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An Analysis of the Determinants of Job Satisfaction when Individuals' Baseline Satisfaction Levels May Differ

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Abstract A growing literature seeks to explain differences in individuals' self-reported satisfaction with their jobs. Most of the accumulated evidence so far has, however, been based on cross-sectional data and when panel data have been used, individual unobserved heterogeneity has been modelled following the random effects approach, namely using the ordered probit model with random effects. This paper makes use of longitudinal data for Denmark, taken from the waves 1995-1999 of the European Community Household Panel, and estimates fixed effects ordered logit models using the estimation methods proposed by Ferrer-i-Carbonel and Frijters (2004) and Das and Van Soest (1999). For comparison and testing purposes a random effects ordered probit is also estimated. Estimations are carried out separately on the samples of men and women for individuals' overall satisfaction with the jobs they hold. We find that using the fixed effects approach (that clearly rejects the random effects specification), considerably reduces the number of key explanatory variables. In addition to wages, good health and being a public sector employee are particularly important in explaining individual differences in job satisfaction. Moreover, the impact of being employed on a temporary contracts or working in the public sector differs between the genders.

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Keywords: Job satisfaction, fixed effects ordered logit model, random effects

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1 Introduction

In recent years economists have taken an increasing interest in the analysis of the subjective well-being of individuals; see Frey and Stutzer (2002) for a recent review. In the field of labour economics, following the seminal papers by Hamermesh (1977), Freeman (1978) and Clark and Oswald (1996), this has spawned a growing number of studies of the determinants and consequences of differences in individuals' reported job satisfaction. Work psychologists have for a long time been arguing that for most people jobs cannot only be characterised by the pay and hours of work associated with them, as standard economic analysis does, but also by job and workplace features like promotion and other career prospects, job security, job content and interpersonal relationships; see Warr (1999) for a comprehensive survey. In fact, when the employees are asked, as in *e.g.* the International Social Survey Programme – see Clark (1999) – they typically rank job security and job interest highest, whilst pay and hours of work are found in the opposite end of the ranking.

Economists have a longstanding tradition of viewing subjective measures of individuals' preferences with considerable scepticism. As described by Wansbeek and Kapteyn (1983): "Utility seems to be to economists what the Lord is to theologians. Economists talk about utility all the time, but do not seem to have hope of ever observing it this side of heaven. In micro-economic theory, almost every model is built on utility functions of some kind. In empirical work little attempt is made to measure this all-pervasive concept. The concept is considered to be so esoteric as to defy direct measurement by mortals. Still, in a different role, viz., of non-economists, the same mortals are the sole possessors of utility functions and can do incredible things with it." The arguments that individuals may differ with respect how they scale feelings and hence communicate their well-being level, that well-being is ordinal (not cardinal), and that subjective feelings may be reflecting their innate personalities are obviously all valid and important objections that should not be swiftly dismissed. This led Stigler and Becker (1977) to conclude: "economists continue to search for differences in prices or incomes to explain any differences or changes in behaviour" (p. 76).[1] On the other hand, the often made claim that much of economic analysis considers intrinsic motivational factors to be unimportant is obviously wrong. Rather, as is eloquently discussed in Lazear (2000), economic analysis is concerned with the study of situations where the intrinsic motives are taken as given and the aim of the analysis is to examine the influence of extrinsic motives and especially the trade-offs economic agents face at the margin.

[1] For a more recent and very useful discussion and summary of the experimental and field data literature on the meaningfulness of answers to subjective questions, see Bertrand and Mullainathan (2001).

And yet, scholars in other social sciences, especially in psychology, have taken individuals' responses to questions about their perceived well-being much more seriously and based much of their empirical evidence on this type of information. A substantial body of research has been built showing that job satisfaction is strongly correlated with several mental physical health indicators. In parallel, a growing number of studies focusing on life satisfaction or financial situation have appeared in economics; see e.g. Frijters et al. (2004) and Bonke and Browning (2003).

It should be noted that the concept of job satisfaction has certainly not been a cornerstone in the economic analysis of the labour market. Rather, in many analyses job satisfaction is more or less absent. Nevertheless, in many countries firms and employers pay close attention to the subjective well-being of their employees and to how these perceive their current jobs. Thus, for instance in Denmark, several of the major companies are regularly carrying out their own worker/job satisfaction surveys, and an employee satisfaction index constructed using identical questionnaires has in recent years been computed for an increasing number of European countries.[2] At the macro-economic level, the European Union has called the member states' attention to the quality aspects of work and has emphasized the importance of improving job quality in order to promote social inclusion and employment (European Commission 2001; 2002).

Job quality is certainly a multi-dimensional concept. Related to this, some authors distinguish between the economic contract and the psychological contract. In the economic contract the focus is on the relationship between effort and reward, while in the psychological contract the interest is mainly in the working conditions. A further distinction is made between extrinsic and intrinsic job characteristics, the former being concerned with financial rewards, working time, work/life balance, job security and opportunities for advancement and the latter with features such as job content, work intensity, risk of ill health or injury and relationships with co-workers and managers. Because of this multi-dimensionality, the possibility of using one dimension to classify jobs according to their quality is often rejected. A similar approach has been taken by the EU. Namely the Employment in Europe (2001) report suggests that "in the absence of a single composite indicator, any analysis of job quality must be based on data on both objective and subjective evaluations of the worker-job match". The EU with the report Employment in Europe 2002 goes further in including job satisfaction in its definition of quality of work, and claims that "in all Member States self-reported job satisfaction is strongly positively correlated with wages, job status and job related skills acquired through training". We therefore think that job satisfaction can be considered, at least to some extent, a good proxy for job quality.

The current paper is concerned with identifying what lies behind differences in people's subjectively reported job satisfaction and changes therein. Earlier research has typically found that consistent with economic theory, pay and work hours are positively and negatively, respectively related to job satisfaction. Other important contributing factors are individual traits, such as age and gender, and some features characterising the individuals' workplaces and jobs. These studies have, however, been mainly based on cross-sectional data, and moreover to a surprisingly high extent on a single data set – the British Household Panel Survey (BHPS). Even when researchers have had access to panel data on employees' job satisfaction levels, with only a few exceptions they have exploited the longitudinal nature of the data in their analyses. In that case, the statistical model predominantly used has been the ordered probit with random effects (see Butler and Moffitt 1982; Frechette, 2002).

In this paper we make use of data for Denmark from the European Community Household Panel (ECHP), more specifically the five waves from the period 1995-99. The waves 1994-1998 have recently been analysed in European Commission's (2002) annual report *Employment in Europe*. Denmark has been shown to have among the most satisfied workers in the world; in Europe only Austria and Ireland reach similar levels (see Blanchflower and Oswald (undated) and Sousa-Poza and Sousa-Poza, 2000). The same trend is confirmed by charts 1 and 2 that report job-satisfaction over the Member States countries for years 1995 and 1999.

The approach in our paper differs from that in previous analyses in that we make use of longitudinal job satisfaction data while at the same time preserving the ordered nature of the information in the fixed effects approach. We estimate fixed effects ordered logit models on two samples, males and females, for the individuals' overall job satisfaction using the estimators recently developed by Ferrer-i-Carbonel and Frijters (2004) and Das and Van Soest (1999). For comparison and testing purposes we also estimate a random effects ordered probit model. We therefore attempt to make use of the panel element of the ECHP to deal with the problem of unobserved heterogeneity. That is, certain reported levels of overall individual job satisfaction may be recorded because underlying unobservable characteristics, which vary across individuals, may increase the probability that a certain level of job satisfaction is reported as opposed to another. For instance, we can assume that an individual's

[2] See www.europeanemployeeindex.com.

emotional state or "mood" may influence positively or negatively his or her reported levels of job satisfaction irrespective of job, industry or other personal characteristics.
We find that using the fixed effects approach (that clearly rejects the random effects specification), considerably reduces the number of key explanatory variables. In addition to wages, good health and being a public sector employee are particularly important in explaining individual differences in job satisfaction. Moreover, the impact of being employed on a temporary contracts or working in the public sector differs between the genders.

The remainder of the paper is organised as follows. Section two provides a brief review of the economic literature on the topic of the paper. Section three outlines the data used and section four discusses the empirical strategy adopted. Section five gives the results. The sixth section summarises our conclusions.

