Discussion Papers No. 406, Feburary 2004
Statistics Norway, Research Department

Zhiyang Jia

Spousal Influence on Early Retirement Behavior

Abstract:
In this paper, we use a binary choice panel data model to analyze married individuals.retirement behavior in Norway when a new option, AFP early retirement becomes available. We focus our study on the influence of the spouse.s characteristics on early retirement behavior. We find the directions of spousal e¤ects are quite symmetric but women seem to have a much stronger response to their spouses' characteristics than men. The comparison of di¤erent specifications indicates that correct modeling of the error term covariance structure in a panel data binary choice model is quite important.

Keywords: Retirement, Spousal Influence, Panel Data, Random Effects.

JEL classification: H55, J26

Acknowledgement: I thank John Dagsvik and Steinar Strøm for helpful comments, Fedor Iskhakov for preparing part of data used in this analysis.Financial support from the Research Council of Norway (project 140082/530) is gratefully acknowledged.

Address: Zhiyang Jia, Statistics Norway, Research Department. E-mail: Zhiyang.jia@ssb.no

Discussion Papers comprise research papers intended for international journals or books. A preprint of a Discussion Paper may be longer and more elaborate than a standard journal article, as it may include intermediate calculations and background material etc.

Abstracts with downloadable Discussion Papers
in PDF are available on the Internet:
http://www.ssb.no
http://ideas.repec.org/s/ssb/dispap.html

For printed Discussion Papers contact:

Statistics Norway
Sales- and subscription service
NO-2225 Kongsvinger

Telephone: +47 62 88 55 00
Telefax: +47 62 88 55 95
E-mail: Salg-abonnement@ssb.no

1 Introduction:

In this paper, we apply a binary choice panel data model to study the early retirement behavior using Norwegian data. The analysis is based on the married individuals who are qualified for a subsidized early retirement scheme hereafter called AFP (a Norwegian abbreviation). We focus on the understanding of how the early retirement decisions are affected by the spousal characteristics such as labor market status, income and wealth etc. Different specifications of the model are discussed and estimated.

Recent studies of retirement behavior have recognized the phenomenon that husbands and wives often coordinate their labor supply at older ages. Among them, Blau (1997) finds "strong associations between the labor force transition probabilities of one spouse and the labor force status of the other spouse." Using data from US Health and Retirement Study (HRS), Johnson and Favreault (2001) find that the employment and health status of the spouse appear to have important effects on retirement decisions for married women and men. Similar patterns have also been found on European data, see for example, Blau and Rihahn (1999) and Jimenez-Martin, Labeeaga, and Granado (1999).

This phenomenon, in turn, has generated lots of literature on possible sources of this coordination in elderly labor supply. However, few discusses another important aspect of this issue, namely whether the husbands' characteristics affect the behavior of wives differently from the way wives' characteristics affect their husbands' behavior, and which one is stronger. Among the limited literature, there are mixed conclusions. Schellenberg (1994) conducts a survey and finds that men are far less influenced by their spouses' situation than women in Canada. On the other hand, Gustman and Steinmeier (2000) reach just the opposite conclusion based on some old US National Longitudinal Survey data from 1970s and 1980s. They state that 'There is some suggestion in the data that the wife's retirement decision is not strongly influenced by the husband's, but the husband's decision is more strongly influenced by the wife's'. This is confirmed by Coile (2003)'s study on much recent data from HRS. She finds that the response of women to their own incentive measures is virtually identical to the response of men. Spill-over effects from the wife are important determinants of husband's retirement, while the spill-over effects from the husband are small and statistically insignificant.

One of the most important reasons for the limited literature on comparing spousal effects might be the lack of proper data due to the low participation rate of the elderly women. However, the labor force participation for elderly women has increased dramatically in most western countries during the last decades. In Norway, the participation rate for women aged 55-66 rose from 40 per cent in 1972 to 54 per cent in 1997 (Dahl, Nilsen, and Vaage (2003)). Similar patterns are documented for US in Coile (2003). This gives us a chance to study spousal effect for not only married men but also married women.

There are several studies of retirement behavior based on Norwegian data. Among them, Dahl, Nilsen, and Vaage (2003) take family characteristics such as spousal income, wealth and labor market status into account when they analyze labor market behavior for elderly men and women. Using a multinomial logit model on pooled yearly data, they mainly focus their study on gender difference and have not paid much attention to the difference in spousal effects.

In this paper, we model the early retirement as a sequence of yearly decisions while taking the spouse's labor market status as exogenous. We take the advantage of the panel structure of the available data, which gives us the possibility to incorporate unobserved heterogeneity across the sample.

Our paper differs from Dahl, Nilsen, and Vaage (2003) mainly on two aspects. First, we have a much richer structure on the unobserved error term. The error term is assumed to be the sum of two parts: an individual-specific time-invariant part and a transitory part. The first part is there to incorporate the unobserved heterogeneity across the sample. In addition, we allow a very general autocorrelation structure on the transitory component. Second, unlike simply pooling the data from different year together as done in Dahl, Nilsen, and Vaage (2003), we take into account the repeated self-selection in the data set to eliminate the selection bias that arises when we allow for unobserved heterogeneity.

The model is estimated on married individuals qualified for AFP between 1994 and 1997. The results show that the employment status and other characteristics of the spouse have important effects on early retirement decision for married men and women. We find the direction of spousal effects are quite symmetric but women have a much stronger response to their spouses' characteristics than men. This indicates that joint modeling of couples' labor market behavior is appropriate. In addition, by comparing different specifications of the panel probit model, our study shows that failure to correctly model the cross individual heterogeneity structure (i.e. the intertemporal covariance structure of the error term) might lead to completely erroneous conclusions.

The rest of this paper proceeds as follows. In section 2, we discuss the model setup. Data and empirical specifications are presented in section 3. Section 4 reports the estimation results. The conclusion is elaborated in section 5.

2 The Model

In this paper, we concentrate on the spousal effects on the early retirement behavior of married men and women, and use a reduced form modeling framework.

There are mainly two reasons why we do not use a structural retirement model for this analysis. Firstly, the labor market decision process of the couples is quite complicated, and we are still in a stage of figuring out how it works. Many approaches on modeling family labor supply have been discussed in the literature. For a survey of these models, see Bergstrom (1997), Blundell and MaCurdy (1999). Roughly, the literature can be divided into two strands, the cooperative and non-cooperative approaches. They differ substantially on the underlying behavioral assumptions. Although there are analyses which compare the empirical performance of some of those competing models such as Hernæs, Jia, and Strøm (2001), the literature is still limited and no consensus has been reached so far. By making a particular behavioral assumption of the decision process, which we are not sure is a correct one, there is a risk that we might have already restricted the spousal effects to have certain counterfactual patterns. Moreover, in a behavioral model, certain points of interest can be obscured in the mist of complicated interactions. For example, in a household labor supply model where the decisions are assumed to be the outcome of a Nash game, the question: 'what will be the effect of an increase in the wife's labor income by 10 per cent on husband's labor decision', cannot be answered without numerical simulations.

Secondly, reduced form models can be thought as approximations of certain labor market decision rules derived from some unknown behavioral models, as Blau and Rihahn (1999) point out. No specific behavioral assumptions need to be made. The reduced models are relatively simpler to implement compared with the structural models, and easier to interpret. Those reduced form models for retirement range from simple probit or logit used by Coile (2003), Dahl, Nilsen, and Vaage (2003), to bivariate duration model developed by An, Christensen, and Gupta (1999), or to dynamic multinomial probit model applied in Blau and Rihahn (1999).

