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## How Stressful is Retirement? New Evidence from a Longitudinal, Fixed-effects Analysis

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# How stressful is retirement? New evidence from a longitudinal, fixed-effects analysis

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The purpose of this paper is to investigate the effect of retirement on psychological well-being. Findings from previous research in this field are inconsistent, as both positive, negative, and sometimes no effect of retirement on well-being is reported. In the paper we suggest that the divergent results may arise from the mixing of cross-sectional and longitudinal studies, problems with the size and quality of existing longitudinal data, and the statistical methods used to analyze the impact of retirement on well-being. In the paper we propose to deploy the fixed-effect estimator whichs provides consistent estimates of the effect of retirement on well-being, even when retirement is correlated with other observed and unobserved explanatory variables. Using a large (N = 4,634) and nationally representative panel data set with elderly Danish respondents, we find that retirement does not have any significant effect on well-being. When estimating separate model for men and women we find indications (p = .06) that men experience a decline in well-being as a consequence of retirement, while women are unaffected by retirement. Our findings for men would substantiate the crisis theory perspective that holds that retirement implies a loss of important social roles associated with labor market participation. Several suggestions for future research are also discussed.

Key words: Consequences of retirement, psychological well-being, gender, methodology, fixed-effect model

Word count: 5,378.

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## 1. Introduction

In the research literature on psychological well-being among elderly people, one of the pertinent research questions concerns the impact of retirement on the well-being of the individual. Does retirement affect well-being and level of stress in a positive or negative way, or is it the case that retirement does not pose any particular challenge to the individual? Whereas the literature has persistently identified significant differences in well-being among elderly people due to socio-economic, demographic, familial, and psychological factors (see Larson 1978; Lohman 1980; George 1992; Diener et al. 1999), no consensus exists on the effect of retirement on well-being (Palmore et al. 1985; Kim and Moen 2001a; 2001b; 2002). In fact, empirical studies report both positive (Ekerdt et al. 1983; Bossé et al. 1991; Thériault 1994; Midanik et al. 1995; Reitzes et al. 1996), negative (Peretti and Wilson 1975; Elwell and Maltbie-Crannell 1981; Bossé et al. 1987; Richardson and Kilty 1991), and sometimes mixed or no effects of retirement on observed well-being (George and Maddox 1977; Stull 1988; Gall et al. 1997).

The objective of this paper is to analyze the effect of retirement on psychological well-being, while at the same time trying to overcome some of the empirical and methodological limitations that we believe may have produced the hitherto contradictory findings. More specifically, we aim to advance existing research by dealing with three important limitations in the literature concerning (1) the type of data used, (2) the quality of this data, and (3) the statistical methods used to analyze the impact of retirement on well-being. By doing so we aim to shed more light on the complex process of retirement adaptation and presenting an empirically consistent estimate of the effect of retirement on well-being

First, an important reason why findings in the literature on the effects of retirement on well-being are sometimes contradictory has to do with the type of data deployed. As has been argued by several scholars (e.g. Palmore et al. 1984; Kim and Moen 2002), cross-sectional data which is often used in this type of research really cannot answer the question of whether the *transition* from work to retirement, when controlling for relevant demographic, socioeconomic, and familial variables, affects well-being or not. Consequently, while cross-sectional studies may identify statistically significant differences in well-being when comparing retirees and non-retirees, then these studies cannot disclose if the observed differences in well-being are attributable to retiring or some other observed or unobserved characteristics of the respondents. To answer this question longitudinal data is required.

However, even when only considering findings from longitudinal studies, the evidence on the effect of retirement on well-being is still mixed. For example, Streib and Schneider (1971), Thériault (1994) and Reitzes et al. (1996) all find moderate positive effects of retirement on well-being (Kim and Moen 2002 find

positive effects for men only), whereas Ekerdt et al. (1987) find a negative effect. In contrast, Palmore et al. (1984), using several data sets, report both positive and negative outcomes of retirement on well-being, while George and Maddox (1977), and Gall et al. (1997) do not find any significant effects of retirement. This points to the second limitation in existing research: the quality of the data used. Among the existing studies that take a longitudinal approach, available data sets containing elderly people tend to be quite small (the sample size is typically below 800 observations) and subject to considerable attrition in panels over time. Furthermore, in most cases a non-random sampling procedure has been used to select respondents (and sometimes only men are included), and the extent to which empirical findings may be generalised to a larger population essentially remains unknown (see appendix table 1 for a summary of methods and findings in major longitudinal studies). Together, problems with the quality of longitudinal data may also account for the diverging results reported in the literature.

Finally, most existing studies do not pay explicit attention to the methodological caveats inherent in attempting to estimate the effects of retirement on well-being with panel data. Notably, selection bias with respect to retirement represents an important source of potentially erroneous results. This bias occurs because the probability of retiring between two waves in a panel depends on a number of factors, e.g. socio-economic position, health, and gender (see Kolodinsky and Avery 1996; Kolodinsky 1997). When this relationship between the retirement transition and other observed and unobserved variables is not explicitly dealt with, estimates of the effect of retirement on well-being may become significantly biased.

To deal with the first and second limitations of existing research, the type and quality of data, we use a large and nationally representative longitudinal data set (N = 4,634) with only very limited attrition to the panel over time. Furthermore, in terms of methodology we suggest to use the fixed-effect estimator whichs provides consistent estimates of the effect of retirement on well-being even when retirement is correlated with other observed and unobserved variables. However, we would like to state that the objective of the paper is not methodological, but rather on estimating the effect of retirement consistently. Given the ambivalence of this issue in the existing research literature, we believe that our focus in this paper on estimating a single coefficient of the effect of retirement in a consistent way is justified.

In the following section we briefly review the main theoretical arguments in the literature on retirement and well-being. Section 3 presents the longitudinal data set used in the paper and discusses the variables included in the analysis. Section 4 describes the fixed-effects model to be deployed, and in section 5 we present the results of the empirical analyses. Finally, in section 6 we discuss the results.

