right) - \left(\frac{1}{N_t^w}\right)^2 \sum_{i \in W_t} i \left(\frac{\widehat{r}_{it} - \widehat{y}_{it}}{\overline{y}_t}\right) \tag{4}$$

where $\overline{r}_t$ and $\overline{y}_t$ denotes average of respective distribution. This index incorporates the differences in means as in traditional decomposition approaches but also includes a term which is the predicted wage gap weighted by a woman's rank in the predicted wage distribution.

The second measure of the unexplained wage gap suggested by Jenkins (1994) is related to C but introduces a new parameter (κ) that incorporates the degree of aversion against the wage gap. The index is defined as:

$$J_t^{\kappa} = \sum_{i \in W_t} \omega_{it}(1 - d_{it}^{-\kappa}) = 1 - \sum_{i \in W_t} \omega_{it} d_{it}^{-\kappa} \tag{5}$$

where

[8] These two measures are similar to two of the measures presented in Jenkins (1994). However, since the wage rate is distributed lognormal, we need to include the estimates of the variances within the exponential brackets. Further, by assumption, the two variance components are independent, hence, $Var(\mu_i + \epsilon_{it}) = Var(\mu_i) + Var(\varepsilon_{it})$.

[9] GLC is obtained by ordering women in ascending order of observed wage and plot the cumulative wage per capita against cumulative sample share. GCC is obtained in a similar fashion except that observed wages are replaced by the predicted reference wages. However, the ordering remains the same as for GLC. For more details, see Jenkins (1994).

[10] However, if $\widehat{y}_{it} > \widehat{r}_{it}$ for any woman i, the distance between the two curves is not an appropriate measure of the unexplained wage gap.

$$
\begin{aligned}
\omega_{it} &= \frac{\widehat{y}_{it}}{N_t^w \overline{y}_t} \\
d_{it} &= 1 + \frac{|\widehat{r}_{it} - \widehat{y}_{it}|}{\overline{r}_t} \\
\kappa &> 0
\end{aligned}
$$

The variable d_{it} is a normalized wage gap (the gap for a woman relative to the mean of the reference distribution) and ω_{it} is woman $i's$ wage share at time t. J_t^κ allows us to aggregate the wage gaps in different ways using different values of κ. Higher values for κ corresponds to a greater degree of aversion against the wage gap. The aversion parameter can be thought to represent the increase in wages required to compensate a woman for a small increase in the wage gap.

In the results section below, we will report the values of these indices using both OLS and HT estimates of the wage parameters. These indices will also be compared with more conventional measures of the unexplained wage gap based on differences in averages.

It should be pointed out that the above analysis, like the traditional wage decomposition, is restricted to the distribution of the gender wage gap that is induced by the variation in the regressors included in the wage equations (1a and 1b) above. In this setting, the wage gap between men and women is explained by two sets of factors: differences in observable characteristics and differences in estimated returns associated with these characteristics. There exist alternative methods for analyzing wage differentials which allow for a more general treatment of the unobservable factors. For instance, Juhn et al (1991 and 1993) suggested a framework that further decomposed the residual wage differential based on percentile rankings. Changes in the residual wage differential are decomposed into two parts: changes in the level of unmeasured skill and changes in the returns to skill. The method has gained popularity because of recent interest in changing wage distributions and rising wage inequalities, both in the U.S. and elsewhere. Studies using this approach applied to the gender wage gap include Blau and Kahn (1996 and 1997) and Edin and Richardson (forthcoming). However, despite the attractive features of this method, it is restrictive in the sense that it relies on a decomposition of the *average* wage gap.

An alternative method to analyze wage differentials was suggested by DiNardo et al (1996). In their paper, they argue that it is important to work with the whole wage distribution instead of focusing on differences in average wages. They focus on the increased wage inequality in the U.S. during the 1980s and their results are based on semi-parametric techniques using kernel densities instead of relying on parametric regression methods. The main advantages of using such techniques, as opposed to parametric methods, include relaxing distributional assumptions and assumptions about the functional form of included covariates.

A parametric alternative to the approach taken by Dinardo et al (1996) is the use of quantile regressions. This framework allows the gender wage gap to be examined in different parts of the wage distribution. A recent application of quantile regression techniques on the gender wage gap is Albrecht et al (forthcoming). They examine the existence of a glass ceiling in Sweden and find that the gender wage gap in Sweden increases throughout the wage distribution, a result they interpret as a glass ceiling effect.[11]

3 Data

The empirical model outlined above is implemented on a sample of labor force participants extracted from the Swedish survey "*Household Market and Non-market Activities (HUS)*". The HUS survey is devoted to the construction of a reliable database which can be used to study the dynamics of household behavior. The first wave of HUS took place in 1984, and provides detailed information on a random sample of about 2,300 Swedish households. Information was gathered in a personal interview, and the questions cover a broad set of information, such as: family background, education, market work experience, family composition, labor market status at the time of the interview, child care arrangements, wages and incomes, and family composition. The second wave took place in 1986, where the respondents from 1984 were contacted again. Also, a new additional sample was questioned. The participants in 1986 were later contacted in 1988, 1991 and 1993. Finally, the most recent wave took place in 1996.[12]

The unbalanced panels used in this paper are composed of data extracted from two or more of the following surveys: 1984, 1986, 1993, and 1996. To be included in our sample, an individual has to be observed in at least two surveys. Since we lack information on some of the included variables for the surveys in 1988 and 1991, they are not included in this study. Individuals for whom information concerning the included variables was missing were excluded. The upper age limit was set to 65 years of age. In addition, we excluded self-employed and non-employed people.[13] After these selections, we were left with 547 males, and 478 females.

We also exclude those who participated in only one of the surveys. There exist a

[11] In this paper, we choose not to pursue any of the alternative methods for the following reasons. First, we believe that it is important to consider the whole wage distribution and not just focus on differences in average wages. This rules out the Juhn et al method. Second, since we are not only interested in comparing wage distributions but also interested in obtaining estimates of the return to human capital, we do not pursue a semi-parametric approach in this paper. Finally, because of the potential endogeneity of some of the regressors and the difficulty of handling endogenous variables in a quantile regression framework, we choose not to use quantile regression techniques.

[12] For more details on coding procedures, see Klevmarken and Olovsson (1993).

[13] This may introduce selection bias that should be controlled for during estimation. We will briefly discuss the robustness of our results towards this potential bias below.

growing literature on how to account for the potential bias this introduces (for a recent survey, see Vella (1998)). It is however beyond the scope of this paper to correct for this potential problem. In Table 1, we present descriptive statistics for those included in our sample but also for those who only participated in one survey (we refer to this group as the attrited sample). From the table we conclude that there are no systematic differences between those who participated in at least two surveys and those who did not. Those individuals who only participated once have on average slightly lower work experience and wages, but apart from that, the samples appear to be similar. We take this as an indication that attrition from the panel may not seriously bias our estimates.

Information on wages and incomes are obtained from two sources, and all refers to the previous year. For about two thirds of the sample, the information was collected from the Swedish tax registers, which ensures a high degree of accuracy. For the remaining one third, this information was collected during the interview. The wage measure used in this paper was constructed as the ratio of annual labor income and annual hours of work.[14] Hours of work include both hours spent on extra work and/or overtime. Moreover, annual hours of vacation are also included in the measure of hours of work. The reason is that vacation is generally paid in Sweden (and then included in labor income) and ignoring this would produce upward biased wage measures.

The education and work experience variables are the reported actual years of each activity. The first time the individuals participate in the survey, they are asked about their accumulated years of schooling and work experience. In subsequent surveys, individuals are asked to report their labor force status month by month since the last interview. Using this information, we can identify possible changes in both years of education and work experience between the survey times.

The remaining covariates in the model are defined as follows: four binary variables concerning the educational status of the parents corresponding to medium schooling (essentially high school equivalent education and vocational schooling), and to high schooling (a university degree or some university education); a binary variable that indicates whether the person's mother worked or not while growing up; two binary variables concerning the work status of the father while growing up (the first is equal to one if the father was a farmer, and the second is equal to one if the father was a blue-collar worker); a binary variable that is equal to one if both parents were born outside Sweden; two binary variables identifying the area of residence (medium-sized city or countryside); a binary variable which is equal one if the person is married or cohabiting. In addition to these variables, time dummies are included to capture macro-economic shocks and general wage inflation.

[14] The hourly wage rate was deflated by consumer price index using 1983 as the base year.

4 Empirical results

In this section, we first discuss model testing and present the estimated coefficients from the wage equations. Special attention is devoted to determine the effect of ability and endogeneity on the returns to human capital. The size and composition of the gender wage gap are discussed in sections 4.2 and 4.3.

4.1 Wage estimates and returns to human capital

Before discussing estimates from the wage regressions, we present the outcome of some model specification tests. We first estimated equations 1a and 1b above using HT under the assumption that years of schooling is the only endogenous variable. However, some of the time-varying variables may also be correlated with the person-specific error component (μ_i) in the regression equations. In this case, the estimates from HT (assuming that only education is endogenous) would be biased and inconsistent whereas the estimates from a fixed-effect (within-person) estimator would be consistent. To determine the validness of the HT estimates, we can use a Hausman-test based on the difference between the HT and the fixed-effect estimates. Specifically, let $q = \widetilde{\theta} - \widehat{\theta}$, where $\widetilde{\theta}$ is the vector of estimates from the fixed-effect model and $\widehat{\theta}$ is the corresponding vector of estimates from HT (assuming that only education is endogenous). Moreover, let $\widehat{\Sigma}$ denote the estimate of the covariance matrix of q. Under the null hypothesis, the statistic $\widehat{TEST} = q'\widehat{\Sigma}q$ is distributed chi-squared. If we fail to reject the null hypothesis of no significant difference between $\widetilde{\theta}$ and $\widehat{\theta}$, this would indicate that instrumentation of the schooling variable is sufficient to remove any correlation between the individual specific terms (ability) and the remaining regressors. The values of the test statistics, $\widehat{TEST} = 5.67$ for males and $\widehat{TEST} = 6.47$ for females, reported in Table 2 imply that we cannot reject the null hypothesis neither for males nor for females.[15]

The estimates from the wage regressions are presented in Table 2. The entries in columns one and three refers to OLS while columns two and four show the HT estimates.

Family background variables (dummy variables indicating parents' education, parents' work status, and whether both parents were born outside Sweden or not) are included because they are believed to capture some of the effects of pre-school investments, and hence ability, on the labor market. For instance, if both parents were born outside Sweden, this may affect the child's possibility to profit from a Swedish education. Similarly, the level of parents' education and their working status may be thought of as indicators of intellectual stimulus received from the parents. However, most of the effects of family background on wages are insignificant both for males and for females, even if there appears to exist a positive relationship between parents' education and wages. One exception is

[15] The associated critical values for this test statistic are: 12.0 at the 10%-level and 14.1 at the 5%-level.

the immigrant status of the parents. If both parents were born outside Sweden, this has a significant and negative effect on male wages, but not on female wages. The estimates in Table 2 show that persons whose fathers were blue-collar workers have lower wages, a result that holds for both males and females. Further, individuals residing outside urban areas (Stockholm, Göteborg, and Malmö) have lower wages, especially those living in the countryside. Being married or cohabiting does not seem to influence male wages, and for females the coefficient is only significant in the OLS case. Finally, the results reveal a significant increase in real wages between 1983 and 1995 for females but not for males.

The estimates on the return to human capital that are reported in Table 2 show a large difference between gender. For males, an additional year of schooling implies a wage gain of about 4.4% (in the OLS case), while this figure is only 2.8% for females. Regarding work experience the gender difference is even larger, an increase with one percent raise male wages with about 0.23% (OLS) and female wages with only 0.14% (OLS).[16] Allowing for schooling to be endogenous has a dramatic effect on the estimated returns to both schooling and work experience. For males, the return to education increases to 6.1%, an increase of 38%. This result is in line with much of the previous work on endogenous schooling and the return to education, which have typically found that OLS underestimates the true return. Interestingly, the return to work experience also increases, from 0.23% to 0.29%. For females, the return to education drops from 2.8% to 0.7%. However, the return to schooling using HT is estimated without precision, and it is not significantly different from the OLS estimate. The return to work experience is also lower using HT (0.11%) compared to OLS (0.14%). One possible reason for the divergence in returns to schooling between males and females is that unobserved ability may be negatively correlated with schooling for males and positively correlated with schooling for females.[17] To illustrate the point, assume two students, which apart from gender, appears to be identical to the econometrician. Given an existing unexplained wage gap between men and women, the opportunity cost for the male student of schooling is higher. Similarly, assume further that there exists segregation in the labor market and that women tend to choose (or be pushed into) occupations that pay less than occupations typically occupied by male workers. If this is the case, the difference in opportunity cost of schooling between males and females is even higher. As a result of this, if a difference in opportunity cost between men and women exist, and is sufficiently large, then OLS would *underestimate* the true return to schooling for men (because of the negative correlation between schooling and ability) and *overestimate* it for women (because of the positive correlation between schooling and ability).

[16] The elasticities are evaluated at 20 years of market work experience.

[17] There is no reason to assume that measurement error in the schooling variable should differ across gender, so this is less likely to explain the difference in returns.

4.2 The size of the gender wage gap

In Table 3 we present summary statistics for the distribution of the unexplained (by the included covariates) wage gap for the years: 1983, 1985, 1992 and 1995. For all years, the gap is higher in the OLS case. The difference in the average estimated gap between OLS and HT range from 60% in 1983 to 12% in 1995. Focusing on the HT results, a closer look at the distribution for 1995 reveals that for some women the unexplained wage differential is as high as 36.6% while for some women in the lower end of the distribution, the differential is negative (-8.6%).[18]

The entries in Table 3 show that the unexplained gap has increased monotonically between 1983 and 1995, regardless of which estimation method we rely on.[19] A possible explanation for this result is the increased wage dispersion Sweden has experienced since the beginning of the 1980s (as documented by Edin and Holmlund (1995)). Blau and Kahn (1992 and 1995) have argued that overall wage inequality is important in explaining international differences in the gender wage gap. For example, their results indicate that the higher level of wage inequality in the U.S. fully accounts for the larger gender wage gap in the U.S. in comparison to countries with relatively small gender differentials (like Sweden).

The result of an increase in the wage gap during the 1980s contrasts the results reported in Blau and Kahn (1997). They showed that the gender wage gap in the U.S. between 1979 and 1988, controlling for race, education and experience, was reduced by 0.12 log points. They found that rising wage inequality slowed down women's progress but that other favorable gender-specific factors were more than sufficient to counterbalance the unfavorable changes in the wage distribution. For instance, during this period, women's relative experience levels increased and their occupational distribution improved compared to those of men. Our results, and the ones reported by Edin and Richardson (forthcoming), indicate that these factors were not sufficient to offset the negative impact of increased wage inequality in Sweden.

In Table 4 we present values of the various indices described in section 2.2. We also report the percentage difference between $\widehat{r}$ and $\widehat{y}$, evaluated at sample means. We refer to this measure as Df_t. This measure is essentially identical to the traditional measure based on differences in average wages (equation 2). We report the values when the wage equations are estimated with OLS and HT. Again, the results are consistent with what we have seen above. OLS yields

[18] This means that, for some women in our sample, $\widehat{y}_{it} > \widehat{r}_{it}$. As noted above, in this case the wage gap measures suggested by Jenkins (1994) are less appropriate. A more suitable way to compare or test differences between wage distributions in this case would be to test for first- and second-order stochastic dominance. However, we believe that the broad range of indices that we present in this paper provides a reasonably accurate description of the wage gap, and how it has evolved over time, in the Swedish labor market.

[19] Edin and Richardson (forthcoming) report a similar finding.

substantially higher measures of the unexplained wage gap, regardless of which index we use (C_t or J_t) and which value on the aversion parameter (κ) we use.

As a last check of the validity of the difference between OLS and HT in terms of the wage gap, we have calculated Generalized Lorenz (GL) and Inverse Generalized Lorenz (IGL) ordinates for the absolute raw wage gap ($|\widehat{r}_{it} - \widehat{y}_{it}|$) for two years, 1983 and 1995. The results are presented in Table 5. These results serve as dominance check and as can be seen, the ordinates for OLS lies everywhere above those of HT. This is true both regarding the GL ordinates as well as for the IGL ordinates. It is also true for both years. From the results presented in Tables 3 to 5, we conclude that OLS may seriously overestimate the unexplained gender wage gap and that the gap has increased substantially between 1983 and 1995.

4.3 The composition of the gender wage gap

In Table 6 we have decomposed the distributional index by different sample subgroups. The objective is to investigate which groups of women are worse off in terms of a gender wage gap and if this has changed between 1983 and 1995. We focus on two measures only: the traditional measure (Df_t) and the J_t^{κ}-index (using $\kappa = 5$). Further, since we have clearly shown the difference between OLS and HT, we only use wages predicted using HT.

As a reference, we include the values of Df_t and J_t for the whole sample. Table 6 contains some interesting results. First, the unexplained gender wage gap is relatively constant across different educational groups. In 1983, the unexplained wage differential is between 3.2% (among workers with a college/university degree) and 5.1% (among workers with a high-school degree). In 1995, the wage gap varies between 12.8% (high-school graduates) and 14.2% (college/university graduates).

Second, the results in Table 6 indicate that the unexplained gender wage gap differs substantially across different values of work experience. The average unexplained wage gap among women with more than 20 years of work experience is 13% in 1983 and 17.8% in 1995. These estimated wage gaps are significantly higher than what applies to women with less than 20 years of market work experience. One explanation for this result is the flatter wage-experience profiles that we find for women compared to men. The profiles imply that women have somewhat higher entry wages in the labor market than men, but a significantly slower wage growth. The difference in the shape of the profiles may be due to labor market discrimination (i.e. women are restricted to slow wage growth jobs) or to differences in preferences (women may foresee a number of labor market interruptions and choose jobs where the penalty for interruptions are lowest).

In an attempt to explore the robustness of our HT results towards possible overidentification, the HT estimates were also obtained using different instrument sets. The results from this sensitivity analysis are available upon request from the authors. In addition to rely on variation in individual means over time of the included regressors (HT) we included indicators for exogenous changes in the Swedish educational system. Using this type of information as potential instruments was suggested by Harmon and Walker (1995). Sweden has experienced a number of changes in the educational system over time, and it is natural to use this information here.[20] The estimates were similar to those reported in Table 2. We have also attempted to explore the sensitivity of our results towards non-random selections (reasons including attrition and non-participation in the labor market). The results, also available upon request from the authors, suggest that the results reported in Tables 2-6 are robust towards both types of potential misspecifications.[21]

5 Conclusions

The purpose of this paper has been to study how the unexplained wage gap is affected by treating years of education in a traditional wage equation as endogenous. There exist a number of recent studies that have documented a significant difference between the OLS and the IV estimates of the return to education. However, it has been proven difficult to find valid instruments. Using panel information on a sample of Swedish labor force participants, we adopted a random effects instrumental variable estimator derived by Hausman and Taylor (1981) that uses variation over time to identify the system of equations. For males, the return to education is significantly higher when we use IV (6.1%) than when we use OLS (4.4%). This tendency of OLS to underestimate the return to education has also been found in previous work. However, for females we found no significant difference between the OLS and the IV estimates. In contrast to the results for males, the IV estimate of the return to education was lower than the corresponding OLS estimate.

In order to assess the impact of biased wage estimates on the unexplained gender wage gap, we compared the distributions of predicted female wages using both male and female estimates, following Jenkins (1994). The results showed that the estimated wage differential was substantially higher (between 12% and 60%) when relying OLS compared to IV. Our results were robust both towards different sets of instruments as well as towards a number of different ways to measure the unexplained wage gap. When we investigated the unexplained wage gap for different subgroups of our sample, we found that the gap is relatively constant across educational groups. However, the gap differs substantially across different values of work experience. Finally, we found that the gap has increased over

[20] Meghir and Palme (2001) evaluates effects of these changes in the Swedish educational system on both educational attainment and wages. They also provide details on how the Swedish educational system has changed since the 1940s.

[21] Participation rates in market work in Sweden are relatively high, both for men and women, and the robustness of our results towards non-random participation was expected.

time, regardless of which estimation method we used, from about 4% in 1983 to 13% in 1995.

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Table 1. Descriptic statistics.
Standard deviations in parantheses.

	Males		Females	
	Used Sample	Attrited Sample	Used Sample	Attrited Sample
Variable	Mean	Mean	Mean	Mean
Family Background Variables:				
% fathers medium schooling	20	19	15	19
% fathers high schooling	11	14	10	12
% mothers medium schooling	18	20	18	22
% mothers high schooling	6	8	5	7
% fathers were farmer	28	36	35	36
% fathers were blue collar	43	36	40	37
% mothers never worked	55	55	55	52
% parents were immigrants	6	11	9	10
Human Capital Variables:				
Education (years)	11.79	11.50	11.51	11.26
	(3.30)	(3.77)	(3.49)	(3.37)
Experience (years)	24.23	22.72	17.16	15.19
	(10.43)	(12.94)	(8.68)	(10.08)
Other Variables:				
% residing in medium-sized cities	55	47	53	49
% residing in countryside	21	22	19	17
% married or cohabiting	88	84	88	84
Wage per hour	57.00	53.71	46.99	44.73
(1983 SEK)	(18.46)	(18.56)	(14.15)	(12.07)
Number of individuals	547	655	478	720
Number of observations	1354	655	1151	720

Table 2. Estimated coefficients of wage equations.
(T-statistics in Parentheses)

	Males		Females	
Variable	OLS	HT	OLS	HT
Intercept	3.182	2.912	3.289	3.591
	(65.79)	(10.06)	(71.41)	(16.79)
Father medium schooling	0.010	-0.017	0.030	0.078
	(0.61)	(0.47)	(1.53)	(1.92)
Father high schooling	-0.035	-0.081	-0.021	0.024
	(1.39)	(1.30)	(0.78)	(0.47)
Mother medium schooling	0.016	0.030	0.036	0.089
	(0.90)	(0.96)	(1.86)	(2.15)
Mother high schooling	0.073	0.058	0.097	0.161
	(2.44)	(1.00)	(3.07)	(2.66)
Father was farmer	-0.030	-0.030	-0.061	-0.078
	(1.68)	(0.97)	(3.09)	(2.26)
Father was blue collar	-0.043	-0.033	-0.044	-0.052
	(2.58)	(1.05)	(2.26)	(1.63)
Mother never worked	0.004	0.003	0.003	-0.003
	(0.30)	(0.11)	(0.24)	(0.13)
Both parents immigrants	-0.125	-0.091	-0.010	-0.004
	(4.82)	(1.84)	(0.43)	(0.10)
Education	0.044	0.061	0.028	0.007
	(20.16)	(2.82)	(11.97)	(0.04)
Experience	0.016	0.022	0.012	0.012
	(6.09)	(7.26)	(4.30)	(3.16)
Experience^2/100	-0.011	-0.019	-0.013	-0.016
	(2.12)	(3.44)	(2.00)	(1.84)
Resides in medium-sized cities	-0.081	-0.061	-0.009	0.026
	(5.36)	(2.76)	(0.54)	(1.13)
Resides in countryside	-0.098	-0.047	-0.048	-0.027
	(5.22)	(1.43)	(2.39)	(0.95)
Married or cohabiting	0.025	0.005	0.055	0.025
	(1.30)	(0.24)	(2.69)	(0.84)
Year 1985	0.044	0.042	0.044	0.046
	(2.70)	(4.21)	(2.45)	(3.65)
Year 1992	0.077	0.063	0.088	0.112
	(4.43)	(2.85)	(4.69)	(5.76)
Year 1995	0.054	0.037	0.029	0.073
	(2.89)	(1.29)	(1.43)	(3.08)
$\widehat{\sigma}_\varepsilon$	0.225	0.122	0.221	0.224
$\widehat{\sigma}_\mu$	-	0.251	-	0.152
Mean of θ_i	-	0.687	-	0.578
$\widehat{\text{TEST}}$	-	5.67	-	6.47
$\overline{\text{R}}^2$	0.376	0.678	0.242	0.603

Table 3. Summary statistics for male/female wage differentials.

Differentials (%)	1983 OLS	1983 HT	1985 OLS	1985 HT	1992 OLS	1992 HT	1995 OLS	1995 HT
Mean	7.2	4.5	8.2	5.9	9.6	8.4	15.3	13.7
Median	6.9	3.4	8.0	5.1	8.9	8.3	14.4	13.3
Minimum	-18.4	-21.9	-18.0	-21.5	-12.9	-13.2	-8.5	-8.6
Maximum	56.5	31.7	58.6	33.4	39.7	31.6	47.0	36.6
Std Dev	8.2	8.4	8.2	8.4	8.2	8.3	8.7	8.5
% positive	82	69	86	75	91	86	96	95

Table 4. Distributional discrimination index estimates.

Index		1983 OLS	1983 HT	1985 OLS	1985 HT	1992 OLS	1992 HT	1995 OLS	1995 HT
Df_t		6.8	4.1	7.9	5.5	9.3	8.1	15.0	13.0
C_t		0.03	0.02	0.03	0.03	0.04	0.04	0.07	0.06
J_t^κ	$\kappa = 0.25$	0.02	0.01	0.02	0.01	0.02	0.02	0.03	0.03
	$\kappa = 5$	0.25	0.13	0.28	0.184	0.33	0.28	0.44	0.41

Table 5. Discrimination dominance checks.

	1983				1995			
	Gen. Lorenz ordinates for $\lvert\widehat{r}_{it} - \widehat{y}_{it}\rvert$		Inv. Gen. Lorenz ordinates for $\lvert\widehat{r}_{it} - \widehat{y}_{it}\rvert$		Gen. Lorenz ordinates for $\lvert\widehat{r}_{it} - \widehat{y}_{it}\rvert$		Inv. Gen. Lorenz ordinates for $\lvert\widehat{r}_{it} - \widehat{y}_{it}\rvert$	
Cumulative sample share	OLS	HT	OLS	HT	OLS	HT	OLS	HT
0.20	0.8	0.7	1.3	1.1	0.6	0.5	1.4	1.0
0.40	1.6	1.3	2.8	2.5	1.3	1.0	3.3	2.6
0.60	2.2	1.8	4.2	3.7	2.2	1.8	4.9	4.0
0.80	2.9	2.3	6.1	5.2	3.0	2.5	6.3	5.2
1.00	3.8	3.1	7.7	6.6	3.8	3.1	7.7	6.6

Table 6. Decomposition of distributional index
by sample subgroup (using HT)

	1983			1995		
Variable	Sample prop.	Df_t	J_t^κ $\kappa = 5$	Sample prop.	Df_t	J_t^κ $\kappa = 5$
All	*100*	4.1	0.13	100	13.0	0.41
Primary education	36	3.6	0.11	20	13.6	0.41
High school	42	5.1	0.17	49	12.8	0.39
College/University	22	3.2	0.10	31	14.2	0.43
Experience < 20	79	1.9	0.05	50	9.1	0.31
Experience over 20	21	13.0	0.40	50	17.8	0.49

7 Appendix

7.1 Robustness of the results

In an attempt to explore the robustness of our HT results towards possible over-identification, the HT estimates were also obtained using different instrument sets. The results from this sensitivity analysis is found in Table A1 in the appendix and the entries in the first column shows the results for males while the second column shows the female results. In addition to rely on variation in individual means over time of the included regressors (HT) we included indicators for exogenous changes in the Swedish educational system. Using this type of information as potential instruments was suggested by Harmon and Walker (1995). Sweden has experienced a number of changes in the educational system over time, and it is natural to use this information here. As can be seen from Table A1, the estimates are similar to those reported in Table 2.