In this study, we are not interested in recovering the utility parameters for any of the spouse. The behavioral assumptions we make in this model should be as weak as possible to avoid the assumption contaminating our results. Given all these concerns, we specify a reduced form model that is in the spirit of Blau and Rihahn (1999). The parameters in our model may represent a certain combination of the preferences of both spouses and some other factors that influence the decision. But since the model can be seen as a decision rule that is the result of some complicated decision processes instead, the parameters do have quite straightforward interpretations.

2.1 The Model Setup

We model the early retirement as a sequence of yearly decisions from the date of the qualification.

We define $t = 1$ at the year of AFP qualification. The available choices at period t for each individual i are either taking out the early retirement ($y_{it} = 1$) or continuing to work ($y_{it} = 0$).

Let y^*_{it} represent the net value of retiring at time t. We will assume a linear approximation to the underlying function that determines y^*_{it}:

$$y^*_{it} = x_{it}\beta + u_{it} \quad i = 1, 2, ..., N; t = 1, 2, ...T_i. \tag{1}$$

Here x_{it} is a vector of observed explanatory variables which include wage, pension, spouse's labor market status etc. β is the unknown parameter vector. u_{it} is the unobserved disturbance. T_i is number of observation periods for individual i.

At time t, the individual chooses to retire only when the net value of retiring at time t is no less than 0, i.e.:

$$y_{it} = \begin{cases} 1 & \text{if } y^*_{it} \geq 0, \text{ and} \\ 0 & \text{otherwise.} \end{cases} \tag{2}$$

We are not able to observe the net value of retirement y^*_{it}, only the retirement behavior y_{it}. Moreover, the individual drops out from the panel once he takes out retirement. Namely, we only observe y_{it} when $y_{it-1} = 0$, since retirement is assumed to be an absorbing state. As a consequence, state dependence in terms of lagged dependent variable is not relevant in our setting, since for all the observations in our sample, we always have $y_{it-1} = 0$.

Modeling of the unobserved disturbance u_{it} is not straight forward. In the present study, the unobserved disturbance is assumed to consist of two part: α_i is a permanent individual specific effect that does not change over time, which is meant to capture the unobserved heterogeneity across the individuals. ε_{it} is a transitory part. Namely:

$$u_{it} = \alpha_i + \varepsilon_{it}. \tag{3}$$

If we assume there is no unobserved heterogeneity, that is $\alpha_i = 0$, and assume that ε_{it} is *i.i.d.* distributed, then the panel data structure is irrelevant and we can simply pool all the data together. The choice probability will have a Logit/Probit structure, depending on the distributional assumption on ε_{it}. This is essentially what has been done in Dahl, Nilsen, and Vaage (2003). Although as we can see later, the assumption that no unobserved heterogeneity across the sample is not valid in our study, we will still estimate of a pooled probit model for comparison reasons.

If we treat α_i as unknown parameters together with β, we will have a fixed effect model. Maximum likelihood estimation method can be used to estimate the fixed effect model. However, the estimators are only consistent in the sense that both the observation number N and the number of time periods T can tend towards to infinity. In practice, we usually only have quite small T, both the parameters of interest and the nuisance parameters (fixed effects) can not be estimated consistently because of the so-called 'incidental parameters problem' pointed out by Neyman and Scott (1948). The general solution is to apply the conditional likelihood model suggested by Chamberlain (1984). But it is not possible in our case, due to the fact that we only observe the transition from working to retirement not vice versa.

For this reason, in the second specification of our model, we take a standard random effect approach and assume that α_i is random across individuals and normally distributed with mean 0 and variance σ_α^2. One important restriction of the standard random effect model is the assumption that the correlation between the disturbance u_{it} for any two decision points is the same regardless of how far apart the decisions are. However, u_{it} may reflect the tastes that gradually change over time, then one should expect that the correlation is bigger with shorter period in between. A typical solution to this problem is to specify a dynamic structure for the transitory error ε_{it} as well. So for the third specification, similar to Hyslop (1999) and Michaud (2003), we assume that ε_{it} follows a stationary AR(1) process with autocorrelation coefficient ρ,

$$\varepsilon_{it} = \rho\varepsilon_{it-1} + v_{it}, \tag{4}$$

where the innovation v_{it} is assumed to be i.i.d. normal distributed over individuals and time with mean 0 and variance σ_v^2

When we consider the specifications discussed above, we notice that even with an AR(1) transitory error term, the intertemporal covariance structure is still quite restrictive. Ideally, we would like to specify that ε_{it} and ε_{is} are freely correlated. As the last specification, we would like to estimate a panel binary choice model with an unconstrained covariance structure.

Let $T = \max_i(T_i)$, and Σ be a $T \times T$ positive definite covariance matrix. We assume that $(\varepsilon_{i1}, \cdots \varepsilon_{iT_i})$ is T_i-variate normally distributed with mean 0 and a covariance matrix $\Sigma_{(i)}$ which is the T_ith leading principal submatrix of Σ. Recall that $\alpha_i \sim N(0, \sigma_\alpha^2)$, and using (3), we see that the total disturbance $(u_{i1}, \cdots u_{iT_i})$ is also T_i-variate normally distributed with mean 0. However, the covariance matrix is now (I_{T_i} denotes the identity matrix of rank T_i):

$$\sigma_\alpha^2 I_{T_i} + \Sigma_{(i)}.$$

It follows immediately that with an unconstrained covariance structure on ε_{it}, we will not be able to identify σ_α^2 separately from Σ. In other words, the cross individual heterogeneity can not be distinguished from the simple cross period correlation. So we can ignore the individual effect α_i in this setting.

One of the most important reasons for the popularity of the simple random effect model is that the cost of evaluating high dimensional integral expressions. However, with the development of hardware as well as the simulation methods such as the GHK simulator and Simulated MLE, this is no longer the case.

3 Data and Empirical Specifications:

3.1 Sample Construction

We use data from the merged administrative registers at the Frisch Centre[1]. The data contain detailed socio-economic information and give an account of the main labor market activities for virtually the whole Norwegian adult population.

We first restrict our sample to married individuals qualified for AFP between 1994 and 1997, when the AFP qualification age is constant at 64. The qualification of AFP requires: 1) currently employed and have earnings higher than the basic pension G[2] , 2) at least 10 years of work experience since the age of 50, 3) at least 3 years' tenure in the present firm, 4) an average of the 10 highest yearly income after 1966 exceeding at least 2G. The observations are then censored upon dissolution of marriage for any reason (death of the principle person or the spouse is the most important factor here) during the period of analysis. We suspect that those who have very high personal income may have different incentives on retirement decisions. So we exclude those who have labor earning higher than 1,000,000 NOK or business income higher than 500,000 NOK. In

[1]The original data have been received from Statistics Norway, and held by the Frisch Centre with permission for research use.

[2]G is a crucial parameter in the Norwegian pension system, used for defining contributions as well as benefits. The amount is adjusted by the Parliament once or more times each year, in accordance with changes in the general income level.

this study, disability retirement is not considered as a voluntary choice, so those who actually took out disability pension during our observation period are also excluded in our sample.

Starting from the year of eligibility, we track the labor market status on yearly bases for all individuals in the data set until they take out the early retirement or the ordinary retirement age of 67 is reached. The maximum number of observed period is 3, since at the starting period the individuals are all 64 years old. For those who retire at the year of eligibility we only have one observation per individual.

The resulted data set for men contains 9971 individuals and 18707 observations, while the data set for women contains 6210 individuals and 11628 observations.

3.2 Empirical Specifications:

We write the deterministic part of (1) as the following:

$$x_{it}\beta = \beta_0 + Z_{it}\beta_z + S_{it}\beta_s, \tag{5}$$

where Z_{it} is a vector of the individual's own characteristics. S_{it} is a vector of the spouse's characteristics.

3.2.1 Own characteristics Z_{it}

For the individual's own characteristics, we include a series of demographic variables: age, years of education, the number of dependent children (children under 18 years old). Personal wealth from tax authority is also included in our analysis.