**2. Theoretical background**

Essentially, two competing perspectives prevail in the literature on retirement adaptation: role theory and continuity theory. Briefly stated, from the perspective of *role theory* (Miller 1965; George 1993) retirement may render individuals vulnerable because leaving the labor market undermines the social role associated with being employed. Consequently, role theory views work and employment relations as an important source of identity, and loss of this role might have negative consequences on the well-being of the individual. *Continuity theory*, on the other hand, emphasizes that individuals tend to preserve their social roles, lifestyles, and values even when going into retirement (Atchley 1976; 1985; 1993). Hence, it does not automatically follow that retirement affects individuals in a negative way, but rather that one would expect largely a status quo in well-being.

While both theoretical perspectives have their merits, they clearly lack contextualisation. As has been argued in recent work (Moen 1996; 1998; Kim and Moen 2001b; 2002), retirement should be seen as one of several transitions in the *life-course* of individuals that is embedded in historical, social, and personal contexts. That is, deciding to retire and the subsequent process of psychological adaptation to retirement are affected by macro-social phenomena (e.g. how pension systems operate and ruling norms in society concerning the "appropriate" timing of retirement; see Moen et al. 1992; Han and Moen 1999), as well as the employment patterns of spouse and other family members. Taken together, theory states that retirement should be viewed as a gradual process in which the individual, interacting with societal norms, may experience rupture or continuity in well-being as a consequence of retirement.

**3. Data and variables**

Data for this study come from the Longitudinal Study of Elderly People (LSEP) (Platz 2000; 2003). The LSEP consists of a representative sample of approximately 5,800 elderly people in Denmark drawn randomly from the 1920, 1925, 1930, 1935, 1940, and 1945 cohorts. The size of each cohort was sampled as to reflect the relative proportion of the cohort in the Danish population as a whole. The first wave of the survey was conducted in 1997 when respondents had their 52-77th birthday (the response rate in the first wave was 70 percent), with a follow up of the original panel carried out in late 2002. In this study, only respondents who participated in both panels are included, yielding an effective sample of 4,636 respondents. Attrition to the sample is very low, as 88 percent of the original 1997-respondents were re-interviewed in 2002. Hence, the LSEP sample is considerably larger than most other data sets used in this type of research and attrition is very low by comparative standards. In terms of contents, the LSEP contains rich information on respondents' labor market careers, family situation, health, social networks, and other issues.

*Dependent variable: Well-being*

Many different empirical applications of the concept of "well-being" have been deployed in the literature, tapping into different aspects of the general notion of well-being (Gerson 1976; Schuessler and Fisher 1985; Ryff 1989; Ryff and Keyes 1995; Higgs et al. 2003). Unfortunately, since the LSEP was not specifically designed for research into well-being, it is not possible to completely reconstruct any of the most commonly used scales in the literature, such as the Life Satisfaction Scale (Neugarten et al. 1961) or the Centre for Epidemiologic Studies-Depression Scale (CES-D) (Radloff 1977). However, the LSEP does contain several items resembling those of the CES-D.

In this paper we use as the dependent variable the factor scores derived from a factor analysis of respondents' answers in 1997 and 2002 to five Likert-type items dealing with well-being and depression. The five items measure the degree to which the respondent was 1) feeling in good form, 2) afraid of certain things, 3) worried, 4) depressed, and 5) feeling lonely.[1] The estimated factor loadings of each of the five items are shown in table 1.

Table 1. Factor loadings for latent scales of well-being

| Year | 1997 | 2002 |
|---|---|---|
| Eigenvalue $\lambda$ | 2.18 (43.6%) | 2.31 (46.1%) |
| Item: | | |
| 1. Feeling in good form | .481 | .492 |
| 2. Afraid of certain things | .494 | .586 |
| 3. Worried | .704 | .715 |
| 4. Depressed | .843 | .864 |
| 5. Lonely | .609 | .655 |
| Comparative fit index | .994 | .959 |
| Tucker-Lewis Index | .992 | .951 |
| Root Mean Square Error Of Approximation | .037 | .099 |

Extraction method: Weighted least squares with robust standard errors. Variance of factors fixed at 1 for identification.

First, as may be discerned from the table, the latent factor of well-being explains 43.6 and 46.1 percent of the variance among the five items in 1997 and 2002. Second, the fit indices reported indicate that the factor models have an excellent fit to the data (Bentler and Bonett 1980; Bentler 1990). Finally, we find that the factor loadings are remarkably stable over time (with slightly higher factor loadings in all

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[1] The available response categories were 1) often, 2) sometimes, 3) rarely, and 4) never. For item 1 ("Feeling in good form") the coding scheme was reversed such that, for all five items, increasing values indicate a higher level of well-being. The factor analysis was based on the polychoric correlation matrix of the items, thereby taking into account the fact that the observed items are categorical variables assumed to represent continuous latent variables.

items in 2002), indicating that the latent variable capturing well-being is clearly identified both in 1997 and 2002. In terms of interpretation, all five items display substantial positive loadings on the factor indicating that respondents who report rarely or never being afraid, worried, depressed, and lonely have a high score on the factor. Additionally, we find a zero-order correlation of .487 (with p < .01) between the factor scores obtained in 1997 and 2002. This suggests that respondents' level of well-being is fairly stable over time (correlation is comparable to that found in Costa Jr. et al. 1987; Gall et al. 1997; Kim and Moen 2002), but also that some change has occurred between the two waves.

*Independent variables*
The literature on well-being has found a wide range of demographic, economic, familial, and social-network-related variables to be significant in predicting level of well-being in elderly people (see Herzog and Rodgers 1981; Kim and Moen 2001a; 2001b). In our analysis we include a number of these factors as control variables when analysing the effects of retirement on well-being. The means and standard deviations of the variables used in the analysis are shown in table 2.