Whenever we use panel data for estimation purposes, we have to be concerned with non-random attrition from the panel. In order to evaluate if those who left the panel are systematically different from those who stayed, we estimated wage regressions only on those who left the panel. The results from these OLS regressions are shown in Table A2. Among those estimates that are significant, the results are virtually identical to those reported in Table 2. This result, together with the similarities in the data that was described in section 3, suggests that those who left the panel, and are excluded from our sample, are not systematically different from those who stayed.

As a final check of the robustness of our results, we have estimated versions of Heckman's two-step method to investigate if the fact that we focus only on working individuals imposes any bias on our estimates. The participation rates in market work in Sweden were relatively high in the 1980's and 1990's, both for men and women. This fact may reduce the importance of adjusting for this type of selection. The results are shown in Table A2.[22] Again, the estimates are very similar to the ones reported in Table 2.

[22] We included age and child dummies in the first stage Probit model while these variables were excluded from the wage regressions.

Table A1. Estimated coefficients of wage equations using minimum school leaving age as additional instrument.
(T-statistics in Parentheses)

Variable	Males	Females
Intercept	2.882 (19.91)	3.667 (19.97)
Father medium schooling	-0.020 (0.64)	0.088 (2.26)
Father high schooling	-0.086 (1.82)	0.035 (0.70)
Mother medium schooling	0.029 (0.96)	0.098 (2.53)
Mother high schooling	0.055 (1.06)	0.173 (2.95)
Father was farmer	-0.030 (0.98)	-0.079 (2.38)
Father was blue collar	-0.032 (1.09)	-0.055 (1.72)
Mother never worked	0.003 (0.11)	-0.006 (0.24)
Both parents immigrants	-0.088 (2.00)	-0.007 (0.17)
Education	0.064 (6.13)	-0.006 (0.39)
Experience	0.022 (7.63)	0.012 (3.11)
Experience^2/100	-0.019 (3.53)	-0.017 (1.93)
Resides in medium-sized cities	-0.061 (2.89)	0.028 (1.20)
Resides in countryside	-0.046 (1.65)	-0.028 (1.00)
Married or cohabiting	0.005 (0.22)	0.020 (0.71)
Year 1985	0.041 (4.37)	0.047 (3.66)
Year 1992	0.061 (3.77)	0.116 (6.15)
Year 1995	0.035 (1.71)	0.078 (3.44)
$\widehat{\sigma}_\varepsilon$	0.123	0.154
$\widehat{\sigma}_\mu$	0.246	0.224
Mean of θ_i	0.681	0.578
$\overline{\mathrm{R}}^2$	0.680	0.593

Table A2. Estimated coefficients of wage equations using OLS.
(T-statistics in Parentheses)

	Males		Females	
Variable	Attrited Sample	Full Sample	Attrited Sample	Full Sample
Intercept	3.167	3.154	3.295	3.283
	(46.92)	(66.85)	(63.86)	(65.63)
Father medium schooling	0.051	0.007	-0.022	0.039
	(1.69)	(0.40)	(0.98)	(1.77)
Father high schooling	0.066	-0.026	-0.001	-0.049
	(1.64)	(1.04)	(0.03)	(1.74)
Mother medium schooling	0.018	0.016	0.019	0.032
	(0.60)	(0.89)	(0.85)	(1.52)
Mother high schooling	0.008	0.080	0.008	0.096
	(0.18)	(2.66)	(0.22)	(2.78)
Father was farmer	-0.038	-0.014	-0.012	-0.037
	(1.33)	(0.76)	(0.50)	(1.70)
Father was blue collar	-0.028	-0.038	0.006	-0.015
	(0.96)	(2.29)	(0.07)	(0.74)
Mother never worked	-0.001	-0.003	0.013	-0.016
	(0.006)	(0.24)	(0.74)	(0.99)
Both parents immigrants	-0.109	-0.116	-0.069	-0.049
	(3.34)	(4.47)	(2.56)	(1.89)
Education	0.035	0.044	0.026	0.029
	(10.78)	(19.79)	(9.16)	(11.12)
Experience	0.023	0.019	0.013	0.011
	(7.19)	(7.33)	(4.71)	(3.72)
Experience^2/100	-0.029	-0.018	-0.014	-0.014
	(4.50)	(-3.56)	(2.22)	(1.93)
Resides in medium-sized cities	-0.012	-0.076	-0.050	-0.008
	(0.51)	(5.04)	(2.77)	(0.48)
Resides in countryside	-0.064	-0.091	-0.097	-0.046
	(2.19)	(4.89)	(4.03)	(2.11)
Married or cohabiting	0.029	0.033	0.045	0.048
	(1.04)	(1.72)	(2.06)	(2.14)
Year 1985	0.042	0.050	0.075	0.050
	(1.81)	(3.05)	(3.88)	(2.53)
Year 1992	0.068	0.079	0.067	0.080
	(2.69)	(4.55)	(3.49)	(3.90)
Year 1995	-	0.064		0.036
		(3.46)		(1.63)
Selection correction term		0.004		-0.090
		(0.10)		(2.02)
$\widehat{\sigma}_\varepsilon$	0.251	0.228	0.209	0.253

Occupational Gender Composition and Wages in Sweden*

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Abstract

We estimate the relationship between wages and occupational gender segregation in Sweden. The results showed that the unexplained gender wage gap is largest in female dominated occupations and smallest in male dominated occupations. The experience-earnings profile is steeper for women in male dominated occupations. The results also indicated that about 30 percent of the gap can be attributed to segregation in the labor market. Ignoring occupational segregation produces significantly higher estimates of the unexplained fraction of the gender wage gap.

Key words: occupational segregation, gender wage gap

JEL classification: J31; J71

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1 Introduction

In the literature there exist a general understanding that occupational segregation is present in the labor market and that women are gathered disproportionally in occupations with lower earnings. However, there is no agreement on the cause of these outcomes and two contradicted theories have been provided in the literature. According to the first theory, women are gathered disproportional in occupations with low earnings due to market discrimination. The second theroy predicts that occupational segregation is the outcome of a self-sorting mechanism. Unfortunately, it has proven very difficult to empirically test these two competing theories.[1]

Despite the difficulties of establishing the reasons for occupational gender segregation, it is still important to assess the impact of this labor market phenomena on wages and wage gaps. Recently, a number of studies devoted to empirically determining the impact of the density of females (FEM) in a certain occupation on individual wages have appeared, see for instance Bayard et al (1999), Macpherson and Hirsch (1995), Sorensen (1990), Sorensen (1989), England et al (1988) and Johnson and Solon (1986) for applications on U.S. data, Baker and Fortin (1999) using Canadian data, Miller (1987) using data from the U.K. and le Grand (1991) using Swedish data. The results from these studies vary but most suggest a negative relationship between proportion of females in a given occupation and wages in that occupation.[2]

A potential problem with most of the previous studies is the assumption that occupational attainment is exogenous.[3] As argued by Macpherson and

[1] The problem is similar to the one of the existence of dual labor markets. Dickens and Lang (1985) presents a model, which they argue is able to test the human capital theory against the dual labor market theory. However, there is no general agreement on the validity of their claim. Their model is able to test differences in wage distributions between a primary and a secondary labor market but not the reasons for the presence of this difference.

[2] A problem which is neglected in many of these studies (i.e. in Bayard et al (1999), Sorensen (1990), le Grand (1991), Miller (1987) and Johnson and Solon (1986)) is the fact that the standard errors from OLS estimation are biased since the error term is correlated across workers within occupations, see Moulton (1990). It is therefore difficult to assess the significance of the results in these studies.

[3] Exceptions include Macpherson and Hirsch (1995), Sorensen (1989) and England et al (1988).

Hirsch (1995), there exist at least two reasons for why the exogeneity assumption may be false. First, if men and women with higher unmeasured skills (captured by the error term in the wage equation) are more likely to be sorted into male jobs and those with lower skills into female jobs, then the exogeneity assumption will obviously be violated.[4] Second, the error term may also capture unobserved taste differences among workers. To illustrate this point, some female workers may foresee future work interruptions due to childbearing and thus prefer part-time jobs or jobs where the wage "penalty" for absence from work is low. Based on this argument, we would observe a concentration of female workers in these types of jobs, which may also pay lower wages. It is again clear that the assumption of no correlation between the density of females in an occupation and the error term can be violated. To avoid the potential problem with endogeneity, Macpherson and Hirsch (1995) use longitudinal data covering the period 1983 to 1993 and apply a fixed-effects estimator.[5] The advantage with such a procedure is that it differences out any time-invariant unobserved (and observed) variables. Under the assumption that only the time-invariant portion of the error term is correlated with FEM, this procedure yields unbiased estimates of the effect of FEM on wages. A serious problem with this approach, however, is the fact that few workers change their occupational status over time and, as a consequence, only a small subsample of occupational movers identify the coefficient. Furthermore, the movers may constitute a non-representative portion of the sample, they may for instance be younger and clustered in low-skill jobs.

Bearing in mind the potential problems with the usage of a fixed-effects estimator in this type of study, we instead suggest the use of a different approach. We aggregate the FEM-variable into three categories depending on the proportion of women in the occupation: male dominated (less than 33 percent females), intermediate (between 33 and 66 percent females) and female dominated (more than 66 percent females). Non-random selection into an occupation is controlled for by estimating an ordered probit model in the first stage and including a selection correction term in the second stage wage equations. The advantage with this approach is that it allows us not only to estimate the wage effect of female density in any given occupation,

[4] Note that this kind of sorting may result from employer discrimination.

[5] England et al (1988) apply a similar strategy on a sample taken from the National Longitudinal Survey.

but it also enables us to estimate the unexplained gender wage gap within a given occupation and how this gap varies across occupations. In addition, we can also test whether the returns to accumulated human capital differ across both gender and occupations. For example, if most female workers choose occupations where the wage "penalty" for work absence is low, we expect flatter age-earnings profiles for women in female dominated jobs. The problems with our approach are finding valid instruments for occupational choices and the importance of aggregation. Concerning the first problem, it is in general difficult to obtain observable characteristics that influence occupational choice but not wages. In this paper, we use information on the number of children and age as instruments.[6] The second concern is how sensitive our results are towards the degree of aggregation we pursue. In order to assess this point, we provide estimates from two different specifications that differ only in the number of occupational groups.

Our main results can be summarized as follows. First, the unexplained gender wage gap varies substantially across occupations. For example, in male dominated occupations, the unexplained wage gap is about 0.018 (and not significant). This is significantly smaller than the estimate (0.121) obtained in female dominated occupations. Second, the results show that the female coefficient for work experience is about 60 percent higher in male dominated occupations compared to female dominated occupations. Thus, the experience-earnings profile is substantially steeper for women in male dominated occupations than for women in female dominated occupations. This result supports indicates that individuals who expect labor force intermittence will choose occupations in which the penalty for intermittence is lowest. Third, our results suggest that ignoring occupational differences in labor market may substantially overestimate the unexplained gender wage gap, as much of this differences can be explained by wage differences *across* occupations combined with occupational segregation.

[6] We expect that number of children is a more valid instrument for women since they are more likely to base their occupational choice on expected number of children than men. This implicitly assumes that there is a strong correlation between expected and actual number of children and that number of children has no impact on wages, conditional upon occupation. We believe that these assumptions are valid. Concerning the use of age as an instrument, we note that, once control for actual work experience is included in the wage equation, there is nothing in human capital theory that predicts age to be a determinant of wages. Overidentification tests reported in the result section suggests that our choice of instruments appear to be valid.

The paper is organized in the following way. Section 2 describes the data and sample used in this study. The empirical specification is presented in Section 3 while the results are presented in section 4. A final section contains a summary of the paper.

2 Data

The data used in the empirical analysis is drawn from a cross-section of the Swedish Household Income Survey (HINK) complemented with information on occupational segregation taken from the 1996 Labor Force Survey. Both of these data sources are supplied by Statistics Sweden. HINK provides information on labor market activities and incomes for a random sample of Swedish households, and approximately 7,000 households are interviewed each year. In this paper we use data from the 1997 survey. An interesting feature of this data set is the possibility of matching individual records with wage information provided by employers. The hourly wage rates obtained in this fashion correspond to the workers' contracted wage and do not suffer from the usual measurement errors that are common in self-reported wages. The wage information is available for all publicly employed workers as well as for the majority of privately employed workers.[7]

We limit the analysis to individuals aged 18 to 65 and we exclude self-employed workers. After these selections we are left with 3,995 females and 3,625 males.

To construct the FEM-variable, which measures the proportion of workers who are women in a given occupation, we used information from the 1996 Labor Force Survey. In HINK we have information regarding individuals' occupation at a two-digit level, and we can distinguish between 38 different occupations in the data.[8] We split occupations into three groups based on

[7] Since wages are not available for all privately employed workers, our sample contains a higher concentration of publicly employed workers than what is observed in the population. To test if our results were sensitive towards this, we estimated models with and without sample weights. We found that there were virtually no difference in the regression estimates and we are therefore confident that our results are not driven by the non-representative nature of our sample.

[8] Details about type of occupation and the proportion of women in each occupation is provided in Table A3 in Appendix.

the proportion of female workers in the occupation (i.e. male dominated, female dominated and integrated) following the convention in the literature, see Hakim (1998) and Jacobs (1995). Specifically, we define occupations with less than 33 percent women as being male dominated occupations and occupations with less than 33 percent men form the female dominated category. The remaining occupations form the integrated occupations category.

Explanatory variables used in the empirical analysis include information on: the highest educational degree each person has obtained, actual years of work experience, area of living (urban areas, medium-sized cities or the countryside), marital status and hours of work. In addition to these variables, we included information on number of children and age, acting as instruments, in the ordered probit model.

In Table 1 we present descriptive statistics for females and males by occupational type. For women, we observe higher average wage rates in male dominated and intermediate occupations than in female dominated occupations. However, women in the latter occupational group have, on average, acquired more education. We also observe a significant difference in the proportion of women working full-time. Among females in male dominated occupations, 77 percent work full-time (more than 1,500 hours per year). Among females in female dominated occupations, this figure is only 48 percent. It is interesting to note that this pattern cannot be observed for male workers confirming the traditional view that male labor supply is less flexible than female labor supply. Finally, Table 1 shows that males in female dominated occupations are highly educated as 50 percent has a college/university degree, compared to only 8 percent in male dominated occupations.

3 Econometric Specification

As was argued in the introduction, there exists plausible reasons for assuming that FEM is endogenously determined. The approach that we adopt in this paper, which controls for this type of potential misspecification, is to estimate a version of Heckman's two-step estimator. In the first stage, we estimate an ordered probit model that determines the probability of choosing a specific type of occupation (that is, male dominated, female dominated or an intermediate occupation). The parameters from the ordered probit are then used to form a selection correction term (similar to Heckman's lambda) that is added to the regression equation in the second stage. Formally, the model can be specified as follows:

$$\begin{aligned} FEM_{ij}^{*} &= \gamma_j Z_{ij} + \eta_{ij} \\ FEM_{ij} &= k \quad if \quad \mu_{k-1} < FEM_{ij}^{*} \leq \mu_k, \\ where\ k &= 0, 1, 2\ and\ \mu_{k-1} < \mu_k. \end{aligned}$$

$$\widehat{\lambda}_{ijk} = \frac{\phi\left(\widehat{\mu}_{k-1} - \widehat{\gamma}_j Z_{ij}\right) - \phi\left(\widehat{\mu}_k - \widehat{\gamma}_j Z_{ij}\right)}{\Phi\left(\widehat{\mu}_k - \widehat{\gamma}_j Z_{ij}\right) - \Phi\left(\widehat{\mu}_{k-1} - \widehat{\gamma}_j Z_{ij}\right)} \tag{1}$$

$$\ln w_{ijk} = \theta_{jk} + \beta_{jk} X_{ijk} + \partial_{jk} \widehat{\lambda}_{ijk} + \varepsilon_{ijk}$$

$$\begin{aligned} &\varepsilon_{ijk} \sim i.i.d.\ N(0,\, \sigma_{\varepsilon}^{2}) \\ &\eta_{ij} \sim i.i.d.\ N(0,1) \end{aligned}$$

where index i denotes individuals, index k denotes occupation and index j denotes gender ($j = w$ or m). Further, ϕ and Φ are the standard normal probability density function and distribution function, respectively. The μ's are unknown parameters to be estimated jointly with γ, and reflect threshold values for moving through the occupational choice decision. It is further assumed that ε_{ijk} and η_{ij} are correlated with correlation coefficient ρ. As is the case in a standard Heckman model, the standard errors of the estimates in the log-wage equation needs to be adjusted.

4 Empirical Results

In Table 2 we present results from the ordered probit. The entries in the first two columns present the results for males, while the last two columns show the results for females. The estimated coefficients indicate that education and occupational choice is strongly correlated and that the probability of working in a female dominated occupation is higher for those with higher education. This result holds for both males and females. As expected, the effect of work experience is the opposite of that of education. However, it is only significant for males. Men living in urban areas (i.e. Stockholm, Göteborg or Malmö) have a higher probability of working in female dominated occupations than other men. For women, we observe the opposite, namely that women who lives in urban areas are more likely to work in male dominated occupations. This suggests that occupational segregation is more significant in the countryside, perhaps because (occupational) traditions are more important there compared to larger cities. Marital status does not have any significant impact on either male or female occupational choice. Finally, we note that labor supply has no impact on occupational choice among men, but a significant effect on females' choices. The negative coefficient implies that women who work full-time are more likely to work in a male dominated occupation.

Regarding the effects of the instruments on occupational choice, we see that number of children has a negative effect for males and a positive effect for females. The estimate for males has a p-value of about 0.13, and implies that, everything else held constant, males with many children are more likely to hold a job in a male dominated occupation. For females, the estimate in column three implies the opposite, namely that females with many children are more likely to hold a job in a female dominated occupation. These results are not surprising. For instance, assuming a flatter age-earnings profile for women in female dominated occupations than in male dominated occupations, the wage penalty of work absence is lower in these jobs. Thus, women with many children (and therefore with more work absence) would prefer these jobs rather than jobs where the wage penalty is bigger (male dominated jobs). As a second instrument we include age. The reason for including this variable is the assumption that occupational segregation may be more pronounced among older cohorts than among younger ones. The results are mixed. For males, we find that older workers are more likely to work in female dominated occupations than younger workers. For females, the estimate suggest that women from older cohorts are more likely to possess jobs in segregated occupations. In the very last portion of Table 2, we report values of likelihood-ratio tests, which clearly reject the null hypothesis that the instruments have no significant impact on occupational choice.

Table 3 contains the wage estimates for females. The results in the first two columns refer to male dominated occupations, while columns three and four show the results for integrated occupations and the last two columns show estimates for females working in female dominated occupations. For all categories, the estimates regarding highest educational attainment (i.e. high-school degree or a college/university degree) are all insignificant. This suggests that, everything else held constant and *conditional on occupation*, wages among women with a high-school degree (or a college degree) are not significantly different from wages among women with less schooling.

It is reasonable to assume that workers who expect labor force intermittence will choose occupations where the penalty for intermittence is lowest (see for instance Polachek (1981 and 1985)). These occupations will have relatively high starting wages and flat earnings profiles. An implication of this is that women (who expect more frequent labor force intermittence) choose female dominated occupations because the penalty for intermittence is lower in these occupations than in male dominated occupations. Whether the earnings profiles are indeed flatter in female dominated occupations is an empirical matter. The results in Table 3 lends some support for this hypothesis since the coefficient for work experience is about 60 percent higher in male dominated occupations compared to female dominated occupations. The experience-earnings profiles for women in these two occupational groups are shown in Figure 1, which clearly shows a steeper earnings profile for women in male dominated occupations than for similar women in female dominated occupations.

Table 4 contains the wage estimates for males. The results regarding the effects of human capital imply higher return to education in male dominated occupations and higher return to work experience in female dominated occupations. Further, there is a significant, negative wage effect of being single in integrated occupations and a significant, positive effect of working full-time.

To test whether our choice of instruments is valid, we report the p-values for these variables when they were included in the wage equation. To achieve identification (without relying on the non-linear nature of the model), we included them separately. For women, the p-values for age suggest that age can serve as an instrument. However, regarding number of children, the p-values are high in both male dominated and female dominated occupations, but not so in intermediate occupations (where the p-value equals 0.039). This may suggest that part of our identification does not rely on a proper set of instruments. Perhaps as a consequence of poor instruments, none of the selection correction terms are significant for females and our results are not sensitive to the inclusion of these terms. For males however, both age and number of children appear to be valid instruments and the selection correction terms are significant in each wage regression.

In Table 5, we present observed, explained and unexplained gender wage differentials in the three occupational groups. The observed gender wage gap is smallest (2.6 percent) in male dominated occupations and largest in female dominated occupations (16.7 percent). As is shown in the last column, most of the observed wage difference is attributed to unobserved factors. In both male and female dominated occupations, about 30 percent of the observed wage gap can be "explained" by differences in observable characteristics (such as accumulated human capital and labor supply) and about 70 percent remains unexplained. It is interesting to observe that there exists substantial heterogeneity in the gender wage differentials across occupational groups. This is an observation that, surprisingly, has received little attention in the literature. Another interesting implication of our results on the gender wage gap is that the unexplained portion of this gap is not smallest in occupations with an equal gender distribution. This would suggest that policies such as affirmative action would have only limited effect on reducing the unexplained wage gap.

4.1 Decomposing the gender wage gap

Using our approach to estimate the gender wage gap enables us to decompose this gap into three mutually exclusive parts: differences in endowments, differences in occupational structure and differences in rewards to endowments. Formally, this can be written as:

$$\begin{aligned}
\overline{\ln w_m} - \overline{\ln w_w} \;=\; & \sum_{k\in K} (\pi_k^m - \pi_k^w) * \{\widehat{\theta}_{wk} + \widehat{\partial}_{wk}\widehat{\lambda}_{wk} + \widehat{\beta}_{wk}\overline{X}_{wk}\} + \\
& \sum_{k\in K} \pi_k^m * \widehat{\beta}_{mk} * (\overline{X}_{mk} - \overline{X}_{wk}) + \\
& \sum_{k\in K} \pi_k^m * \{(\widehat{\theta}_{mk} + \widehat{\partial}_{mk}\widehat{\lambda}_{mk}) - (\widehat{\theta}_{wk} + \widehat{\partial}_{wk}\widehat{\lambda}_{wk})\} + \\
& \sum_{k\in K} \pi_k^m * \overline{X}_{mk} * (\widehat{\beta}_{mk} - \widehat{\beta}_{wk})
\end{aligned}$$

where the first term on the right hand side measures differences in occupational structure $(\pi_k^m - \pi_k^w)$, the second measures differences in endowments $(\overline{X}_{mk} - \overline{X}_{wk})$, and the last two measures the gap due to unexplained factors. The π_i's are proportions of workers (men or women) in occupation k, and there is a total of K different occupations.

The results of this decomposition is reported in Table 6 for four different specifications. In the first column, we show the results from a model (estimated by OLS) which does not allow occupational structure to affect wages. In this case, 30 percent of the observed (log) wage gap between men and women can be explained by differences in endowents (primarily education and experience) and 70 percent of the gap is left unexplained. The entries in the second column are obtained by estimating separate wage regressions for three occupational groups. The results from this shows that the unexplained wage gap drop substantially, from 70 percent to about 40 percent. Hence, it is important to control for occupational differences when making inference about the gender wage gap, and neglecting to do so might yield overestimated unexplained wage gaps. Columns three and four shows results when non-random selection into occupations are controlled for and they show a similar picture as the results in column two.

4.2 Robustness of the results

In an attempt to explore the robustness of our results towards the assumption of aggregation of the FEM-variable we estimated a model in which we aggregated the FEM-variable into four groups instead of three.[9] The results from this sensitivity analysis are found in Table A2 in Appendix. This table shows the observed, explained and unexplained (log) gender wage gap across occupations. The entries should be compared to the ones in Table 5. The results in Table A2 show, in general, the same pattern as in Table 5. The wage gap is smallest in male dominated occupations and largest in female dominated occupations.

[9] Specifically, we define occupations with less than 25 percent women as being male dominated occupations and occupations with less than 25 percent men form the female dominated category. The remaining occupations form two "semi-integrated" occupations, one consisting of occupations with 25-50 percent men and one consisting of 25-50 percent women.

5 Conclusions

This paper studies the effects of occupational segregation on wages and on the gender wage gap. Specifically, we studied how the gender wage gap varies across different occupations. The data indicated that the gap is smallest in male dominated occupations and largest in female dominated occupations. Based on estimated wage equations, we decomposed the observed wage gap into explained (by differences in endowments) and unexplained parts. The results showed that the unexplained gap is largest (and significant) in female dominated occupations and smallest (and insignificant) in male dominated occupations.

Our results also showed that the female coefficient for work experience is about 60 percent higher in male dominated occupations compared to female dominated occupations. The experience-earnings profile for women in male dominated occupations is steeper than the profile in female dominated occupations. This result is in line with one of the theories explaining occupational segregation, which argues that individuals who expect labor force intermittence will choose occupations in which the penalty for intermittence is lowest.

Finally, we used our model to investigate the fraction of the observed wage gap that can be attributed to occupational segregation. The results, which were robust towards different model specifications, indicated that about 30 percent of the gap can be attributed to segregation in the labor market. Ignoring occupational segregation produces significantly higher estimates of the unexplained gender wage gap. Hence, it is important to include information on the occupational structure of the labor market when estimating the size of the unexplained wage gap.

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Figure 1. Experience-earnings profiles for women in female and male dominated occupations.