We include incomes for both continuing to work and taking out AFP retirement in the analysis. Since the individuals can only be observed in one state, either working or retired, we need to impute potential AFP pension or potential wage income. For wage income, we exploit the fact that all the individuals in our sample have been working before entering the sample and specify a simple autoregressive process for the $\log(wage_t)$ as follows:

$$\log(w_{t+1}) = \gamma_1 + \gamma_2 \log(w_t) + \gamma_3(\log(w_t))^2 + \gamma_4 age_t + \gamma_5 age_t^2 + \xi_t, \tag{6}$$

where age_t denotes the individual's age at period t, and ξ_t is $i.i.d.$ normal distributed with mean 0 and variance σ_ξ^2. The quadratic specification allows for an age income profile.

The regression is done separately for men and women. Table (1) shows the results from the estimation of this model. Both regressions have quite high R^2 (>90%), which

		Men		Women	
Variable	Parameter	Estimate	S.E.	Estimate	S.E.
constant	$\hat{\gamma}_1$	0.175	0.82	-18.189	11.35
$\log(w_t)$	$\hat{\gamma}_2$	-0.871	0.009	0.439	0.17
$\log^2(w_t)$	$\hat{\gamma}_3$	0.727	0.004	0.022	0.007
age	$\hat{\gamma}_4$	0.03	0.51	0.667	0.35
age^2	$\hat{\gamma}_5$	-2.4e-4	1.9e-4	-0.005	0.002
	R^2	0.94		0.92	
	$\hat{\sigma}^2_\xi$	7.1e-3		8.0e-3	
Number of observations		9617		6260	

Table 1: Estimation results for wage regression equations, men and women

indicates a very good fitting. The estimates then are used to impute the wage income when the individual is observed not working at time t, based on the wage income of last year using a Markovian updating formula:

$$w_{t+1} = \exp(\hat{\gamma}_1 + \hat{\gamma}_2 \log(w_t) + \hat{\gamma}_3(\log(w_t))^2 + \hat{\gamma}_4 age_t + \hat{\gamma}_5 age_t^2 + \hat{\sigma}^2_\xi/2). \tag{7}$$

The AFP pension is calculated using detailed pension rules. Both of these income variables are net of taxes using detailed tax rules with all the spousal characteristics considered. (Both the pension rules and the tax rules can be found in Haugen (2000))

We are also interested in whether the employer specific variables influence the retirement behavior. We include a dummy on whether the individual is working in the private sector. We also try to construct a firm specific employment reduction dummy to proxy firm's labor demand. However, due to some data problems on identifying firms throughout time, the dummy variable we generated is not of high accuracy. We need to bear it in mind when we interpret the results.

One important factor which influences the retirement behavior is the health condition. Based on the data we have, we construct a yearly sick-leave ratio, which measures the fraction of sick leave during the year prior to the year of decision.

Social norms have been considered to be an important factor in the retirement decision making process in recent years. We use a county level gender specific retirement ratio as a measure of the social norm. It is defined as the fraction of pensioners of the same gender in the county where the individual lives.

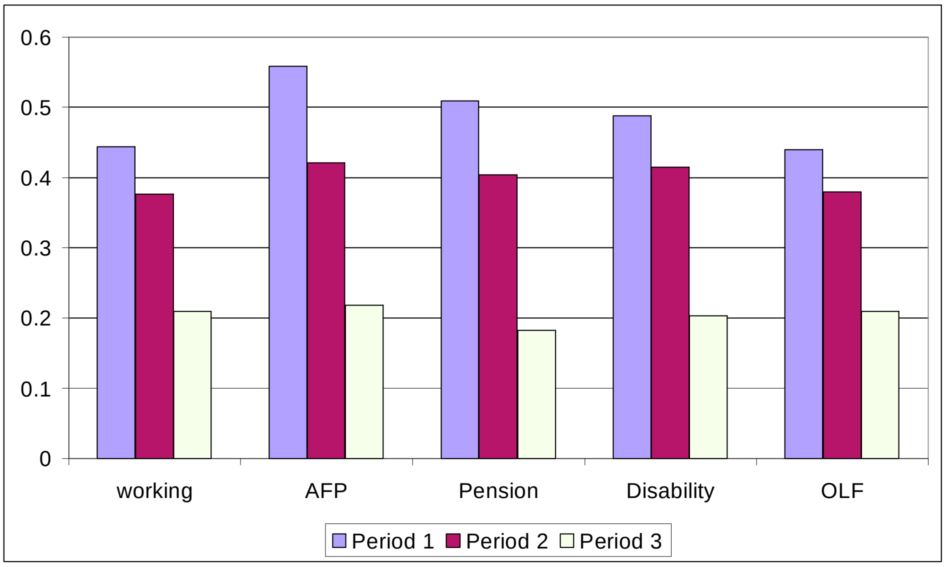

Figure 1: Men's retirement hazard over the three periods, grouped by their spouses' labor market status

3.2.2 Spouse's characteristics S_{it}

The most important variable for the spouse is the labor market status. Spouses are categorized into five states: work, early retirement, on disability pension, other pensions (including ordinary retirement at age 67), and out of labor force. The spouses are classified as in the working state if they can be found in the work register file or their annual labor earnings from tax file are more than 50,000 NOK. The states, 'early retirement' and 'on disability pension', are easy to identify, given that we have detailed register files on those activities. The state 'other pensions' includes those on ordinary pension, and those who are observed to have pension income greater than 80,000 NOK in the tax file. Those who are not in any of four states above are then classified as 'out of labor force'.

We also include spouse's age, actual income and personal wealth in our analysis.

Unfortunately, we are not be able construct a similar health indicator for the spouse. The reason is simple — only those who are working have the sick leave data. On disability pension might be used as a proxy for bad health. But we need to be very careful, since the eligibility criteria have been lax for older workers.

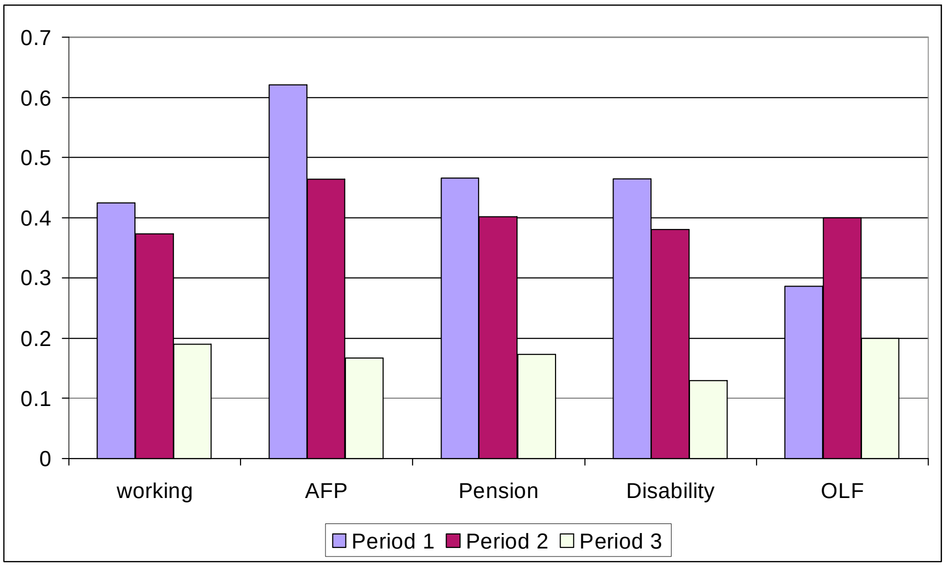

Figure 2: Women's retirement hazard over the three periods, grouped by their spouses' labor market status

3.3 Some Summary Statistics

The summary statistics for both data sets are given in table (2). We see from the table that although on average men and women have similar years of education (12.6 years for men versus 11.5 years for women), women earn much less and have accumulated less wealth than men. The average AFP pension replacement ratio for both men and women are around 60 per cent. The redistribution effect of the AFP scheme is apparent. The standard deviation of the AFP pension income is only around 30 per cent of that of wage income for men, and 40 per cent for women. Husbands are, on average, 3 years older than wives. This explains the different patterns of spouse's labor market status between men and women.