Table 2. Means and standard deviations (SD) of variables used in the analysis

| Variable | 1997 | | 2002 | |
|---|---|---|---|---|
| | Mean | SD | Mean | SD |
| Retired | 0.48 | 0.50 | 0.66 | 0.47 |
| Non-retired | 0.52 | 0.50 | 0.34 | 0.47 |
| Retired betw. 1997 and 2002 | - | - | 0.18 | 0.38 |
| Age | 61.29 | 8.27 | 66.29 | 8.27 |
| Age at retirement | 58.61 | 7.19 | 58.65 | 8.27 |
| Time since retirement in years | 8.08 | 6.48 | 8.93 | 13.85 |
| Spouse retired betw. 1997 and 2002 | - | - | 0.13 | 0.34 |
| Gender (=male) | - | - | 0.47 | 0.50 |
| Subjective health | 4.00 | 0.97 | 3.84 | 0.94 |
| Physical health | 0.94 | 2.07 | 1.24 | 2.60 |
| Cognitive capability | 2.44 | 2.42 | 2.44 | 2.40 |
| Divorced | 0.11 | 0.31 | 0.10 | 0.30 |
| Widow | 0.14 | 0.34 | 0.19 | 0.39 |
| Never married | 0.05 | 0.21 | 0.05 | 0.21 |
| Married | 0.70 | 0.46 | 0.66 | 0.47 |
| Moved into sheltered/nursing home | - | - | 0.01 | 0.12 |
| Contact w. grandchildren | 2.74 | 1.36 | 3.07 | 1.31 |
| Contact w. other family | 4.04 | 0.95 | 4.07 | 1.00 |
| Contact w. friends | 3.60 | 0.77 | 3.59 | 0.80 |
| Involuntarily alone | 1.47 | 0.82 | 1.45 | 0.83 |

*Work/retirement status and transition*. In 1997 48 percent of the sample was retired. In 2002, some five years later, this figure has risen to 66 percent, reflecting

the ageing of the panel and the gradual withdrawal from the labor market. The definition of retirement used in the study is being active (part- or full-time) in the labor market in 1997 and having fully withdrawn from the labor market in 2002. In line with previous studies (Gall et al. 1997; Kim and Moen 2002), we define three subpopulations in the data: (a) respondent who retired between 1997 and 2002 (18 percent of the sample), (b) respondents who were continuously retired both in 1997 and 2002 (47 percent), and finally (c) respondents who were working both in 1997 and 2002 and not retired throughout the period (35 percent).

Furthermore, several other factors relating to the retirement transition are included. First, age at retirement and the length of the retirement spell are included as controls. As has been argued theoretically (Atchley 1976; 1982), and to some extent verified empirically (e.g. George and Maddox 1977; Ekerdt et al. 1985; Gall et al. 1997), retirees may experience a "honeymoon phase" following retirement during which well-being increases for a shorter period of time, and after which it tends to decrease and stabilise. Thus, in predicting well-being, the length of the retirement period constitutes an important control variable. Second, several authors have argued that timing of, and adaptation to retirement in many cases is a 'couple phenomenon' (Smith and Moen 1998; Moen et al. 2001; Szinovacz 2002; Pienta and Hayward 2002), and that the labor market careers of partners should be analyzed together. Accordingly, in the empirical analysis we include a dummy variable indicating if the repondent's spouse (also) retired between the two waves.[2]

Additional variables included in the analysis are gender, subjective and objective health condition, and cognitive ability. First, earlier research has found gender to be a central variable in explaining the effects of retirement on well-being. Men and women have qualitatively different labor market careers, and consequently they are also very likely to experience the retirement transition in qualitatively different ways (Matthews and Brown 1987; Szinovacz and Washo 1992; Calasanti 1996; Moen 1996; Kim and Moen 2002). With respect to preparation for and adaptation to retirement, studies find that women tend to prepare less for, and adjust more poorly to retirement than men (Krause 1991; Kim and Moen 2001b). Consequently, to control for gender effects we carry out the empirical analyses both on a pooled data set as well as separately for men and women.

Second, bad health has been found to have significant negative impact on well-being among elderly people (Bossé et al. 1987; Dorfman 1995; Midanik et al. 1995; Bennett 1998; Schulz et al. 1998; Dwyer and Mitchell 1999). In the analysis we include, first, an index on physical mobility comprised of 7 items with higher

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[2] Unfortunately, in the LSEP we have no information on the timing of retirement of the respondent's spouse.

values indicating poorer health (α in 1997 and 2002 = 0.82/0.88).[3] Second, we include a single-item variable measuring repondents' self-rated health condition on a 5-point scale. The available response categories are "very poor" (1), "poor" (2), "acceptable" (3), "good" (4), and "very good" (5). Finally, cognitive ability is evaluated by index variable made up of 8 items (α = 0.79/0.80) that measures, among other aspects, the ability to remember events, dates, and short- and long-term memory in general. Higher values indicate more problems with memory and cognitive skills.

Familial and social characteristics among elderly people have also been found to influence well-being. First, marital status has consistently been shown to be related to well-being among elderly people (Ferraro and Wan 1986; Barer 1994; Hilbourne 1999; Moen et al. 2001), as well as in the general population (Haring-Hidore et al. 1985; Kurdek 1991). Studies find that married or cohabitating people display higher levels of well-being compared to divorcees, widowers and people who never married.[4] In our analysis we control for the effect of changing marital status from 1997 to 2002 with respects to 3 situations: (1) if respondents have been divorced, (2) (re)married, or (3) widowed. Scarce evidence exists on situations 1 and 2, whereas it is well-documented that widowhood or loss of significant family members or friends has a negative impact on well-being (Thompson et al. 1984; Lubben 1988; Morgan 1989; Umberson et al. 1992; Bennett 1998; van Baarsen et al. 1999). Finally, we control for the effect on well-being of moving from one's own home and into a sheltered or nursing home.[5] As reported in e.g. Evans et al. (2000; 2002), elderly people's sense of 'belongingness' to their place of residence is significantly related to well-being.