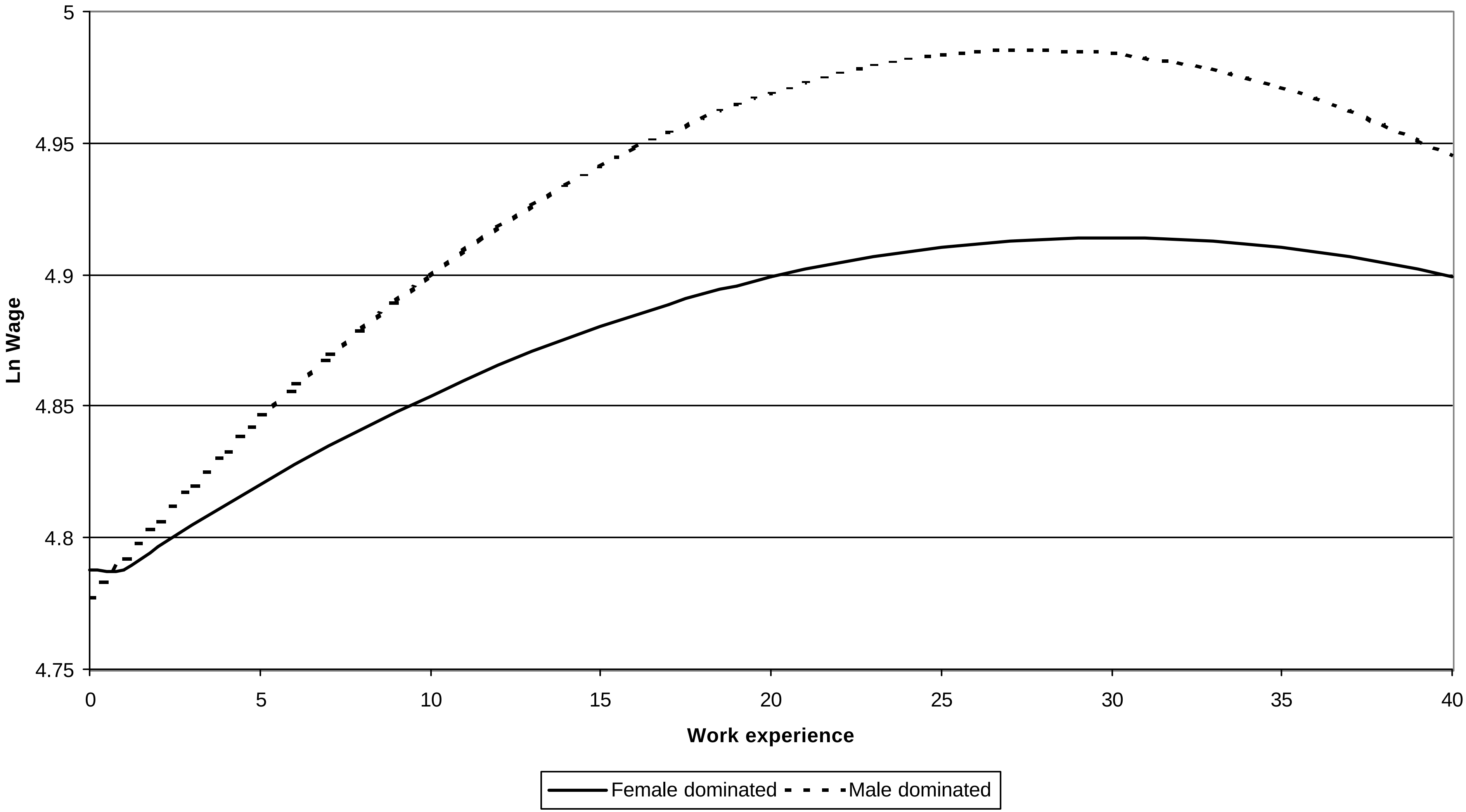

Table 1. Descriptive Statistics by Occupational Type.

WOMEN

Characteristics	Male Dominated Occupation Mean	Male Dominated Occupation Std	Intermediate Occupation Mean	Intermediate Occupation Std	Female Dominated Occupation Mean	Female Dominated Occupation Std
Wage per hour (1997 SEK)	107.7	28.4	106.1	31.6	99.3	22.2
High-School	0.53	-	0.60	-	0.67	-
College/University	0.13	-	0.16	-	0.20	-
Work experience	18.9	10.1	21.3	10.6	20.4	10.1
Living in urban areas	0.31	-	0.39	-	0.29	-
Living in medium-sized cities	0.37	-	0.37	-	0.39	-
Prop. single	0.18	-	0.15	-	0.15	-
Prop. working full-time	0.77	-	0.67	-	0.48	-
Number of children	0.8	1.0	0.7	1.0	0.9	1.1
Age	40.3	10.7	43.1	10.8	42.5	10.8
Number of observations	332		1650		2013	

MEN

Characteristics	Male Dominated Occupation Mean	Male Dominated Occupation Std	Intermediate Occupation Mean	Intermediate Occupation Std	Female Dominated Occupation Mean	Female Dominated Occupation Std
Wage per hour (1997 SEK)	111.0	32.2	124.8	49.2	121.5	46.9
High-School	0.65	-	0.58	-	0.42	-
College/University	0.08	-	0.24	-	0.50	-
Work experience	23.2	12.2	22.5	11.9	21.1	11.3
Living in urban areas	0.27	-	0.43	-	0.34	-
Living in medium-sized cities	0.41	-	0.37	-	0.40	-
Prop. single	0.14	-	0.13	-	0.19	-
Prop. working full-time	0.76	-	0.82	-	0.78	-
Number of children	0.8	1.1	0.8	1.0	0.8	1.0
Age	41.8	11.0	42.8	10.8	43.1	11.1
Number of observations	2073		1155		397	

Table 2. Ordered Probit Estimates.

	Males		Females	
Variable	Est.	Std.err.	Est.	Std.err.
High-School	0.215	0.053	0.506	0.048
College/University	0.919	0.072	0.686	0.063
Experience	-0.040	0.009	-0.003	0.009
Experience^2/100	-0.005	0.015	0.002	0.017
Living in urban areas	0.148	0.053	-0.142	0.048
Living in medium-sized cities	0.048	0.051	-0.024	0.047
Single	-0.071	0.062	0.014	0.053
Working full time	-0.033	0.050	-0.536	0.040
Number of children	-0.033	0.022	0.050	0.021
Age	0.047	0.006	0.006	0.004
μ_1	-1.548	0.159	1.131	0.123
μ_2	1.165	0.030	1.459	0.031
N	3,625		3,995	
Ave. Log-Likelihood	-0.8576		-0.8730	
LR-test[1]	63.0 (0.0001)		6.8 (0.033)	

Notes: The dependent variable takes on three values: 0 if male dominated occupation, 1 if intermediate and 2 if female dominated.
LR-test[1]: value of the LR-statistic when testing the instruments in the selection equation, p-value in parenthesis (truncated for males).

Table 3. Wage Equation Estimates for Females, by Occupation.

Variable	Male dominated Est.	Male dominated Std.err.	Intermediate Est.	Intermediate Std.err.	Female dominated Est.	Female dominated Std.err.
Constant	4.737	0.737	4.170	0.160	4.650	0.298
High-School	0.082	0.172	-0.086	0.133	-0.042	0.107
College/University	0.339	0.245	0.071	0.187	0.109	0.146
Experience	0.015	0.007	0.009	0.005	0.009	0.004
Experience^2/100	-0.027	0.016	-0.016	0.010	-0.015	0.009
Living in urban areas	0.101	0.065	0.133	0.049	0.054	0.039
Living in medium-sized cities	0.040	0.041	0.012	0.033	0.010	0.025
Single	-0.097	0.045	-0.048	0.034	-0.027	0.029
Working full time	-0.048	0.193	0.280	0.146	0.138	0.110
Lambda	0.178	0.412	-0.313	0.308	-0.350	0.322
Adj. R^2	0.220		0.269		0.273	
s_ε	0.257		0.353		0.326	
P-value for number of children[a]	0.946		0.039		0.874	
P-value for age[b]	0.972		0.347		0.646	

Note: The dependent variable equals the logarithm of hourly wage rates. [a]: P-value when number of children was included in the wage equations. [a]: P-value when age was included.

Table 4. Wage Equation Estimates for Males, by Occupation.

Variable	Male dominated Est.	Male dominated Std.err.	Intermediate Est.	Intermediate Std.err.	Female dominated Est.	Female dominated Std.err.
Constant	4.214	0.063	4.532	0.132	4.882	0.391
High-School	0.038	0.025	0.014	0.046	-0.013	0.096
College/University	0.163	0.084	0.011	0.117	0.061	0.185
Experience	0.010	0.003	0.016	0.005	0.015	0.007
Experience^2/100	-0.015	0.006	-0.025	0.010	-0.020	0.016
Living in urban areas	0.013	0.026	0.061	0.044	0.012	0.063
Living in medium-sized cities	0.020	0.022	0.007	0.041	-0.016	0.055
Single	-0.009	0.028	-0.092	0.046	-0.071	0.060
Working full time	0.136	0.022	0.231	0.039	0.090	0.056
Lambda	-0.294	0.104	-0.288	0.106	-0.275	0.170
Adj. R^2	0.265		0.279		0.387	
s_ε	0.320		0.388		0.354	
P-value for number of children[a]	0.531		0.579		0.540	
P-value for age[b]	0.181		0.308		0.587	

Note: The dependent variable equals the logarithm of hourly wage rates. [a]: P-value when number of children was included in the wage equations. [a]: P-value when age was included.

Table 5. Observed, Explained and Unexplained Gender Wage Gaps.

Variable	Observed Gap	Explained Gap	Unexplained Gap
Male Dominated	0.026	0.008	0.018
	(0.013)	(0.016)	(0.025)
Intermediate	0.136	0.037	0.099
	(0.011)	(0.013)	(0.023)
Female Dominated	0.167	0.047	0.121
	(0.016)	(0.013)	(0.023)

Note: Standard errors are reported in brackets.
The explained wage gap is calculated as: $(\overline{x}_{md} - \overline{x}_{fd})\widehat{\beta}_{fd}$ where $\overline{x}_{md}$ equals average characteristics in male dominated occupations, and $\overline{x}_{fd}$ equals average characteristics in female dominated occupations.

Table 6. Decomposing the Gender Wage Gap.

	OLS^1	OLS^2	Selection Corrected (3 groups)	Selection Corrected (4 groups)
Observed log wage gap	0.107	0.107	0.107	0.107
Gap due to differences in endowments	0.032 (30%)	0.031 (29%)	0.021 (20%)	0.040 (37%)
Gap due to occupational segregation	-	0.031 (29%)	0.031 (29%)	0.021 (20%)
Gap due to unobserved factors	0.075 (70%)	0.045 (42%)	0.055 (51%)	0.046 (43%)

OLS^1: Includes no control for occupational segregation.
OLS^2: Includes control for occupational segregation.

Appendix:

Table A1. Observed, Explained and Unexplained Gender Wage Gaps.

Variable	Observed Gap	Explained Gap	Unexplained Gap
Male Dominated (0-25% women)	0.005 (0.017)	0.012 (0.023)	-0.007 (0.047)
Intermediate I (25-50% women)	0.123 (0.015)	0.012 (0.023)	0.111 (0.037)
Intermediate II (50-75% women)	0.110 (0.011)	0.036 (0.017)	0.074 (0.023)
Female Dominated (75-100% women)	0.147 (0.030)	0.060 (0.078)	0.087 (0.098)

Note: These figures are estimates based on a model with FEM divided into four different groups. The explained wage gap is calculated as: $(\overline{x}_{md} - \overline{x}_{fd})\widehat{\beta}_{md}$ $\overline{x}_{md}$ equals average characteristics in male dominated occupations, and $\overline{x}_{fd}$ equals average characteristics in female dominated occupations

Table A2. Occupational Specification and Proportion of Women in each Occupation.

Occupation	Prop. of women
Science: Technical	0.111
Science: Chemical and Biological	0.612
Medicine, Health and Nursing	0.880
Education	0.703
Law	0.299
Religion, Journalist, Artist	0.581
Administration: Government and Business	0.465
Administration: Accounting, Clerical	0.906
Administration: Other	0.448
Sales: (business services, purchase, goods)	0.385
Sales: Other	0.542
Agriculture, Horticulture, Forestry: Management	0.117
Agriculture, Horticulture, Forestry: Workers	0.295
Wildlife Protection, Hunting and Fishing	0.038
Mining	0.027
Transport and Communication: Air, Sea, Other	0.189
Transport and Communication: Drivers, Delivery	0.077
Transport and Communication: Postal Servcie, Telecommunication	0.579
Manufacturing: Textile	0.658
Manufacturing: Iron and Metal	0.071
Manufacturing: Precision-tool	0.400
Manufacturing: Workshop and Construction	0.104
Manufacturing: Electrical	0.147
Manufacturing: Wood	0.146
Manufacturing: Painting and Varnishing	0.027
Manufacturing: Other Construction and Building	0.003

Table A2. Continued.

Occuaption	Prop. of women
Manufacturing: Graphics	0.262
Manufacturing: Glass, Pottery, Tile	0.292
Manufacturing: Dairy	0.286
Manufacturing: Chemical Processing	0.256
Manufacturing: Material Handling	0.049
Manufacturing: Packing and Storage	0.296
Manufacturing: Other	0.317
Services: Civilian Protection	0.182
Services: Lodging and Catering	0.767
Services: Caretaking and Cleaning	0.603
Services: Military	0.037
Services: Other	0.677

Household Labor Supply and Welfare Participation in Sweden

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Abstract:

The purpose of this paper is to analyze the joint effects of the tax and benefit systems on household labor supply. We formulate a discrete structural model where labor supply and participation in social assistance programs are jointly determined. The estimates from the structural model yielded small wage and income elasticities. A tax simulation showed that reducing the progressivity in the Swedish tax system may have considerable welfare effects. It also showed that these effects might differ substantially between poor households and rich households. The effect on working hours from the reform was quite small and tax revenues were predicted to drop significantly. Finally, the evaluation of a change in the welfare system showed that the stigma effect had a substantial impact.

Key words: Labor Supply, Social Assistance Participation, Unobserved Heterogeneity, Factor Loading, Tax Simulation.

JEL Classification: J22, I38

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1 Introduction

Our understanding of economic incentives plays a crucial role in economic policy design. Sweden, a small open economy with a large public sector, faces important challenges regarding the design of the tax and transfer systems. Increased labor mobility in the European community makes it costly to maintain tax systems that differ substantially across nations. Since progressive tax systems normally tax skilled workers' income more than low skilled workers' and since skilled workers are presumably more mobile, tax system competition could lead towards a common proportional European tax system. One purpose of this paper is to study effects of moving from a progressive income tax system, such as the Swedish, towards a less progressive system. A second challenge refers to the long term financing of the welfare system. The recession in the early 1990s created large budget deficits and the funding of the welfare system was at stake. This, combined with a dramatic increase in unemployment, created a need for a restructure of the whole Swedish welfare system. During the first half of the 1990s several reforms were implemented that reduced the generosity of the welfare system. As a consequence there has been a dramatic increase in the expenditure on social assistance in Sweden during the last decade. According to the National Board of Health and Welfare, total real expenditures between 1983 and 1997 increased from 4.4 billion Swedish kronor (SEK) to 12.4 billion SEK.[4] The second main objective of this study is to analyze the effect on social assistance on household labor supply.

To accomplish this, we need to specify a structural model of labor supply that allows for non-convexities in the budget set. The traditional way to model labor supply assumes that the decision variable, hours of work, is continuous and unconstrained. However, it has been shown that this framework need to impose restrictive conditions in order to be statistical coherent, see for instance MaCurdy et al (1990). Further, one underlying assumption in the traditional labor supply model is that the individual (or household) budget set is convex. Hence, to estimate such a model, a number of important simplifications of the income tax and transfer system must be imposed.

[4] 1 Euro $\approx$ 9 SEK

As an alternative to the continuous hours of work model, van Soest (1995), Hoynes (1996), Keane and Moffitt (1999) and Blundell et al (1999) has suggested the use of a discrete choice model instead. In this framework, it is straightforward to include as many details as possible regarding the budget set. It further extends naturally into a household model, where husbands and wives jointly determine their labor supply. A disadvantage with this approach is the introduction of a classification error in hours of work. This error arises because of the aggregation of a continuum of hours of work into a finite number of classes. However, by using a multiplicative classification error specification, following MaCurdy et al (1990) and Hoynes (1996), we can reduce this problem.[5]

In this study, we specify a structural model of discrete household labor supply along the lines described above. We also incorporate the decision of whether or not to participate in a social assistance program into the decision set.

The empirical part is based on a sample of households drawn from the Swedish Household Income Survey (HINK), which contains very detailed income information supplied by the tax registers. As a consequence, this study differs from most previous studies since almost all relevant details in the tax and transfer systems are considered.

To evaluate the budget set at different combinations of hours of work in the household, we use a micro simulation model (FASIT) developed by the Statistical Sweden and the Swedish Ministry of Finance. The hourly wage rates in this study are of an exceptional high quality. To the best of our knowledge this is the first labor supply study based on register information on hourly wage rates. Thus, the problem of endogeneity or measurement errors should be less severe in this study.

We choose to present our results in terms of simulated wage and income responses, but we also use the estimates from the model to evaluate the effects of changes in taxes and social assistance. The first experiment involves a reduction in the marginal tax rate among high-income earners, and in the second experiment the level of social

[5] Note also that there is no simple way to determine the appropriate number of classes. But according to both van Soest (1995) and Hoynes (1996), the main results seem to be rather insensitive regarding number of classes.

assistance is changed. Both these experiments will be evaluated using a micro simulation method. In addition to reporting the effects on hours of work and on consumption, we also report the overall welfare effects of the tax cut using an equivalent variation measure.

The result from the policy simulation indicates that moving from a progressive income-tax system towards a proportional system may have considerable welfare effects. Because of the structure of the reform, the welfare effects differ substantially across households, with the largest effects found for rich households. However, the predicted effect on hours of work is quite small and tax revenues are predicted to decrease significantly. The effects of changing the level of social assistance clearly demonstrate the importance of the stigma effect. For instance, without stigma an increase of the income level required for social assistance (the norm) by 25% implies a reduction in wife's labor supply by 17% for the poorest 10:th percent households. The corresponding result for the specification that incorporates the stigma effect is 6%.

The remainder of this paper is organized as follows. Section 2 gives a detailed description of the budget set and the relevant benefit programs. Section 3 presents the economic model and Section 4 describes the data used in the analysis. In Section 5 we present the results, while the conclusions are found in Section 6.

2 The Budget Set

A static model of household labor supply is assumed where spouses determine hours of work and consumption by maximizing a utility function $U(C, h_h, h_w)$ subject to the following budget constraint:

$$C = C_h + C_w + B$$

where C_h and C_w are husband's and wife's after tax income, respectively, and B is the amount of household specific means-tested benefits/subsidies. The individual components to total consumption are given as:

$$C_i = W_i h_i + Y_i + V_i - t(I_i) \qquad \text{i = h, w}$$

where

W_i	=	Gross wage per hour
h_i	=	Hours of market work per year
Y_i	=	Taxable non-labor income per year
V_i	=	Non-taxable non-labor income per year
t	=	Taxes determined by the function $t(\cdot)$
I_i	=	Taxable income per year, $I_i = W_i h_i + Y_i - D_i$
D_i	=	Deductions per year

The two major transfer programs included in B are: *housing allowance* (B_h) and *social assistance* (B_w).[6] Housing allowance is determined by nationwide rules and is mainly directed toward families with children. About 17 percent of all households are eligible for housing allowance. The amount received by a household is determined by: net household income, housing expenditures, number of children and the ages of the spouses.

The rules determining social assistance are based on rather complicated systems and they also differ across municipalities. For each municipality and each type of family, we calculated a "norm" (the minimum level of disposable income to qualify for social assistance) based on information provided by the Swedish municipalities. The amount of social assistance a family receives is simply the difference between the norm and the household's disposable income.

A detailed treatment of the income tax and benefits systems generally results in non-convex budget sets. This is also the case in Sweden, and to illustrate this, we show the household budget sets for two typical households in Figure 1. The budget sets are evaluated at 49 discrete points (seven for each spouse) ranging from 0 to 3,000 hours per year. In the upper left-hand panel shows the budget sets for the husband conditional on different hours for his wife, while the lower left-hand panel shows similar information for the wife. The non-convexity of the budget sets at lower hours of work is apparent. The return to low hours of work (from 0 to 500 hours) varies

[6] When constructing the budget sets, we also include information about childcare costs.

substantially depending on spouse hours. If the spouse does not work, the budget set is flat for the household. The reason for this is that, at low earnings, the household is entitled to social assistance and there is a 100% marginal tax rate on social assistance. For the wife there is a very small return from an increase to 1,000 hours per year if the husband work few hours. The main reason for this pattern is the reduction in housing allowance associated with the increase in household income.

The budget set for a high-income family looks quite different. The non-labor income of this family is too high to enable them for social assistance even if none of them work. However, the shape of the budget set is affected by both childcare costs and housing allowance. In the case where none of the spouses work, this family is entitled to 32,000 SEK in housing allowance and the cost of childcare (2 children) is 6,500 SEK. Due to a high wage rate, the housing allowance is reduced to zero already at 1,000 hours for the wife, regardless of the husband's hours of work. The cost of childcare reaches its maximum (18,000) at a household income of about 300,000 SEK. From these illustrations it follows that it is mainly low-income households who face non-convex budget sets and high marginal effects.

The main source for the non-convex budget sets is the transfer system designed to equalize the income distribution. However, income-tax system also produce non-convexities, but not as large as those produced by the benefit systems. The structure of income taxes is presented in figure 2. In principle there are two different levels, below the breaking point there the marginal tax rates depends on the level of municipal tax rate as well as a basic deduction and above the break point there the marginal tax rate is the municipal rate plus a 20% governmental rate.

3 Economic Model and Empirical Specification

We assume that each household chooses husband's hours of work (h_h), wife's hours of work (h_w), consumption (C), and social assistance (d_w=1 if the household receives social assistance and 0 otherwise) by maximizing a utility function given the budget set in (1). Following van Soest (1995), a translog specification of the direct utility function is used, and for any specific household we have:

(2)

$$
\begin{aligned}
U\left(C,h_h,h_w\right) = {} & \beta_C \log\left(C\right) + \beta_h \log\left(TE-h_h\right) + \beta_w \log\left(TE-h_w\right) + \\
& \beta_{CC}\left(\log(C)\right)^2 + \beta_{hh}\left(\log\left(TE-h_h\right)\right)^2 + \beta_{ww}\left(\log\left(TE-h_w\right)\right)^2 + \\
& 2\beta_{Ch}\log(C)\log\left(TE-h_h\right) + 2\beta_{Cw}\log(C)\log\left(TE-h_w\right) + 2\beta_{hw}\log\left(TE-h_h\right)\log\left(TE-h_w\right) \\
& -\phi d_w
\end{aligned}
$$

where it is assumed that the disutility from social assistance participation (d_w) is separable from the utility of leisure and consumption (Moffitt (1983) and Hoynes (1996)). The disutility from social assistance is included to account for nonparticipation among eligible families.

The Total Endowment of time (TE) is set to 4,000 hours/year.[7] As mentioned above, the husband and wife are assumed to choose among seven different working states, respectively, ranging from zero up to 3,000 hours/year. Hence, for the household there are altogether 49 different hour's combinations.

The flexible specification in equation (2) does not automatically fulfill the quasi-concavity conditions. However, these conditions can be tested *ex post*. This contrasts a continuous model in which quasi-concavity has to be imposed *a priori* in order to guarantee model coherency.

In order to implement the model, we also have to specify the nature of heterogeneity in household preferences and the stochastic disturbances. Heterogeneity in preferences for leisure is introduced as

$$\beta_h = \sum_{i=1}^{k} \beta_{hi} x_{hi} + \theta_h$$

(3)

$$\beta_w = \sum_{i=1}^{k} \beta_{wi} x_{wi} + \theta_w$$

[7] TE can also be regarded as a parameter that can be estimated together with all other parameters. van Soest (1995) reports that the results are insensitive towards the choice of TE.

where the x-variables consist of observed individual and family characteristics, such as age, children, and born in Sweden or not. The θ´s represents unobserved variables that affect preferences for leisure. It is reasonable to assume that an important source for population heterogeneity in terms of preferences for leisure is unobserved. In order to account for this, we formulate a finite mixture model, which allows for unobserved heterogeneity in a very flexible way without imposing a parametric structure. This idea of incorporating unobserved heterogeneity origins from Heckman and Singer (1984) and there exist a number of applications in duration data (Ham and Lalonde (1996)), count data (Deb and Trivedi (1997)), and labor supply (Hoynes (1996)). Heckman and Singer (1984) also showed that estimation of finite mixtures might provide a good discrete approximation even if the underlying distribution is continuous.

To be specific, we assume that there exist M different (θ_{hj}, θ_{wj}) pairs that determine the spouse's preferences, each observed with probability π_j (where $\pi_j > 0$ and $\Sigma\pi_j = 1$). This specification allows for arbitrary correlation between the husbands and wives labor supply. The interpretation of these unobserved heterogeneity parameters are straightforward, and a high value simply implies a high preference for leisure.

The specification of social assistance participation takes the form

(4) $$\phi = \mu + \sigma_h \theta_{hj} + \sigma_w \theta_{wj} \qquad j=1,\ldots,M$$

where μ are a function of observed individual and family characteristics, such as number of children, educational attainment and country of birth. The σ´s are parameters to be estimated. This specification is very general and allows for correlation between the spouses' preference for work and social assistance. This way of allowing for correlation across alternatives is based on factor loading technique, see for instance Ham and Lalonde (1996).

Adding an additive error term to the utility function in equation (2), drawn from the extreme value distribution, results in the conditional logit model.[8] The contribution to the likelihood function for a given household (i',j',k') becomes

$$(5) \qquad (p \mid \theta_h, \theta_k)_{i'j'k'} = \frac{\exp(U_{i'j'k'})}{\sum_{i,j,k} \exp(U_{ijk})}$$

where i and j indicates husbands and wife's hours, respectively, and k indicates social assistance participation. This expression simply denotes the probability that the utility in the observed state is the highest amongst all possible hours and social assistance combinations.

In our specification of classification error, we follow MaCurdy et al (1990) and Hoynes (1996), and assume a multiplicative classification error structure. Let H_h and H_w denote reported hours and h_h and h_w optimal (discrete) hours. The multiplicative classification error specification is given as

$$(6) \qquad H_i = h_i e^{\varepsilon_i} \qquad \text{with } \varepsilon \sim N\left(-\tfrac{1}{2}\sigma_i^2, \sigma_i^2\right) \qquad \text{for i=h,w}$$

Thus, zero hours are observed with certainty but when optimal hours are positive they differ from reported hours by a factor of proportionality.