Figure (1) and (2) give the early retirement hazard rates over the three observed periods, grouped by their spouses' labor market status. We see a general trend that the retirement hazard rate is decreasing regardless of their spouses' status when t increases. And the differences between the different groups seem to diminish over time.

	Men		Women	
	mean	std	mean	std
Own characteristics Z_{it}				
Age	64.64	0.76	64.65	0.76
Working in private sector	0.68	0.47	0.49	0.50
Number of dependent children	0.03	0.22	0.001	0.02
Work income (net of tax)	1.92	0.57	1.33	0.39
AFP income (net of tax)	1.10	0.16	0.82	0.16
Sick ratio in last year	0.03	0.10	0.04	0.12
Years of education	12.57	3.54	11.49	2.84
Share of retirees in same county	0.25	0.05	0.29	0.05
Firm downsizing	0.15	0.36	0.13	0.33
Wealth	4.77	8.04	2.08	2.94
Spousal characteristics S_{it}				
Age	61.02	4.50	67.81	4.20
Wealth	1.41	2.91	3.80	7.14
Income	1.00	0.58	1.47	0.73
AFP retired	0.02	0.13	0.03	0.17
Pension	0.07	0.26	0.59	0.49
On disability	0.20	0.40	0.09	0.28
Out of labor force	0.15	0.35	0.01	0.07

Note: all the income variables and wealth variables are in 100000 NOK

Table 2: Descriptive Statistics

4 Estimation Results

In this section we present the results for several different specifications of the early retirement decision model discussed in section 2.1. All the specifications are estimated separately for men and women.

We first present estimates from a pooled probit model which ignores the panel structure of the data. We use it mainly as a benchmark for comparison purpose. Following this, estimates for several panel data models are reported and discussed. These panel data models include a static random effect model, a dynamic panel model with AR(1) transitory error terms and a dynamic panel model with no restriction on the covariance structure of the transitory error terms.

We then provide a discussion of possible determinant factors of early retirement behavior based on these results.

4.1 Pooled Probit Model

The simplest way to deal with panel data on discrete choice models might be simply pooling the data together and assume independence over both time horizon and individuals, instead of using only one of the cross-sections. This leads to a pooled probit/logit model. Both Coile (2003) and Dahl, Nilsen, and Vaage (2003) apply this framework.

In most cases, compared with the single cross section, the pooled estimators are usually more efficient due to the increased sample size. Even when there are correlations among the error terms u_{it} in the net value formula (1), Maddala (1987) argues that ignoring these correlations among the errors and using a standard Probit estimation method with pooled data produces consistent (though inefficient) estimates. However, this is not true for studies of retirement when retirement is considered as an absorbing state. In this case, those who retire at time t will drop out sample from $t+1$. This is essentially equivalent to a repeated self-selection. Simply pooling data together when there are individual specific effects and repeated self-selection will lead to inconsistent estimates of the parameters.

To make this point clear, denote the pooled sample log-likelihood function as:

$$S_n(b) = \sum_t \sum_i \ln(\Pr(x_{it}\beta + u_{it} \geq 0)^{y_i} + \Pr(x_{it}\beta + u_{it} < 0)^{1-y_i}), \tag{8}$$

and the maximum likelihood estimator $\widehat{\beta}_n$ is defined as $\widehat{\beta}_n = \arg\max_b(S_n(b))$.

For $\widehat{\beta}_n$ to be consistent, we need to have the conditional mean restriction

$$Mean(\alpha_i + \varepsilon_{it}|x_{it}, y_{it-1} = 0) = 0, \qquad \text{for all } t.$$

However, in our case, this mean condition cannot be satisfied due to the self-selection. The individuals who remain in the sample at $t > 1$ have smaller values of α_i than those who retire at $t = 1$. Consequently, $Mean(y_{it}|x_i, y_{it-1} = 0) < 0$ for $t > 1$. This violates the conditional mean condition and thus leads to inconsistency of the estimator $\widehat{\beta}_n$.

For comparison reasons, we include the pooled probit estimates for both men and women in table (3). A striking result for both men and women is that the effects of their own age on the retirement decision are both significantly negative. It suggests that the older the individuals are, the less likely they will take out retirement when other variables are controlled for. When we notice that in our sample, all the individuals are 64 years old at year of eligibility $t = 1$, we see immediately that it is due to the phenomenon that the larger t is, the lower is the hazard rate of retirement. Recall

	Men		Women	
Variable	Estimate	S.E.	Estimate	S.E.
Constant	17.4531	0.9224	21.6149	1.1465
Own characteristics Z_{it}				
Age	-0.2716	0.0150	-0.3364	0.0188
Working in private sector	0.2101	0.0218	0.1123	0.0241
Number of dependent children	-0.1059	0.0479	0.1123	1.0839
Work income (net of tax)	-0.3031	0.0283	-0.7721	0.0539
AFP income (net of tax)	0.0970	0.1122	0.8022	0.1336
Sick ratio in last year	0.7987	0.0930	0.6926	0.0998
Years of education	-0.0271	0.0037	0.0099	0.0052
Share of retirees in same county	0.5156	0.1976	-0.1240	0.2502
Firm downsizing	-0.0196	0.0266	0.0034	0.0364
Wealth	-0.0000	0.0013	-0.0004	0.0044
Spousal characteristics S_{it}				
Age	0.0064	0.0026	-0.0013	0.0041
Wealth	0.0096	0.0036	0.0045	0.0019
Income	-0.0494	0.0279	0.0173	0.0214
AFP retired	0.1331	0.0762	0.3848	0.0742
Pension	-0.0544	0.0432	0.1056	0.0374
On disability	-0.0318	0.0275	0.0607	0.0489
Out of Labor Force	-0.0874	-0.0443	-0.2305	0.1938
Loglikelihood	-11798.32		-7351.06	
Number of Cases	9971		6210	

Note: the standard errors are calculated using BHHH method and have been corrected for panel data.

Table 3: Pooled Probit Model of Early Retirement Behavior, Men and Women

that with individual specific effects, the individuals who remain in the sample at t have higher attachment to the labor market (lower α_i) than those who retire at $t-1$. The significant negative parameter estimates for age might just be a consequence of ignoring the individual specific effect and the repeated self-selection.

4.2 Random Effect Probit Models

In table (4), two versions of the random effect model are presented for both men and women.

The first version is a standard random effect model, with assumption that α_i is $i.i.d.$ normal distributed with mean 0 and variance σ_α^2 across the individuals.

The second version allows an AR(1) structure on the transitory error terms ε_{it}. Namely $\varepsilon_{it} = \rho\varepsilon_{it-1} + v_{it}$, where v_{it} is assumed to be $i.i.d.$ normal distributed with mean 0 and variance σ_v^2. Similar specification can be found in Hyslop (1999) and Michaud (2003).

Like other discrete choice models, some normalizations are required for identification. For both versions, we normalize the variance of the overall error term u_{it} to 1, i.e. $var(u_{it}) = 1$. The main reason for using this normalization is that under this normalization both estimates will have the same scale, which makes the comparison of those two models quite straight forward. Furthermore, under this normalization the estimated variance of α_i has the interpretation as the share of unobserved heterogeneity in the total variance of the stochastic component u_{it}. In the AR(1) specification, given $var(\alpha_i) = \sigma_a^2$ and the AR(1) coefficient ρ, it implies that $\sigma_v^2 = (1-\sigma_a^2)(1-\rho^2)$.