The last set of control variables deal with repondents' social networks. Previous research has found a positive association between extensive family and friendship networks and well-being (Jerrome 1981; Thomas et al. 1985; Morgan 1989; Bar-

[3] The 7 items deal with the extent to which respondents have problems with (1) cooking, (2) buying and transporting groceries, (3) doing laundry, (4) heavy domestic work (e.g. cleaning, vacuuming), (5) cutting one's toe nails, (6) walking on stairs, and (6) walking outside. In each item we assign a "0" if the respondent reports having no problems with performing the given task, a "1" if the respondent is able to perform the task alone but with difficulty, and a "2" if the respondent cannot perform the task without assistance. The index then summarizes the respondent's total score.
[4] Furthermore, recent studies have shown that not only marital status, but also the quality of marriage affects well-being (Cliff 1993; Ward 1993; Booth and Johnson 1994; Kulik 1999; Moen et al. 2001). However, as the LSEP does not tap into this issue we cannot control for changes in marital quality over time.
[5] In Denmark, shelted housing functions as an alternative to nursing homes for elderly people with special care needs. In practice, residents live in rented homes (apartments or small houses) to which public care staff is associated. Eligibility requirements are the same as for nursing homes, i.e. health problems and substantive need for personal care.

Tur and Levy-Shiff 1998) and lower mortality (Seeman et al. 1987; Hanson et al. 1989; Jylhä and Aro 1989; Olsen et al. 1991). In the analysis we control for changes between 1997 and 2002 in the extent to which respondents report having contact with (1) grandchildren, (2) family members, and (3) friends and relatives. In all three social network variables the response categories range from "does not have any" (1), "does not have contact" (2), "less than once per month" (3), "one or several times per month" (4), and "one or several times per week" (5). Additionally, we include a variable that captures if the respondent is sometimes alone even when wanting to be with somebody else. This variable measures involuntary loneliness on a 4-point scale with the outcomes "never" (1), "rarely" (2), "sometimes" (3), and "often" (4).

While our control variables embrace a range of important issues, then some limitations in the LSEP database should be mentioned. First, we only have information on respondents' incomes in 1997 (and not in 2002), thus not allowing for an empirical test of the effect of changes in income on well-being.[6] Previous studies suggests that the economic situation surronding retirement is significant in explaining adaptation and well-being (Fillenbaum et al. 1985; Leon 1985; Morgan 1992; George 1992; Holden and Hsiang-Hui 1996). However, Arendt (2003), using the 1997 wave of LSEP and a different measure of well-being than the one used in this analysis, finds only a weak effect of income on well-being. This suggests that income may not be a particularly important predictor of well-being in this data set, although we cannot estimate the effect of *changes* in income on well-being. Furthermore, studies suggest that other socio-economic factors such as social class and level of education are significant predictors of well-being among elderly people (Kessler and Cleary 1980; Witter et al. 1984; Dahl and Birkelund 1997). With the fixed-effect framework used in the paper, we cannot estimate the effect of social class and level of education (since these variables are time-invariant). However, initial analyses (not shown) regressing a social class variable (using the class scheme from Erikson and Goldthorpe 1992) and a 5-fold ordered categorical classification of level of education on well-being indicates that these variables are not significant when other control variables are included.

## 4. Methodological considerations
In this section we present the statistical model for the analysis of change in well-being from retirement, based on the two panels of interviews. As stated in the introductory section, the major problem in analyzing the effect from retirement on well-being in this context is that we suspect a number of factors (e.g. health) to influence both selection into retirement and well-being. If some of these factors are not observed, we cannot condition on these factors in an OLS panel regression

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[6] The LSEP was designed to integrate survey data with administrative registers, in which information on income is available. Unfortunately, so far data on income is only available for the 1997-wave.

framework. Failing to condition on these factors might induce a spurious correlation between retirement and well-being. To our knowledge, none of the existing longitudinal studies that use the lagged endogenous variable model deal explicitly with selection effects and other sources of unobserved heterogeneity (e.g. George and Maddox 1977; Reitzes et al. 1996; Gall et al. 1997; Kim and Moen 2002).

The consequence of this correlation is that the true and estimated effect of retirement on well-being will differ, see Woolridge (2002). To take this potential source of bias into account, we take advantage of the panel structure of the data and use the fixed-effect estimator to remove any dependency between the error term (unobserved variables) and the event of retirement. Then, our model is

$$
\begin{aligned}
y_{it} &= \beta x_{it} + \gamma_1 d_{it} + \gamma_2 age_{it} + \gamma_3 age_i^d d_{it} + \gamma_4 \left(age_{it} - age_i^d\right) d_{it} + \alpha_i + \varepsilon_{it} \\
&= \beta x_{it} + \gamma_1 d_{it} + \gamma_2 age_{it} + \gamma_3 age_i^d d_{it} + \gamma_4 s_{it}^d d_{it} + \alpha_i + \varepsilon_{it}; t = 1, 2
\end{aligned}
$$

where $y_{it}$ is the measure of well-being for the $i$'th respondent at time $t$, $x_{it}$ is a vector of time dependent explanatory variables (health, social networks etc.), $age_{it}$ is the age of the respondent at time $t$, $age_i^d$ is the age when the respondent retired (if she or he did during the sample period), $s_i^d = age_{it} - age_i^d$, $d_{it}$ is a binary indicator of retirement. It is equal to one if the respondent is retired at time $t$ and zero otherwise. $\beta$ and $\gamma_j$; $j = 1, 2, 3, 4$ are parameters of interest to be estimated. In particular, $\gamma_1$ measures the effect of retirement on well-being, $\gamma_2$ measures the effect of age on well-being, $\gamma_3$ captures a potential age dependent effect from retirement on well-being, and finally $\gamma_4$ captures duration effects from retirement on well-being (e.g. the "honey moon" effect). $\varepsilon_{it}$ is an idiosyncratic error term and $\alpha_i$ is an unobserved individual effect absorbing all unobserved time-invariant variables. These might be correlated with both $x_{it}, d_{it}$ and $\varepsilon_{it}, t = 1, 2$. The reason that we can allow this type of flexibility is that the fixed-effect estimator only uses differences in time-dependent variables. Hence, $\alpha_i$ disappears out of the estimation problem. More specifically, our estimator of $\beta$ and $\gamma_j$; $j = 1, 2, 3, 4$ is derived as