In presence of unobserved heterogeneity and classification errors, the contribution to the likelihood is given by

$$(7) \qquad l = \sum_{m=1}^{M} \pi_m \left((p \mid \theta_{hm} \theta_{wm})_{i'j'k'} g_h g_w \right) \delta_{i'j'k'}$$

[8] Alternatively, we could assume that the errors were drawn from a normal distribution. However, this would require evaluation of high-dimensional integrals, which would be intractable in our framework. Recall that we assume that each household chooses among 98 different state combinations. We also believe that the restrictiveness with the extreme value distribution is smaller when we incorporate unobserved heterogeneity for reasons already discussed.

where $\delta_{i'j'k'}$ is an indicator for the observed state for each household, and g_h and g_w are densities for classification error for the husband and wife. The assumptions presented in (6) implies

$$(8) \qquad g_i = \begin{cases} 1 & \text{if } \mathrm{H_i} = 0 \text{ or } \mathrm{h_i} = 0 \\ \frac{1}{\sigma_i}\phi\left(\dfrac{\left[\log(H_i) - \log(h_i)\right] + \frac{1}{2}\sigma_i^2}{\sigma_i}\right) & \text{else} \end{cases} \qquad \text{i=h,w}$$

4 Data

The data used in the empirical analysis are drawn from the 1993 cross-section of the Swedish *Household Income Survey* (HINK) supplied by Statistics Sweden. HINK provides information on labor market activities and incomes for a random sample of Swedish households.

In order to obtain the sample of interest, several selections have been imposed. To start with there is 3,078 households of married/cohabitant spouses with children 0-12 years of age. From this the following exclusions have been done; spouses younger than 18 or older than 64, students, early retired or own employed and finally a few extreme outliers in hourly wages. After these selections the resulting sample size is 1,603 households.

Information about yearly hours of work is based on survey questions but the hourly wage rate is defined based on register information. Information about full-time equivalent monthly earnings are available. In order to obtain hourly wage rates monthly earnings are divided by 165, i.e. full time monthly hours of work. Thus, this definition avoids the division bias problem in wage rates. However, unfortunately we do not have monthly earnings for the full sample. In the data used for this study there are information available for all individuals employed in the public sector but for the private sector the information is missing for about half of the sample. Hourly wage rates for individuals with missing information have been imputed using an earnings equation. Non-labor income contains income from capital gains and public transfers such as unemployment insurance and different allowances.

Non-labor income is divided in two parts, taxable and non-taxable. Taxable non-labor income consist of: car or expense allowance, job-related injury compensation, rehabilitation compensation, training allowance for labor market training, daily allowance in the case of unemployment, cash labor-market support and other taxable transfers. The main component in non-taxable non-labor income is child allowance, which every family with a child below the age of 16 receives.

Deductions consist of several components: deductions for business expenses, general deductions for retirement insurance, general deductions for periodical supports and loss related deductions. The precision in this variable is a good illustration of the advantage of using register data. It is difficult to obtain a reliable measure of deductions from a survey. Of course all errors in the income variables would lead to errors in the imputed budget set. It is therefore crucial to have income data of a high quality in studies of labor supply and taxes.

In Table 1 we present descriptive statistics for the sample used in this study. Hours of work refer to annual hours and the reported average values are 1,904 for males and 1,438 for females. It is a well-established fact that the participation rates in Sweden are high, both for men and women. This is confirmed in our data where 96 percent of the men performed market work and 93 percent of the women. The distribution of working hours is presented in Figure 3. The husband's hours are concentrated at 40 hours per week whereas there is much more variation in the wife's hours.

The mean hourly wage is 91 SEK for males and 83 SEK for females. For non-workers as well as part of the privately employed the wage rates were imputed using regression methods. A standard Mincer-type of wage equation was estimated separately for males and females.[9]

The level of education is quite similar for both spouses. About 60 percent have a high school degree and about 15 percent have a university degree. There are almost as many households living in large cities (Stockholm, Gothenburg or Malmo), as in smaller cities and only 14% of the households live on the countryside. Of the

[9] The regression results are available from the authors on request.

husbands 91 percent are born in Sweden compared to 89 percent of the wife's. In about 6 percent of the households both spouses are born outside Sweden. In our sample about 5 percent received social assistance during 1993. A household was defined as a social assistance recipient if it received some assistance for at least one month during the year. It should be noted that most of the households that received social assistance in this sample only received it for a short period. Of all the social assistance recipients, about 50 percent received it for three months or less and about 20 percent for more than seven months. We can see from Table 1 that the social assistance norm for this sample ranges from 59,000 SEK to 216,000 SEK. To be eligible to social assistance the household income must be below this norm and if that's the situation the household will get the difference between the household income and the norm in social assistance.

5 Results

The estimated parameters of our structural model are presented in Table 2. We present estimates for two specifications, one that excludes social assistance and one where social assistance participation is modeled jointly with the household labor supply. From Table 2 it follows that the estimates for both specifications are similar, and consequently, we only discuss the results based on the full structural model with social assistance. At these estimates the utility function fulfills the conditions for quasi-concavity for almost all households (only 8 household out of 1603 did not fulfill the condition), evaluated at observed hours and consumption. Since there is a fair amount of variation in both hours and consumption, this means that the utility function is concave over a large region. Further, since the estimated utility function fulfills the theoretical requirements it can be used for predictions and simulation.

The first set of estimates in Table 2 refers to husband's preference for leisure and the second set to the wife's. As expected, presence of children has a strong negative effect on female work preferences and no effect on males. The effect of young children is similar for males and females and is estimated with relatively high precision for husbands and wives who have children between 3-6 years of age.

Estimated age effects are as expected. However, the effect is not significant for males but is significant for females. Weather the husband or wife are born in Swedish or not have a strong significant effect for both spouses. Most β-estimates with respect to consumption and leisure are estimated with a high precision.

The first estimated pair of support points (θ_{h1} = 55.88 and θ_{w1} = 17.82) identifies households where the husband has a high preference for leisure and the wife a low preference. Thus, estimated probability (π_1 = 0.13) indicates that about 13 percent of the sample belongs to this category. The majority, 82 percent, of the households belongs to the second group where both spouses have a low preference for leisure. The third group is households where the husband has a low preference for leisure and the wife a high preference and about 5 percent of the sample belongs to this group.

The last set of results reported in Table 2 refers to the disutility of social assistance. The constant, μ, indicates that there is a positive and significant stigma effect. Thus, social assistance participation lowers the utility level of the household. The other variables included in the welfare equation have a negative and significant effect.

The estimated loading parameters indicate a negative correlation between social assistance and unobserved elements of work effort. Similar to the results reported in Hoynes (1996), this correlation is also higher for the females work effort. The estimated negative covariance between social assistance and labor supply can be taken as support for the hypothesis of self-selection into social assistance.

A well-known problem in labor supply models is poor ability to fit observed distribution of hours of work. One approach to improve the fit of these models is to include controls for fixed costs of work, see Kapteyn et al (1990) and van Soest (1995). In our approach, the estimated support points are used in the calculation of predicted hours of work. This produces a distribution of hours rather similar to the observed one as can be seen from Figure 3. The upper panel shows the observed and predicted distributions for husbands. The results indicate a close replication of the frequency of non-workers as well as for hours below 2,500. The peak in the distribution, around 2,500 hours per year, is however overestimated. About 88 percent of the husbands belong to this category according to our predictions, whereas only 77 percent are observed in that class.

The lower panel of Figure 3 displays the corresponding distribution of hours of work for women. As expected, there is more variation in working hours. Our model is actually able to capture this variation quite well. However, lower hours are underestimated and the peak at 2,000 hours is overestimated.

The effects of wage and income changes are assessed using simulations. Specifically, income and wages were increased by 1 percent and the resulting changes in predicted working hours were calculated. The results in Table 3 imply that working hours are quite insensitive for income and wage changes, especially for males. For instance, an increase in husbands wages by 1 percent (everything else constant) do increase his working hours with 0.01 while wives hours of work increases with 0.02 percent. The corresponding results for wives show an increase in hours by 0.32 percent, but no effect on husband's hours. The estimated income effects are positive but close to zero for females and no effect for males.

Next the structural model with stigma is used for a simple simulation experiment. In order to evaluate the effect of moving from a progressive tax system towards a proportional one, we simply drop the federal tax rate of 20 percent above the break point (see figure 2). As a result of the simulated tax change, working hours increase on average by 2.4 percent for wives and by 0.1 percent for husbands. The resulting increase in disposable income is 4.4 percent and the decrease in tax revenues is almost 6 percent. Thus, despite the fact that relatively few females have earnings above the breakpoint the change in female hours is still larger than for males. In fact only 0.4 percent of the males and 8.4 percent of the females change their working hours. This is a natural consequence of the discrete approach of modeling labor supply where the dominating prediction is no change in working hours.

A more detailed listing of the result for the whole sample is given in Table 4. This table also presents the welfare effects of the tax reform. We chose equivalent variation (EV) as our money metrics of a welfare change. EV is measured as the amount of money added or subtracted from the households' disposable income under the initial tax rules in order to make the household indifferent between the initial and the alternative tax system. As such, EV summarizes the household's net welfare change associated with behavioral responses. As mentioned above, in our simulation the majority of the household members do not change their working hours and in these cases EV just measure the change in disposable income before and after the tax change.

The average EV for the whole sample is 11,725 SEK. However, there is a substantial variation across the households. Table 5 lists EV for different levels of household disposable income. All EV-values are non-negative, which suggests that there are welfare gains from the tax change. However, there are dramatic differences in EV depending on the level of household income and the estimated average EV for the poorest 10 percent is 3,907 SEK/year compared to 52,982 SEK/year for the richest 10 percent. The gain from the simulated tax reform is quite small evaluated for all households below the median.

As mentioned above, the average effects of the tax cut on working hours were relatively small. However, this does not imply that the effects for all income groups have to be small. Table 6 lists the predicted changes in working hours for different incomes. The results suggest a relatively strong increase in wife's hours in high-income households.

To summarize, reducing the progressivity in the Swedish tax system has considerable welfare effects. The difference in these effects between poor households and rich households is substantial. However, the effect on working hours is quite small and there will be a sharp decline in tax revenues.

The final experiment evaluates the effects of a change in the level of social assistance (the norm). First, in Table 7, the results are presented for an increase by 25% and secondly, in Table 8, a decrease by 25%. For this experiment it is interesting to compare the predictions of the two specifications, with and without stigma. The effects of changing the level of social assistance clearly demonstrate the importance of the stigma effect. Our results for the model without stigma shows that labor supply for wives in low-income household are quite sensitive. An increase in the norm by

25% reduce wife's hours in the lowest income group by 17%, whereas a decrease in the norm by the same amount produce an increase in hours by about 7%. The corresponding numbers for the model with stigma are 6% percent respectively 1%. Thus, the results indicate that the stigma effect affect individuals behavior when we change the norm for social assistance.

6 Conclusions

In this paper, we used a sample of Swedish households with detailed information on incomes and benefits and estimated a structural household labor supply model. We formulated a model where labor supply and participation in social assistance programs were jointly determined. Further, the labor supply and social assistance participation decisions were treated as a discrete choice problem, and we assumed that these choices follow a simple conditional logit rule. We used a micro simulation model to evaluate consumption bundles at different hours of work combinations. In addition, we allowed for unobserved individual-specific effects and also for these effects to be correlated across alternatives. The unobserved effects were assumed to be drawn from a discrete distribution, and the correlation across alternatives was modeled using factor-loading techniques. Classification error in hours was allowed for by using a multiplicative classification error specification.

The estimates from the structural model yielded small wage and income elasticities, especially for the husbands. A tax simulation showed that reducing the progressivity in the Swedish tax system may have considerable welfare effects. It also showed that these effects might differ substantially between poor households and rich households. The effect on working hours from the reform was quite small and tax revenues were predicted to drop significantly. Finally, the evaluation of a change in the welfare system showed that the stigma effect had a substantial impact.

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Appendix

Figure 1. Household budget sets.

Low Income **High Income**

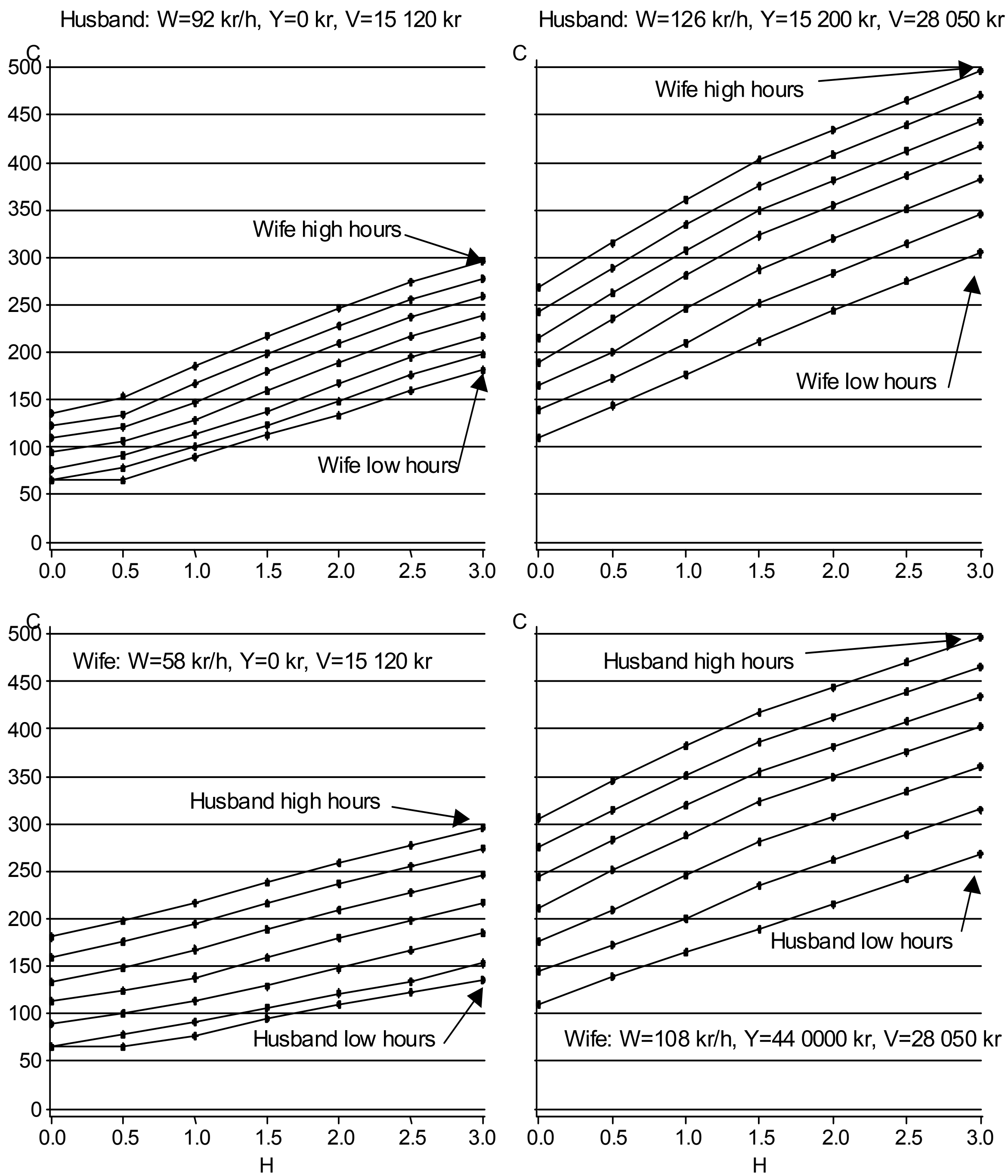


Figure 2. Income taxes in Sweden 1996

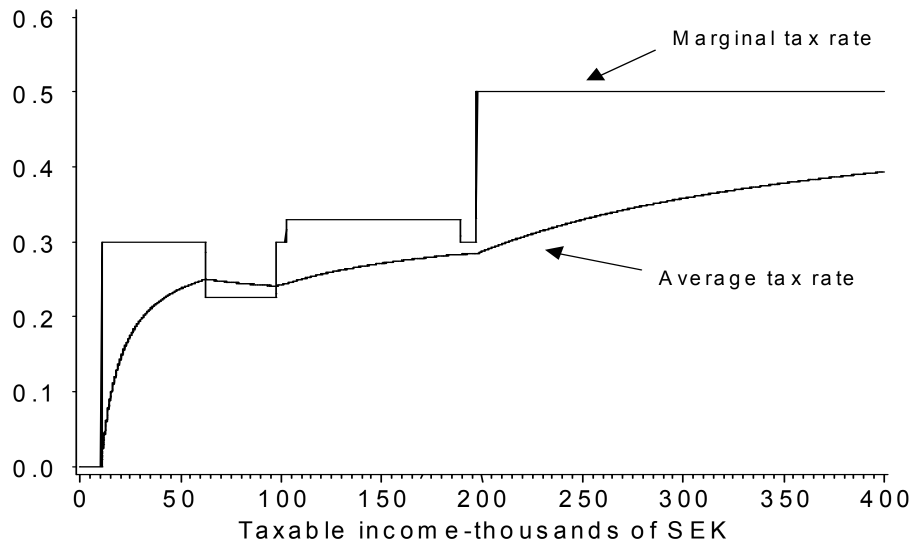
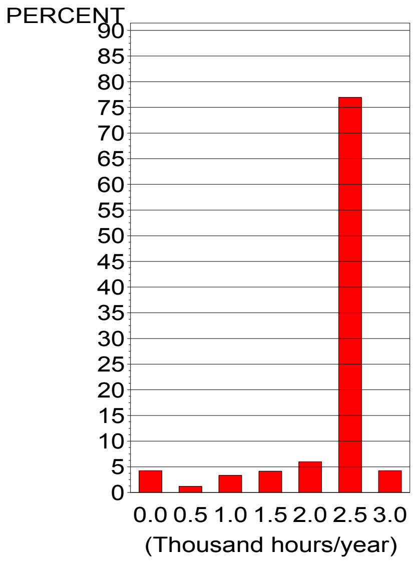
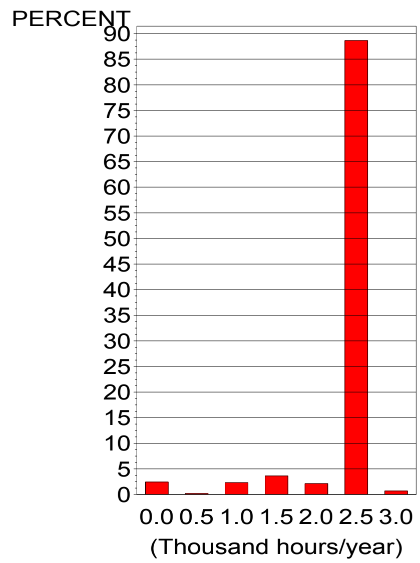


Figure 3. Observed and predicted hours of work

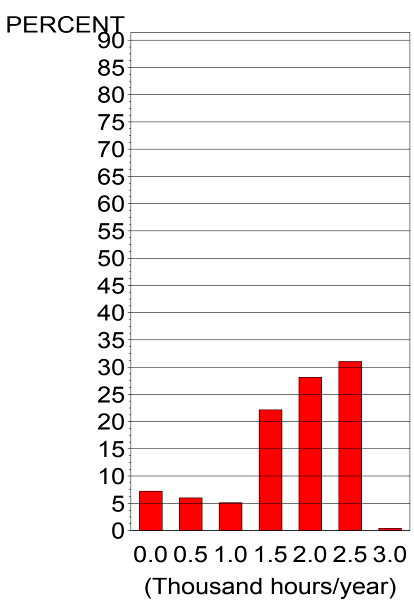
Husband observed



Husband predicted



Wife observed



Wife predicted

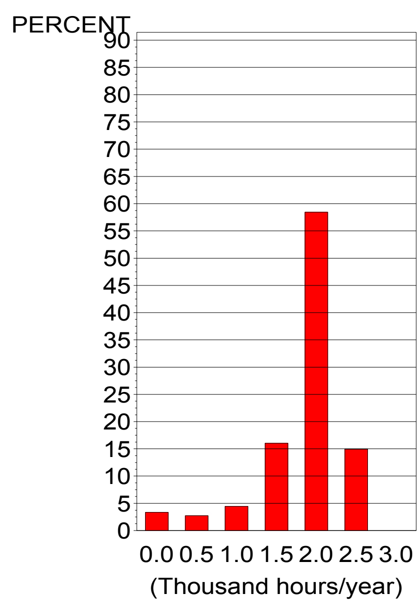


Table 1. Sample Statistics of variables used for estimation.

Variables	Mean	Minimum	Maximum
Husband:			
Age	38	21	62
Education, primary 1=yes, 0=no	0.23	0	1
Education high school 1=yes, 0=no	0.57	0	1
Education university 1=yes, 0=no	0.20	0	1
Working hours per year	1,904	0	3,840
Working 1=yes, 0=no	0.96	0	1
Wage/hour SEK	91	36	305
Taxable non-labor income SEK per year	10,795	0	198,010
Non-taxable non-labor income SEK per year	10,006	375	68,625
Deductions SEK per year	6,728	0	134,500
Swedish, 1=yes, 0=no	0.90	0	1
Wife:			
Age	35	19	54
Education, primary 1=yes, 0=no	0.22	0	1
Education high school 1=yes, 0=no	0.63	0	1
Education university 1=yes, 0=no	0.15	0	1
Working hours per year	1,438	0	3,012
Working 1=yes, 0=no	0.93	0	1
Wage/hour SEK	83	59	221
Taxable non-labor income SEK per year	10,883	0	291,218
Non-taxable non-labor income SEK per year	10,006	375	68,625
Deductions SEK per year	3,648	0	146,200
Swedish, 1=yes, 0=no	0.89	0	1
Household:			
Number of children 0-12 years old	2.01	1	6
Children 0-2 years old 1=yes, 0=no	0.40	0	1
Children 3-6 years old 1=yes, 0=no	0.30	0	1
Large cities 1=yes, 0=no	0.41	0	1
Medium cities 1=yes, 0=no	0.45	0	1
Country side 1=yes, 0=no	0.14	0	1
Social assistance 1=yes, 0=no	0.05	0	1
Social Assistance Norm	108,148	59,856	215,899
Both immigrants, 1=yes, 0=no	0.06	0	1

Table 2. Estimates of the parameters of the utility function.

Variables	Coefficient	Estimates	Standard errors	Estimates	Standard errors
Husband:					
Number of children, 0-12 years old	β_{h1}	-0.1661	0.2707	-0,3142	0,2732
Children 0-2 years old 1=yes, 0=no	β_{h2}	0.7554	0.6524	0,8508	0,6510
Children 3-6 years old 1=yes, 0=no	β_{h3}	1.8498	0.6534	1,9606	0,6474
Age Husband	β_{h4}	-0.2093	0.3068	-0,2550	0,3138
Age husband squared/100	β_{h5}	0.3032	0.4012	0,3606	0,4076
Swedish, 1=yes, 0=no	β_{h6}	-3.2194	0.7861	-3,1574	0,7864
Wife:					
Number of children, 0-12 years old	β_{w1}	0.3653	0.1419	0,2937	0,1411
Children 0-2 years old 1=yes, 0=no	β_{w2}	0.2516	0.3433	0,3104	0,3408
Children 3-6 years old 1=yes, 0=no	β_{w3}	0.7951	0.3106	0,7852	0,3084
Age Wife	β_{w4}	-0.2683	0.1556	-0,2732	0,1546
Age Wife squared/100	β_{w5}	0.4166	0.2182	0,4250	0,2171
Swedish, 1=yes, 0=no	β_{w6}	-1.0528	0.3613	-1,0495	0,3571
Consumption	β_C	8.8453	1.8214	9,5891	3,7378
Consumption squared	β_{CC}	5.6892	0.5483	5,9879	1,3181
Husband hours squared	β_{hh}	-19.5550	0.7706	-19,2561	0,7738
Wife hours squared	β_{ww}	-3.6276	0.4740	-3,2337	0,4700
Husband hours times consumption	β_{Ch}	0.3525	0.5231	0,6823	0,6117
Wife hours times consumption	β_{Cw}	-1.0197	0.3358	-1,0028	0,5076
Husband hours times wife hours	β_{hw}	0.2032	0.5041	-0,3466	0,5076
Classification error, Husband	ε_h	0.1235	0.0022	0,1235	0,0022
Classification error, Wife	ε_w	0.2542	0.0046	0,2542	0,0046
Heterogeneity, Husband:	θ_{h1}	55.8776	6.1409	56,2249	6,3953
	θ_{h2}	23.5811	5.9968	23,8462	6,2340
	θ_{h3}	24.2124	6.3535	24,8505	6,6820
Heterogeneity, Wife:	θ_{w1}	17.8197	3.2353	17,8205	3,4227
	θ_{w2}	17.0572	3.0386	17,1675	3,2332
	θ_{w3}	26.5687	3.4762	26,9780	3,7336
Heterogeneity Probabilities:	π_1	0.1289	0.0087	0,1303	0,0088
	π_2	0.8182	0.0229	0,8240	0,0229
	π_3	0.0529	---------	0,0457	---------
Social assistance participation:					
Constant	μ_0	11.9431	2.0220	---------	---------
Number of children, 0-12 years old	μ_1	-0.4059	0.1433	---------	---------
Education Husband, primary school, 1=yes, 0=no	μ_2	-0.8245	0.3078	---------	---------
Education Wife, primary school, 1=yes, 0=no	μ_3	-0.9737	0.3143	---------	---------
Both immigrants, 1=yes, 0=no	μ_4	-1.0339	0.3807	---------	---------
Covariance husband hours, social assistance	σ_h	-0.0860	0.0123	---------	---------
Covariance wife hours, social assistance	σ_w	-0.2277	0.0833	---------	---------
Log of Likelihood Function		-3191,72		-2989,93	

Table 3. Change in working hours as wage and income change 1%.

	Male hours Percentage change	Female hours Percentage change
Male wage increase 1%	0.015	0.019
Female wage increase 1%	0	0.319
Household non-labor income increase 1%	0	0.019

Table 4. Tax simulation: Before and after tax changes for the whole sample

	Mean	Minimum	Maximum	Variance
Husband:				
Working hours before tax change	2,114	0	2,750	207
Working hours after tax change	2,116	0	2,750	208
Wife:				
Working hours before tax change	1,600	0	2,250	264
Working hours after tax change	1,638	0	2,250	288
Household:				
Disposable income before tax change	296,277	19,794	1,162,010	7,987,800
Disposable income after tax change	309,300	19,794	1,190,917	11,192,430
Taxes paid before tax change	69,238	0	335,978	1,258,568
Taxes paid after tax change	65,026	0	249,511	715,619
Equivalent variation	11,725	0	142,634	365,383

Table 5 Tax simulation: Equivalent variation for different income levels

	Mean	Minimum	Maximum	Variance
Poorest 10:th percent	3,907	0	80,206	251,622
Poorest 25:th percent	1,912	0	80,206	107,955
Below the median	2,331	0	80,206	70,322
Richest 25:th percent	33,561	0	142,635	582,594
Richest 10:th percent	52,982	0	142,635	603,056

Table 6 Tax simulation: Predicted changes in hours of work for different income levels.