We see from table (4) that after allowing for the individual heterogeneity and autocorrelation in the transitory term, the coefficient for age no longer has the significant negative sign. This confirms our hypothesis that the wrong sign of the parameters is mainly due to the failure to take into account of the individual specific effect and the repeated self-selection. This also gives a nice empirical illustration for the inconsistency problem we discussed in the last section.

For both men and women, the $AR(1)$ coefficients are significantly different from zero. In comparison to the pure random effect model, the addition of a serial correlated transitory component in error term improve the model fitting substantially. However, the negative sign of the $AR(1)$ coefficients seems to be somewhat counter-intuitive, since we expect that the shorter the period apart from each other, the stronger the correlation will be. Interestingly, both Hyslop (1999) and Michaud (2003) find the similar results in their study of dynamic probit models when the state dependence is present.

	Men				Women			
	Random		AR(1)		Random		AR(1)	
Variable	Estimate	SE	Estimate	SE	Estimate	SE	Estimate	SE
Constant	4.8738	3.309	-7.2833	5.071	16.8679	4.114	0.5766	3.291
Own characteristics Z_{it}								
Age	-0.0746	0.052	0.1145	0.079	-0.2621	0.065	-0.0076	0.052
Working in private sector	0.2350	0.028	0.2376	0.060	0.1229	0.004	0.1416	0.033
Number of children <18	-0.1077	0.051	-0.1009	-0.056	0.1298	1.155	0.2501	1.445
Work income (net of tax)	-0.3528	0.041	-0.3738	0.093	-0.8048	0.095	-0.8340	0.119
AFP income (net of tax)	0.1391	0.122	0.1860	0.126	0.8295	0.155	0.8304	0.158
Sick ratio in last year	0.8165	0.102	0.7306	0.165	0.7005	0.111	0.6618	0.125
Years of education	-0.2567	0.004	-0.0352	0.009	0.0098	0.005	0.0079	0.005
Share of retirees	0.8165	0.102	0.7305	0.165	-0.1031	0.259	-0.0082	0.270
Firm downsizing	-0.0189	0.032	-0.0197	0.024	0.0058	0.036	0.0189	0.033
Wealth	0.0004	0.001	0.0008	0.001	-0.0003	0.004	0.0008	0.001
Spousal characteristics S_{it}								
Age	0.0071	0.002	0.0080	0.003	-0.0013	0.004	-0.0018	0.004
Wealth	0.0106	0.004	0.0109	0.004	0.0045	0.001	0.0008	0.004
Income	-0.0523	0.029	-0.0528	0.031	0.0187	0.022	0.0236	0.023
AFP retired	0.1572	0.077	0.1422	0.076	0.3855	0.079	0.3573	0.082
Pension	-0.0482	0.045	-0.0509	0.044	0.1093	0.039	0.1237	0.040
On disability	-0.0243	0.029	-0.0176	0.029	0.0667	0.050	0.0949	0.051
Out Labor Force	-0.0894	0.047	-0.0929	0.050	-0.2480	0.198	-0.2656	0.202
Covariance Parameters								
SD of random effect	0.5342	0.1029	0.7891	0.2224	0.3263	0.1970	0.7266	0.1419
AR(1) coefficient			-0.7250	0.1022			-0.7128	0.0618
loglikelihood	-11782.9		-11722.2		-7349.87		-7293.92	
Number of Cases	9971		9971		6210		6210	

Table 4: Random Effect Model with independent transitory error term, and AR(1) transitory error term

4.3 A Panel Probit Model with unconstrained covariance structure

As discussed in section 2.1, with a freely specified covariance structure of the transitory disturbances ε_{it}, we will not be able to identify the individual specific effect α_i in the random effect modeling framework. So we make assumption directly on the total disturbance. We assume that $(u_{i1}, \cdots, u_{iT})$ is normally distributed with mean 0 and an unconstrained covariance matrix Σ. As shown in Greene (2002), this setting can be interpreted as a special case of 'random efficient probit/mixed probit' with only a random constant term which embodies the latent heterogeneity.

Similar to the standard random effect model discussed in last section, normalization is required to ensure identification. However, normalization of all the diagonal elements is unnecessary because the slope vector is time invariant. So only one main diagonal element of the covariance matrix is required to be normalized. We will simply normalize $var(u_{i1}) = 1$. To ensure the positive definiteness of the covariance matrix, instead of directly specifying the parameters for the covariance matrix Σ, we choose to specify the Cholesky decomposition of Σ. Namely, we assume that $\Sigma = LL'$ where L is a lower-triangular matrix which is defined as

$$L = \begin{pmatrix} 1 & & \\ L_{21} & L_{22} & \\ L_{31} & L_{32} & L_{33} \end{pmatrix}. \tag{9}$$

The elements L_{ij} of this Cholesky factor are the parameters to be estimated in the model.

Estimation results for this unconstrained covariance panel probit model are reported in table (5).

From the estimates, we see that the log-likelihood value improves considerably. The hypotheses that $var(u_{it}) = 1$, $t = 2, 3$ are both rejected at 1% level. The estimated covariance matrices for ordinary random effect, AR(1) and unconstrained covariance specification are reported in table (6)

We see from table (6) that for both men and women, the estimated covariance matrices for these three different specifications vary a lot. For men, the unconstrained covariance matrix actually shows that the shorter the time distance, the stronger the correlation between the error terms, despite that autocorrelation coefficient estimates from the AR(1) are significantly negative. However, this is not the case for women. Anyway, this result suggests that the negative autocorrelation parameter reported in Hyslop (1999) and Michaud (2003) might be due to the restriction that the error terms

	Men		Women	
Variable	Estimate	S.E.	Estimate	S.E.
Constant	-0.5511	10.4183	44.3187	33.5199
Own characteristics Z_{it}				
Age	0.0074	0.1630	-0.6941	0.5238
Working in private sector	0.3305	0.0287	0.2111	0.0320
Number of children below 18	-0.0975	0.0575	0.0690	1.4931
Work income (net of tax)	-0.5655	0.0417	-1.1688	0.0770
AFP income (net of tax)	0.5543	0.1626	1.3376	0.1828
Sick ratio in last year	1.1060	0.1197	1.0569	0.1392
Years of education	-0.0426	0.0050	0.0022	0.0068
Share of retirees in same county	0.5961	0.2538	0.2948	0.3253
Firm downsizing	-0.0078	0.0320	0.0290	0.0478
Wealth	0.0032	0.0017	-0.0026	0.0065
Spousal characteristics S_{it}				
Age	0.0089	0.0034	-0.0004	0.0055
Wealth	0.0145	0.0048	0.0063	0.0020
Income	-0.0443	0.0367	0.0343	0.0292
AFP retired	0.2888	0.1125	0.5486	0.0925
Pension	-0.0057	0.0578	0.1613	0.0476
On disability	-0.0011	0.0356	0.1312	0.0598
Out Labor Force	-0.0876	0.0556	-0.3716	0.2292
Covariance structure L				
L_{21}	0.4868	0.1692	-0.0088	0.4819
L_{22}	1.7691	0.2456	3.1564	0.8711
L_{31}	0.2852	0.4229	1.1464	0.4527
L_{32}	1.9091	0.6315	-0.0157	1.3755
L_{33}	1.3968	0.4252	2.3367	0.6543
Loglikelihood	11665.1		7252.356	
Number of Cases	9971		6210	

Table 5: Random Effect Model with Unconstrained Covariance Matrix

	Standard Random Effect	AR(1)	Unconstrained Covariance
Men	$\begin{pmatrix} 1 & & \\ 0.29 & 1 & \\ 0.29 & 0.29 & 1 \end{pmatrix}$	$\begin{pmatrix} 1 & & \\ 0.35 & 1 & \\ 0.82 & 0.35 & 1 \end{pmatrix}$	$\begin{pmatrix} 1 & & \\ 0.49 & 3.37 & \\ 0.29 & 3.51 & 5.67 \end{pmatrix}$
Women	$\begin{pmatrix} 1 & & \\ 0.11 & 1 & \\ 0.11 & 0.11 & 1 \end{pmatrix}$	$\begin{pmatrix} 1 & & \\ 0.19 & 1 & \\ 0.77 & 0.19 & 1 \end{pmatrix}$	$\begin{pmatrix} 1 & & \\ -0.01 & 9.96 & \\ 1.15 & -0.06 & 6.77 \end{pmatrix}$

Table 6: Estimated Covariance Matrices from the three specifications of errors

have equal variance across time.