$$
\begin{aligned}
y_{i2} - y_{i1} &= \left(\beta x_{i2} + \gamma_1 d_{i2} + \gamma_2 age_{i2} + \gamma_3 age_i^d d_{i2} + \gamma_4 s_{2i}^d d_{i2} + \alpha_i + \varepsilon_{i2}\right) \\
&\quad - \left(\beta x_{i1} + \gamma_1 d_{i1} + \gamma_2 age_{i1} + \gamma_3 age_i^d d_{i1} + \gamma_4 s_{1i}^d d_{i1} + \alpha_i + \varepsilon_{i1}\right) \Rightarrow
\end{aligned}
$$

$$
\Delta y_i = \beta \Delta x_i + \gamma_1 \Delta d_i + \tilde{\gamma}_2 + \gamma_3 age_i^d \Delta d_i + \gamma_4 \left(s_{2i}^d d_{i2} - s_{1i}^d d_{i1}\right) + \Delta \varepsilon_i \Rightarrow
$$

$$
\Delta y_i = \beta \Delta x_i + \gamma_1 \Delta d_i + \tilde{\gamma}_2 + \gamma_3 age_i^d \Delta d_i + \gamma_4 \Delta s_i^d d_{i2} + \Delta \varepsilon_i \Rightarrow
$$

$$
\left(\tilde{\gamma}_2 \quad \hat{\beta} \quad \hat{\gamma} \quad \hat{\gamma}_3 \quad \hat{\gamma}_4\right)' = \left(z'z\right)^{-1} z'\Delta y
$$

where

$$z=\begin{pmatrix} 1 & \Delta x & \Delta d & age^d\Delta d & \Delta s^d d_2 \end{pmatrix}',1=(1,...,1)',\Delta x'=(\Delta x_1,..,\Delta x_n),\Delta d'=(\Delta d_1,..,\Delta d_n),$$

$$\left(age^d\Delta d\right)'=(age_1^d\Delta d_1,...,age_n^d\Delta d_n),\left(\Delta s^d d_2\right)'=(\Delta s_1^d d_{12},..,\Delta s_n^d d_{n2}),\tilde{\gamma}_2=\gamma_2\Delta age_i=\gamma_2\Delta age,$$

where $\Delta age_i = \Delta age$ is constant by individuals, as all respondents were interviewed at the same time (in 1997 and 2002), and finally $n$ is the sample size. From this we find that $\hat{\gamma}_1$ is a consistent estimate of the effect of retirement on well-being when any spurious correlation between unobserved variables and the event of retirement has been taken into account.[7] As has been stated previously, the drawback of the fixed-effect estimator is its inefficiency in that it does not allow for estimation of the effect of time-invariant variables.

## 5. Results
In this section we present the findings from the empirical analysis of the effect of retirement on well-being. In table 3 below, results from the fixed-effect panel regression are shown. Models were estimated for the entire sample as well as separately for men and women. For comparison, we also present results from the lagged endogenous variable model, as this model is commonly used in the literature.

Beginning with the fixed-effects model for the entire sample $I_{FE}$, we find that the variable of primary interest, the average effect of retiring between 1997 and 2002 on change in well-being, is estimated at .03 and insignificant. This finding suggests that, in the entire sample, retirement does not seem to have any impact on well-being.[8] Among the other variables of primary interest in $I_{FE}$ relating to the retirement transition, $\hat{\gamma}_{3-4,}$ none are significant. Turning towards the control variables, our findings mostly replicate existing findings. Improved subjective

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[7] Note finally that we cannot estimate the effect of age on well-being directly but must retrive it from the estimate on $\gamma_2\Delta age$, where $\Delta age$ is the common time-span between the two panels of interview (i.e. 5 years).

[8] To elaborate this finding, we tested whether the type of occupation from which the respondent retired was significant in predicting adaptation. This was done done by including dummy variables in the model indicating interaction between type of occupation held in 1997 (unskilled worker, skilled worker, self-employed, lower level professional, higher level professional; omitting one category for reference) and having retired between 1997 and 2002. This analysis revealed no significant results and is not reported here.

health between 1997 and 2002 is significantly associated with higher well-being, whereas increased problems with physical health and cognitive ability reduces well-being. Changes in marital status are not significant in our analysis, but moving from one's own home to a sheltered or nursing home significantly reduces well-being. Finally, while none of the social network variables are significant, then increases in involuntary loneliness are a highly significant predictor of reduced well-being.

Results from the subsamples of men and women $II_{FE}$ and $III_{FE}$ reveal interesting differences. For men, we find a strong negative, albeit just insignificant effect of retirement on well-being (estimate -1.07, p = .06), whereas for women the estimate is .00 and insignificant. Our analysis, consistent with crisis theory, then suggests that in this sample men would tend to experience a deleterious effect of retirement on their well-being. This finding is opposite to that reported in Kim and Moen (2002), in which men were found to gain in well-being as a consequence of retirement, whereas for women our finding of no effect of retirement is similar. Furthermore, our results for women is contrary to those found in other studies in which women are found to have a poorer adjustment to retirement than men (Szinovacz and Washo 1992; Kim and Moen 2001b).

Table 3. OLS panel and fixed-effect regression of well-being on retirement transition and control variables. Standard errors in parentheses