	Husband	Wife
Poorest 10:th percent	0%	0.40%
Poorest 25:th percent	0%	2.54%
Below the median	0.09%	1.29%
Richest 25:th percent	0.17%	3.17%
Richest 10:th percent	0.27%	3.10%

Table 7 Social assistance simulation: Predicted changes in hours of work for different income levels when the norm increases with 25%.

	Husband		Wife	
	Stigma	**No stigma**	**Stigma**	**No stigma**
Poorest 10:th percent	0.01%	-2.71%	-6.05%	-16.98%
Poorest 25:th percent	-1.15%	-1.71%	-4.39%	-7.28%
Below the median	-0.53%	-0.79%	-1.89%	-3.38%
Richest 25:th percent	0%	0%	0%	0%
Richest 10:th percent	0%	0%	0%	0%

Table 8 Social assistance simulation: Predicted changes in hours of work for different income levels when the norm decreases with 25%.

	Husband		Wife	
	Stigma	**No stigma**	**Stigma**	**No stigma**
Poorest 10:th percent	0.48%	4.57%	1.01%	9.52%
Poorest 25:th percent	0.23%	1.22%	0.51%	2.73%
Below the median	0.11%	0.58%	0.23%	1.18%
Richest 25:th percent	0%	0%	0%	0%
Richest 10:th percent	0%	0%	0%	0%

Labor Supply and Welfare Participation of Single Mothers in Sweden

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Abstract:

In this paper we analyze the effects of changes in income taxes, cost of childcare and social assistance on labor supply for single mothers households. We specify and estimate a discrete structural labor supply model where labor supply participation in welfare is determined jointly. Our results show that there is a positive and significant stigma-effect associated with welfare participation. Fixed costs of working is an important factor in a single mother's decision to enter the labor market. We find a negative covariance between social assistance and labor supply, which implies self-selection into welfare. The results from the micro simulation show rather small average incentive effects, but there are some substantial effects for low-income groups. Welfare effects from the childcare reform are quite similar across all income deciles, even if predicted increases in hours of work are substantial for the poorest single mother households.

Key words: Single mothers, Labor Supply, Welfare Participation, Unobserved Heterogeneity, Fixed Costs of Working, Tax Simulation.

JEL Classification: J22, I38

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1 Introduction

High marginal effects caused by the interaction between social security transfers and taxation characterizes Sweden and many other western countries. For unemployed single parents the net benefit from working is almost negligible. According to a report from the Swedish Ministry of Finance (Eklind et. al., 1997) the net replacement rate for this group is about 90%. The question addressed in this paper is what the size of the incentive effects for individuals facing such high marginal effects in Sweden is. To answer this question a discrete labor supply model is specified, estimated and simulated. Specifically, we analyze the effects of changes in income taxes, cost of childcare and social assistance on labor supply for single mothers households. Two recent papers addressing similar questions in the US and the UK are Hu (1999) and Duncan and Giles (1998).

The data used for this study is the 1996 wave from the Swedish LINDA database. This data is completely based on register information and provides us with a large sample size together with an unusually high quality of tax and income data. Another attractive quality of this data is that even hourly wage rates are obtained from the registers.

One distinctive feature of this study is the detailed construction of the budget sets that incorporates the most relevant tax and benefit system. Not surprisingly, this results in non-convex budget sets and these non-convexities are considered in the estimation of the model. We assume that individuals choose their working hours from seven different discrete working states, ranging from not working at all to a maximum of 3,000 hours per year. An important advantage of the discrete approach compared to the continuous specification is that the coherency conditions do not have to be imposed *a priori* (Van Soest, 1995).

Special attention is devoted to the effects of fixed working costs related to the labor market participation. When individuals make a decision to participate in the labor force they face certain entry costs. The effects of these costs on the decision to participate and on hours of work among workers, are obvious. These costs consist of costs of transportation and costs of childcare. The fixed costs increase the reservation wage (and reservation hours) that is observed by the fact that very few work more than zero hours but less than 750 hours per year.

We find that there is a positive and significant stigma-effect associated with welfare participation. Fixed costs of working is an important variable in a single mothers' decision to enter the labor market. The estimated negative covariance between social assistance and labor supply implies self-selection into welfare. Our results from the micro simulation give rather small average incentive effects, but there are some substantial effects for different income groups. Welfare effects from the childcare reform are quite similar across all income deciles, even if predicted increases in hours of work are considerable for the poorest single mother households.

The paper is organized in the following way. We start the discussion in Section 2 by introducing the economic status of single mothers in Sweden. Section 3 gives a detailed description of the budget set and the relevant benefit programs. Section 4 presents the economic model and Section 5 describes the data used in the analysis. In Section 6 we present the results, while the conclusions are found in Section 7.

2 Single mothers in Sweden

The motivation for studying labor supply decisions of single mothers is first of all, the increase in the number of single-headed households in Sweden. Today, almost 25 % of all households with children, and about 90 % of one-parent households, are female headed. Secondly, the main trends in the Swedish labor market consist of a strong rise in the female labor force participation rate and a slight long-run decline in both working hours and participation rates of men. Especially during recent years, the labor force participation among single mothers has been higher (over 80 %) than among cohabiting mothers (76 %). It has also been slightly higher than among cohabiting or married fathers (79 %). Thirdly, since single mothers are the net receivers of transfer payments, they are also the most vulnerable group when there are cuts in the benefit system, meaning that they might also be the most sensitive in their labor supply responses (Statistic Sweden, 1997).

A large part of the total disposable income of single parent households consists of different social benefits, on average 30 %. Also, most of the publicly provided social services (for example, childcare) are offered for free or at discounted rates to these low-income households. In Sweden, the major family benefit programs are housing allowance, parental allowance, child allowance and social assistance. More than 70 % of single mothers received housing allowance and almost one third received social assistance in 1996. Of all social welfare participants, 15 % were single mothers. In the group of single mothers, the young, the immigrants and the mothers with only a low level of formal education have the lowest income (Socialdepartementet, 1996).

Even though the single parent families are relatively poor compared to other types of households in Sweden, they are much better off in an international comparison. About 4 % of these households in Sweden lie below the poverty line (having less than half of the median income), whereas in the US, Great Britain and Germany the figures are 60 %, 14 % and 22 %, respectively (Hobson and Takahashi, 1996). Of all single mothers in Sweden, 85 % are working, either part-time or full-time. One half of the mothers with young children are working part-time.

During the economical recession, starting in the beginning of the 1990s, single mothers suffered more from unemployment than any other group in the labor market. In 1995, 39 % of single mothers who had young children were unemployed, which is twice as high as before the recession. Another indicator of the economic hardship of this group is given by income information from Statistics Sweden. From 1989 to 1997 the median disposable income for single parents decreased by more than 10 % (Statistics Sweden, 1999).

3 Budget Constraints

The empirical model takes as a starting point a static model of labor supply. It is assumed that individuals determine their hours of work and disposable income by maximizing a utility function U(C, h), subject to the following budget constraint:

$$(1) \qquad C = \mathrm{Wh} + \mathrm{Y} + \mathrm{V} - \mathrm{t(I)} + \mathrm{B}$$

where C is the income after taxes, W is the gross wage per hour, h is hours of market work per year and Y and V are taxable and non-taxable non-labor income per year, respectively. Income taxes are determined by the tax function $t(\cdot)$, where the argument is taxable income. Finally, B is the amount of household specific means-tested subsidies.

The three major programs included in B are housing allowance, social assistance and cost of childcare. Housing allowance is determined by nationwide rules and is mainly directed to families with children. The amount of housing allowance a household is entitled to is determined by the net household income, housing expenditures, the number of children and the ages of the parents.

The rules determining social assistance are based on a rather complicated system and they also differ across municipalities. For each municipality and each type of family, we calculate a Welfare Eligibility Limit (WEL) (the maximum level of disposable income to qualify for social welfare) based on the information provided by the Swedish municipalities. The amount of social assistance a family receives is simply the difference between the WEL and the household's disposable income.

Childcare payment schemes vary with municipality, the household's income and the number and ages of the children (the rules in 1996). The program for calculating taxes and benefits used here is a modified version of FASIT, a tax-benefit program designed at the Statistics Sweden and the Ministry of Finance. FASIT includes most of the municipalities and their fee structures for childcare. For municipalities that are not included, different methods of imputation have been used. We have added an assumption to FACIT regarding the use of childcare given different hours of work.

Figure 1 illustrates the non-convexity of the budget set for a low-income household with two children. The budget set is evaluated at seven discrete points ranging from 0 to 3,000 hours of work per year. Non-convexity of the budget set at lower hours of work is visible. There is no return at all from an increase in labor supply from 0 to 500 hours, as is shown in the lower panel. The reason for this is a 100 % decline in social assistance for the earned income. Very few employed mothers in our sample are eligible for social assistance. Thus, the effect of social assistance is basically on the decision to work only a few hours or not to work at all. A second kink can be observed at 1,500 hours of work. The reason for this kink is primarily the housing allowance that is ceased and to some extent the income taxation. Income taxes are of course important, but for low-income earners they do not contribute much to the non-convexity of the budget set. The tax system is displayed in Figure 2 and, as shown, the main progressive part of the income tax comes from the increase in earnings at about 200,000

SEK.[4] Not many single mothers obtain this high of an income.

Cost of childcare is an important component in the budget of a household since it typically increases over the whole hours of work interval, although the payment scheme differs across municipalities. The most typical scheme is one in which the fees increase progressively with income.

To summarize, the main source of the non-convexity in budget sets is the generous and sharply declining transfer system designed to equalize the income distribution. Also, the progressive income-tax system produces non-convexities, but they are not as large as those produced by the benefit systems.

4 Economic Model and Empirical Specification

The analysis is based on a static model of labor supply. Individuals determine their hours of work (or rather leisure) and disposable income by maximizing a utility function, U(C,h), subject to a budget constraint. We assume that every individual maximizes her utility choosing the hours of work, h; disposable income, C; and welfare participation (d=1, if the household receives welfare, 0 otherwise), subject to the specific budget constraint (1) and the total Time Endowment, *TE*, which is set at 4,000 hours. The econometric model used here closely follows the model used in Van Soest (1995 and 2000) and Flood, Hansen and Wahlberg (1999). Preferences for disposable income and leisure are described by a direct translog utility function, specifically

$$\begin{aligned} U(C,h) = {} & \beta_Y \log(C) + \beta_h \log(TE-h) + \beta_{YY}(\log(C))^2 + \beta_{hh}(\log(TE-h))^2 \\ & + 2\beta_{Yh}\log(C)\log(TE-h) - \varphi d + \varepsilon_h \end{aligned} \tag{2}$$

The utility function is assumed to be increasing in disposable income, and decreasing with respect to hours of work and welfare participation. We assume that there is a negative side effect associated with receiving social assistance. The disutility from welfare participation, d, is assumed to be separable to allow for non-participation among eligible households. If the disutility from welfare participation is assumed to be separable, it affects the decision to

[4] 1 Euro ≈ 9 SEK.

participate in welfare, but not the labor force participation decision conditional on welfare receipt. We assume the unobserved preference component, ε_h, relating to the particular hours alternative, has an extreme value distribution. The question of the effects of different types of government transfer payments on labor supply has been examined in many studies, for example in Moffitt (1983, 1992), Hagstrom (1996), Hoynes (1996), and Keane and Moffitt (1999).

Fixed costs of working, FC, are incorporated in the utility function. We assume that the argument log (C_i) in the utility function for employed individuals should be replaced by log (C_i) – log (FC_i).[5] Since the utility increases with income, positive fixed costs decrease the utility of the employed but do not affect the utility of those not employed. The level of fixed costs depends on the family composition and some other family and societal characteristics. Following van Soest (2000), we model the costs log-linearly as
$\log(FC_i) = \gamma_1 z_1 + \gamma_2 z_2 + \ldots + \gamma_k z_k$.

The flexible specification in Equation (2) does not automatically fulfill the quasi-concavity conditions. However, these conditions can be tested *ex post*. This contrasts a continuous model in which the quasi-concavity has to be imposed *a priori* in order to guarantee the coherency of a model.

In order to implement the model, we have to specify the nature of heterogeneity in household preferences and the stochastic disturbances. Heterogeneity in preferences for leisure is introduced as

(3) $$\beta_h = \sum_{i=1}^{k} \beta_{hi} x_{hi} + \theta_h$$

where the x variables consist of observed individual and family characteristics. θ's represents unobserved variables that affect preferences for leisure. It is reasonable to assume that an important source of population heterogeneity in terms of preferences for leisure is unobserved. In order to take this into account, we formulate a finite mixture model, which allows for unobserved heterogeneity in a very flexible way without imposing a parametric structure. This idea of incorporating unobserved heterogeneity originates from Heckman and

[5] An alternative would be to correct for fixed cost in levels but this led to a nonzero probability of negative income, which is not allowed by our translog utilty function.

Singer (1984), and there is a number of applications in duration data (Ham and Lalonde, 1996), count data (Deb and Trivedi, 1997), and labor supply (Hoynes, 1996). Heckman and Singer (1984) also show that estimation of a finite mixture model might provide a good discrete approximation even if the underlying distribution is continuous.

To be specific, we assume that there are M different θ-parameters that determine the preferences for leisure. Each one is observed with probability π_j (where $\pi_j > 0$ and $\Sigma\pi_j = 1$). The interpretation of these unobserved heterogeneity parameters are straightforward, a high value simply implies a high preference for leisure.

The specification of welfare participation takes the form

(4) $\varphi = \mu + \sigma\theta_j \quad j=1,\ldots,M$

where μ and σ are parameters to be estimated. We define μ as a linear function of age, education and citizenship. The σ parameter is included in order to allow for correlation across alternatives. This idea is based on a factor loading technique (see, for instance, Ham and Lalonde, 1996).

The contribution to the likelihood function for a given household (i',k') becomes

(5) $$(p|\theta_h, \theta_w)_{i'k'} = \frac{\exp(U_{i'k'})}{\sum_{i,k} \exp(U_{ik})}$$

where i is index hours and k is welfare participation. This expression simply denotes the probability that the utility in the observed state is the highest amongst all of the possible hours and welfare combinations.

In our specification of measurement errors or classification errors, we follow MaCurdy et. al. (1990) and Hoynes (1996) and assume a multiplicative classification error structure. Let H denote reported hours and h denote optimal (discrete) hours, the multiplicative classification error specification is given as

(6) $H = he^{\varepsilon} \quad \text{with} \quad \varepsilon \sim N\left(-\tfrac{1}{2}\sigma_\varepsilon^2, \sigma_\varepsilon^2\right)$

Thus, zero hours are observed with certainty, but when optimal hours are positive they differ from reported hours by a factor of proportionality.

In the presence of unobserved heterogeneity and classification errors, the contribution to the likelihood is given by

(7) $$l = \sum_{m=1}^{M} \pi_m \left((p|\theta_{hm}, \theta_{wm})_{i'k'} g\right)\delta_{i'k'}$$

where $\delta_{i'k'}$ is an indicator for the observed state for each household, and g is the density for classification errors. The assumptions presented in (6) imply

(8) $$g = \begin{cases} 1 \text{ if } \mathrm{H} = 0 \text{ or } \mathrm{h}_i = 0 \\ \text{else} \\ \frac{1}{\sigma_\varepsilon}\varphi\left(\frac{[\log(\mathrm{H}) - \log(\mathrm{h})] + \frac{1}{2}\sigma_\varepsilon^2}{\sigma_\varepsilon}\right) \end{cases}$$

5 Data

The empirical analysis of single mother labor supply is based on the Longitudinal Individual Data for Sweden (LINDA), using a cross section drawn from 1996. LINDA is a register-based longitudinal representative data of the Swedish population (since 1960). The data consists of a large panel of individuals and their household members. There is also incremental register data for monthly wages and hours of duty (fraction of full working hours). In total, data for 1996 contains information on approximately 300,000 individuals, which provides us with a large enough sample of single mothers.

Our sample of single parent households consists of females who have at least one child 1-12 year old living with them. Early retirees, students, mothers of infants and mothers under age 18 or above 60 are excluded from the sample.

Sample statistics are shown in Table 1. Our selected sample includes 7,172 single mother households. The mean age of the mothers is 35 and they each have 1.83 children on average. In the sample, 88 % of the mothers are native Swedes, 5 % of them are natives of other Nordic countries, while 1 % and 6 % are natives of other western countries and refugee countries, respectively.

Hourly wage rates are constructed using information from a supplementary register. Nevertheless, this information is not available for the full sample. We have information on wages and hours of duty for all of those who work in the public sector, but only for about half of those working in the private sector. Thus, this might introduce a selectivity problem in our data and we have not tried to correct for this in the estimation. In effect, this might be a smaller problem in our sample of single mothers. A comparison of earnings for the publicly and privately employed exhibits only a small difference, and the same is true when comparing hours of work. Since missing observations on wages for employed females reduce the sample of working single mothers, we also adjusted the sample of non-working single mothers in order to keep the participation rate unchanged. This was accomplished by randomly deleting a sub-sample of non-working single mothers.

In order to impute hourly wage rates for the employed, monthly full-time earnings were divided by the "standard" full-time monthly working hours, i.e. 165 hours. Yearly hours of work, h, is then defined as total labor earnings divided by the hourly wage rate. Note that our definition of hourly wage rate is quite different from the common definition that is obtained by dividing observed earnings by observed hours, and thus has a tendency to include measurement errors. The definition that is used here is not subject to the same problem of measurement error. A characteristic feature of the wage rate used in this study is that the variation is small. Moreover, as a consequence of this there is no need to truncate extreme values. As shown in Table 1, mean hourly wage rate is 90 SEK and mean yearly hours of work is 1,283.

A remaining problem is, as usual, that wage rates are missing for non-working individuals. Here we predict missing values for hourly wage rates by estimating a wage equation. A standard Heckit-approach is used for the estimation and results are shown in Table 2. The wage increases at a decreasing rate with age (approximating work experience) and education level. As expected, in bigger cities the wage level is higher than in smaller towns. Eventually, the labor supply model is estimated using these selectivity-corrected predicted wage rates for non-workers and actual wage rates for workers.

The total income of a household consists of labor and non-labor income. Main components of non-labor income are income from capital gains and public transfers such as unemployment benefits and different social security allowances. Non-labor income is further divided into two parts, taxable and non-taxable. Taxable non-labor income consists of occupational car or expense allowance, job-related injury compensation, rehabilitation compensation, allowance for labor market training, unemployment benefits, cash labor-market support and other taxable transfers. The largest type of non-taxable non-labor income is child allowance, which every family with a child below the age of 16 is entitled to.

Tax deductions consist of several components, such as deductions for business expenses, general deductions for voluntary retirement insurance, general deductions for periodical supports and capital loss related deductions. The precision in this variable is a good illustration of the advantage of using register data. It is difficult to obtain a reliable measure of deductions from a survey. Of course, all errors in the income variables lead to errors in the calculated budget set. It is therefore crucial to have income data of high quality in studies of labor supply and taxes.

Housing allowance is determined by nationwide rules and is mainly directed to families with children. The amount of benefit received by a household is determined by economic and demographic factors such as net household income, housing expenditures, number of children and the ages of spouses. Unfortunately, there is no information about housing expenditures in LINDA. Therefore, the cost of housing is imputed using information from an alternative data source, the Swedish *Household Income Survey* (HINK) supplied by Statistics Sweden. In the imputation of costs of housing, we have used the method of minimum distance using age, number of children, earnings, place of residence and citizenship as classification variables.

As mentioned earlier, single mother households are heavily supported by the public sector. Both monetary transfers and publicly provided services are to a large extent targeted to the single parent households. About 16 % of the households in the sample receive social assistance and about 70 % receive housing allowance.

Unfortunately, we do not know exactly how many households in our sample are using municipal (subsidized) childcare. In spite of this, the cost of childcare can still be calculated, since the rules are known and this information will be utilized in the construction of the budget sets. Thus, the budget set has been constructed assuming that all mothers are using municipal childcare.

6 Results

Three versions of the model have been estimated, and the results are presented in Table 3. Model 1 is estimated without welfare participation and fixed costs of working, while Model 2 includes welfare participation but not fixed costs of working, and finally, Model 3 utilizes both factors.

As discussed above, the utility function is not restricted to being quasi-concave, instead the concavity can be tested after the estimation (for more details, see Van Soest, 1995).

Regarding Models 1 and 3, almost no individual violates the Slutsky-condition and in Model 2 about 700 violate it, i.e. about 10 %. In the following discussion we concentrate on Model 3, since it is the most general, but simulation results are presented for the alternative models as well. The results from the simulations are based on individuals' satisfying the Slutsky-constraint. Note however, that it might well be the case that individuals, who do not violate the constraints at their observed state before the applied change, may do so after the change.

According to the Model 3 results in Table 3, all variables reflecting observed heterogeneity have a significant effect on the preference for leisure. As expected, there is a strong and positive effect of younger children on the preference for leisure. The lower level of education also has a strong positive effect. Age has a negative effect and age squared a positive. Finally, living in a big city has a negative effect on the preference for leisure.

Fixed costs of working is an important variable for a household's decision making. The estimated intercept indicates that the costs increase the preference for leisure. Number of children and travel cost deductions decrease the preference for leisure. The negative sign for deductions can be interpreted as it picking up the effect of individuals who can deduct travel costs having lower costs of working. Surprisingly, the number of children has a significant negative effect on the fixed costs. One interpretation of this is that it is an age effect.

The estimated support points and the accompanying probabilities provide the unobserved heterogeneity in individuals' preferences. From the estimated support points, we can see that the first estimate of θ indicates a relatively strong preference for leisure. The corresponding π-parameter suggests that about 17 % of the individuals belong to this group (compared to the sample information, 19 % are not working). Consequently, the second group is identified as having a low preference for leisure.

The welfare equation includes a constant and three dummy variables to explore the significance of the stigma-effects on different types of households. These estimates indicate that there is, in fact, a positive and significant stigma-effect associated with welfare participation. Our results support the fact that not all households utilize the benefits for which they are eligible. The stigma-effect is weaker in young and in low educated groups of single mothers. Furthermore, the effect is higher among native Swedes than among others.

The estimated covariance shows a negative correlation between welfare participation and unobserved elements of work effort. Similar results are reported in Hoynes (1996) and Flood, Hansen and Wahlberg (1999), where the estimated correlation is higher for the work effort of women than of men. A negative covariance between welfare and labor supply implies self-selection into welfare, the higher the preference for leisure, and the smaller the influence of the stigma-effect.

A well-known problem of labor supply models is their poor ability to fit the observed distribution of hours of work. One way to improve the fit of these models is to include controls for fixed costs of working, as in Kapteyn et. al. (1990) and Van Soest (1995). In our approach, the estimated support points serve a similar purpose. Thus, our procedure produces a distribution of hours rather similar to the observed distribution shown in Figure 3. In the first upper and lower panel we compare the frequency of observed hours with the predicted hours. The results indicate an underestimation of hours below 1,250 and an overestimation at 1,750 hours of work per year. This results in an overestimation of the total mean value of about 180 hours per year. In the second upper and lower panel we have adjusted for this difference in levels by adjusting predicted values so that the mean values of observed and predicted hours are the same.

The method of predicting hours of work, conditional on unobserved heterogeneity, is straightforward. The estimated parameters and the utility function are used to calculate a probability for each discrete choice, 1 - 14 (seven classes for hours of work times two for the welfare participation choice). The maximum value of the probability then gives the predicted choice of hours of work and welfare participation. If the maximum probability falls among the first seven alternatives, no welfare participation is predicted. Consequently, welfare participation is predicted for the household if the probability falls among the last seven alternatives.

The procedure described above is repeated twice for both values of θ. Thus, for each individual, two predicted alternatives of hours (and welfare) are obtained. Then, one of them is chosen. In order to choose a group, the weighted values are calculated as

$$r_j = \frac{p_j \pi_j}{\sum_i p_i \pi_i} \quad j = 1, 2$$

where p_j is the maximum probability of all 14 alternatives evaluated at θ_j, and π_j is the estimated probability of a group j (j = 1, 2).

The predicted group is chosen by calculating the max (r_j). Thus, if r_1 is the maximum, then the individual is predicted as belonging to group 1 (a small preference for work). Finally, once an individual has been assigned to a group, she always belongs to this group.

As argued above, this method underestimates low hours of work. The reason for this is that group 1 is under predicted. There is a tendency in multinomial models to over predict the most typical outcome. It must be remembered that discrete unobserved heterogeneity serves the important task of increasing the variance in the predictions. Nevertheless, the variance is still not enough, which shows in the overestimation of hours of work.

As an attempt to correct for this overestimation we also use an alternative method, a sequential approach to predict group belonging. First we draw π_1 % of the highest values of r_1, and these individuals are assigned to group 1. Then all the remaining individuals are assigned to group 2.

There are four policy simulations implemented with Models 1, 2 and 3. These policy experiments are a 10 % wage increase, a cost reduction in childcare fees, a tax reduction and a 25% decrease in the WEL. In the following simulations, both methods (the mean adjusted and mean unadjusted) are considered.

The first set of results, presented in Table 4, shows the effects of a gross wage increase of 10 percent. Model 3 implies larger incentive effects compared to Models 1 and 2. Regardless of the specification of the model, the incentive effects seem to be small. For instance, the implied wage elasticity of Model 3 is around 0.1 - 0.2. The value varies slightly depending on the method of prediction. However, the main result does not change the wage increase has only a small effect on supplied hours of work. In fact, this is illustrated by the second row in Table 4. At most (Model 3, unadjusted mean), about 4 % of the individuals in the sample change their hours of work class as a result of the wage increase.

Since there is such a small change in labor supply due to a 10 % wage increase, the following results are mainly comparisons of direct effects before and after a change. For instance, the mean disposable income increases by about 4 %, and received housing allowance drops by 8 – 14 % depending on the specification. The wage increase has no effect on social assistance receipts. It seems that the change in wages by 10 % is too small to predict a change in welfare participation.[6] Cost of childcare increases by about 9 – 11 % and tax duties by about 12 – 14 %. Finally, the welfare effects are calculated using Equivalent Variation (EV) as a measure. EV is the amount of money added or subtracted from a household's disposable income before a reform is applied, in order to make the household indifferent between the situation before and after the reform. As such, EV summarizes the change in net welfare of a household associated with the behavioral responses due to the reform. As pointed out earlier, in our simulation the majority of the single mother households do not change their labor market behavior, and in these cases EV measures just the change in disposable income before and after the tax change (caused by the wage increase). In Table 4, these EV measures are divided by disposable income before the change. Thus, the level of compensation varies from 2 % to 5 % compared to the pre-change disposable income.