2 Previous research

As was already mentioned above, economists' interest in job satisfaction is of relatively recent date, whereas sociologists and work psychologists have a considerably longer and hence more extensive experience of examinations of the determinants and impact of job satisfaction; for an excellent summary; see Warr (1999). This literature differs in at least three respects from how economists have approached essentially the same data sets and closely related questions. First, the dependent variable in analyses aiming at understanding the factors underlying differences in job satisfaction across individuals has usually been constructed by averaging the ordinal responses to the questions concerning satisfaction. Thus, satisfaction is implicitly assumed to be cardinal.

Second, the vast majority of the non-economists' investigations employ ordinary least squares as their estimation technique. This follows naturally from the implicit assumption that the job satisfaction responses are cardinal. Moreover, they do not account for the fact that the dependent variable is bounded. Typically, the literature has little discussion both of measurement errors in the dependent variable and of what is subsumed in the error term. Most of the psychological as well as the economic research have been based on cross-sections. As a consequence, little attention has been paid to the importance of individual differences in baseline job satisfaction levels, which in a longitudinal framework could be modelled as individual-specific fixed effects. Clark and Oswald (2002) discuss the role of fixed effects in studies of well-being. The method used in their application is however OLS, that is, cardinality is implicitly assumed. Other factors that can affect self-reported job satisfaction measures and which should be taken into consideration in thinking about what goes into the error term include circumstances – local or of business cycle type – and aspirations. The latter may for instance be cohort-specific. For an analysis of business cycle influences, see Gerlach and Stephan (1996). A previous study recognising the potential importance of controlling for fixed effects is Winkelmann and Winkelmann (1998), in which the effects of changes in labour force status on life satisfaction are examined. But in order to enter fixed effects, they collapsed the satisfaction variable into a binary variable and used Chamberlain's (1980) conditional logit estimation technique. The same approach is adopted by Hamermesh (2001).

Finally, as pointed out in Clark and Oswald (1996), unlike economists, scholars in psychology and related fields have not entered working hours as an explanatory variable. This reflects the economists' notion of the satisfaction equation as an empirical counterpart of a utility function in which income and leisure are the natural arguments.

In view of the fact that there are many surveys available that contain job satisfaction questions[3] (which furthermore are quite similar across surveys), the data of which are frequently used by economists, surprisingly few economic studies have been carried out. This could, at least at a certain extent, reflect the economists' suspicion towards variables that measure what people say rather than what they do. Next, we give a brief review of the work carried out by economists in the area. Some of the key characteristics of the studies surveyed are collected in *Table 1*.

The early contributions to the economic job satisfaction literature are from the late seventies. Hamermesh (1977) is to the best of our knowledge the first to develop and test a theory of overall job satisfaction, whereas Freeman (1978) and Borjas (1979) examined the relationship between unionism and job satisfaction where the latter is adopted with the motivation that it is a measure that captures other aspects of the workplace, which are not reflected by conventional objective variables. This line of research of the effects of unionism has been picked up in several later studies (see *e.g.,* Sloane and Bender, 1998), which find trade union membership being associated with lower reported job satisfaction.[4] The same reasoning concerning the nature of the subjective satisfaction measure led

[3] A non-exhaustive list of examples includes: in the United States: NLS and PSID; in Europe: ECHP (15 different EU countries), BHPS for Britain and GSOEP from Germany.
[4] A recent paper by Bryson, Capellari and Lucifora (2003) account for the endogeneity of union membership and find this to significantly reduce the dissatisfaction of unionised workers.

Freeman (1978) to include it as an additional regressor in his otherwise standard model for explaining the quit behaviour of employees. This line of research has also been followed up more recently in studies by Clark, Georgellis and Sanfey (1998), and Lévy-Garboua, Montmarquette and Simmonet (2002), in which job satisfaction information is used for testing alternative theories of individual labour market outcomes like quits and wages.

The 1990s witnessed a renewed interest in job satisfaction research among economists spawned by a series of papers by in particular, Clark and Oswald. Clark (1996) makes use of three different measures of job satisfaction obtained from the first wave of the British Household Panel Survey (BHPS) and carries out an ordered probit analysis of the importance of individual and workplace characteristics in explaining reported differences. His main findings, several of which have been replicated in later studies using additional and later waves from the BHPS are that being male, in the thirties or older, a union member, well-educated, working longer hours and being employed in larger establishments, all lower the individual's level of job satisfaction. The data used by Clark and Oswald (1996) also come from the first wave of the BHPS (1991). The satisfaction model is estimated by ordered probit for two dependent variables: overall job satisfaction and pay satisfaction. The focus in their paper is on relative versus absolute income as a determinant of job satisfaction and in the case of the former on different comparison groups. Specifications with relative income clearly outperform those with absolute income as a regressor. They use the same variables as in Clark (1996) – obtaining similar results – plus two new ones: dummies for whether the employee is employed on a temporary contract and for whether she is in a managerial or supervisory position. The dummies attach negative and positive coefficients, respectively.

A third paper published the same year, using the same data set and methods is Clark, Oswald and Warr (1996), which contains a more detailed analysis of the relationship between the age of the employee and her job satisfaction level. The relationship is found to be U-shaped, reaching its lowest level at about the age of forty. In the fourth study using the 1991 wave of the BPHS, Clark (1997), the focus is on gender differentials in job satisfaction, and in particular on why women's satisfaction levels exceed those of male employees. This is a rather surprising observation in view of women's disadvantaged position in labour market with respect to earnings and promotions. Estimation is mainly by ordered probit, but the robustness of the results is checked by principal components, too. When estimated separately for women and men, most of the results of the other studies from the same data source are replicated except for marital status, managerial status, hours of work and union membership which are statistically significant for women only (the first two with a positive, and the other two with a negative sign). Clark's preferred explanation for why female employees are more satisfied with their jobs is that women's jobs have improved relative to their expectations. (See also Sloane and Williams, 2000 for similar conclusions from a study based on data from the British academic labour market[5]). His conclusions is supported by Sanz de Galdeano (2002), who uses the waves 1991-98 of the BHPS and employs both Heckman selection models and propensity score methods in order to correct for differences in personal and job characteristics by gender as well as for potential sample selection problems.

All four above mentioned studies are based on a single cross-section from the BHPS. A more recent study by Gardner and Oswald (2001) makes use of the panel data for the years 1991-99 that can be obtained from the BHPS. The aim of their analysis is to explain the behaviour of two dependent variables: one is the GHQ12, which is a widely used measure of subjective well-being with a considerable weight put on mental health and the other is a simple index running from 1 to 6 based on answers to the question about overall job satisfaction. The GHQ12 scores are analysed by OLS regression, whereas the job satisfaction data are once again modelled by ordered probit (as a matter of fact by ordered logit in the update). The main focus of the study is on time-series changes in subjective well-being/satisfaction and in particular on differences between the private and public sectors therein. Only the analysis of the GHQ12 makes use of the longitudinal character of the data by including individual fixed effects. In the job satisfaction analysis where the dependent variable is ordered, the data are treated as annual cross-sections. The ordered logit estimates with year dummies and individual and job/workplace characteristics as regressors show that job satisfaction is positively related to pay and public sector employment, and negatively to: hours, educational level, being male, of ethnic origin, workplace size, being in a temporary job, and union recognition in the workplace. The relationship is U-shaped with respect to age and job tenure. The authors also find a discernable negative time trend in job satisfaction, which is particularly pronounced among public sector employees. This is quite remarkable considering the fact that the nineties was a period of strong economic growth.

Two recent papers from the UK have utilized different data sources. In a study on self-employment, Blanchflower and Oswald (1998) estimate ordered probit models on two samples; one from the 1981 National Child Development Study when the respondents were 23 years of age, and the other from the same source ten years later. In 1981 the question is about satisfaction with the respondent's current job

[5] See also Leontaridi and Sloane (2002).

"as a whole". The question asked ten years is about "satisfaction with the way life has turned out". They find that females, married and part-time workers were more satisfied than males, non-married and full-time employees. Union membership changed from having a negative impact in 1981 to a positive one in 1991. In both observation years the self-employed are observed to be significantly more satisfied than wage earners.