| Model | $I_{FE}$ | $II_{FE}$ | $III_{FE}$ | $I_{PA}$ | $II_{PA}$ | $III_{PA}$ |
|---|---|---|---|---|---|---|
| Variable | Total sample | Men | Women | Total sample | Men | Women |
| Retired b. 1997 and 2002 ($\gamma_1$) | .03 (.03) | -1.07 (.57)#b | .00 (.05) | -.98 (.33)** | -1.36 (.48)** | .00 (.04) |
| Retired 1997 and 2002 | - | - | - | -.03 (.02) | -.04 (.04) | -.04 (.03) |
| Age ($\gamma_2$) | - | - | - | -.00 (.00) | -.00 (.00) | -.00 (.00) |
| Age at retirement ($\gamma_3$) | .01 (.01) | .01 (.01)#b | .00 (.01) | .02 (.01)** | .02 (.01)** | .01 (.00) |
| Time since retirement ($\gamma_4$) | .00 (.02) | .04 (.03) | -.03 (.02) | -.00 (.02) | .01 (.02) | -.02 (.02) |
| Spouse retired b. 1997 and 2002 | -.03 (.04) | -.04 (.05) | -.02 (.05) | -.00 (.03) | -.06 (.05) | .00 (.04) |
| Control variables: | | | | | | |
| Gender (= male) | - | - | - | .18 (.02)*** | - | - |
| Well-being | - | - | - | .37 (.01)*** | .34 (.02)*** | .40 (.02)*** |
| Subjective health | .09 (.01)*** | .12 (.02)*** | .07 (.02)*** | .07 (.01)*** | .06 (.02)** | .08 (.02)*** |
| Physical mobility | -.04 (.01)*** | -.04 (.01)*** | -.04 (.01)*** | .00 (.01) | .01 (.01) | .00 (.01) |
| Cognitive ability | -.05 (.01)*** | -.05 (.01)*** | -.05 (.01)*** | -.02 (.00)*** | -.02 (.01)*** | -.02 (.01)** |
| Divorced | -.06 (.13) | -.08 (.21) | -.06 (.17) | -.08 (.03)* | -.14 (.05)** | -.02 (.05) |
| Widowed | .01 (.05) | .06 (.13) | .04 (.06) | .02 (.03) | -.07 (.07) | .07 (.04)* |
| Never married | - | - | - | -.05 (.05) | -.05 (.07) | -.06 (.07) |
| (Re)Married | .15 (.10) | .06 (.13) | .27 (.15)# | 0a | 0a | 0a |
| Moved into sheltered/nursing | -.19 (.10)* | -.33 (.16)* | -.09 (.13) | - | - | - |
| Contact w. grandchildren | -.02 (.01) | -.03 (.01)* | -.00 (.02) | .02 (.01) | .02 (.01)# | -.00 (.01) |
| Contact w. other family | .01 (.01) | .00 (.02) | -.01 (.01) | .01 (.01) | .01 (.02) | .01 (.01) |
| Contact w. friends and relatives | .00 (.01) | -.01 (.02) | .01 (.02) | -.01 (.01) | -.01 (.02) | -.01 (.02) |
| Involuntarily alone | -.17 (.01)*** | -.19 (.02)*** | -.16 (.02)*** | -.08 (.01)*** | -.07 (.02)** | -.09 (.02)*** |
| Intercept | -.00 (.01) | -.01 (.02) | .01 (.02) | -.23 (.06)*** | -.04 (.09) | -.27 (.08)*** |
| $R^2$ | c | c | c | .27 | .21 | .28 |
| N | 4,531 | 2,145 | 2,384 | 4,545 | 2,149 | 2,395 |

*** $p < 0.001$, ** $p < 0.01$, * $p < 0.05$, # $p < 0.10$. Coefficients for non-significant variables prior to removal shown. a reference group, b p = 0.06, c $R^2$ not shown as it is not comparable to other models.

Furthermore, for men age at retirement $\hat{\gamma}_3$ is estimated at .01 (p = .06) indicating that the later men retire then the negative impact of retirement is (albeit weakly) reduced. Comparing the effects of the control variables for men and women we find similar effects of health variables and involuntary loneliness, but also that only men experience the negative impact of moving into a sheltered home, whereas women gain in well-being from remarrying between 1997 and 2002.

In the OLS panel regression models $I_{PA}$, $II_{PA}$, and $III_{PA}$ results are somewhat different. The substantive statistical reasons why estimates from the OLS panel regression are inconsistent are decribed in the appendix. The major difference is that in the pooled model $I_{PA}$, we obtain a significant negative estimate of the effect of retirement of -.98 compared to an insignificant estimate of .03 in $I_{FE}$. This difference indicates that the negative effect of retirement found in the entire sample $I_{PA}$ disappears when selection effects and unobserved heterogeneity have been taken into account. Furthermore, for men the negative effect of retirement is slightly aggrevated, as the estimate of $\hat{\gamma}_1$ of -1.36 in $II_{PA}$ is higher than that of –1.07 in $II_{FE}$. For women, both $III_{PA}$ and $III_{FE}$ predict no effect of retirement on well-being.

As a further illustration of the adverse effect of selection bias and unobserved heterogeneity on estimates of the effect of retirement on well-being, in appendix table 2 we present the results from cross-sectional OLS regressions explaining well-being in 1997 and 2002 using the same data and control variables as those in the present analysis. In this case we compare levels of well-being between retired and non-retired respondents in 1997 and 2002; the only difference being the common time period of 5 years between the two waves. Both in the 1997 and 2002 cross-sections for the entire sample we find that retired respondents do not differ significantly in well-being compared to non-retired respondents. However, in the 2002 cross-section we now observe large gender differences, in that male retirees display higher well-being (.11, p < .01) compared to male non-retirees while women exhibit significantly lower well-being (-.80, p < .001). For men, this result is opposite to that found in the fixed-effect model in which men tended to experience a negative effect of retirement on well-being. Among women, it would appear from the cross-sectional analysis that retired women exhibit substantially lower well-being than non-retired respondents. These findings illustrate, first, that cross-sectional studies are highly unsuited for this type of research, and, second, that selection bias with respect to who retires produces spurious empirical results when not taken into account.

## 6. Discussion
The purpose of this paper was to investigate the effect of retirement on psychological well-being. Existing empirical findings in this field are somewhat inconsistent, as both positive, negative, and sometimes no effect of retirement on

well-being is reported. In the paper we suggest that the divergent results may arise from the mixing of cross-sectional and longitudinal studies, problems with the quality of existing longitudinal data, and finally inadequate methodologies.