Table 5 reports the simulated effects of a recently suggested reform regarding the cost of childcare. This reform reduces the cost of childcare by introducing an upper ceiling. For one child the maximum monthly cost is 1,150 SEK, or 3 % of household income, for the second

[6] Additional simulations, not reported, indicate that it is necessary to increase wages by about 30% to find a

child the corresponding cost is 767 SEK, or 2 % of household income and for the third the cost is 383 SEK, or 1 %. The results show only a small effect of a reduction of childcare costs on hours of work. Since the behavioral responses are small, the reported numbers are again very similar to comparisons of disposable incomes before and after the reform. Not surprisingly, the main effect is the reduction in childcare cost by up to 31 % – 37 %. The reduced fees for childcare result in an increased disposable income of around 1 % – 3 %. There are small changes in housing allowances and taxes reflecting the small change in hours of work.

The third policy experiment is a simple tax reform, which reduces the progression in the income tax system. The maximum marginal tax rate is decreased down to the level of the municipal tax rate (on average, slightly above 30 %). The breaking point at a yearly taxable income of around 200,000 SEK (see Figure 2) is now deleted. Concentrating on Model 3, the main effects of the tax reduction is an increase in hours of work of about 3 %, and an increase in disposable income of about the same size. There is an expected decrease in tax revenues, but due to the increase in hours of work, this reduction is small. Since the housing allowance receipts decrease and costs of childcare increase, the total effect on the government revenues is quite small. The EV measure implies a small welfare increase (about 2 %) as a result of the tax reform.

The fourth experiment is a change in the WEL. The WEL is decreased by 25 %, which implies that the number of households eligible for social assistance must decrease. As expected, the total effect on labor supply is quite small. There is only a 0.3 % increase in labor supply. An interesting effect is the change in social assistance. Model 3 without mean correction predicts a decrease in social assistance of 40 %, while the corresponding result with mean correction is a decrease by only 2 %. The main explanation for this is that the uncorrected predictions underestimate the share of non-workers. Since almost all of the recipients of social assistance are not employed, this implies that the uncorrected predictions also underestimate the share of households on social assistance. In order to obtain a correct prediction of this share, we have to correct our predicted values.

In the discussion above, we have concentrated only on the average effects, while the

noticeable change in social assistance.

distributional effects are not considered at all. Even though the reported mean values are small, there still might be a large effect for various income groups. As an illustration, we continue with Model 3, mean corrected, and examine the distributional effects of the tax and childcare reforms more closely. In the tax simulation the estimated average EV for all of the households is about 1,700 SEK per year. However, this reform has a completely different impact on different households belonging to different income groups. Ordering the households into income deciles, according to the predicted disposable income before the reform, confirms that the EV measure is zero for the poorest deciles compared to about 6,800 SEK per year for the richest deciles. Consequently, there is almost no gain from the tax reform for the households below the median income.

A closer look at the distributional effects of the childcare reform shows a distinct result the welfare effects are quite similar over all of the income deciles. The average EV for the whole sample is slightly above 1,900 SEK per year, and the corresponding measures for the lowest and highest quartiles are 1,300 SEK and 1,400 SEK, respectively. The predicted increases in hours of work and disposable income for the lowest quartile are 18 % and 2 %, respectively, while the corresponding results for the highest quartile are 0 % and 1 %.

To conclude, despite the rather small average incentive effects, there can be some substantial effects for specific income groups. The micro simulation approach enables us to consider both the average and distributional effects of a change in the tax and benefit systems.

7 Conclusions

The question addressed in this paper is what the size of the incentive effects for individuals facing high marginal effects is. To answer this question we use a sample of Swedish single mother households with detailed information on incomes and benefits and estimate a structural labor supply model. We formulate a model where labor supply and participation in welfare are determined jointly. Furthermore, the labor supply and welfare participation decisions are treated as discrete choice problems, and we assume that these choices follow a simple conditional logit rule. In addition, we allow for unobserved individual-specific effects and also for these effects to be correlated across alternatives. The unobserved effects are assumed to be drawn from a discrete distribution, and the correlation across alternatives is modeled using factor-loading techniques. Classification error in hours is allowed for by using a multiplicative classification error specification. Special attention is devoted to the effects of fixed working costs related to the labor market participation.

We find that there is a positive and significant stigma-effect associated with welfare participation. Fixed costs of working is an important variable for single mothers deciding whether or not to enter the labor market. The estimated negative covariance between social assistance and labor supply implies self-selection into welfare. Our result from the micro simulation shows rather small average incentive effects, but there are some substantial effects in different income groups. The tax simulation shows that there is a difference of about 6,800 SEK per year between the poorest and richest single mother household. Welfare effects from the childcare reform are quite similar across income deciles, even if predicted increases in hours of work are considerable for the poorest single mother households.

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Appendix

Table 1. Description of the Sample of the Single Mother Households (N=7 172).

Variables	Mean	Minimum	Maximum
Age	35	18	60
Education (the highest)			
- Basic school	0.26	0	1
- High school	0.65	0	1
- University	0.10	0	1
Working hours per year	1,283	0	5,351
Participating the labor market	0.81	0	1
Wage/hour SEK	89	56	411
Taxable non-labor income SEK per year	18,929	0	290,706
Non-taxable non-labor + net capital income	19,468	0	859,543
Deductions SEK per year	3,058	0	143,528
Number of children 1-12 years old	1.83	1	4
Place of residence			
- Big cities	0.37	0	1
- Medium sized cities	0.45	0	1
- Rural cities	0.19	0	1
Nationality			
- Swedish	0.88	0	1
- Nordic countries	0.05	0	1
- Western countries	0.01	0	1
- Refugee countries	0.06	0	1
Welfare participant	0.16	0	1
Social assistance per year, SEK	4,276	0	159,108
Housing costs per month, SEK	4,734	1,374	8,337
Receiving housing allowance	0.71	0	1
Housing allowance, SEK per year	14,830	0	46,800

Table 2. The Estimated Parameters of the Probit Model of the Labor Force Participation and the Wage Equation for Single Mothers.

Variables	Estimate	Standard error
PARTICIPATION EQUATION:		
Intercept	-1.4701	0.8066
Age	0.1912	0.0321
Age Squared / 100	-0.2416	0.0356
Education, Basic school	-1.7034	0.6203
Education, High school	-0.4733	0.6039
Number of children	-0.0959	0.0244
Medium sized city	-0.0979	0.0428
Rural area	-0.1348	0.0532
Having 1 year old children	-0.7339	0.0834
Having 2-5 year old children	-0.1682	0.0443
Age * Basic school	0.0165	0.0157
Age * High school	-0.0042	0.0151
Immigrant, Nordic	-0.1596	0.0862
Immigrant, Western	-0.7193	0.1957
Immigrant, Refugee	-1.1546	0.0633
WAGE EQUATION (log wage):		
Intercept	3.7853	0.0943
Age	0.0368	0.0045
Age Squared / 100	-0.0319	0.0054
Education, Basic school	0.1512	0.0541
Education, High school	0.1294	0.0446
Number of children	-0.0155	0.0030
Medium sized city	-0.0319	0.0046
Rural area	-0.0390	0.0059
Age * Basic school	-0.0095	0.0013
Age * High school	-0.0080	0.0011
Immigrant, Nordic	-0.0207	0.0096
Immigrant, Western	0.0162	0.0274
Immigrant, Refugee	-0.0761	0.0192
Lambda	0.0135	0.0303
$\overline{R}^2$	0.25	

Table 3. The Estimates of the Parameters of the Utility Function.

Variables	Coefficient	Model 1 Estimates (Std.)	Model 2 Estimates (Std.)	Model 3 Estimates (Std.)
Observed heterogeneity, β_h:				
Children 1 years old, 1=yes, 0=no	β_{h1}	6.4917	7.0501	4.2271
		(0.4269)	(0.4585)	(0.2975)
Children 2-5 years old, 1=yes, 0=no	β_{h2}	1.5448	1.6448	1.1923
		(0.1918)	(0.1972)	(0.1541)
Children 6-9 years old, 1=yes, 0=no	β_{h3}	0.7015	0.7935	0.4918
		(0.1749)	(0.1789)	(0.1437)
Education basic-school, 1=yes, 0=no	β_{h4}	1.3496	1.4773	1.7969
		(0.2354)	(0.2385)	(0.2072)
Education high-school, 1=yes, 0=no	β_{h5}	0.5034	0.6198	0.6983
		(0.2017)	(0.2048)	(0.1843)
Region big cities, 1=yes, 0=no	β_{h6}	-0.3800	-0.5134	-0.1713
		(0.1280)	(0.1312)	(0.1007)
Age	β_{h7}	-0.6564	-0.5027	-0.3135
		(0.0918)	(0.0934)	(0.0689)
Age Squared / 100	β_{h8}	0.8264	0.6292	0.3967
		(0.1232)	(0.1254)	(0.0935)
Utility arguments:				
Consumption	β_C	11.5114	9.9925	11.5926
		(1.0140)	(0.7342)	(0.7764)
Consumption squared	β_{CC}	4.8611	2.9738	1.0460
		(0.6312)	(0.2049)	(0.2045)
Hours squared	β_{hh}	-3.5082	-5.7845	-7.7602
		(0.3363)	(0.3321)	(0.2477)
Hours times consumption	β_{Ch}	-0.3455	-1.1674	-2.1280
		(0.3740)	(0.2551)	(0.2034)
Fixed cost of working:				
Intercept	γ_1	--------	--------	2.3094
				(0.1196)
Number of children	γ_2	--------	--------	-0.0468
				(0.0171)
Travel deductions, 1=yes, 0=no	γ_3	--------	--------	-0.5041
				(0.0675)
Classification error	ε	0.1510	0.1510	0.1510
		(0.0014)	(0.0014)	(0.0014)
Unobserved Heterogeneity:	θ_1	58.4698	53.2106	20.2581
		(2.9190)	(2.3366)	(1.3461)
	θ_2	23.9896	22.6299	17.6851
		(1.8373)	(1.7998)	(1.3102)
Heterogeneity Probabilities:	π_1	0.1522	0.1720	0.1746
		(0.0055)	(0.0055)	(0.0215)
	π_2	0.8478	0.8280	0.8254
Welfare participation:				
Intercept	μ_1	--------	3.6493	36.2623
			(0.2024)	(5.1911)
Age 18-29, 1=yes, 0=no	μ_2	--------	-0.7474	-1.4840
			(0.0939)	(0.2284)
Education basic-school, 1=yes, 0=no	μ_3	--------	-0.6010	-1.7633
			(0.0889)	(0.2970)
Swedish, 1=yes, 0=no	μ_4	--------	0.7167	1.7992
			(0.1076)	(0.2146)
Covariance hours, welfare	σ_{hw}	--------	-0.0518	-1.8347
			(0.0047)	(0.2596)
Log Likelihood function		-1.1474	-1.4641	-1.3823

Table 4. An Increase in Gross Wages by 10 Percent.

	Model 1		Model 2		Model 3	
	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %
Hours	0.00	0.00	0.05	0.04	1.22	2.10
Changed Hours	0.00	0.00	0.10	0.10	2.70	4.10
Disposable income	3.68	4.03	3.50	3.90	4.29	4.89
Housing allowance	-11.02	-13.99	-7.76	-11.67	-12.28	-14.30
Social assistance	-------	-------	0.00	0.00	0.00	0.00
Cost of childcare	8.25	7.93	7.99	8.01	9.88	10.91
Taxes	11.88	12.04	13.16	13.21	13.56	14.37
Equivalent variation/ disposable income before change	3.27	3.83	2.30	3.02	3.92	4.95

Table 5. A New Child Day-care Payment Scheme.

	Model 1		Model 2		Model 3	
	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %
Hours	0.00	0.00	0.05	0.04	1.35	2.29
Changed Hours	0.00	0.00	0.10	0.10	1.80	3.10
Disposable income	1.63	2.01	1.67	1.93	2.10	2.75
Housing allowance	0.00	0.00	-0.05	-0.06	-1.12	-1.75
Social assistance	------	-------	0.00	0.00	0.00	0.00
Cost of childcare	-32.84	-36.92	-32.26	-34.33	-31.39	-32.13
Taxes	0.00	0.00	0.06	0.05	1.32	2.09
Equivalent variation/ disposable income before change	1.47	2.00	0.80	1.34	1.93	3.16

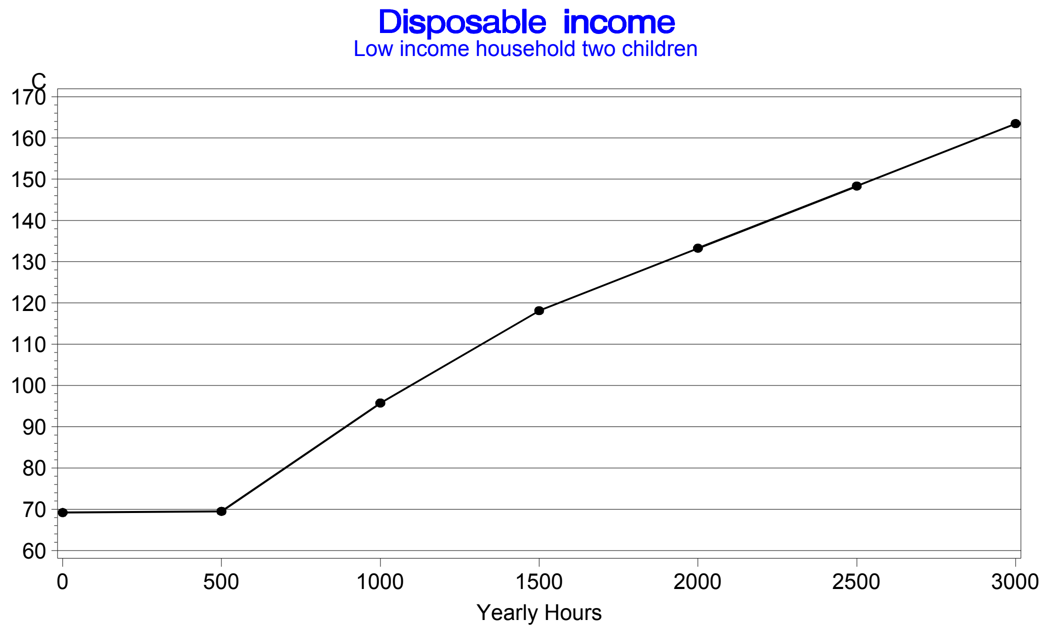
Table 6. No Governmental Income Tax.

	Model 1		Model 2		Model 3	
	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %
Hours	0.00	0.00	0.04	0.03	2.50	2.59
Changed Hours	0.00	0.00	0.00	0.00	4.40	4.70
Disposable income	1.08	1.17	4.19	4.08	2.55	2.59
Housing allowance	0.00	0.00	-0.02	-0.02	-1.89	-2.15
Social assistance	------	-------	0.00	0.00	0.00	0.00
Cost of childcare	0.00	0.01	0.06	0.04	2.35	2.39
Taxes	-3.44	-3.49	-11.68	-10.55	-1.66	-1.36
Equivalent variation/ disposable income before change	0.62	0.72	1.89	2.13	1.77	1.96

Table 7. Decrease in Social Assistance Norm by 25%.

	Model 1		Model 2		Model 3	
	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %	Mean corrected %	Not corrected %
Hours	0.15	0.16	0.02	0.02	0.33	0.35
Changed Hours	0.20	0.20	0.00	0.00	0.50	0.40
Disposable income	-1.31	-0.36	0.02	0.02	0.22	0.20
Housing allowance	0.00	-0.01	0.00	0.00	-0.01	-0.06
Social assistance	------	------	-37.58	-37.77	-2.16	-40.00
Cost of childcare	0.22	0.22	0.03	0.02	0.41	0.49
Taxes	0.10	0.11	0.02	0.02	0.26	0.26
Equivalent variation/ disposable income before change	-2.87	-0.95	-0.01	-0.01	0.41	0.36

Figure 1.


Disposable income
Low income household two children
C
170
160
150
140
130
120
110
100
90
80
70
60
0
500
1000
1500
2000
2500
3000
Yearly Hours

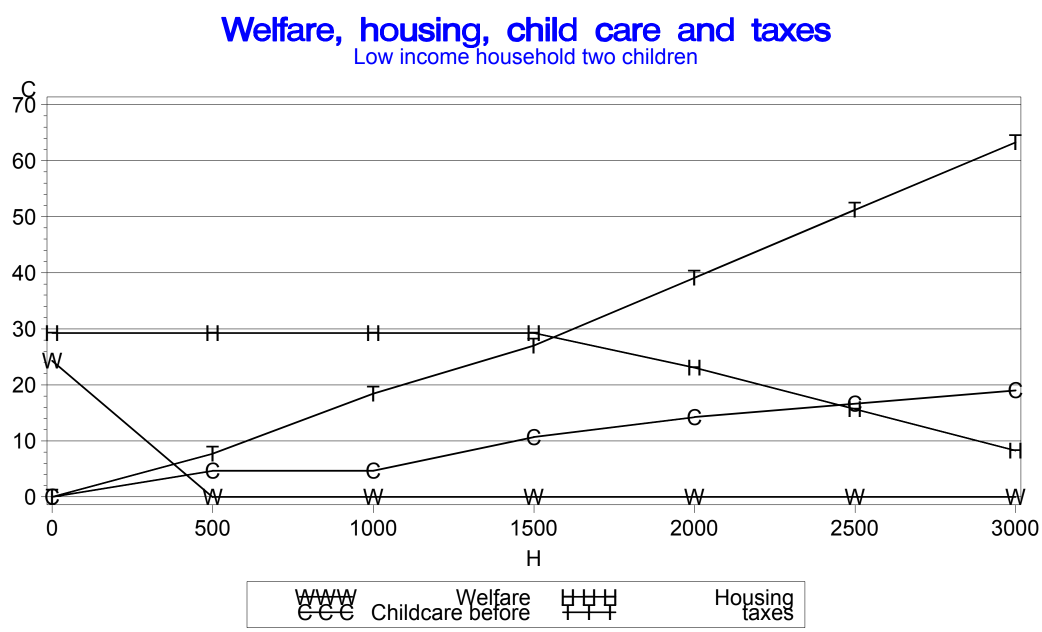

Welfare, housing, child care and taxes
Low income household two children
C
70
60
50
40
30
20
10
0
0
500
1000
1500
2000
2500
3000
H
Welfare
Housing
Childcare before
taxes

Figure 2. Income taxes in Sweden 1996

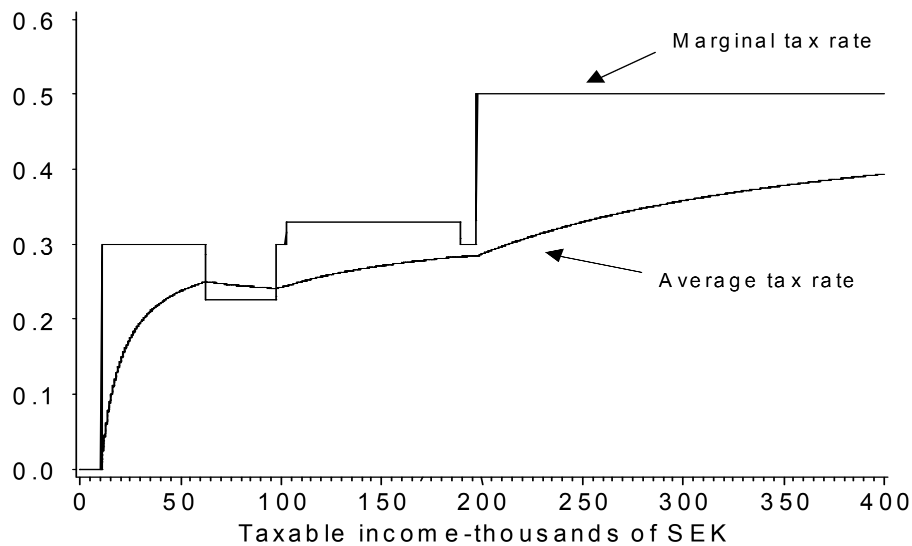
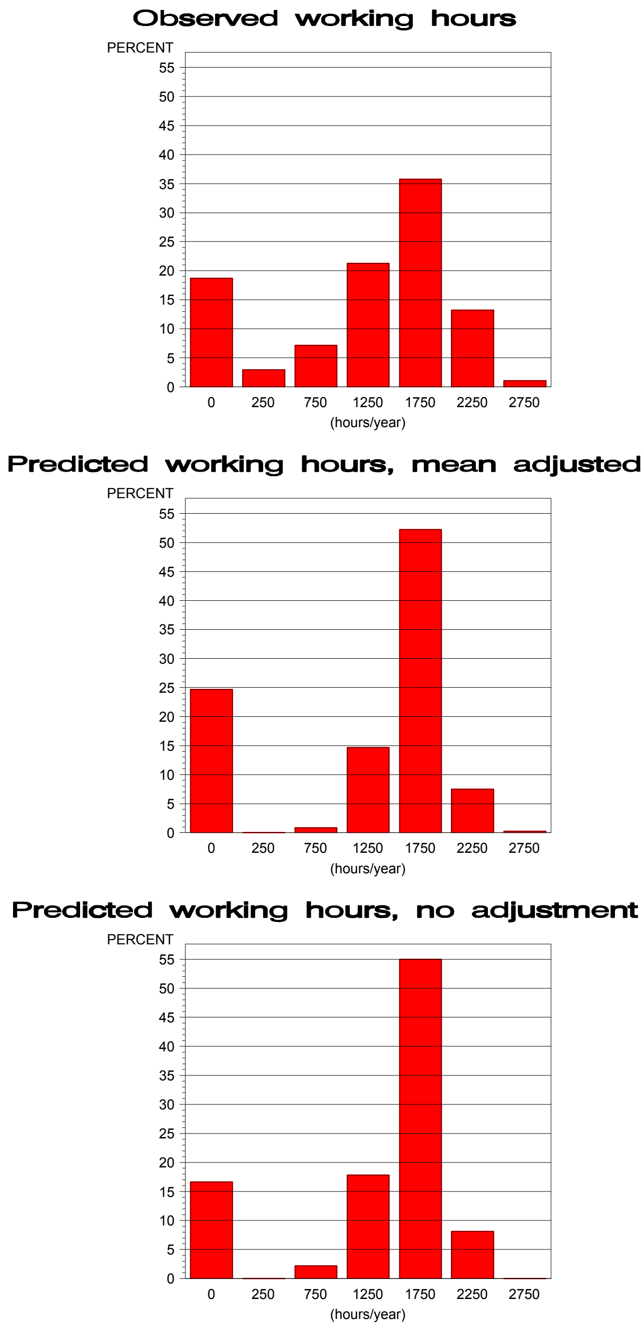


Figure 3. Distribution of Hours of Work



Early Retirement in Sweden

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Abstract:

In this paper we analyze the determinants of early retirement from the Swedish labor market for both males and females. We estimate a panel random parameter logit model with three labor market states: working, part time pension and full early age retirement. The result shows that there is heterogeneity in the underlying preference structure and that the probability of a complete early withdrawal from the labor market increases with age. Blue collar workers have lower probability to take part time pension and full early age retirement than workers from other occupational schemes. Finally, we find that economic incitements affect the decision of Swedish workers to leave the labor market.

Key words: Early Retirement, Occupational Pensions, Random-Parameter Logit Model

JEL Classification: C23, C25, J26,

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1 Introduction

The Swedish retirement, income tax and benefit systems have gone through several changes during the last decades. One of the purposes of these policy reforms has been to sustain and increase the labor market participation among older Swedish workers. However, despite these tax reforms and changes in the Swedish retirement system, Sweden has experienced a strong tendency towards early exit from the labor market.[2] At the same time the proportion of elderly in the population is increasing and it will most likely continue to do so in the future.

In Figure 1, we can see how the observed labor market participation rate change between age 59 and 66 for males and females. At age 60, the participation rate is 67% for males and 60% for females. The older individuals become, the lower the participation rate for both males and females. At age 64, one year before the mandatory retirement age in Sweden, the participation rate is 37% and 27%, respectively. There are few Swedish workers who work past the mandatory retirement age. The participation rate at age 66 is 8.8% and 3.9%, respectively.

In Figure 2 and Figure 3, we have broken up the Swedish labor market into four different labor market states: full early age retirement (early), part time pension (part), in the labor force (work) and disability pension (disability). Women are more likely to be in full early age retirement than men. At age 64, 31% of women have taken full early age retirement, compared to 23% for men. Part time pension is used more by men than women, at age 64, 12% of men have part time pension while this figure is only 7% for women. At age 60, 30% of women and 26% of men have disability pension. These numbers increase by age and at age 64, the corresponding numbers are around 40% for both men and women. Disability pension is the most common exit route from the Swedish labor market.

In Figure 4 to Figure 7, we present different labor market states at age 61 and 64 for males and females during the 1993-1999 period. The most striking feature of the Swedish labor market during the last decade is the decrease in part time pension for both males and females. For the younger cohort, disability pension decreases over time while disability pension is constant for the older cohort.

[2] Wadensjö (1996) describes why Sweden, in comparison with many other countries, has a relatively late exit from the labor market and why the trend is towards earlier retirement.

From these figures it seems that the changes in the Swedish retirement, income tax and benefit systems has not been sufficient to change the behavior of Swedish workers. Therefore, to increase our general understanding of the labor market and to be able to evaluate effects of policy reform it is essential that we study the determinants of early retirement from the Swedish labor market for both males and females in Sweden.

To analyze early retirement of older Swedish males and females we estimate a panel random parameter logit model with three states: working, part time pension and full early age retirement.[3] The random parameter logit model generalizes the multinomial logit by allowing the parameters of the explanatory variables to fluctuate randomly across individuals. Estimation explicitly considers the fact that the variation in coefficients across individuals induces a correlation in unobserved utility over time by the same individual.

For every individual and year we have calculated a hypothetical income for the three labor market states: working, part time pension and full early age retirement. When we calculate this hypothetical income we have taken into account the rules of the Swedish pension system and the rules of the occupational pension schemes.

We have included benefit accrual and social security wealth as measures of economic incentives. These two variables allow us to measure the income and substitution effects of the retirement decision.

The empirical analysis of early retirement in Sweden is based on the Swedish Longitudinal Individual Data set (LINDA). LINDA is a register-based longitudinal representative data of the Swedish population since 1960, and consists of a large panel of individuals and their household members.