Lydon and Chevalier (2001) examine two cohorts (from 1985 and 1990) of graduates from UK higher education institutions using data from the Higher Education Funding Council for England Survey carried out in 1996. At the time of the survey cohort members were on average 34-35 and 31 years of age, respectively. The key question addressed in the paper is the effect of the potential endogeneity of wages in job satisfaction studies. In effect, the authors find that the direct wage effect is doubled once endogeneity is controlled for.[6] Job satisfaction is estimated with ordered probit models. According to the estimates, pay, managerial status and the number of children have a significant and positive impact on the individual's job satisfaction, whereas the number of weekly working hours, public sector employment, clerical job, workplace size, age and being a male has the opposite effect. The employee's educational level and months as employed turned out insignificant.

All of the more recent studies of the factors underlying differences in individual job satisfaction we have discussed so far have used data from Britain. This is no coincidence as the bulk of research has been carried out in that country. Outside Britain, investigations on the topic have been rather thin on the ground. For the U.S., Hamermesh (2001) have carried out an analysis of changes in the distribution of young men's self-reported job satisfaction between years 1978, 1988 and 1996 using logit models. The dependent variable being there retrieved from the dummy reported in the NLSY "likes the job very much". The article also contains a corresponding analysis of German data (from the GSOEP) for the period 1984-96. The GSOEP data have 11 satisfaction levels and their determinants are analysed by OLS. For both countries Hamermesh only enters wage variables as explanatory variables in his logit models.

One important source of information regarding job satisfaction is the European Community Household Panel (ECHP), which for a large number of countries contains the same seven questions concerning the respondent's level of satisfaction with respect to different aspects of her current job. This battery of questions is identical to that in the much used BHPS. The Employment in Europe

[6] Other explanatory variables in job satisfaction analyses that plausibly can be conceived of as endogenous are hours, part-/full-time job, job tenure, and marital status.

(chapter 3, 2002) presents a pooled ordered probit analysis of the reported overall job satisfaction measures for 14 countries and 4 years (1995-98). The model is estimated under the assumptions that the effects of the individual and job characteristics are the same across countries, whereas there may be differences between countries and years. The two differences are picked up by year and country dummies. The results are quite similar to those reached by the studies on the BHPS. The same applies also for the results when the model is estimated separately for men and women.

Another study using the ECHP for years 1995-97 and with a special focus on the impact of different employment constellations – more precisely, the terms of the employment contract (fixed or permanent) and the length of the working day – is Kaiser (2002), which makes use of data for Denmark, Germany, the Netherlands, Portugal and the United Kingdom. Kaiser estimates probit models for high versus low satisfaction on both pooled and individual country data. He finds that in most of the countries under study fixed-term contracts are associated with lower reported job satisfaction levels. Satisfaction levels appear to differ little between employees working part- and full-time. Of course, both findings beg the question whether these features of the employment contracts are exogenous. An interesting result in Kaiser's study is that the higher job satisfaction of female workers found by Clarke (1997) could not be replicated for Denmark, the Netherlands or Portugal. In fact, the level is found to be lower and statistically significantly so for the two latter countries.

3 Econometric analysis

In order to analyse overall job satisfaction we use the random effects ordered probit model and the fixed effects ordered logit estimator recently proposed by Ferrer-i-Carbonel and Frijters (2004)[7]. For comparison purposes and as a robustness check we also use the estimator proposed by Das and Van Soest (1999).

Considering cross-sectional data, the ordered probit/logit model arises when considering an independent sample of data $\{y_i, x_i\}$ where the dependent variable y_i has M possible outcomes with a "natural" ordering. Consider a latent variable $y_i^* = x_i'\beta + u_i \ for\ i = 1,...,n$ where x are explanatory variables and u is the error term.

[7] The fixed effects ordered logit model is used also in the companion paper by Frijters, Haisken-DeNew and Shields (2004).

Define the following observability criterion $y_i = \lambda \ if \ \lambda_{k-1} \leq y_i^* \leq \lambda_k \ for \ k = 1,....,K$. Let $\lambda_0 < \lambda_1 < ... < \lambda_M$ and $\lambda_0 = -\infty$ and $\lambda_K = \infty$. The conditional probability of observing $y_i = m$ is

$$\begin{aligned}\Pr(y_i = k \mid x_i) &= \Pr(\lambda_{k-1} \leq y_i^* \leq \lambda_k) \\ &= \Pr(\lambda_{k-1} \leq x_i^{'}\beta + u_i \leq \lambda_k)\end{aligned}$$

Rearranging terms gives

$$\begin{aligned}\Pr(y_i = k \mid x_i) &= \Pr(\lambda_{k-1} - x_i^{'}\beta \leq u_i \leq \lambda_k - x_i^{'}\beta) \\ &= \Pr(u_i \leq \lambda_k - x_i^{'}\beta) - \Pr(u_i \leq \lambda_{k-1} - x_i^{'}\beta)\end{aligned}$$

For $k=1,\ldots,K$

Assuming a standard normal distribution for the error term yields the ordered probit model. Symmetrically assuming a logistic distribution leads to the ordered logit model

Both these models have been often used with cross-sectional data in analyses about well-being and satisfaction. The ordered probit model has also been used in longitudinal studies. In that case unobserved heterogeneity has been dealt within the random effects approach. The fixed effects approach has been rarely followed owing to the lack of suitable econometric methods. However, some authors have adopted it by transforming the ordinal variable into a binary one that takes the value of one above (or under) a specific threshold. Assuming that the error term is logistic yields the ordinary logit model that can be estimated using standard likelihood methods (see Andersen, 1970; Rasch, 1970 and Chamberlain, 1980).

Two recent studies have proposed new estimations methods that can handle the original rankings of the dependent variable in the fixed effects approach. These are Das and Van Soest (1999) and Ferrer-i-Carbonel and Frijters (2004)[8]. Those models have the particularly appealing property that no particular correlation is assumed between the fixed individual effects and the error term. Moreover, while the random effects ordered probit model assumes "ordinal comparability", i.e. that satisfaction is interpersonally comparable, implying that if $S_i > S_j \ then \ W_i > W_j$ (S, standing for "satisfaction" and W, standing for "well-being"), the fixed effects ordered logit does not.

In our approach, we assume that (see Ferrer-i-Carbonel and Frijters, 2004)

[8] See also Frijters et al. (2004), Ejrnaes and Pörtner (2002).

1. Job satisfaction (JS) is a positive monotonic transformation of an underlying concept called well-being: if $JS_{it} > JS_{is}$ then $W_{it} > W_{is}$ $(t \neq s)$.
2. Both time-invariant, υ_i, and time-varying unobserved factors, ε_{it}, are present. However, while the former are related to observed factors, the latter are not, i.e. $\operatorname{cov}(\varepsilon_{it}, x_{it}) = \operatorname{cov}(\upsilon_i, \Delta x_{it})$ *and* $\operatorname{cov}(\upsilon_i, x_{it}) \neq 0$

The first assumption implies that there is a correspondence between what is measured, JS_{it} (for an individual *i=1,…N* and for a time period *t=1,….,T*) and what we are interested in, namely a form of well-being, i.e. W_{it}. Several studies have shown that reported general satisfaction levels are likely to perform well in predicting the underlying concept of welfare. This in turn implies that self-reported job satisfaction levels can be used as proxies for the well-being in the job sphere. The second assumption is more strictly related to the statistical properties of our model. Through it, we assume that all relevant time-varying factors are observed and the remaining fixed unobserved factors affect the levels of other variables and not their changes. An example of such factors is "personality traits"; see Diener and Lucas (1999) and Argyle (1999).

Our dependent variable *JS– job satisfaction --* $\in \{1,...,6\}$ is an ordinal indicator of the individual's overall satisfaction in his/her main activity. Since the data available are longitudinal, we dispose of this measure for a number of individuals i=*1,..,N* over a given time-period indexed by $t = 1...T$. More precisely, we observe a sample of Danish workers over the years 1995-1999. In addition to their self-reported job satisfaction levels, the data set includes many individual and job-related characteristics for each survey year, some of which will be used as explanatory variables in our analysis.