Using likelihood ratio tests, all the other specifications can be rejected with quite high level of significance.

4.4 Determinant factors of early retirement behavior

We base our discussion of the estimation results on the unconstrained covariance structure specification (table (5)).

4.4.1 Effects of own characteristics

When controlling for other variables, own age seems to have no significant effect on early retirement decision for both men and women. Given the fact that all the individuals in our sample are qualified for AFP, and the relative small range of variation (64-66), it is not unreasonable.

The effects of working income and benefits are in the expected direction. We expect that higher wages and lower benefits corresponding to a strong attachment to the labor force. The results support our hypothesis. Increased working income significantly reduces the probability of taking out early retirement, while increased AFP pension income has an opposite effect. Similar to Blau and Rihahn (1999) and Dahl, Nilsen, and Vaage (2003), we also find that the response for women is much stronger compared to men, which is consistent to the generally higher labor supply elasticities for women than men found in the literature. The influences of own wealth on early retirement are less obvious. The estimated parameters are not significant for both men and women.

We expect that the length of education will have a negative effect on the early retirement decision, for those who with higher investment on human capital will be less inclined to exit the labor market. We do find the expected effect for men, while the effect for women has wrong sign but with quite high standard error, thus is far from significant.

Since husbands generally are older than their wives, there are more men than women with dependent children (younger than 18 years) in our sample. Similar to Dahl, Nilsen, and Vaage (2003), we find that having dependent children tends to reduce the probability of early retirement for man. The effect for women is negligible since there are too few women with dependent children in our sample (less than 0.1%).

We specify our model to include a proxy of health condition: the sick leave ratio during the year before eligibility to AFP. It may not be a precise measurement of the health condition. But we expect that it has a high correlation with the actual health condition. The estimates do show a quite high propensity to early retirement for those with higher sick leave ratios during the last year. The same pattern holds for both men and women.

Sociologists who study retirement behavior are interested in factors like norms, family features etc, which have not been discussed much in the economic literature. We include a crude measure of social acceptance of retirement in our study, namely the share of retirees in the county where the individual lives in. Interestingly, we find that the retirement behavior seems to be positively influenced by this share. In other words, keeping other variables constant, individuals who live in the county with higher share of retirees are more likely to take out the early retirement. This holds for both men and women, although it is not significant for women.

Working in the private sector has a positive effect on taking out early retirement. And the effect is bigger for men. It might be due to the fact that working life in the private sector is tougher than that in public sector, or more importantly, due to the fact that we are not be able to observe the occupational pension in the private sector. The finding is in line with what found in Hernæs, Jia, and Strøm (2001). Rather out of our expectation, belonging to a downsizing firm doesn't have a significant effect on the retirement behavior. One reason for this could be that the data quality of this variable is poor as we discussed in earlier section.

4.4.2 Spousal spill-over effects

Turning to the spouse's characteristics, we find that spouse's age seems to have a positive effect on the retirement for men, while a non-significant negative effect for women. For men, similar patterns are also documented in Hernæs, Jia, and Strøm (2001).

High spousal income implies a reduction of the probability of early retirement for man, while the opposite for women. However, none of these effects are significant. The spouse's wealth does have a significant positive effect for both men and women. This might be due to the argument that increased wealth will improve the possibility of early retirement through the increased ability of self-support, as mentioned by Dahl, Nilsen, and Vaage (2003).

Using working spouse as the reference group, for men, wife being a AFP pension receiver has a quite strong positive effect on early retirement. With a wife who is classified as out of labor force, the probability for the husband to take out early retirement is reduced. Two hypotheses are usually discussed in the literature, the added worker effect and the complementary of leisure effect. The added worker effect says that when one spouse is not working and has limited resources, we would expect compensating behavior from the other, namely non-working wife may correlate with a low probability of early retirement. While the complementarity of leisure hypothesis stating that the couples put high value on their joint leisure than leisure enjoyed just by one of them, we would expect that the probability of taking out early retirement will be higher if he has a non-working wife. No conclusion can be made in general on which is dominant from our estimation results. The reduced early retirement probability for husband with wife who is out of labor force supports the added worker hypothesis. And the strong positive effect of an AFP wife on the probability can be seen as an evidence of the complementarity of leisure hypothesis. Note that the state 'out of labor force' is corresponding to quite low income or zero income, it is no surprising the added worker effect dominates. When a non-working state corresponding to enough income such as pension etc, the budget constraints are less important, thus the complementarity of leisure effect dominates.

Similar picture can be found for women, but these effects are more sharply determined and much stronger than those we find for men. For women, non-working spouse increases the probability of retirement except the case when the husband are observed to be out of labor force.

An interesting point to note is that the spouse being in the state of AFP retired seems to have the highest effect on probability of early retirement, for both men and women. Possible explanations for this are as follows. On one hand, early retirement through AFP is fully voluntary, so planning of retirement through this path way is possible. On the other hand, the relative generous AFP pension makes the budget constraints not so important. These two factors combined together greatly increase the freedom of coordination of the retirement within the family.

In general, the direction of spousal effects are quite symmetric but women seem to have a much stronger response to their spouses' characteristics than men. It is quite different from what has been found in Gustman and Steinmeier (2000) and Coile (2003) on US data, but to some extent agrees with the findings of Schellenberg (1994) on Canadian survey data.

5 Conclusion

The aim of this paper is to use the panel data structure to model the individuals' early retirement behavior when a new option, AFP early retirement becomes available. Several specifications of the binary choice panel data model are estimated on Norwegian data. Main focus of this study is the comparison of spousal effect for men and women.

The main findings in the paper are two folds. First, we find that the employment status and other characteristics of the spouse have important effects on early retirement decision for married men and women. This suggests that joint modeling of the couple's labor market behavior is appropriate. The direction of spousal effects are quite symmetric but women seem to have a much stronger response to their spouses' characteristics than men. Second, the comparison of different specifications of the panel probit shows that it is important to correctly model the error term covariance structure or equivalently the cross individual heterogeneity structure. This point has been overlooked in the most practice of retirement behavior modeling when panel data are available. One possible reason is the heavy computational burden associated with a general correlation setting. With the development of simulation methods such as simulated maximum likelihood, simulate score method and relative fast algorithms for high dimensional integral, computation is no longer a constraint.

However, some reservations have to be made. First, we need to keep in mind that the effects we find in this paper are not really causal effects, since we do not model the joint labor market behavior of husband and wife simultaneously like in Gustman and Steinmeier (2000) and Hernæs, Jia, and Strøm (2001). Moreover, our results are based on a sample of individuals who have selected themselves into quite strong attachment to the labor participation (the requirement for AFP qualification). We need to be careful when we want to generalize our results. Finally, although we try to specify a general covariance structure for the error term, we still made a strong distributional assumption that u_{it} is normally distributed. It would be ideal if we could specify a semi-parametric estimator which can relax this assumption. However, the attempt to apply a revised version of maximum score estimator suggested by Mayer *et al.*(2002) was not successful. The numerical optimization for such an estimator, when the number of parameters is large, is proven to be too difficult to handle, even with the help of genetic optimization algorithm, simplex method and simulated annealing algorithm.