In the paper we aimed to advance research by dealing with some of these empirical and methodological limitations. Using a high-quality and nationally representative panel data set with elderly Danes, we used a fixed-effects framework to investigate the effect of retirement on reported well-being, thereby taking selection effects with respect to retirement and unobserved heterogeneity into account. From the empirical analysis, we find that, net of the effect of a range of control variables, retiring does not have any significant impact on well-being. This finding is similar to that found in several previous studies (e.g. George and Maddox 1977; Gall et al. 1997). Estimating models seperately for men, we find that men seem to experience a decline in well-being as a consequence of retirement (with $p = .06$), whereas women do not seem to suffer any consequences from retirement. For men, our results substantiate the crisis theory perspective suggesting that loss of the work role impacts negatively on men's well-being. Given the fact that in the LSEP cohorts male respondents were comparatively more likely than female respondents to be engaged in full-time employment as "breadwinners", this interpretation seems plausible.

Several other findings deserve attention. First, our analyses indicate that empirical results are sensitive to the quality of data and the methodological approach used. As was demonstrated, the fixed-effects model is superior to the lagged endogenous variable model when selection effects into retirement and unobserved heterogeneity exist. Furthermore, we found that cross-sectional models yield unreliable results and should not be used in this type of analysis. Second, we do not find any evidence of a "honey moon" effect in adaptation to retirement, as has been identified in several previous studies (Ekerdt et al. 1985; Thériault 1994; Gall et al. 1997).[9] This may be the case because no such effect may be identified in this data, but existing studies indicating the presence of a "honey moon" effect might suffer from a bias known in labor market economics as Ashenfelter's dip or "justification bias" (Ashenfelter 1978). This bias of downwardly biased well-being occurs because respondents who are close to retiring (for known or unknown reasons) report a "falsely" low level of well-being compared to those not expecting to retire for some time. As a consequence, the observed increase in well-being following retirement interpreted as a "honey moon" effect may reflect repondents returning to their 'true' level of well-being rather than experiencing a positive effect of retirement. Our approach does not deal with this problem, but future research into the "honey moon" effect should take this issue into consideration.

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[9] Several non-linear formulations of the "honey moon" effect were also tested empirically, e.g. quadratic and cubic effects. None were found to be significant.

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**Appendix**

*A brief description of why the lagged endogenous variable model is inconsistent.*

The OLS linear panel model for the first period is

(1) $y_{it} = \beta x_{it} + \gamma_1 d_{it} + \gamma_2 age_{it} + \gamma_3 age_i^d d_{it} + \gamma_4 s_{it}^d d_{it} + e_{it}; e_{it} = \alpha_i + \varepsilon_{it},$

where $\alpha_i$ captures unobserved time-invariant explanatory variables, and $\varepsilon_{i1}$ is an idiosyncratic error term uncorrelated with any explanatory variables and uncorrelated with past and future idiosyncratic error terms (ruling out time-dependent unobserved explanatory variables). Usually we do not want to rule out selection effects from retirement on to well being *a priori*, that is that some unobserved explanatory variables, captured by $\alpha_i$ affects both the decision to retire as well as well-being, causing spurious correlation between retirement and well-being.

The standard framework to account for selection into retirement caused by well-being is to include the lagged endogenous variable into the model for the second period:

(2) $y_{i2} = \delta y_{i1} + \beta x_{i2} + \gamma_1 d_{i2} + \gamma_2 age_{i2} + \gamma_3 age_i^d d_{i2} + \gamma_4 s_{i2}^d d_{i2} + e_{i2}; e_{i2} = \alpha_i + \varepsilon_{i2}.$

However, as $\text{cov}(y_{i1}, e_{i1}) \neq 0$, by construction, and $\text{cov}(e_{i1}, e_{i2}) \neq 0$, as both terms depend on the same unobserved effect, $(\alpha_i)$, we get $\text{cov}(y_{i1}, e_{i2}) \neq 0$, a violation of a standard requirement for OLS (on (2)) to yield unbiased estimates. This is because a regression variable, ($y_{i1}$), is correlated with the error term (through $\alpha_i$).

Note that we do not need $\alpha_i$ to be correlated with any of the other explanatory variables to achieve inconsistency of OLS on (2). The requirement for OLS on (2) to be unbiased in this case, is thus the absence of fixed effects in the models.

There are alternative formulations to (1) which allow the unobserved effect to vary by time, Lee (2002), and which also leads to a more similar models compared to the traditional approach with lagged dependent variables. One is:

(3) $y_{it} = \beta x_{it} + \gamma_1 d_{it} + \gamma_2 age_{it} + \gamma_3 age_i^d d_{it} + \gamma_4 s_{it}^d d_{it} + \alpha_{it} + e_{it}; \alpha_{it} = \rho \alpha_{it-1} + \varepsilon_{i1},$

where $\rho$ is a auto-correlation parameter. This model allows the unobserved effect to vary between time periods. By taking the quasi-difference $y_{it} - \rho y_{it-1}$ We obtain:

$$y_{it} = \rho y_{it-1} + b z_{it} - \rho b z_{it-1} + \varepsilon_{it}; z_{it} = \left( x_{it}, d_{it}, age_{it}, age_i^d d_{it}, s_{it}^d d_{it} \right); b = \left( \beta, \gamma_1, \gamma_2, \gamma_3, \gamma_4 \right)',$$

where lagged values of *y* appear on the right hand side of the equation and where the regression is on levels of *y*, not differences. But note that the unobserved effect is swept out of the regression by the quasi-difference and that the regression coefficients are not allowed to vary freely. By the transformation restrictions on the coefficients appears that yields estimates of only seven coefficients with *13* right hand variables. Breaking this restriction amounts to deviate from the transformation that sweeps out the unobserved effect, and would result in biased parameter estimates.