3.1 Random-Effects Ordered Probit

Our reference model is the ordered probit model with individual random effects:

$$
\begin{aligned}
&JS^{j^*}_{i,t} = x^{'}_{i,t}\beta + v_i + \varepsilon_{it} \\
&JS^{j}_{i,t} = k \Leftrightarrow JS^{j^*}_{i,t} \in [\lambda_k, \lambda_{k+1})
\end{aligned}
\qquad (1)
$$

where JS^*_{it} is latent overall satisfaction in job while JS_{it} is the observed satisfaction level declared at the survey's date; λ_k is the *k*-th cut-off point (increasing in *k*) for the categories; x_{it} are observable individual characteristics; v_i is an individual random characteristic, normally-distributed, fixed over time and orthogonal to *x* with unknown variance; and finally ε_{it} is a time-varying error-term, normally-distributed, orthogonal to all x. As in the binary choice model, the underlying variance $\sigma^2 = \sigma_v^2 + \sigma_\varepsilon^2$ is not identified. We adopt the normalization $\sigma_\varepsilon^2 = 1$. With it $\rho = \frac{\sigma_v^2}{\sigma_v^2 + \sigma_\varepsilon^2} = \frac{\sigma_v^2}{1 + \sigma_v^2}$, is estimated directly, and from it we can recover $\sigma_v = \sqrt{\frac{\rho}{1-\rho}}$

To account for the presence of the random effect, the CDF is in this case computed from:

$$
\begin{aligned}
\Pr(y_{it}^* \le k \mid x_{it}) &= \Pr(x'_{it}\beta + \varepsilon_{it} + v_i \le \lambda_k) \\
&= \Pr(\varepsilon_{it} + v_i \le \lambda_k - x'_{it}\beta) \\
&= \Pr\left(\frac{\varepsilon_{it} + v_i}{\sqrt{1+\sigma_v^2}} \le \frac{\lambda_k - x'_{it}\beta}{\sqrt{1+\sigma_v^2}} \right) \\
&= \Pr\left(\eta_{it} \le k^* - x'_{it}\beta^* \right)
\end{aligned}
$$

$$\textit{with}\ \eta_{it} \sim N(0,1)$$

The assumption on the normality of the error terms yields an ordered probit model. The model is again built around a latent regression model with the λ being the unknown parameters that are estimated along with the β . Individual heterogeneity is unobserved; therefore to obtain the unconditional log-likelihood we need to integrate the conditional log-likelihood. The integration is done with the Gauss-Hermite quadrature (25 points were chosen); see Frechette, (2001a), (2001b), Butler and Moffit (1982), and Greene (2003).

3.2 Fixed-effects ordered logit

3.2.1 The Ferrer-i-Carbonel and Frijters (2004) estimator

Some unobserved individual characteristics may affect a particular job dimension. In that case a spurious correlation between that dimension and those unobserved characteristics may arise and thereby bias the estimated coefficients. While the random effects ordered probit can to a certain extent indicate the direction of the effects of some determinants of job satisfaction, the above-mentioned spurious correlation is most likely to be present. In that case a fixed effects approach seems to be more appropriate.

The estimator proposed by Ferrer-i-Carbonel and Frijters (2004) derives from an extension of the idea of Chamberlain (1980) to a fixed-effect ordered logit framework. The model is:

$$\begin{aligned} JS^{j^*}_{it} &= x_{it}\beta + f_i + \varepsilon_{it} \\ JS^{j}_{it} &= k \Leftrightarrow JS^{j^*}_{it} \in [\lambda_k, \lambda_{k+1}) \end{aligned} \qquad (2)$$

where again JS^*_{it} is latent overall job satisfaction; JS_{it} is the observed satisfaction level; f_i is an individual fixed effect; ε_{it} is the error term logistically distributed and orthogonal to all *x*. This model is an ordered logit model with fixed individual effects and individual specific thresholds λ^i_k.

The model assumes that the intercepts are increasing i.e. $\lambda^i_k < \lambda^i_{k+1}$. However, it does *not* assume ordinal comparability.

The statistic of interest is:

$$\Pr\left[I(\mathbf{JS}_{i1} > k_i),....,I(\mathbf{JS}_{iT} > k_i) \mid \sum_t I(JS_{it} > k_i) = c_i \right]$$

$$= \frac{\dfrac{\prod_{t=1}^{T}\left\{1 + I(\mathbf{JS}_{it} > k_i)(e^{-\lambda^i_{k_i} + (x_{it}\beta + f_i)} - 1)\right\}}{\prod_{t=1}^{T}\left\{1 + e^{-\lambda^i_{k_i} + (x_{it}\beta + f_i)}\right\}}}{\sum_{JS \in S(k_i,c)} \dfrac{\prod_{t=1}^{T}\left\{1 + I(\mathbf{JS}^j_{it} > k_i)(e^{-\lambda^i_{k_i} + (x_{it}\beta + f_i)} - 1)\right\}}{\prod_{t=1}^{T}\left\{1 + e^{-\lambda^i_{k_i} + (x_{it}\beta + f_i)}\right\}}}$$

$$= \frac{e^{\sum_{t=1}^{T} I(\mathbf{JS}^j_{it} > k_i) x_{it}\beta}}{\sum_{JS \in S(k_i,c)} e^{\sum_{t=1}^{T} I(\mathbf{JS}^j_{it} > k_i) x_{it}\beta}} \qquad (3)$$

This statistic implies that all the individuals whose satisfaction scores vary over time are included in the estimation procedure (see Frijters and al. 2004). It is important to note that the last expression in (3) is the likelihood of observing the job satisfaction levels that are above the cut-off point, given that there are *c* satisfaction levels that are higher than *k* for each individual *i*. Thus, $S(k_i, c)$ represents the set of all possible combinations of job satisfaction (in each of the j dimensions considered) that satisfy:

$$\sum_t I(JS_{it} > k_i) = c\,.$$

Evidently, one advantage of this estimation method is that it avoids loosing a huge amount of information: any individual whose job satisfaction level changes can indeed be used. The model is estimated by maximum likelihood.[9]

3.2.2 The Das and Van Soest (1999) estimator

Das and Van Soest (1999) have developed another method that exploits the Chamberlain estimator to build ordered logit models with fixed effects. Their estimator is based on a weighted average of the Chamberlain estimator for each *k*. Hence, in their method, the authors obtain an estimate of *k* based on those individuals for which $T > \sum_{t=1}^{T} I(JS_{it} > k) > 0$ for each *0<k<K*. The clear advantage of this estimator is that it accounts for all possible individuals' *k*'s and hence uses more information. Its disadvantage is that there may not be enough data in each category k in order to estimate β_k. This implies that when there is not enough variation over the categories, those thresholds cannot be used, and the corresponding categories should be dropped. This happens in our estimation for the low satisfaction values reported by men (only values higher than 2 could be used for them). Moreover, it should be noted that this estimator requires stricter regularity conditions than the one examined in the previous section.

The Das and van Soest estimator is based on the following procedure. First one transforms each individual satisfaction's vector $\{JS_{i1},....,JS_{iT}\}$' into a set of *K* vectors, such that $\{(JS_{i1} > k);\ :::;\ (JS_{iT} > k)\}'$ for *k=0* to *K-1* with *K* being the number of categories that the ordinal variable may take. For each *k*, one estimates the parameters of interest applying the Chamberlain estimator to data for individuals for whom $T > \sum_{t=1}^{T} I(JS_{it} > k) > 0$. This allows one to obtain a consistent estimator implying $\sqrt{n_k}(\beta_k - \beta) \to N(0, \sum_{kk}^{-1})$ k=0,..,K-1. The final Das and Vas Soest estimator is obtained through a minimum distance step

$\hat{\beta} = \underset{\beta}{\arg\min} \frac{1}{2} \left[\begin{pmatrix} \beta_0 \\ .. \\ \beta_K \end{pmatrix} - \begin{pmatrix} \beta \\ .. \\ \beta \end{pmatrix} \right]' \Omega^{-1} \left[\begin{pmatrix} \beta_0 \\ .. \\ \beta_K \end{pmatrix} - \begin{pmatrix} \beta \\ .. \\ \beta \end{pmatrix} \right]$ with Ω being the weighting matrix with entries $\omega_{a,b} = \Sigma_{aa}^{-1}\Sigma_{ab}^{-1}\Sigma_{bb}^{-1}$ with *a,b=0,…,K-1*. In order to make the estimator operational, the unknown matrices are replaced with their sample analogues. Since, asymptotically, $\Sigma_{kk} \to E\left\{ \frac{\partial \ln L}{\partial \beta_k} \frac{\partial \ln L}{\partial \beta_k}' \right\}$ it is also possible to improve the estimator by using the sample Hessian instead of $\frac{1}{n_a}\sum_i E\left\{ \frac{\partial \ln L}{\partial \beta_k} \frac{\partial \ln L}{\partial \beta_k}' \right\}$(see Ferrer-i-Carbonel and Frijters, 2004).