Our sample used in this study consists of information for the years 1992-1998 together with a supplementary register from the national social insurance board of pensions points for those listed in the 1998 LINDA. The sample used in this study consists of 14,301 observations of 4,613 males, and 13,420 observations of 4,370 females all between the ages of 60 and 64. We find that there is heterogeneity in the underlying preference structure. The probability of

[3] We will not consider disability pension in this study. See Andrén (2001) and Skogman Thoursie (1999) for

full early age retirement and part time pension increases with age. Blue collar workers have lower probability to take part time pension and full early age retirement than workers from other occupational schemes. We simulated the income distribution and found that the probability of part time pension decreases over the income distribution. Our simulation results show that an increase in the hypothetical income leads to positive direct effects and negative cross effects. An increase (decrease) in benefit accrual leads to a decreasing (increasing) probability of a total withdrawal from the labor market. Finally, we found that a rise (reduction) in social security wealth has a positive (negative) effect on the retirement decision. Thus, economic incitements affect the decision of Swedish workers to leave the labor market either gradually or totally.

The paper is organized in the following way. Section 2 describes the rules for the Swedish pension system and the rules for the different occupational pension schemes. Previous studies on Swedish data are presented in Section 3. In Section 4 we describe the data and sample used in this paper. The econometric model is presented in Section 5, while the results are presented in Section 6. Section 7 concludes the paper.

2 The Swedish Pension System

The Swedish pension system is financed by employer contributions and consists of three parts: basic pension, supplementary pension and part time early pension.

The basic pension is given to all Swedish citizens and all individuals residing in Sweden. All individuals receive the same amount, although the amount decreases if an individual has not been living in Sweden for at least 40 years, or has had labor income for less than 30 years. The basic pension is connected to a Basic Amount (BA) that follows the Consumer Price Index (CPI).[4] Even though the BA is connected to the CPI, the amount is still decided by the Swedish Government.

Individuals with a work history receive not only basic pension, but also a Supplementary Pension (ATP). The condition that must be met in order to receive ATP is that an individual must have had pension-right income for at least three years between the ages of 16 and 65.

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analysis of disability and work in Sweden.

[4] BA93=34,400 SEK, BA94=35,200 SEK, BA95=35,700 SEK, BA96=36,200 SEK, BA97=36,300 SEK.

Pension-right income is an income that is above 1 BA per year but less than 7.5 BA per year. Included in the retirement-based income are incomes from work, sickness, unemployment benefits, parental cash benefits and part time pension.

The amount a worker receives from ATP is based on pension points, the 15 year rule and the 30 year rule. Every individual has a record of ATP points for each year from age 16 to age 65. To calculate ATP points for a given year, divide the pension-right income for that year by the corresponding BA. The maximum number of ATP points an individual can receive in one year is 6.5. The 15 year rule states that the best 15 years of ATP points should be used to determine the ATP. Accordingly to the 30 year rule, a person must have ATP points for at least 30 years to receive full pension from ATP. If an individual has less than 30 years of ATP history the pension income is reduced proportionally to the number of missing years. An individual's ATP pension income (YATP) is then calculated as:

$$YATP_i = 0.6 * AP_i * \min\left(\frac{N_i}{30}, 1\right) * BA$$

where AP_i denotes average pension points for the best 15 years and N_i is number of years earning ATP points.

Swedish workers can claim basic pension and ATP in advance from age 60, or postponed them up to age 70. For each month an individual chooses to withdraw early from the labor market, the monthly benefit is reduced by 0.5%. Every month an individual postpones retirement past age 65, result in a permanent increase of the yearly pension income by 8.4%.

Those who have no or low supplementary pension are entitled to a special pension increase. The special pension increase is 55.5% of the BA and is reduced on a one to one basis against the supplementary pension. A single retired individual with no supplementary pension receives 151.5% of BA.

At age 61, workers are allowed to take partial retirement pension.[5] To get partial retirement pension a worker has to decrease working hours by at least 5 hours/week, and must work at least 17 hours but not more than 35hours/week after the reduction. Before 1994, the benefit was 65% of the difference in earnings before and after part time pension. The benefit was

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In August 2002 1 Euro ≈ 9 SEK.

lowered to 55% in 1994.

A complement to the ordinary pension system is occupational pension schemes. Occupational pension schemes are a result of agreements between unions and employer confederations, and are financed through employer contributions. Essentially, there are four different occupational plans: blue collar workers in the private sector, white collar workers in the private sector, central government employees and local government employees.[6]

The blue collar pension scheme (STP) cannot be postponed or claimed before the 65th birthday, and the amount received is 10% of the average yearly income of the best three years between ages 55 and 59. If workers have contributed less than 30 years after age 28, the size of the occupational pension decreases proportionally, and to get any occupational pension at all blue collar workers have to have contributed at least three years between ages 55 and 59. STP follows BA, and individuals earning more than 7.5 BA per year are not given any STP.

A white collar worker in the private sector receives ITP and ITPK. The amount of ITP received depends on earnings the year before a white collar worker retires, 10% of earnings up to 7.5 BA, 65% of earnings between 7.5 and 20 BA and 32.5% between 20 and 30 BA. As in STP, workers have to be 28 years of age before they can start earning qualification years in ITP. If the qualification years total are less than 30, ITP is decreased proportionally to the number of missing years. The white collar pension scheme could be postponed or claimed in advance of the 65th birthday. If ITP is claimed (postponed) ITP is reduced (increased).[7] Normally the yearly IPTK pension is around 3.5% to 4% of the earnings the year before retiring.[8]

Central government employee pension is earned from age 28 and workers have to work for 30 years to get full occupational pension. Otherwise, the occupational pension is reduced proportionally. The average yearly earning of the last five years before retiring decides the size of the occupational pension, and amounts up to 30BA are considered. Full occupational pension is 10% of this amount up to 7.5 BA. Thereafter the occupational pension is calculated

[5] Before 1994 it was 60 years of age.
[6] See Kangas and Palme (1989) for more information about occupational schemes in Sweden.
[7] From age 60 up to age 70 the reductions / improvements are: 0.739, 0.783,0.831,0.884,0.942,1,1.076, 1.154,1.241,1.338 and 1.448.

in the same way as ITP, with 65% of the amount between 7.5 and 20 BA and 32.5% of the amount between 20 and 30 BA. A central government employee could claim early retirement starting at age 60, but with a permanent 0.4% reduction per retired month before the 65th birthday. Central government employees have a supplement pension similar to ITPK, around 2-3% of average yearly earnings of the five years preceding the year a worker withdraws from the labor market.

Local government employees have to work 30 years in the local governmental sector from age 18 to 65, otherwise the occupational pension decreases gradually. Average yearly earnings of the five best years of the seven years preceding the year of retirement decides the size of the occupational pension. The local government worker receives 96% of this amount below 1BA, 78.5% between 1 and 2.5 BA, 60% between 2.5 and 3.5 BA, 65% between 7.5 and 20BA and finally, 32.5% between 20 and 30 BA. If a local government employee decides to retire early, the occupational pension is permanently decreased by 0.5% per retired month between age 60 and 62, 0.4% per retired month in the age interval 62 to 63 and 0.3% per retired month between ages 63 and 65. For every month the retirement is postponed after the 65th birthday, the occupational pension increases by 0.1%. The occupational pension a local government employee receives depends on basic pension and ATP, the amount decreases the more basic pension and ATP the individual has.

In January 1999 a new pension system was introduced in Sweden. Essentially, the main differences are that lifetime earnings are now taken into account, the pension income follows the general wage development in the Swedish economy and that changes in life expectancy affect the yearly pension income. Thus, a lower economic growth and higher life expectancy decreases individuals pension income at a given retirement age.

[8] Ståhlberg (1995), page 32.

3 Previous Studies on Swedish Data

There is vast empirical literature on the decision to retire. Most of the research has been done on the United States; see Gustman and Steinmeier (1986), Stock and Wise (1990a), Berkovec and Stern (1991), Blau (1994), Rust and Phelan (1997), Blau (1998), French (2000) and Blau and Gilleskie (2000). For other studies see Meghir and Whitehouse (1997), Suen (1997), Herneas et al (2000), Dahl et al (2000) and Heyma (2001). However, there have been few econometric studies of early retirement in Sweden.

Hansson-Brusewitz (1992) analyzes how socioeconomic variables like the health of an individual and economics factors like individual wage rate, the tax system and the pension system, affect the labor supply of elderly. He estimates an econometric model where hours of work and labor force participation are determined simultaneously. To evaluate the effects of various tax and pension reforms on labor supply, Hansson-Brusewitz simulates an abolishment of the partial pension system and a replacement of the ATP system with a pension benefit, which is equal to 60 percent of the average lifetime earnings. The data used is the Level of Living Survey. His sample consists of married men between 55 and 70 years of age in 1973 or 1980. A total of 595 observations are used in the estimation. Hansson-Brusewitz finds that abolishment of the partial pension system had a small but positive effect on the lifetime hours of work, while replacing the ATP system with the pension benefit had no effect in the long run. The main conclusion is that the diminishing labor supply of elderly is to a large degree due to aging as such, and poor health may step up the process.

Sundén (1995) examines how the introduction of the partial retirement program has affected retirement behavior among older workers in Sweden. She estimates a multinomial logit model with four choices: 1) early full retirement at age=60, 2) early retirement with disability insurance, 3) early partly retirement at age 60, and 4) no early retirement before age 65. Sundén used the Level of Living Survey for the years 1974 and 1981. The data from 1974 gives the retirement pattern before partial retirement was introduced and the data from 1981 gives a picture of the retirement behavior after the introduction of partial retirement. Her sample includes about 5000 individual aged 16 and older. The Sundén result shows that the number of individuals collecting disability pensions decreased and the majority who chose early partial retirement were men who had previously worked full time. She finds that most of the changes in retirement behavior is due to changes in preferences.

Palme and Svensson (1997) give a summary of the Swedish social security system and its influence on individual retirement behavior. They use data from the Swedish Labor Force Survey and the 1994 Household Income Survey. Palme and Svensson find that the male labor force participation decreased among all age cohorts. In contrast to men, women increased their labor force participation up until the 1991 recession. After the 1991 recession, female labor force participation decreased slightly. The share of women older than 55 who actually receive old age or disability pension is larger than the share of men in the same age category. The most common reason for leaving the labor market is disability. Palme and Svensson simulate the social security outcome for a representative blue collar worker in the private sector and find that economic incitement generated by the age pension scheme affects the retirement behavior.

Palme and Svensson (2001) analyze the consequences of economic incentives built into the social security system, and how the mandatory old age pension system influences the retirement behavior. They use the Swedish longitudinal individual data set for the years 1983 to 1997. In their econometric analysis they use a probit model with 15,619 men with a total of 127,390 observations, and 14,820 females with a total of 123 979 observations. One control variable in the probit model is net social security wealth. It is included to measure the income effect. Palme and Svensson's result shows that economic incentives matter for the retirement behavior in the Swedish labor market, and their simulation shows that there may be a significant effect on labor force participation from changing the economic incentives of retirement.

The national insurance board (RFV) (2001a) tries to figure out what makes us work until age 65. They use a logistic regression, and the data used in this study comes from a survey done in 2000, "Enkät till individer om arbetsförhållande, hälsa och pension",[9] together with information from the Swedish Longitudinal Individual Data set (LINDA). This inquiry is a random sample of 5,100 individuals from the 1998 LINDA aged 35-70 in the year 2000. The result shows that workers in a psychosocial work environment characterized by high demands have a higher probability of leaving the labor market in advance. Individuals belonging to an

[9] "Inquiry to individuals about working environment, health and pension."

occupational scheme have a positive probability of leaving the labor market early compared to those not belonging to an occupational scheme. Age has a negative effect on leaving the labor market early, and the explanation for this contradiction is a selection problem in the sample. Married men have a higher probability of leaving the labor market early.

RFV (2001b) analyzes how the flexibility in the new Swedish pension system affects the pension income. Their analysis is based on constructed type cases that differ with respect to age, time of retirement, the level of take-out, degree of gainful employment and level of income. A longer working life gives higher retirement income irrespectively of the income level. Individuals who combine partial pension with a partial gainful employment will not affect their retirement income much. Low income or high income individuals gain relatively less by working more. Their main conclusion is that generally there is not an optimum choice with respect to combination of work, the level of pension take-out and retirement age.

4 Data

The empirical analysis of this study is based on the Swedish Longitudinal Individual Data set (LINDA). LINDA is a register-based longitudinal representative data of the Swedish population since 1960 and consists of a large panel of individuals and their household members. It contains information on approximately 300,000 individuals each year.[10]

Our sample used in this study consists of information from the years 1992-1998, together with a supplementary register from the national social insurance board of pensions points for those included in the 1998 LINDA.

We include males and females between 60 and 64 years of age for every year from 1993 to 1997. The self-employed are excluded. Individuals who are out of the labor force either voluntarily or due to disability are not included in the sample. We have three different labor market states: working, part time pension and full early age retirement. An individual has full early age retirement if labor income is less than 1 BA. If part time pension income is larger than zero the individual has part time pension. People who are not retired, do not have partial pension and who are still in the labor market are then defined as working.

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[10] See Edin and Fredriksson (2000) for a more detailed description of the LINDA database.

There is no information about occupational schemes in LINDA. However, by using information from the 1998 LINDA, we take for those who worked in 1998 and use occupation and sector, to decide occupational schemes belonging. To create this variable for individuals who had already retired in 1998, we use pension income from different occupational schemes. Individuals who had their highest occupational pension from the blue collar scheme are defined as Blue Collar workers (BC). Those individuals who are not BC are defined as White Collar workers (WC) if white collar scheme pensions were larger than potential municipality and government occupational pension incomes. Further, those individuals who are not BC or WC and have a municipality pension higher than potential pension from the government occupational scheme are defined as members of the Municipality occupational scheme (M). An individual not belonging to BC, WC or M having a government occupational pension larger than zero is then defined as belonging to the Government occupational scheme (G). We exclude those for whom we do not have any 1998 occupational pension information.

Included in the econometric model are three dummy variables for educational attainment. The reference group consists of individuals with a basic education as the highest attained education (nine years or less of schooling) in comparison to individuals with a high school degree (12 years of schooling) or a university degree (more than 12 years of schooling). Educational attainment is included since there may be different retirement behavior among different educational groups. Individuals with basic education may not afford to take part time pension or full age pension before the mandatory retirement age of 65.

Dummy variables for marriage, wealth and being born in Sweden are included in the econometric model as well. The dummy variable for marriage is included since we expect that people who are married may afford to decrease their working hours in advance. Wealth is a variable that is equal to one if the taxable wealth is at least 900,000SEK.[11] This variable may reflect that individuals with a wealth may withdraw from the labor market, either gradually or totally, to a greater extent than those who do not have a wealth of 900,000SEK or higher. We have also included a dummy variable for being born in Sweden to see if there are any differences in the retirement pattern between Swedish born individuals and foreign born individuals. Yearly dummies are included to capture macro-economic shocks. Age is

[11] 800,000 SEK for the years 1993 and 1994.

controlled for by including age dummies, since we expect that the closer an individual comes to the 65th birthday the greater the possibility of withdrawing early from the labor market. To control for occupational belonging, we have included dummy variables for the different occupational schemes. Blue collar worker is the reference group. We have included this variable since we believe that workers from separate occupational schemes have different retirement behavior and expect that white collar workers may decrease their labor supply more than other occupational schemes.

In period *t* a hypothetical income from working, part-time pension and full early age retirement is calculated for each individual. LINDA has an annual structure, which means that an individual who retires during year *t* could have income from both work and retirement. Therefore, we use pension income from *t+1* and labor income from *t-1* in period *t*. Individuals who retire in period *t* will get pension income from *t+1* and labor income from *t-1* in period *t*. RFV (2001c) find that the average reduction of labor hours for part time retired individuals was 23%. Therefore, we assume that individuals who have part time pension reduce their labor supply by 25%. This means that part time income in time *t* is 91.25 %[12] (or 88.75% after 1994) of full labor income in period *t-1*. For those who work, we use the register of pensions points from the national social insurance board to calculate supplementary pension and basic pension. We use labor income in *t-1*as a benchmark for all occupational schemes. The sign of this hypothetical income is expected to be positive since if income increases, then the probability of staying in that labor markets state should also increase.

Finally, we have included two variables that measure economic incentives of the retirement decision, benefit accrual and social security wealth. Benefit accrual is the rise in pension wealth from a person working another year and then retiring. Formally, benefit accrual is defined as

$$BAC(t) = \sum_{s=t+1}^{MaxAge} \gamma^{s-t} E_t B(s,t+1) - \sum_{s=t+2}^{MaxAge} \gamma^{s-t} E_t B(s,t+2)$$

where *E* denotes the expectations operator, B(*s*, *t*) benefits received at age *s*, *t* is the retirement year and γ is the discount factor. If we hold expected benefits constant the benefit accrual will be $BAC(t) = B_t \gamma$. The benefit accrual is the retirement benefit in year *t* times the discount factor. Each month of early retirement reduces the monthly benefit by 0.5%, six

percentage points per year, before the worker's 65th birthday. Therefore, the benefit accrual at age *t* could be calculated as $B_t \gamma B$ where B_t is the change in retirement income from working one additional year. We assume that the effect of additional pension right of working an additional years is low. If the worker does not receive a higher labor income in period *t+1* than in period *t*, this could be true. This is a myopic behavior even if it has some forward looking elements. More forward looking approaches are the "option value" model (Stock and Wise, 1990a, 1990b; Lumsdaine et al, 1992; Harris, 2001) and the dynamic programming model (Rust, 1987,1989,1990; Daula and Moffitt, 1995; Rust and Phelan, 1997; French, 2000; Heyma, 2001). However, Palme and Svensson (2001) did not find that the "option value" approach performed better than the myopic method. In a survey in 1990 (Overbye, 1991) only
39 percent of the responded workers answered yes to the question if they had right to occupational pension, then everybody belonged to an occupational pension scheme. It seems that workers do not have a forward looking behavior as the "option value" and the dynamic programming method assume. Therefore, we believe that our myopic approach is appropriate to use in this study.

Social security wealth is the present discounted value of social security wealth when a person is to retire at a given age. We define social security wealth as

$$SSW(t) = \sum_{t}^{MaxAge} \gamma^{t-1} B_t$$

and assume that all individuals have the same discount rate, 1/1.03. For each age we have used life expectancy tables to calculate the social security wealth.

We have included benefit accrual and social security wealth to account for income and substitution effects of the retirement decision. The sign is expected to be positive for social security wealth since higher pension wealth should lead to an increased demand for leisure. If the return of working another year is large enough, the substitution effect leads to the individual continuing to work. Therefore we expect that the sign for benefit accrual should be negative.

Sample statistics is shown in Table 1. The sample used in this study consists of 14,301

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[12] 0.75+0.65*0.25=0.9125

observations of 4,613 males and 13,420 observations of 4,370 females between the ages of 60 and 64. To be included in the sample, an individual has to be observed for at least two years during the period 1993 to 1997.

More females than males have a high school education as their highest attained education, while more men have a university degree. Men take part time pension to a larger extent than women, but women take full early age pension to a greater degree. The full early age retirement income is around 72% of working income for males and around 59% for females. The white collar occupational scheme is the most common in our male sample, while the local government occupational scheme is the most frequent occupational scheme for women. One reason for the low share of blue collar workers in the sample may be that they, to a greater extent, have disability pension. In the sample, 92 % of males and 91% of females are native Swedes, 86 % of the men are married while 76 % of the women are married. Around 19 % of males and females have wealth above 900,000SEK.

5 Econometric specification

To analyze early retirement of older Swedish males and females, we estimate a panel random parameter logit model (see Bhat, 2000; Revelt and Train, 1997; Train, 1997; Brownstone and Train, 1999; McFadden and Train, 2000) with three labor market states: working, part time pension and full early age retirement. An individual chooses among three possible labor market states at the beginning of each year from the ages of 60 to 64. The utility individual *i* receive from state *j* at time *t* is

$$U_{ijt} = \alpha_j + \theta_j^{'} z_{it} + \gamma_i^{'} x_{ijt} + \varepsilon_{ijt}$$

where i =1… m denotes individual i, j = 1…J denotes choice alternative j and t = 1…T denotes time period t. In the utility function, α_j is an alternative specific constant that may be fixed or random; θ_j is a vector of fixed coefficients; z_{it} is a set of choice invariant individual characteristic; x_{ijt} is a vector of choice varying attributes of choices; γ_i is a vector of coefficients that is unobserved for every individual and that varies across individuals, representing each individual's preferences; and ε_{ijt} is the stochastic part of the utility function. The variance in γ_i induces correlation in utility across labor market states and time. Thus, the parameter vector for each individual, γ_i, can be expressed as $\gamma_i^{'} x_{ijt} = \gamma^{'} x_{ijt} + \lambda_i^{'} x_{ijt}$, where γ is the population mean and λ_i is the stochastic deviation that corresponds to the individual's taste relative to the average tastes of the population. The utility function takes the form

$$U_{ijt} = \alpha_j + \theta_j^{'} z_{it} + \gamma^{'} x_{ijt} + \lambda_i^{'} x_{ijt} + \varepsilon_{ijt}$$

where $\lambda_i^{'} x_{ijt} + \varepsilon_{ijt}$ is the new stochastic part of the utility function. This unobserved term is correlated across labor market states and time due to the common influence of λ_i.

Let coefficient vector λ_i vary in the population with density $f\left(\lambda_i / \xi\right)$, where ξ are the true coefficients of the distribution and the unobserved error component, ε_{ijt}, has an extreme value distribution. Then the conditional probability that individual *i* choose alternative *j* in period *t* is a standard logit:

$$\Pr_i(jt \mid \lambda_i) \frac{\exp(\alpha_j + \theta_j^{'} z_{it} + \gamma^{'} x_{ijt} + \lambda_i^{'} x_{ijt})}{\sum_{k=0}^{J} \exp(\alpha_j + \theta_j^{'} z_{it} + \gamma^{'} x_{ijt} + \lambda_i^{'} x_{ijt})}, j = 1, 2, \ldots .J$$

If we know the value of λ_i, the probability of individual *i*'s observed sequences of choices is the product of standard logit probabilities:

$$S_i\left(\lambda_i\right) = \prod_t \Pr(jt \mid \lambda_i)$$

The unconditional probability of a sequence of choices for individual *i* is then the integral over all possible values of λ_i

$$S_i\left(\xi\right) = \int S_i(\lambda_i) f(\lambda_i / \xi) d\lambda_i \qquad (1)$$

The integral cannot be calculated analytically, and exact maximum likelihood estimation is not possible to use. Therefore, the probability is approximated through simulation.[13]

6 Results

In Table 2, we present results from the panel random parameter logit specification. We have assumed that the hypothetical income variable is random and is normally distributed with a mean and a standard deviation. These results are based on 500 replications using Halton drawings.[14]

The mean parameter of the hypothetical income is positive and highly significant for both

[13] See Stern (1997) for an introduction to simulations based estimation and Stern (2000) for simulation based inference in econometrics.
[14] See Train (1999) for an introduction to Halton draws.

males and females. We can see that the standard deviation is 0.533 and 0.524, respectively, and very significant. This means that there is heterogeneity in the underlying preference structure. Our model is highly significant and the pseudo adjusted R^2 is 0.31 and 0.48, respectively.

People with a university degree have a lower probability of taking full early age retirement but a higher probability of taking part time pension. Wealth has a positive effect on part time pension for males. Otherwise wealth is not significant. Swedish born is not significant, except for part time pension where Swedish born males have a higher probably of taking part time pension than foreign born men. The probability of taking part time pension decreases throughout the period, while the probability of taking full early age retirement increases throughout the years. Although the probability of full early age retirement and the probability of part time pension increase as workers get closer to the mandatory retirement age of 65. Blue collar workers have a lower probability of taking part time pension and full early age retirement than workers belonging to other occupational schemes.

Benefit accrual and social security wealth are significant (except for women and part time retirement) and have expected signs. The probability of leaving the labor market increases with social security wealth while the probability of a total or a gradual withdrawal from the labor market decreases with benefit accrual.

Table 3 shows predicted probabilities of determining variables. We can conclude that the simulated probabilities show the same pattern for both males and females. Simulated probabilities for the part time labor market state decrease across the income distribution. The model predicts that 13.6% of the richest 10 percent men should choose part time pension while 32.3% of the poorest 10 percent men should have part time pension. Our simulations show that the probability of belonging to the work state increases across the income distribution. Labor force participation goes down by age. Part time pension decreases over time while the probability of full time early age pension increases over the period.

The next question we would like to answer is how the probability of being in a specific labor market state differs among different occupational groups. Occupational probabilities are shown in Table 4. If everyone in the male sample had their characteristics evaluated using blue collar worker parameters instead of having them evaluated at the coefficients relevant to their occupational group then 64.9% of the total sample would be in the work state, 25.2% would be in the part times state and 32.9% would be in the full early age retirement state. However, as the sample proportions shows, only 4.3% of blue collar workers had retired totally from the labor market. The sample proportions of blue collar, white collar, government and municipality workers in the different labor market states will generally be different from the occupational probabilities. This reflects that blue collar workers and white collar workers differ not just in terms of how they are treated, but also in terms of their characteristics. Thus, blue collar workers have both an occupational and a characteristic based disadvantage. The sum of these two disadvantages is referred to as the overall disadvantage (Borooah, 2002).

The occupational disadvantage experienced by blue collar workers is given by $\psi^B = \frac{\hat{p}_j^B}{\hat{p}_j^W}$, where j is the labor market state, and white collar workers are the reference group. There is no occupational disadvantage if this ratio is equal to 1. However, if the ratio is less than one, there is blue collar occupational disadvantage in labor market state j. Sample proportions are measures of overall disadvantage, $\varphi^B = \frac{s_j^B}{s_j^W}$. If the ratio is equal to 1, there is no overall disadvantage. Thus, if the ratio is less than 1, blue collar workers experience overall disadvantage in labor market state j. A measure of the characteristics disadvantage is given by $\vartheta^B = \frac{\varphi_j^B}{\psi_j^W}$. If this ratio is equal to 1, blue collar workers do not face a characteristics disadvantage, which they do if the ratio is less than 1.

The figures in Table 5 (in the column "Occupational Disadvantage" for men) show that if male blue collar and white collar workers were assigned a common set of characteristics, then the probability of a blue collar worker being in full early age retirement is 76% of the corresponding white collar worker probability. The corresponding figure is 157% for government workers and 254% for municipality workers. For full early age retirement, municipality workers have a characteristics disadvantage of 45% while government and blue collar workers have characteristic disadvantages of 6% and 58%, respectively. In the part time state blue collar workers have 81% of the corresponding white collar worker probability. The corresponding numbers are 56% for government workers and 60% for municipality workers.