References

An, M. Y., B. J. Christensen, and N. D. Gupta (1999): "A Bivariate Duration Model of the Joint Retirement Decisions of Married Couples," Center for Labor Market and Social Research Working Paper 99-10.

Bergstrom, T. C. (1997): "A Survey of Theories of the Family," in *Handbook of Population and Family Economics Volumn 1A*, ed. by M. R. Rosenzweig, and O. Stark, chap. 1. Elsevier Science, Amsterdam.

Blau, D. M. (1997): "Social Security and the Labor Supply of Older Married Couples," *Labor Economics*, 4, 373–418.

Blau, D. M., and R. T. Rihahn (1999): "Labor Force Transistions of Older Married Couples in Germany," *Labor Economics*, (6), 229–251.

Blundell, R., and T. MaCurdy (1999): "Labor Supply: A Review of Alternative Approaches," in *Handbook of Labor Economics 3A*, ed. by O. Ashenfelter, and D. Card, pp. 1560–1695. Elsevier Science, Amsterdam.

Chamberlain, G. (1984): "Panel Data," in *Handbook of Econometrics, Vol II*, ed. by Z. Griliches, and M. D. Intriligator, chap. 22. Elsevier Science, Amsterdam.

Coile, C. C. (2003): "Retirement Incentives and Couple's Retirement Decisions," NBER Working Paper Series 9496.

Dahl, S., O. A. Nilsen, and K. Vaage (2003): "Gender Differences in Early Retirement Behavior," *European Sociological Review*, (19), 179–198.

Greene, W. (2002): "Convenient Estimator for the Panel Probit Model: Further Results," Manuscript, Department of Economics, Stern School of Business, New York University.

Gustman, A. L., and T. L. Steinmeier (2000): "Retirement in Dual-Career Families: A Structure Model," *Journal of Labor Economics*, 18(3), 503–545.

Haugen, F. (2000): "Insentivvirkninger av Skatte-og Pensjonsregler," Frisch Center of Economic Research, Report 4/2000.

Hernæs, E., z. Jia, and S. Strøm (2001): "Retirement in Non-cooperative and Cooperative Families," Cesifo Working Paper 476.

Hyslop, D. R. (1999): "State Dependence, Serial Correlation and Heterogeneity in Intertemporal Labor Force Participation of Married Women," *Econometrica*, 67(6), 1255–1294.

Jimenez-Martin, S., J. Labeeaga, and M. Granado (1999): "Health Status and Retirement Decisions for Older European Couples," IRISS Working Paper Series No 1999-01.

JOHNSON, R. W., AND M. M. FAVREAULT (2001): "Retiring Together or Working Alone: The Impact of Spousal Employment and Disability on Retirement Decisions," CRR Boston College Working Paper 2001-1.

MADDALA, G. S. (1987): "Limited Dependent Variable Models Using Panel Data," *Journal of Human Resources*, XXII, 307–338.

MAYER, W. J., T. BLACKSTONE, B. SHANG, AND K. N. WOMER (2002): "Maximum Score Estimation of Binary Choice Panel Data Models with Sample Selection: An Application to Navy Officer Retention," University of Mississippi Department of Economics Working Paper.

MICHAUD, P.-C. (2003): "Joint Labour Supply Dynamics of Older Couples," CentER Working Paper 2003-9.

NEYMAN, J., AND E. L. SCOTT (1948): "Consistent Estimates Based on Partially Consistent Observations," *Econometrica*, 16, 1–32.

SCHELLENBERG, G. (1994): *Road to Retirement: Demographic and Economic Changes in the 90s.* The Canadian Council on Social Development, Ottawa.

Recent publications in the series Discussion Papers

314 K. E. Rosendahl (2002): Cost-effective environmental policy: Implications of induced technological change

315 T. Kornstad and T.O. Thoresen (2002): A Discrete Choice Model for Labor Supply and Child Care

316 A. Bruvoll and K. Nyborg (2002): On the value of households' recycling efforts

317 E. Biørn and T. Skjerpen (2002): Aggregation and Aggregation Biases in Production Functions: A Panel Data Analysis of Translog Models

318 Ø. Døhl (2002): Energy Flexibility and Technological Progress with Multioutput Production. Application on Norwegian Pulp and Paper Industries

319 R. Aaberge (2002): Characterization and Measurement of Duration Dependence in Hazard Rate Models

320 T. J. Klette and A. Raknerud (2002): How and why do Firms differ?

321 J. Aasness and E. Røed Larsen (2002): Distributional and Environmental Effects of Taxes on Transportation

322 E. Røed Larsen (2002): The Political Economy of Global Warming: From Data to Decisions

323 E. Røed Larsen (2002): Searching for Basic Consumption Patterns: Is the Engel Elasticity of Housing Unity?

324 E. Røed Larsen (2002): Estimating Latent Total Consumption in a Household.

325 E. Røed Larsen (2002): Consumption Inequality in Norway in the 80s and 90s.

326 H.C. Bjørnland and H. Hungnes (2002): Fundamental determinants of the long run real exchange rate:The case of Norway.

327 M. Søberg (2002): A laboratory stress-test of bid, double and offer auctions.

328 M. Søberg (2002): Voting rules and endogenous trading institutions: An experimental study.

329 M. Søberg (2002): The Duhem-Quine thesis and experimental economics: A reinterpretation.

330 A. Raknerud (2002): Identification, Estimation and Testing in Panel Data Models with Attrition: The Role of the Missing at Random Assumption

331 M.W. Arneberg, J.K. Dagsvik and Z. Jia (2002): Labor Market Modeling Recognizing Latent Job Attributes and Opportunity Constraints. An Empirical Analysis of Labor Market Behavior of Eritrean Women

332 M. Greaker (2002): Eco-labels, Production Related Externalities and Trade

333 J. T. Lind (2002): Small continuous surveys and the Kalman filter

334 B. Halvorsen and T. Willumsen (2002): Willingness to Pay for Dental Fear Treatment. Is Supplying Fear Treatment Social Beneficial?

335 T. O. Thoresen (2002): Reduced Tax Progressivity in Norway in the Nineties. The Effect from Tax Changes

336 M. Søberg (2002): Price formation in monopolistic markets with endogenous diffusion of trading information: An experimental approach

337 A. Bruvoll og B.M. Larsen (2002): Greenhouse gas emissions in Norway. Do carbon taxes work?

338 B. Halvorsen and R. Nesbakken (2002): A conflict of interests in electricity taxation? A micro econometric analysis of household behaviour

339 R. Aaberge and A. Langørgen (2003): Measuring the Benefits from Public Services: The Effects of Local Government Spending on the Distribution of Income in Norway

340 H. C. Bjørnland and H. Hungnes (2003): The importance of interest rates for forecasting the exchange rate

341 A. Bruvoll, T.Fæhn and Birger Strøm (2003): Quantifying Central Hypotheses on Environmental Kuznets Curves for a Rich Economy: A Computable General Equilibrium Study

342 E. Biørn, T. Skjerpen and K.R. Wangen (2003): Parametric Aggregation of Random Coefficient Cobb-Douglas Production Functions: Evidence from Manufacturing Industries

343 B. Bye, B. Strøm and T. Åvitsland (2003): Welfare effects of VAT reforms: A general equilibrium analysis

344 J.K. Dagsvik and S. Strøm (2003): Analyzing Labor Supply Behavior with Latent Job Opportunity Sets and Institutional Choice Constraints