Appendix table 1. A summary of methods and findings in longitudinal studies of the effect of retirement on well-being

| Study | Data collected in year(s) | Sample size | Sampling procedure | Country | Gender division | Age span in sample | Definition of retirement | Definition of well-being | Impact of retirement on well-being |
|---|---|---|---|---|---|---|---|---|---|
| George and Maddox (1977) | 1960, 1966 | 467(1), 58(2) | Non-random, based on selection of a variety of occupational | US | Unclear, probably only men | Not specified | Not specified | Modified Kutner Morale Scale (Kutner et al. 1956) | None |
| Palmore et al. (1984) | Varies, uses 6 data sets (3 national and 3 local); earliest 1961, latest | Varies, smalles t 467, largest 11,153 | Varies, 3 data sets are nationally representative and 3 are not | US | Men and women | Varies | 3 definitions used: 2 based on working time and 1 subjective | Not specified | Mixed positive and negative, no conclusive results |
| Ekerdt et al (1987) | 1978, 1981: Re-interviews with men from original sample (1963) who were working | 297 | Non-random, based on health and occupational status | US | Men only | 50-70 | Not specified | Life Satisfaction Scale (Neugarten et al. 1961) | Negative but non-linear, Evidence of honey moon effect |
| Thériault (1994) | 1979-1980, 3 waves: (1) 6 months prior to retirement, (2) one month after | 39 (17 in experi men-tal + 22 in control | Non-random, based on membership of labor union | Montreal, Canada | French-Canadian male workers only | All respondent s approachin g 65 years of age at | Not specified, all subject to mandatory retirement at age 65 | (1) IPAT Anxiety Scale, (2) Life Satisfaction Index (Neugarten et al. 1961) | (1) Positive effect, (2) None |
| Reitzes et al. (1996) | 1992(1), 1994(2) | 826(1), 757(2) | Random sampling of middle-aged, working men and women | US | Men and women | 58-64 | Not specified | (1) Rosenberg's (1965) Self-esteem scale, (2) Centre for Epidemiologic Studies-Depression Scale (Radloff 1977) | (1) Small positive effect, (2) Moderate positive effect |
| Gall et al. (1997) | 1980-83: 3 waves (1) 2-4 months prior to retirement, (2) 1 year after retirement, and (3) 6-7 years | 117 | Non-random, voluntary participation and recruitment | Ontario, Canada | Men only | 61-75 | Not specified | (1) Retirement Descriptive Index (Rotter 1966), (2) Single-item: "In general, how satisfying do you find the way you're spending your life today?" w. 4 point Likert-type response | None, some indications of honeymoon effect in early retirement |
| Kim and Moen (2002) | 1994-1995 (1) 1996-1997 (2), 1998-1999 (3) | 458 | Non-random, subsample of married respondents | US | Men and women | 50-72 | Receiving pension from career employer or Social Security | (1) Philidelphia Geriatric Center Morale Scale (Lawton 1975), (2) Centre for Epidemiologic Studies-Depression Scale (Radloff 1977) | Positive effect for men, no effect for women |

Appendix table 2. OLS regression of well-being 1997 and 2002. Standard error in parentheses

| | Cross-section 1997 | | | Cross-section 2002 | | |
|---|---|---|---|---|---|---|
| Model | $I_{CS}$ | $II_{CS}$ | $III_{CS}$ | $VI_{CS}$ | $V_{CS}$ | $VI_{CS}$ |
| Variable | Total sample | Men | Women | Total sample | Men | Women |
| Retired ($\gamma_1$) | .06 (.03)#[b] | .05 (.04) | .05 (.04) | -.05 (.03) | .11 (.03)** | -.80 (.17)*** |
| Non-retired | 0[a] | 0[a] | 0[a] | 0[a] | 0[a] | 0[a] |
| Age ($\gamma_2$) | .01 (.01)*** | .01 (.00)*** | .01 (.00)*** | .01 (.00)*** | .10 (.07) | .00 (.01) |
| Age at retirement ($\gamma_3$) | .00 (.00) | .00 (.01) | -.00 (.01) | .00 (.00) | .00 (.00) | .01 (.00)*** |
| Time since retirement ($\gamma_4$) | -.00 (.00) | .00 (.01) | -.00 (.00) | .00 (.00) | .00 (.00) | .01 (.00)*** |
| Control variables: | | | | | | |
| Gender (= male) | .30 (.02)*** | - | - | .26 (.02)*** | | |
| Subjective health | .16 (.01)*** | .18 (.02)*** | .15 (.02)*** | .14 (.01)*** | .17 (.02)*** | .13 (.02)*** |
| Physical mobility | -.03 (.01)*** | -.01 (.01) | -.04 (.01)*** | -.02 (.00)*** | -.01 (.01) | -.03 (.01)*** |
| Cognitive ability | -.06 (.00)*** | -.06 (.01)*** | -.07 (.01)*** | -.07 (.00)*** | -.07 (.01)*** | -.07 (.01)*** |
| Divorced | .01 (.03) | .00 (.05) | .03 (.05) | -.04 (.03) | -.19 (.05)** | .06 (.05) |
| Widowed | .01 (.03) | -.03 (.07) | .05 (.04) | .03 (.03) | -.07 (.05) | .08 (.04)* |
| Never married | .07 (.05) | .16 (.07)* | -.00 (.08) | .01 (.05) | .05 (.07) | -.04 (.08) |
| Married | 0[a] | 0[a] | 0[a] | 0[a] | 0[a] | 0[a] |
| Contact w. grandchildren | .01 (.01) | -.00 (.01) | .02 (.01) | -.00 (.01) | .02 (.01) | -.01 (.01) |
| Contact w. other family | .00 (.01) | -.00 (.02) | .01 (.02) | .00 (.01) | -.01 (.02) | .01 (.01) |
| Contact w. friends and relatives | -.03 (.01)* | -.04 (.02)* | -.02 (.02) | -.04 (.03) | -.02 (.02) | .01 (.02) |
| Involuntarily alone | -.30 (.01)*** | -.32 (.02)*** | -.29 (.02)*** | -.30 (.01)*** | -.27 (.02)*** | -.32 (.02)*** |
| Intercept | -.80 (.12)*** | -.66 (.18)*** | -.76 (.16)*** | -.97 (.11)*** | -.02 (.08) | -.05 (.09) |
| $R^2$ | .28 | .24 | .27 | .29 | .23 | .29 |
| N | 4,280 | 2,143 | 2,137 | 4,521 | 2,143 | 2,378 |

*** $p < .001$, ** $p < .01$, * $p < .05$, # $p < .1$. Coefficients for non-significant variables prior to removal shown. [a] reference group, [b] $p = .07$.