Female blue collar workers have 96% of the white collar workers probability in the total withdrawal state. The corresponding figure is 146% for government workers and 255% for municipality workers. Municipality workers have a characteristics disadvantage of 37% and blue collar workers have a characteristics disadvantage of 72%.

The conclusion for both males and females is that blue collar workers have a disadvantage with respect to part time pension and full early age retirement, compared to other occupational schemes.

In Table 6, we present simulated response changes for changes in income, benefit accrual and social security wealth. The first simulation is made by increasing income by 10% in each of the three labor market states. If income in the totally retired state increases by 10% for men, the probability of belonging to that state increases by 3.5 % while the probability of working decreases by 3.5%. The effect of a pay increase is strongest in the part time state. A 10% income increase in the part time pension state implies a 14.7% and a 19% increased probability of taking part time pension for males and females, respectively. The direct effect of a 10% increase in income is positive and the cross effects is negative for both males and females.

In the second simulation we simulate a change in benefit accrual of ± 5%. We get the expected results even if men and women respond differently. When we increase the benefit accrual by 5% for men, the probabilities for part time and full early retirement decrease. An increase in benefit accrual of 5% gives that the probability of working decreases by 10.7%. Both part time and full early age retirement are affected in the same direction by a change in benefit accrual for men. Women show another pattern when we simulate changes in benefit accrual. Work and part time pension change in the same direction, a 5% increase in benefit accrual implies a 6% decreased probability of full early age retirement. The result is of the same magnitude, but positive, when we simulate a decrease in benefit accrual.

In the last simulation we change social security by ±10%. Increasing social security wealth by 10% for men leads to a negative 19% change in probability in the working state. On the other hand, a decrease in social security wealth by 10% implies a positive 15% change in the working state. The probabilities for both male part time pension and male full early age retirement go in the same direction. Increasing women's social security wealth by 10% gives a positive 13% change in probability for full early retirement, while the probability decreases by 11% if social security decreases by 10%.

7 Conclusions

In this study we have used a panel random parameter logit model to analyze early retirement for both males and females in Sweden. We have three market states: work, part time pension and full early age retirement. For every individual and year we have calculated a hypothetical income for these three labor market states. When we calculate this hypothetical income, we have taken into account the rules of the Swedish pension system and the rules of the occupational pension schemes. We have included benefit accrual and social security wealth as measures of economic incentives. These two variables allow us to measure the income and substitution effects of the retirement decision.

We found that the income variable is significant, which means that there is heterogeneity in the underlying preference structure. The probability of taking part time pension decreases throughout the period, but the probability of taking full early age retirement increases. Furthermore the probability of full early age retirement and the probability of part time pension increase with age. Blue collar workers have a lower probability of taking part time pension and full early age retirement, compared to other occupational schemes. The probability of leaving the labor market increases along with social security wealth. Part time pension and full early age retirement decrease with an increased benefit accrual. We simulated the income distribution and found that the probability of part time pension decreases across the income distribution. An increase in the hypothetical income leads to positive direct effects and negative cross effects. An increase (decrease) in benefit accrual leads to decreasing (increasing) probabilities of total withdrawal from the labor market. Social security wealth has a positive effect on the retirement decision. To conclude, economic incitements affect the decision of Swedish workers to leave the labor market either gradually or totally.

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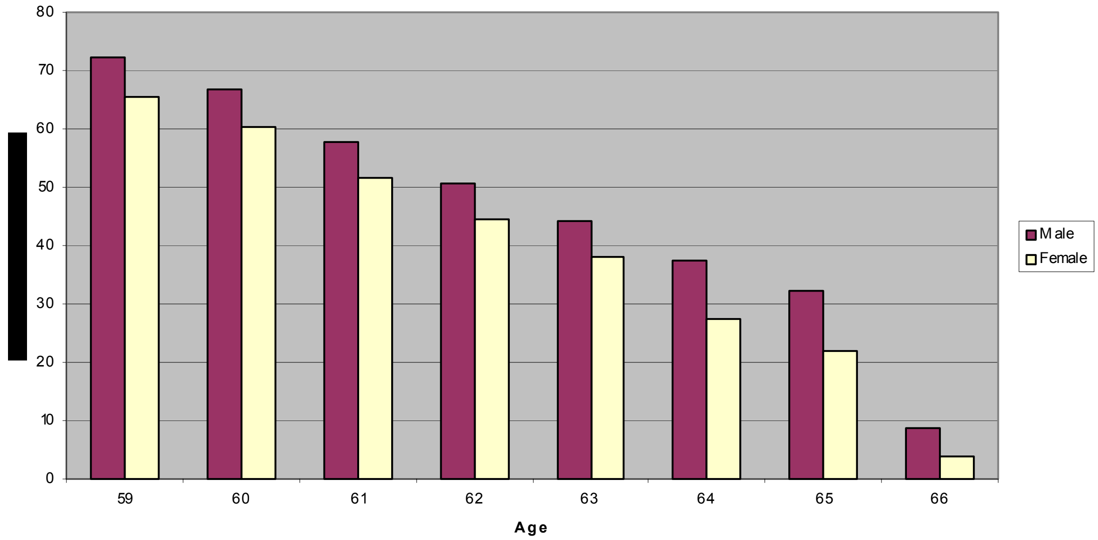
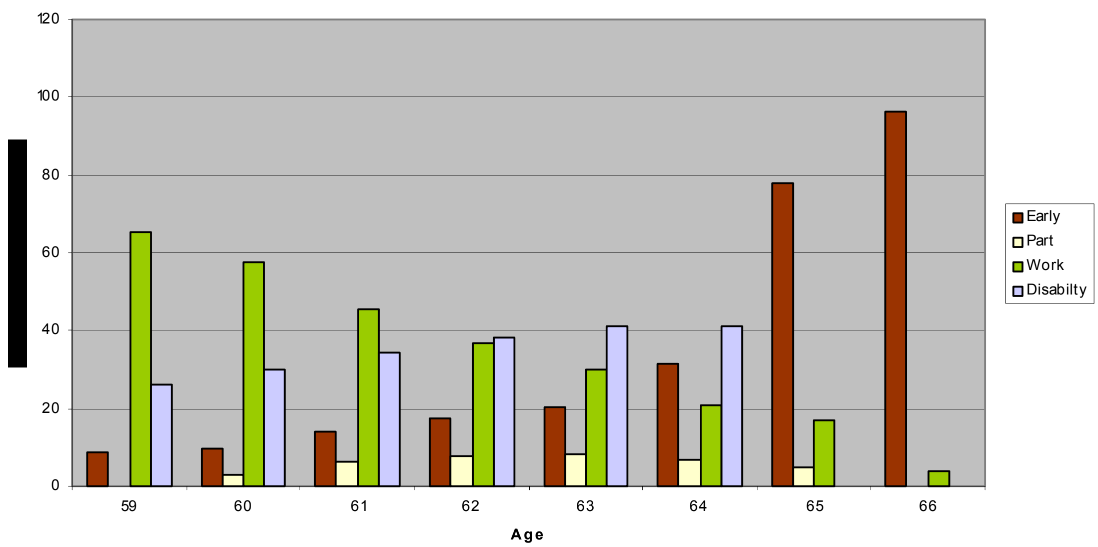
Fig.1. Labor Market Participation, Males and Females.



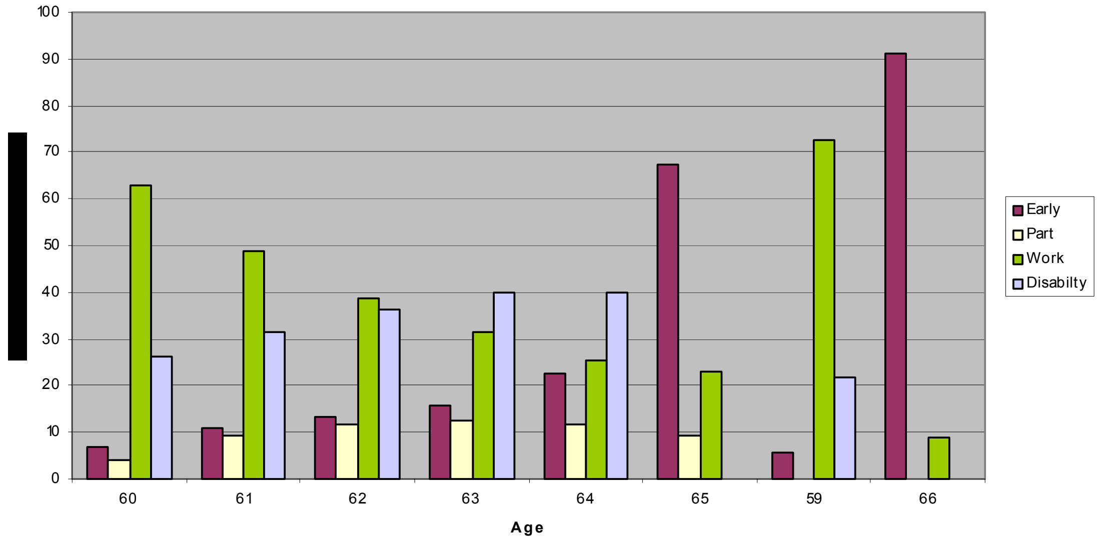
Source: Own calculations based on LINDA (1993-1999).

Fig.2. Labor Market Status, Women.



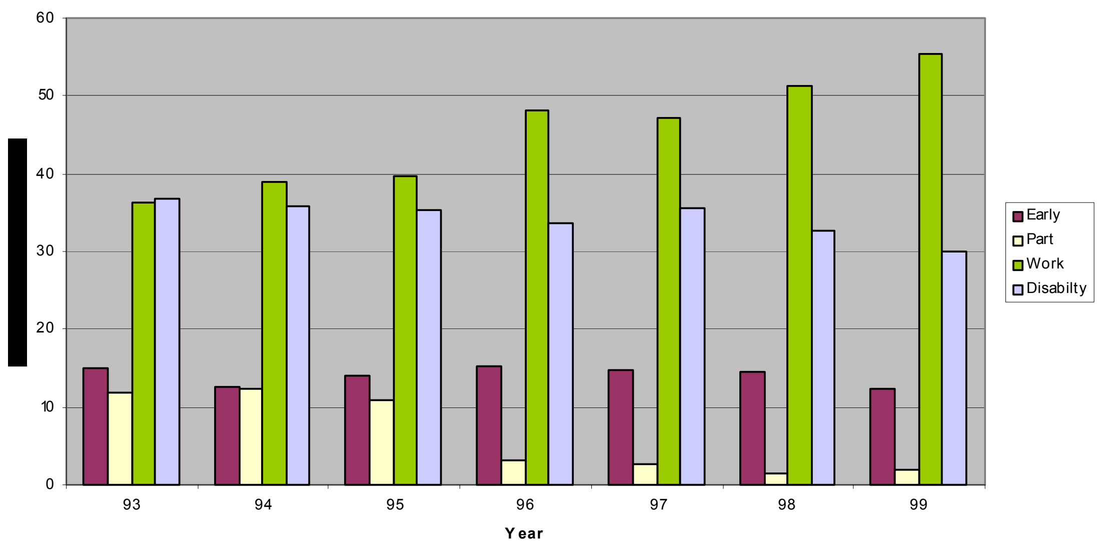
Source: Own calculations based on LINDA (1993-1999).

Fig.3. Labor Market Status, Men.



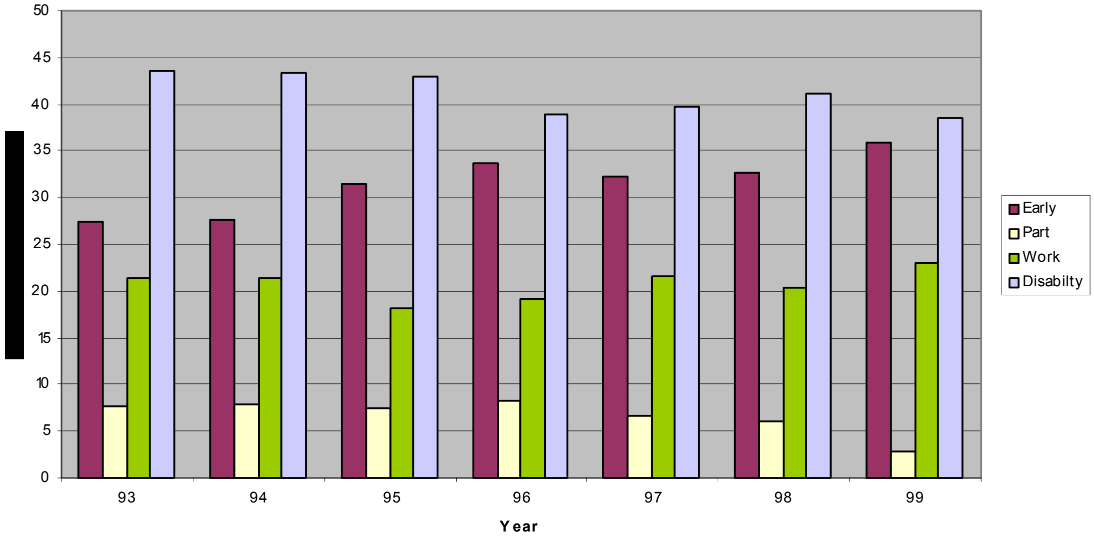
Source: Own calculations based on LINDA (1993-1999).

Fig.4. Labor Market States, Women 61.



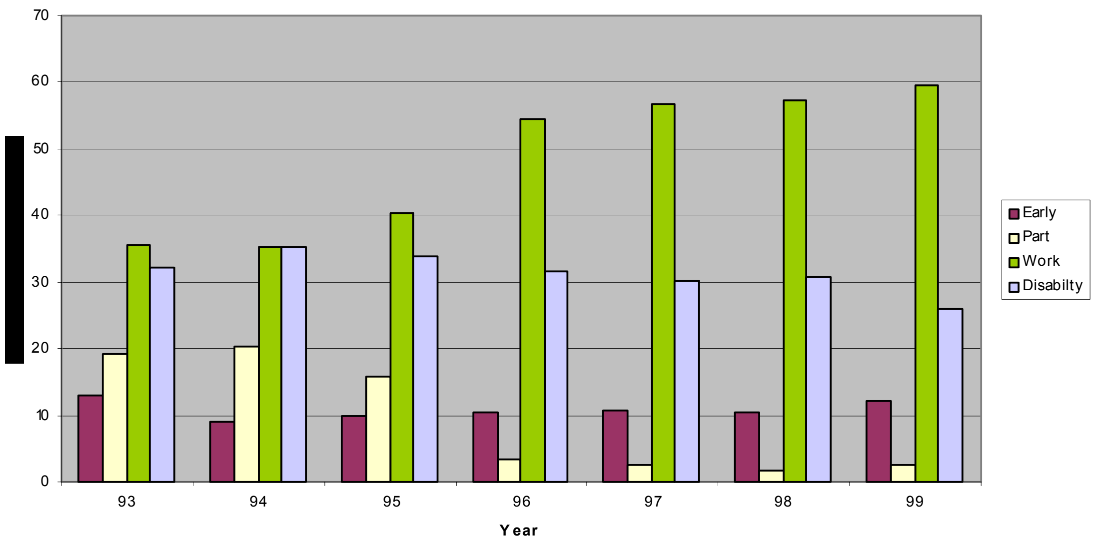
Source: Own calculations based on LINDA (1993-1999), females age 61.

Fig.5. Labor Market States, Women 64.



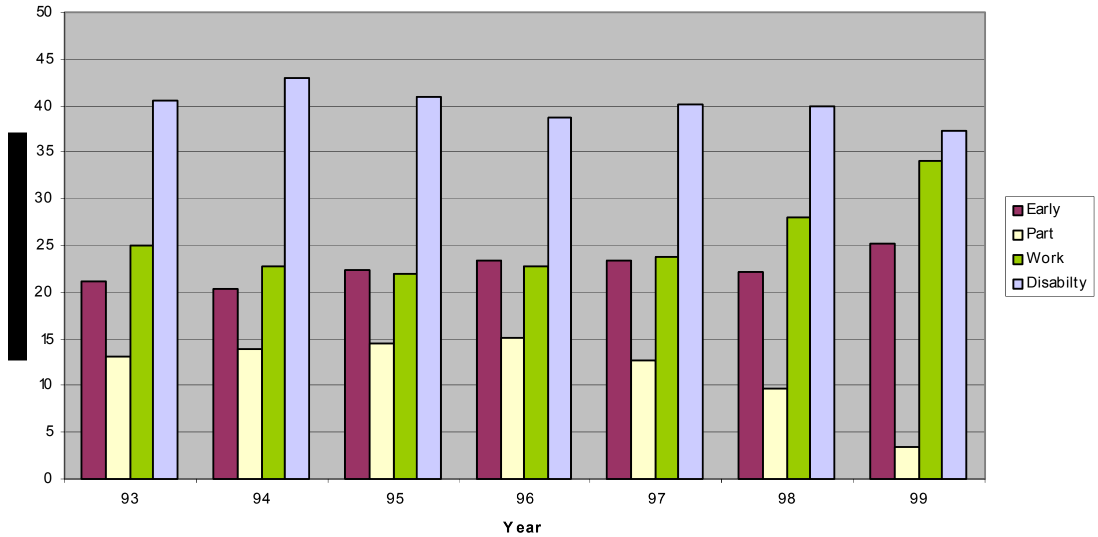
Source: Own calculations based on LINDA (1993-1999), females age 64.

Fig.6. Labor Market States, Men 61.



Source: Own calculations based on LINDA (1993-1999), males age 61.

Fig.7. Labor Market States, Men 64.



Source: Own calculations based on LINDA (1993-1999), males age 64.

Table 1. Descriptive statistics of men and women

	Men		Women	
Variables:	Mean	Std	Mean	Std
Education (highest attained):				
-Basic school	0.41	0.49	0.37	0.48
-High school	0.33	0.47	0.38	0.49
-University	0.26	0.44	0.25	0.43
Labor market status:				
-Full exit	0.13	0.33	0.21	0.41
-Part time pension	0.27	0.44	0.16	0.37
-Working	0.60	0.49	0.63	0.48
Income in different states:				
- Working	244 782	133 566	155 027	58 968
-Part time pension	219 875	119 664	139 177	52 858
- Full exit	177 152	103 784	90 638	48 052
Occupational pension schemes:				
-Blue collar, private	0.27	0.44	0.15	0.35
-White collar, private	0.39	0.49	0.14	0.35
-Central government	0.15	0.36	0.11	0.31
-Local government	0.19	0.39	0.61	0.49
Married	0.86	0.35	0.76	0.43
Wealth	0.19	0.39	0.18	0.39
Swedish born	0.92	0.27	0.91	0.29
Year=93	0.20	0.40	0.18	0.39
Year=94	0.24	0.43	0.23	0.42
Year=95	0.22	0.41	0.22	0.41
Year=96	0.20	0.40	0.21	0.41
Year=97	0.14	0.35	0.15	0.36
Age=60	0.18	0.38	0.19	0.39
Age=61	0.23	0.42	0.24	0.43
Age=62	0.22	0.37	0.22	0.42
Age=63	0.21	0.40	0.20	0.40
Age=64	0.16	0.43	0.14	0.35
Social security wealth	2,629,307	17,974,464	1,532,435	801,361
Benefit accrual	187,782	1,243,356	96,076	50,935
Number of individuals	4,613		4,370	
Number of observations	14,301		13,420	

Table 2. Parameter estimates from panel random parameter logit model

	Men		Women	
Variables:				
Income	Mean coefficient:	0.499 (0.021)	Mean coefficient:	0.922 (0.037)
	Standard deviation:	0.533 (0.023)	Standard deviation:	0.524 (0.028)
	States:			
	Part time pension	Full Early Exit	Part time pension	Full Early Exit
Intercept	-1.016 (0.183)	-5.252 (0.391)	-1.815 (0.186)	-8.262 (0.499)
Basic School	-0.398 (0.082)	0.428 (0.139)	-0.858 (0.089)	0.586 (0.183)
High School	-0.126 (0.077)	0.688 (0.120)	-0.739 (0.085)	0.867 (0.170)
Married	0.135 (0.075)	0.061 (0.136)	0.264 (0.072)	2.204 (0.158)
Wealth	0.353 (0.070)	-0.053 (0.114)	0.084 (0.078)	-0.553 (0.137)
Swedish	0.709 (0.102)	0.510 (0.185)	0.152 (0.110)	-0.049 (0.200)
Year=94	0.037 (0.065)	0.802 (0.136)	0.343 (0.073)	0.113 (0.146)
Year=95	-0.184 (0.072)	0.915 (0.145)	0.274 (0.086)	0.222 (0.154)
Year=96	-0.598 (0.079)	1.084 (0.150)	0.410 (0.099)	0.608 (0.162)
Year=97	-0.848 (0.094)	1.266 (0.163)	0.630 (0.126)	1.081 (0.184)
Age=61	1.299 (0.091)	1.345 (0.162)	0.513 (0.119)	0.815 (0.181)
Age=62	2.118 (0.120)	2.001 (0.228)	0.572 (0.119)	1.676 (0.227)
Age=63	2.885 (0.162)	2.978 (0.302)	0.684 (0.120)	3.382 (0.294)
Age=64	3.103 (0.172)	3.742 (0.313)	0.855 (0.126)	4.982 (0.389)
White collar	0.535 (0.078)	0.967 (0.187)	0.214 (0.109)	-0.044 (0.256)
Government	-0.293 (0.098)	1.774 (0.200)	0.360 (0.119)	1.519 (0.264)
Municipality	0.294 (0.086)	3.479 (0.205)	-0.592 (0.098)	3.709 (0.252)
Benefit accrual	-0.665 (0.075)	-1.092 (0.107)	0.023 (0.186)	-2.388 (0.247)
Social security wealth	0.042 (0.005)	0.075 (0.008)	0.012 (0.012)	0.169 (0.016)
Observations	14,301		13,420	
Individuals	4,613		4,370	
Log likelihood	-10,818		-8,109	
Prob > $\chi 2(40)$	0.000000		0.000000	
Pseudo $\overline{R}^2$	0.31		0.45	

Table 3. Predicted Probabilities (%) of Determining Variables

Predicted Probabilities at mean	Predicted probability of being in:		
Values of determining variables	Work	Part-time	Full-early
	Men		
Base case:	58.4	25.0	16.6
Poorest 10 percent	55.6	32.3	12.1
Poorest 25 percent	58.9	29.9	11.2
Below the median	60.7	28.2	11.1
Richest 25 percent	71.7	18.0	10.3
Richest 10 percent	75.0	13.6	11.4
Age=60	82.2	8.0	9.8
Age=64	24.0	38.3	37.7
Year=1993	57.7	30.7	11.6
Year=1997	63.2	15.9	20.9
Basic school	60.1	22.6	17.2
High school	56.1	25.7	18.1
University	56.7	29.4	13.9
	Women		
Base case:	62.0	15.6	22.4
Poorest 10 percent	46.9	25.1	28.0
Poorest 25 percent	54.4	26.0	19.6
Below the median	59.6	25.2	15.2
Richest 25 percent	74.9	14.3	10.8
Richest 10 percent	77.0	11.7	11.3
Age=60	75.9	10.6	13.5
Age=64	47.4	14.0	38.6
Year=1993	61.2	18.8	20.0
Year=1997	63.1	10.0	26.9
Basic school	63.9	13.2	22.9
High school	61.9	13.9	24.2
University	58.9	22.2	18.9

Table 4. Predicted Probabilities (%) of Different Occupational Schemes

Predicted Probabilities at mean	Predicted probability of being in:		
Values of determining variables	Work	Part-time	Full-early
	Men		
Blue Collar:	64.9	25.2	9.9
White Collar:	55.6	31.3	13.1
Government:	62.1	17.4	20.6
Municipality:	47.8	18.8	33.4
Sample proportions:			
Blue Collar:	62.9	32.8	4.3
White Collar:	58.4	28.2	13.3
Government:	62.7	17.8	19.5
Municipality:	57.8	23.5	18.8
	Women		
Blue Collar:	66.9	22.0	11.1
White Collar:	71.4	17.1	11.6
Government:	63.0	20.1	16.9
Municipality:	57.8	12.7	29.6
Sample proportions:			
Blue Collar:	78.0	17.9	4.2
White Collar:	62.6	21.7	15.7
Government:	53.2	23.7	23.1
Municipality:	61.3	13.4	25.3

Table 5. Estimates of Occupational, Characteristic, and Overall Disadvantage

		Occupational Disadvantage (%)	Characteristic Disadvantage (%)	Overall Disadvantage (%)
Disadvantage calculated from predicted probabilities at mean values of determined variables				
			Men	
Blue collar/ White collar				
Work:		1.17	0.92	1.08
Part time:		0.81	1.06	0.86
Full-early:		0.76	0.42	0.32
Government/ White collar				
Work		1.12	0.96	1.07
Part time		0.56	1.12	0.63
Full-early		1.57	0.94	1.47
Municipality/ White collar				
Work		0.86	1.15	0.99
Part time		0.60	1.38	0.83
Full-early		2.54	0.55	1.41
			Women	
Blue collar/ White collar				
Work:		0.94	1.33	1.25
Part time:		1.29	0.64	0.83
Full-early:		0.96	0.28	0.27
Government/ White collar				
Work		0.88	0.96	0.85
Part time		1.18	0.92	1.09
Full-early		1.46	1.01	1.47
Municipality/ White collar				
Work		0.81	1.21	0.98
Part time		0.74	0.84	0.62
Full-early		2.55	0.63	1.61

Note: Advantage if value > 1. Calculated from the figures in Table 4

Table 6. Simulation Responses Changes in Percent for Changes in Income, Benefit Accrual and Social Security Wealth

Predicted changes at mean	Predicted changes being in:		
values of determining variables	Work	Part-time	Full-early
Simulations:	Men		
Income (Early)*1.10	-3.5	0.0	+3.5
Income (Part)*1.10	-0.1	+14.7	-14.6
Income (Work)*1.10	+9.4	-7.3	-2.1
Benefit accrual*1.05	+9.2	-5.0	-4.2
Benefit accrual*0.95	-10.7	+4.9	+5.8
Social security wealth*1.10	-19.3	+7.1	+12.2
Social security wealth*0.90	+15.2	-8.2	-7.0
	Women		
Income (Early)*1.10	-2.1	-1.1	+3.2
Income (Part)*1.10	-17.7	+19.5	-1.8
Income (Work)*1.10	+10.9	-8.1	-2.8
Benefit accrual*1.05	+3.0	+3.0	-6.0
Benefit accrual*0.95	-3.5	-2.6	+6.1
Social security wealth*1.10	-9.7	-3.8	+13.4
Social security wealth*0.90	+8.1	+3.3	-11.4