3.3 Specification Testing: Random or Fixed-Effects

To assess the comparative advantage of estimating a fixed-effects ordered logit model (based on the first estimator) compared to a random effects ordered probit, we implement the test presented in Frijters *et al.* (2004). In doing so, we use the variables that at the same time are present in both the fixed effects and the random effects models. We define $\tilde{\beta}_{RE}$ to be the coefficients of the variables that are present in both models but resulting from the estimation of the random effects ordered probit. In the absence of effects related to fixed individual characteristics, we expect that the coefficients should be very similar. Under the null-hypothesis that there are no FE, therefore

H0: $$\beta^{FE} = \alpha \tilde{\beta}_{RE} \tag{4}$$

where α is an unknown positive constant originating from the different normalizations assumed in the estimation of the FE and RE models[10]. Notice that the $\tilde{\beta}_{RE}$ only contains the coefficients of those

[9] Regarding the properties of the estimator, the reader is referred to the article of Ferrer-i-Carbonel and Frijters (2004).
[10] This vector of coefficients is obtained through the estimation of the random effects ordered probit on the whole sample and has var(ε_{it})=1. Conversely, when using the fixed-effects estimator only a sub-sample of individuals is used. Thereby, these two models do not share the same normalization. See Frijters et al. (2004)

variables that are present at the same time in both the fixed effects and the random effects models. To simplify the exposition, we write $\gamma_{RE} = \alpha \tilde{\beta}_{RE}$

Under the null hypothesis, we can use the following standard likelihood ratio test:

$$-2*\left[L(\tilde{\beta}_{ML}{}^{FE}) - L(\tilde{\gamma}_{ML}{}^{RE}) \right] \sim \chi(k) \qquad (5)$$

where $(\tilde{\beta}_{ML}^{FE})$ is the coefficients vector obtained from the maximum likelihood estimation of the unrestricted fixed-effects ordered logit and $L(\tilde{\beta}_{ML}^{FE})$ the value of the corresponding (log) likelihood; k denotes the number of restricted parameters; $L(\tilde{\gamma}_{ML}^{RE})$ is the likelihood of the fixed-effects model when the value of the parameters are equal to $\alpha \tilde{\beta}_{RE}$. As has been pointed out by Frijters *et al.* (2004), there are at least two problems that make this testing procedure less easy than it seems at first glance. First, one needs to re-estimate the model to re-fit the unrestricted parameters of the model. Second, and not less importantly, α is unknown. A way to avoid the last problem is noticing that

$$2L(\tilde{\beta}_{ML}^{FE})\text{-}2L(\alpha \tilde{\beta}^{RE}) > 2L(\tilde{\beta}_{ML}^{FE})\text{-}\max_{\hat{\alpha}} \left\{ 2L(\hat{\alpha} \tilde{\beta}^{RE}) \right\} \qquad (6)$$

The inequality (6) implies that a lower bound for $2L(\tilde{\beta}_{ML}^{FE})\text{-}2L(\alpha \tilde{\beta}^{RE})$ is attained by using the $\hat{\alpha}$'s, which maximize $L(\hat{\alpha} \tilde{\beta}^{RE})$. Consequently, rejecting the null at the lower bound implies that the true statistic will also reject it.

4 Data description

The data used in this paper are extracted from the European Community Household Panel. The data are collected annually on several issues regarding family structure, family and family members' incomes and employment. Thanks to the panel character of the data, they provide unique information about the dynamics of social change and individual behaviour. The data used in the following empirical analysis are taken from the waves 1995 to 1999. Concerning non-response and attrition the reader is referred to the paper by Nicoletti and Peracchi (2002). In general the non-response rates in the satisfaction question are found to be very low.

The job satisfaction questions in the ECHP ask the individuals to give an integer response on a scale from 1 to 6 which best describes how satisfied or dissatisfied they are with specific job facets: wages, job security, type of work, working hours, working time, working conditions and commuting time. In addition they are asked about the overall ("all things considered") satisfaction in their main job or activity. The reported "overall" job satisfaction is likely not to be merely the average of the already reported satisfaction levels for the different aspects of jobs, but may be capturing some additional aspects of the jobs held or reflecting the differences in the weights each employee attaches to the individual job facets[11]. By selecting only people employed in the survey year, we can see their reported satisfaction level as an overall evaluation of the satisfaction in multiple job spheres. We will focus on that in our analysis. A response of 1 represents the lowest level of satisfaction and 6 the highest.

Table 2 and *Chart 3* show the means for our dependent variable for men and women, separately. We may note that the means are remarkably constant during the five year-period under study. Furthermore, the scores are high: close to or slightly below 5.

The high persistency in mean job satisfaction levels masks the fact that there are quite frequent year to year changes in individuals' reported levels of job satisfaction; see *Table 3*. In each year, only about 38 (35) per cent of the male (female) respondents report the same levels as in the preceding year. Though the period considered– 5 years – is too short to shed light on whether changes are temporary (blips or dips), or of a more permanent character, this result is especially noteworthy in view of the fact that our subsequent econometric analysis exploits the longitudinal aspects of the data. Although there is no trend in the average satisfaction levels, there is for each year a larger proportion of individuals whose satisfaction level is higher than in the previous year than there are employees whose satisfaction level has decreased compared to the year before. Finally, we can se the patterns are very similar for the male and female employees.

[11] We have carried out some simple cross-tabulations of «overall satisfaction" and the seven different facets of jobs. These show indeed that they are positively correlated, but the correlations are far from perfect.

A detailed description of variables have been used in the empirical analysis can be found in Appendix A.1. These include socio-demographic characteristics (like age, education, marital/cohabiting status, newborn child, health status), job characteristics (sector and kind of occupation, working in the public sector, holding a temporary contract, working part-time, the number of working hours, the (log) gross monthly wage, the fact of having received training, the job status, the amount of experience), and two indicators for changes in the macroeconomic environment (the unemployment rate and growth rate of real GDP, respectively). The trend of the two structural indicators over the observation period along with average satisfaction levels for men and women respectively are reported in Chart 4.

Only intrinsically[12] time-varying variables have been introduced in the fixed effects ordered logit. More specifically they are: the respondents' age, the health indicators, holding a temporary job, having a "newborn child", working part-time, working in the public sector, having been trained, the current gross (log) wage, the unemployment and the real GDP growth rate.

Before plunging in on the empirical estimates we refer to *Tables 4* and *5* which give some descriptive information about the data used. The period under study is one of relatively high and sustained economic growth and hence also one of declining overall unemployment; see *Table 6*. Due to the high labour force participation rates of Danish women, the age structures of the male and female workforces are quite similar. From the tables we can see that women are more educated, more likely to work in the public sector and in the service industries and to work part-time.

To account for potential differences across genders all the estimations have been carried out for males and females, separately. The two samples include 3,936 women and 4,227 men, respectively.

5 Results

Turning now to the estimates, which are set out in *Tables 7* and *8* for males and females, respectively, we may first note that the test of random effects versus fixed effects described in section 4, decisively rejects the former. The random effects ordered probit estimates are in *Tables A2.1 in* Appendix A.2. A comparison of these with the preferred fixed effects model estimates reveals some interesting patterns.

The first thing worth noting is that the key economic explanatory variables like income from work, training and temporary jobs attach quite similar coefficient estimates. Thus, previous job satisfaction models have not been far from the mark in this respect. It is worth remarking, however, that the coefficient to wage income for women is positive albeit insignificant in all specifications. As the data on working hours are crude, making a distinction between full- and part-time work only, the insignificant signs to this dummy variable should not worry us much.[13] Other similarities are found for those explanatory variables the estimated coefficients of which are insignificant; that is, when a variable does not differ from zero in the random effects model, it does not in the fixed effects models, either. There is one exception, however. The age of the employee is consistently negatively related to job satisfaction in models allowing for the baseline satisfaction to differ between individuals. The negative age effect has also found in a number of previous studies; see Table 1.

A second observation that shows up in a comparison is that macroeconomic variables obtain quite different estimates in the different specifications. For both genders the unemployment rate attaches a positively signed and significant coefficient in the random effects model, whereas it changes sign while retaining its significance in the fixed effects models. The random effects model moreover yields a positive and significant effect of the growth rate on men's as well as women's overall job satisfaction. Reconciling the two positive effects is hard, indeed. In the fixed effects models only the unemployment rate survives and is negative, which is consistent with less need of employers to invest in various facets of job quality when jobs are scarce.