345 A. Raknerud, T. Skjerpen and A. Rygh Swensen (2003): A linear demand system within a Seemingly Unrelated Time Series Equation framework

346 B.M. Larsen and R.Nesbakken (2003): How to quantify household electricity end-use consumption

347 B. Halvorsen, B. M. Larsen and R. Nesbakken (2003): Possibility for hedging from price increases in residential energy demand

348 S. Johansen and A. R. Swensen (2003): More on Testing Exact Rational Expectations in Cointegrated Vector Autoregressive Models: Restricted Drift Terms

349 B. Holtsmark (2003): The Kyoto Protocol without USA and Australia - with the Russian Federation as a strategic permit seller

350 J. Larsson (2003): Testing the Multiproduct Hypothesis on Norwegian Aluminium Industry Plants

351 T. Bye (2003): On the Price and Volume Effects from Green Certificates in the Energy Market

352 E. Holmøy (2003): Aggregate Industry Behaviour in a Monopolistic Competition Model with Heterogeneous Firms

353 A. O. Ervik, E.Holmøy and T. Hægeland (2003): A Theory-Based Measure of the Output of the Education Sector

354 E. Halvorsen (2003): A Cohort Analysis of Household Saving in Norway

355 I. Aslaksen and T. Synnestvedt (2003): Corporate environmental protection under uncertainty

356 S. Glomsrød and W. Taoyuan (2003): Coal cleaning: A viable strategy for reduced carbon emissions and improved environment in China?

357 A. Bruvoll T. Bye, J. Larsson og K. Telle (2003): Technological changes in the pulp and paper industry and the role of uniform versus selective environmental policy.

358 J.K. Dagsvik, S. Strøm and Z. Jia (2003): A Stochastic Model for the Utility of Income.

359 M. Rege and K. Telle (2003): Indirect Social Sanctions from Monetarily Unaffected Strangers in a Public Good Game.

360 R. Aaberge (2003): Mean-Spread-Preserving Transformation.

361 E. Halvorsen (2003): Financial Deregulation and Household Saving. The Norwegian Experience Revisited

362 E. Røed Larsen (2003): Are Rich Countries Immune to the Resource Curse? Evidence from Norway's Management of Its Oil Riches

363 E. Røed Larsen and Dag Einar Sommervoll (2003): Rising Inequality of Housing? Evidence from Segmented Housing Price Indices

364 R. Bjørnstad and T. Skjerpen (2003): Technology, Trade and Inequality

365 A. Raknerud, D. Rønningen and T. Skjerpen (2003): A method for improved capital measurement by combining accounts and firm investment data

366 B.J. Holtsmark and K.H. Alfsen (2004): PPP-correction of the IPCC emission scenarios - does it matter?

367 R. Aaberge, U. Colombino, E. Holmøy, B. Strøm and T. Wennemo (2004): Population ageing and fiscal sustainability: An integrated micro-macro analysis of required tax changes

368 E. Røed Larsen (2004): Does the CPI Mirror Costs.of.Living? Engel's Law Suggests Not in Norway

369 T. Skjerpen (2004): The dynamic factor model revisited: the identification problem remains

370 J.K. Dagsvik and A.L. Mathiassen (2004): Agricultural Production with Uncertain Water Supply

371 M. Greaker (2004): Industrial Competitiveness and Diffusion of New Pollution Abatement Technology – a new look at the Porter-hypothesis

372 G. Børnes Ringlund, K.E. Rosendahl and T. Skjerpen (2004): Does oilrig activity react to oil price changes? An empirical investigation

373 G. Liu (2004) Estimating Energy Demand Elasticities for OECD Countries. A Dynamic Panel Data Approach

374 K. Telle and J. Larsson (2004): Do environmental regulations hamper productivity growth? How accounting for improvements of firms' environmental performance can change the conclusion

375 K.R. Wangen (2004): Some Fundamental Problems in Becker, Grossman and Murphy's Implementation of Rational Addiction Theory

376 B.J. Holtsmark and K.H. Alfsen (2004): Implementation of the Kyoto Protocol without Russian participation

377 E. Røed Larsen (2004): Escaping the Resource Curse and the Dutch Disease? When and Why Norway Caught up with and Forged ahead of Its Neughbors

378 L. Andreassen (2004): Mortality, fertility and old age care in a two-sex growth model

379 E. Lund Sagen and F. R. Aune (2004): The Future European Natural Gas Market - are lower gas prices attainable?

380 A. Langørgen and D. Rønningen (2004): Local government preferences, individual needs, and the allocation of social assistance

381 K. Telle (2004): Effects of inspections on plants' regulatory and environmental performance - evidence from Norwegian manufacturing industries

382 T. A. Galloway (2004): To What Extent Is a Transition into Employment Associated with an Exit from Poverty

383 J. F. Bjørnstad and E.Ytterstad (2004): Two-Stage Sampling from a Prediction Point of View

384 A. Bruvoll and T. Fæhn (2004): Transboundary environmental policy effects: Markets and emission leakages

385 P.V. Hansen and L. Lindholt (2004): The market power of OPEC 1973-2001

386 N. Keilman and D. Q. Pham (2004): Empirical errors and predicted errors in fertility, mortality and migration forecasts in the European Economic Area

387 G. H. Bjertnæs and T. Fæhn (2004): Energy Taxation in a Small, Open Economy: Efficiency Gains under Political Restraints

388 J.K. Dagsvik and S. Strøm (2004): Sectoral Labor Supply, Choice Restrictions and Functional Form

389 B. Halvorsen (2004): Effects of norms, warm-glow and time use on household recycling

390 I. Aslaksen and T. Synnestvedt (2004): Are the Dixit-Pindyck and the Arrow-Fisher-Henry-Hanemann Option Values Equivalent?

391 G. H. Bjønnes, D. Rime and H. O.Aa. Solheim (2004): Liquidity provision in the overnight foreign exchange market

392 T. Åvitsland and J. Aasness (2004): Combining CGE and microsimulation models: Effects on equality of VAT reforms

393 M. Greaker and Eirik. Sagen (2004): Explaining experience curves for LNG liquefaction costs: Competition matter more than learning

394 K. Telle, I. Aslaksen and T. Synnestvedt (2004): "It pays to be green" - a premature conclusion?

395 T. Harding, H. O. Aa. Solheim and A. Benedictow (2004). House ownership and taxes

396 E. Holmøy and B. Strøm (2004): The Social Cost of Government Spending in an Economy with Large Tax Distortions: A CGE Decomposition for Norway

397 T. Hægeland, O. Raaum and K.G. Salvanes (2004): Pupil achievement, school resources and family background

398 I. Aslaksen, B. Natvig and I. Nordal (2004): Environmental risk and the precautionary principle: "Late lessons from early warnings" applied to genetically modified plants

399 J. Møen (2004): When subsidized R&D-firms fail, do they still stimulate growth? Tracing knowledge by following employees across firms

400 B. Halvorsen and Runa Nesbakken (2004): Accounting for differences in choice opportunities in analyses of energy expenditure data

401 T.J. Klette and A. Raknerud (2004): Heterogeneity, productivity and selection: An empirical study of Norwegian manufacturing firms

402 R. Aaberge (2005): Asymptotic Distribution Theory of Empirical Rank-dependent Measures of Inequality

403 F.R. Aune, S. Kverndokk, L. Lindholt and K.E. Rosendahl (2005): Profitability of different instruments in international climate policies

404 Z. Jia (2005): Labor Supply of Retiring Couples and Heterogeneity in Household Decision-Making Structure

405 Z. Jia (2005): Retirement Behavior of Working Couples in Norway. A Dynamic Programming Approch

406 Z. Jia (2005): Spousal Influence on Early Retirement Behavior