The third conclusion that can be extracted from the fixed effects estimates is that female public sector employees are more satisfied with their jobs than their colleagues in the private sector. This is not completely unexpected as there is a negative wage premium for Danish public sector employees (Pedersen et al. (1990)) but at the same time more working time flexibility and less pressure on doing overtime work in the public sector.[14]

One interesting explanatory variable, which due to its time-invariant character does not enter the fixed effects-estimations, is education. Some previous studies (e.g., Clark and Oswald, 1995) have found that employees with higher education are less satisfied with their jobs, a finding that is not, however, replicated by Kaiser (2002) in his five-country study. We find opposite effects for men and women –

[12] The information is retrieved from questions concerning "the last twelve months".
[13] Although insignificant, they change sign for both genders, as fixed effects are enter into the model.
[14] In corresponding estimations for six different facets of job satisfaction for males and females separately (but not reported here), we find that public sector employees are more satisfied with their working times and working hours but less satisfied with their earnings than private sector employees.

more educated male (female) employees are less (more) satisfied – but for both, the effect does not differ significantly from zero.

The coefficients estimates obtained using on one hand the Ferrer-i-Carbonel and Frijters (2004) and on the other hand the Das and Van Soest (1999) estimation strategy are very similar. The standard errors are in general smaller for the latter. Being employed on a temporary contract and a higher number of nights spent in hospital both obtained negative and significant coefficients for male employees. The first observation indicates that temporary, fixed-term contract jobs are considered as bad. The second variable is a proxy for health status which is plausibly negatively related to job satisfaction as individuals in a good physical and psychic condition are likely to be able to earn more, to feel relatively more certain of their continued employment, to be more able to choose and carry out the type of work they like, and to have less difficulties with the number of working hours, placement of working hours or with working conditions.

6 Concluding remarks

In recent years data on employees' satisfaction with their jobs, and various aspects of these, have become increasingly available to researchers. This information is typically of ordered character and some of the more frequently used data sets are panels. Given the latter, relatively few analyses have actually made use of the longitudinal character of the data. This is in particular surprising as not only the levels of, but also the changes in job satisfaction, and factors underlying these, are potentially very interesting. Prior to this paper another weakness of the literature has been that the possibility that individuals differ with respect to their baseline satisfaction levels, or in the jargon of panel data econometrics, individual fixed effects are not allowed for in the estimations. The main novel feature of this paper is that we apply new statistical methods for estimating an ordered logit model with fixed effects to panel data on job satisfaction.

Entering individual fixed effects do indeed make a difference: both the estimated coefficients of time-varying explanatory variables, and their significance, change as we allow for individuals to have different baseline satisfaction levels. Moreover, we find that poor macroeconomic conditions, as measured by the unemployment rate, have a negative effect on employees' satisfaction. The main differences between the genders are found with respect to the influence of the individual's own wage, holding a temporary job and working in the public sector. This suggests that different factors are important determinants of men's and women's reported satisfaction. In particular it seems that while, especially for women, monetary factors matter less; working conditions and (at least to a certain extent) a higher degree of flexibility carry a larger weight in the job preferences of female employees.

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Table 1. Summary of previous results

Explanatory Variable:	Clark (1996)	Clark (1997) Males	Clark (1997) Females	Clark & Oswald (1996)	Clark, Oswald & Warr (1996)	Blanchflower & Oswald (1998)	Lydon & Chevalier (2001)	Gardner & Oswald (2001)	Employment in Europe (2002)	Kaiser (2002)
Pay				+	+		+	+	+	
Hours	-	0	-	-	-	-	-	-		
Part-time						+			+	0
Male	-			-	-	-	-	-	-	0
Age	-	U	U		U		-	U	-	U
Married			+			+			+	+
# children							+			-
Ethnic								-		0
Health		+	+							+
Education, Skilled	-	-	-	-			0	-	-	inv U
Job tenure								U	U	
Temporary contract				-				-	-	-
Manager/super-Visor		0	+	+			+		+	+
Public sector							-	+	+	+
Union membership	-	0	-			-		-		
Size of firm, establishment	-		-					-	-	
Data source:	BHPS, 1991	BHPS, 1991	BHPS, 1991	BHPS, 1991	BHPS, 1991	NCDS, 1981 and 1991	HFECES, 1985 and 1990	BHPS, 1991-99	ECHP, 1995-98	ECHP, 1995-97 5 countries; pooled
Estimation method:		Ordered probit	Ordered probit	Ordered probit		Ordered probit	Ordered probit	Ordered probit	Ordered Probit	Probit

Table 2. Mean job satisfaction by year and facet of job, males

	Avg. job satisfaction	
year	**Men**	**Women**
95	4.990	4.999
96	4.927	4.905
97	4.935	4.993
98	4.967	4.939
99	4.865	4.853
Total	4.937	4.938

Table 3. Changes in job satisfaction levels compared to previous year, (shares in %)

	MEN			**WOMEN**		
year	*Lower*	*Higher*	*Same*	*Lower*	*Higher*	*Same*
96	0.273	0.281	0.446	0.269	0.305	0.426
97	0.234	0.346	0.420	0.216	0.377	0.408
98	0.241	0.268	0.491	0.295	0.276	0.429
99	0.286	0.163	0.551	0.290	0.201	0.509
Total	0.207	0.412	0.381	0.214	0.432	0.354

Table 4: Descriptive statistics: Averages for period 1995-99 -MEN

	Mean	Std.Dev	Skewn.	Kurt.	Min	Max
Log of monthly wage	7.969	0.358	-0.377	8.051	5.178	9.501
Age	40.960	10.329	0.076	2.101	17	64
Part-time	0.017	0.129	7.519	57.539	0	1
Nights in hospital	0.347	3.140	20.251	596.053	0	120
Training	0.583	0.493	-0.337	1.114	0	1
Newborn child	0.059	0.236	3.746	15.035	0	1
Temporary job	0.101	0.301	2.656	8.056	0	1
Working in the public sector	0.286	0.452	0.947	1.896	0	1
Experience	22.880	11.215	0.208	2.161	0	50
Main job working hours	39.590	6.731	2.073	13.047	10	96
Sector: Agriculture	0.031	0.174	5.390	30.048	0	1
Sector: Manufacturing	0.371	0.483	0.533	1.284	0	1
Sector: Services	0.598	0.490	-0.398	1.158	0	1
Occupation: Managers	0.088	0.283	2.919	9.518	0	1
Occupation: Professional	0.201	0.401	1.490	3.218	0	1
Occupation: Technicians and associate professionals	0.169	0.375	1.763	4.107	0	1
Occupation: Clerical occupations (reference group)	0.069	0.253	3.405	12.597	0	1
Occupation: Service workers and shop and market sales workers	0.049	0.217	4.156	18.273	0	1
Occupation: Skilled agricultural and fishery workers	0.017	0.129	7.464	56.712	0	1

Occupation: Craft and related trades workers	0.190	0.392	1.584	3.510	0	1
Occupation: Plant and machine operators and assemblers	0.132	0.339	2.174	5.726	0	1
Occupation: Elementary occupations	0.085	0.279	2.977	9.865	0	1
Higher education	0.356	0.479	0.601	1.361	0	1
Secondary education	0.466	0.499	0.136	1.018	0	1
Primary education	0.178	0.383	1.684	3.837	0	1
Job status: Supervisory	0.205	0.404	1.462	3.138	0	1
Job status: Intermediate	0.150	0.357	1.966	4.863	0	1
Job status: Non supervisory	0.646	0.478	-0.609	1.370	0	1
Married or cohabiting	0.803	0.398	-1.523	3.319	0	1
Experience squared	649.257	555.522	0.950	3.066	0	2500

Table 5: Descriptive statistics - Women

	Mean	Std.Dev.	Skewness	Kurtosis	Min	Max
Log of monthly wage	7.698	0.354	-2.013	19.207	3.779	8.872
Age	41.014	10.013	-0.009	2.183	17	64
Part-time	0.148	0.355	1.984	4.935	0	1
Nights in hospital	0.563	4.146	16.017	340.828	0	120
Training	0.643	0.479	-0.598	1.358	0	1
Newborn child	0.052	0.223	4.020	17.158	0	1
Temporary job	0.086	0.280	2.956	9.736	0	1
Working in the public sector	0.580	0.494	-0.322	1.104	0	1
Experience	22.508	11.009	0.164	2.173	0	50
Main job working hours	34.653	6.445	-0.623	5.196	5	65
Sector: Agriculture	0.007	0.081	12.180	149.353	0	1
Sector: Manufacturing	0.131	0.337	2.192	5.806	0	1
Sector: Services	0.863	0.344	-2.109	5.447	0	1
Occupation: Managers	0.031	0.173	5.436	30.553	0	1
Occupation: Professional	0.166	0.372	1.799	4.234	0	1
Occupation: Technicians and associate professionals	0.263	0.440	1.077	2.159	0	1
Occupation: Clerks (reference group)	0.219	0.414	1.357	2.841	0	1
Occupation: Service workers and shop and market sales workers	0.201	0.401	1.492	3.227	0	1
Occupation: Skilled agricultural and fishery workers	0.004	0.060	16.676	279.076	0	1
Occupation: Craft and related trades workers	0.011	0.105	9.298	87.444	0	1
Occupation: Plant and machine operators and assemblers	0.035	0.184	5.055	26.551	0	1
Occupation: Elementary occupations	0.071	0.256	3.351	12.231	0	1
Higher education	0.397	0.489	0.422	1.178	0	1
Secondary education	0.440	0.496	0.244	1.059	0	1
Primary education	0.164	0.370	1.818	4.306	0	1
Job status: Supervisory	0.092	0.289	2.818	8.942	0	1
Job status: Intermediate	0.156	0.363	1.896	4.594	0	1
Job status: Non supervisory	0.752	0.432	-1.166	2.358	0	1
Married or cohabiting	0.814	0.390	-1.610	3.591	0	1
Experience squared	627.791	531.561	0.938	3.131	0	2500

Table 4: Unemployment and growth rates over the observation period

year	**unemployment rate**	**growth**
95	6.7	2.8
96	6.3	2.5
97	5.2	3
98	4.9	2.5
99	4.8	2.6
Total	5.58	2.68

Table 7: Fixed effects ordered logit estimates -- Men

	Ferrer-i-Carbonel and Frijters (2004)		**Das and Van Soest (1999)**	
	Coeff.	Std.Err.	Coeff.	Std.Err.
Age	-0.6310**	(0.1198)	-0.5896**	(0.1058)
Unemployment rate	-0.9187**	(0.206)	-0.7979**	(0.1808)
Growth rate	-0.1141	(0.2515)	-0.1946	(0.2147)
Newborn child	-0.1656	(0.1708)	-0.1132	(0.1399)
Temporary contract	-0.2490	(0.1893)	-0.3451**	(0.1468)
Nights spent in hospital	-0.0338	(0.0195)	-0.04145**	(0.0185)
Training	0.2546**	(0.1009)	0.1976**	(0.0823)
Part-time	-0.6603	(0.5521)	-0.1619	(0.4886)
Public sector	0.0192	(0.2876)	-0.1762	(0.2067)
Log current gross wage	0.7163**	(0.3294)	1.0021**	(0.2739)
α	-0.1271**	(0.0264)		
Log-likelihood ratio test	-39.3129			

Table 8: Fixed effects ordered logit estimates – Women

	Ferrer-i-Carbonel & Frijters (2004)		**Das and Van Soest (1999)**	
	Coeff.	Std.Err.	Coeff.	Std.Err.
Age	-0.4432**	(.1155)	-0.4208**	(.1014)
Unemployment rate	-0.6066**	(.2007)	-0.5688**	(.1748)
Growth rate	0.1144	(.2551)	0.3399	(.2163)
Newborn child	0.2546	(.1758)	0.2153	(.1419)
Temporary contract	0.2422	(.1899)	0.2264	(.1572)
Nights spent in hospital	-0.0050	(.0100)	-0.0128	(.0107)
Training	0.1387	(.1045)	0.1883*	(.0854)
Part-time	-0.0419	(.2044)	-0.0789	(.1569)
Public sector	0.6703*	(.3133)	0.4354*	(.2147)
Log current gross wage	0.2070	(.3222)	0.2128	(.2507)
α	-0.1143**	(0.0234)		
Log-likelihood ratio test	-31.5141			

Charts:

Chart 1:

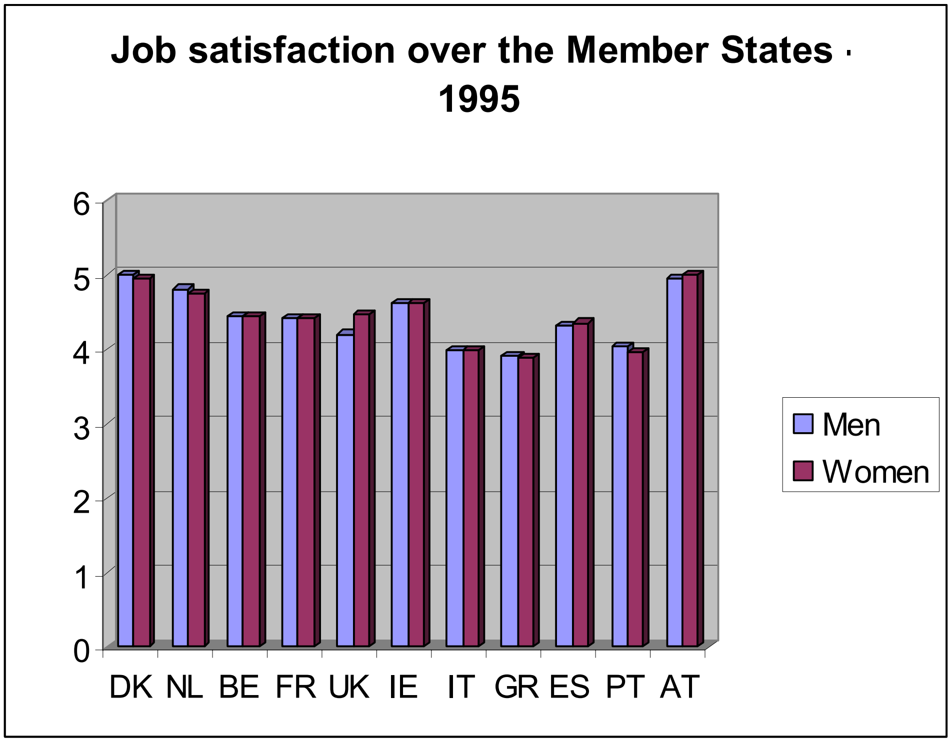


Chart 2:

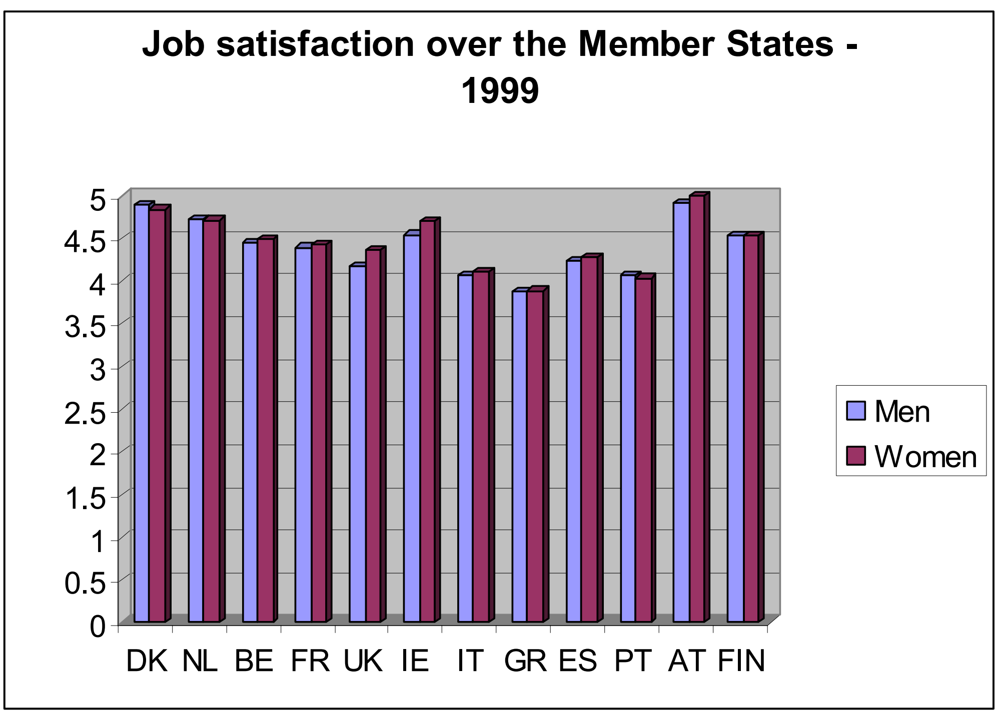


Chart 3: Job satisfaction in Denmark over the observation period

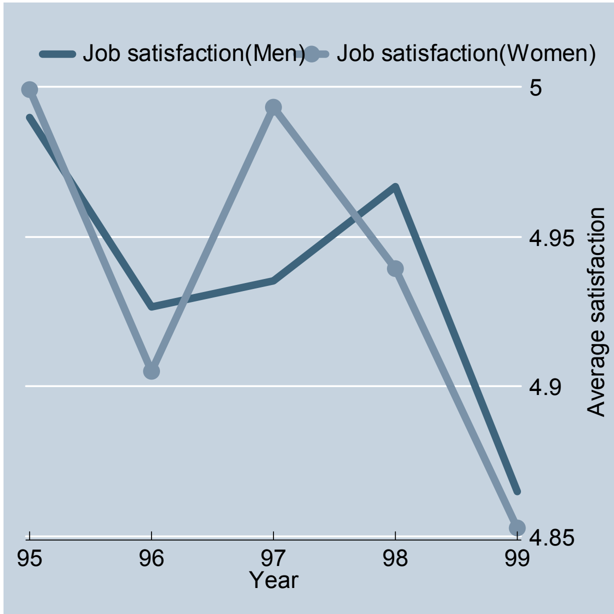
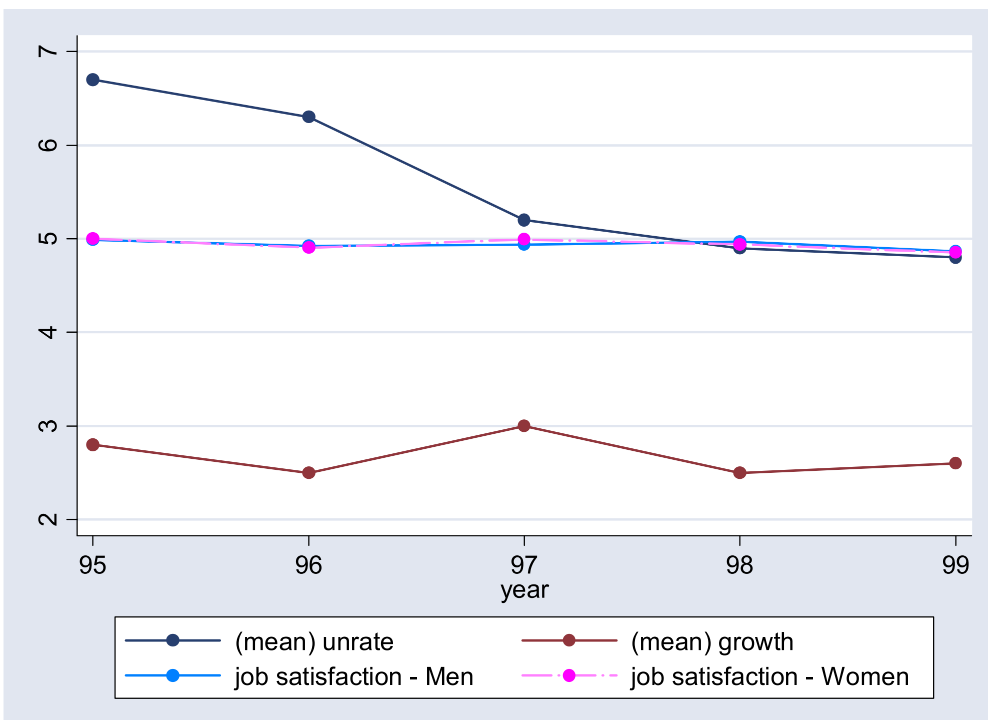


Chart 4. Average satisfaction, unemployment rates and growth rates 1995-1999



Appendix A.1: Data description

The following variables have been used in the empirical analyses:

1) <u>Macroeconomic indicators</u>
    a. Unemployment Rate
    b. Growth rate
2) <u>Job-related factors</u>
    a. Log of current monthly gross wage
    b. Part-time in the last year (1: Yes)
    c. Training in the last year (1: Yes)
    d. Temporary job (1: Yes)
    e. Working in the public sector (1: Working in the public sector)
    f. Main job working hours
    g. Sector: Agriculture
    h. Sector: Manufacturing (reference)
    i. Sector: Services
    j. Occupation: Legislators
    k. Occupation: Professional
    l. Occupation: Technicians and associate professionals
    m. Occupation: Clerks (reference group)
    n. Occupation: Service workers and shop and market sales workers
    o. Occupation: Skilled agricultural and fishery workers
    p. Occupation: Craft and related trades workers
    q. Occupation: Plant and machine operators and assemblers
    r. Occupation: Elementary occupations
    s. Job status: Supervisory
    t. Job status: Intermediate
    u. Job status: Non supervisory (reference)
3) <u>Socio-demographic factors</u>
    a. Age
    b. Higher education
    c. Secondary education
    d. Primary education (reference)
    e. Experience squared
    f. Experience: defined as the age when entering the labour market and current age
    g. Newborn child (1: Having a child in the last year)
    h. Married or cohabiting (1: Married)
    i. Nights in hospital in the last year

Appendix A-2

Table A2.1 Random effects ordered probit estimates

	MEN		WOMEN	
	Coeff.	Std.Err.	Coeff.	Std.Err.
Unemployment Rate	0.0844**	(.02393)	0.0953**	(.0262)
Growth rate	0.2423**	(.1347)	0.3509**	(.1192)
Log of current monthly wage	0.2937**	(.1063)	0.0357	(.1131)
Age	-0.0111	(.0117)	0.0113	(.0120)
Part-time	0.1496	(.1916)	0.0398	(.1023)
Nights in hospital	-0.0223**	(.0065)	0.0007	(.0039)
Training	0.1265**	(.0451)	0.0996**	(.0489)
Newborn children	-0.0435	(.0840)	0.0779	(.0937)
Temporary jobs	-0.1957**	(.0693)	0.1030	(.0820)
Public sector	0.0736	(.0718)	0.0223	(.0700)
Experience	-0.0404**	(.0148)	-0.0225	(.0166)
Working hours	0.0022	(.0041)	-0.0064	(.0064)
Sector: Agriculture	0.3223	(.2184)	0.3394	(.4340)
Sector: Services	-0.1316	(.0738)	-0.0273	(.1097)
Legislators	0.4058**	(.1549)	0.2495	(.2167)
Professionals	0.2753*	(.1319)	0.1387	(.1124)
Technicians	0.2624*	(.1263)	0.0404	(.0903)
Service workers	0.3431*	(.1425)	0.1464	(.0974)
Skilled agricultural	0.6575*	(.3069)	-0.4639	(.3177)
Craft and trade workers	0.1063	(.1359)	-0.1977	(.3214)
Assemblers	0.1309	(.1402)	0.0571	(.1782)
Elementary occupations	0.1978	(.1279)	-0.3025**	(.1329)
Higher education	-0.0240	(.0981)	0.0129	(.0978)
Secondary education	0.0079	(.0768)	0.0131	(.0871)
Supervisory job status	-0.0835	(.0723)	0.3683**	(.0990)
Intermediate job status	-0.1654**	(.0672)	-0.0427	(.0715)
Married or cohabiting	0.0907	(.0656)	0.1204	(.0699)
Experience squared	0.0012**	(.0002)	0.0005*	(.0003)
λ(01)	0.0364	(.8976	0.3586	(.9358)
λ (02)	0.5749**	(.0667)	0.7604**	(.0860)
λ (03)	1.2722**	(.0756)	1.5143**	(.0893)
λ (04)	2.3255**	(.0784)	2.5529**	(.0909)
λ(05)	3.9953**	(.0800)	4.0838**	(.0937)
rho	0.8976**	(.0340)	0.8948**	(.0356)

Standard errors in brackets
* significant at 5%; ** significant at 1%