THE EFFECT OF UNPLANNED CHANGES IN MARITAL AND DISABILITY STATUS: INTERRUPTED TRAJECTORIES AND LABOR FORCE PARTICIPATION

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ABSTRACT

This study explored the effect of unplanned changes in disability and marital status on labor force participation for a sample of just under six thousand men and women born between 1931 and 1941. It was based on wave 1 (1992) through wave 4 (1998) of the Health and Retirement Study (HRS) data. Binomial hierarchical linear models were used to evaluate the change in the probability of working. Unplanned changes in disability and marital status had effects on labor force participation over and above the effects of the statuses themselves. These findings highlight the need for employer and government policies that minimize the stress that exists with unplanned events. Such policies might encourage higher labor force participation among workers who experience unplanned events that prompt them to exit the labor force earlier than they otherwise would have, with potentially adverse consequences for their subsequent socioeconomic status.

THE EFFECT OF UNPLANNED CHANGES

IN MARITAL AND DISABILITY STATUS:

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An increasing amount of the recent literature on retirement decisions has focused on the role of life course trajectories. By treating retirement as a transition in the life course, this research has challenged the primacy of economic factors as determinants of labor force participation and the inflexibility of retirement as a life course event. In particular, the interdependence of family and work trajectories has come to the fore in research on joint retirement (Henretta, O'Rand, and Chan 1993a; 1993b). Concurrently, the growing flexibility of the transition to retirement has been highlighted in research on post-retirement work (Hayward, Hardy, and Lui 1994; Mutchler, Burr, Pieta and Massagli 1997) and bridge jobs (Quinn, Burkhauser, and Myers 1990). This research treats diverse trajectories through work, family, and other domains as essential determinants of retirement decisions.

In this research, we focus on factors influencing retirement timing among individuals ages 51 to 61 in 1992. In particular, we discuss how unexpected events can derail apparently stable life course trajectories, leading to very different retirement timing than would otherwise be expected for a substantial minority of individuals. For the purposes of this research, we choose to focus on unplanned changes in health and marital status during the retirement planning years, when many workers will have begun to plan their age of retirement or have already retired.[1] The incidence of unplanned events often differs by race, class, and gender. For instance, research has consistently found the health of blacks to be worse than that of whites. While these unplanned events are more common for certain subpopulations, the occurrence of these events serves to

differentiate individual life courses from each other. For the purpose of this research, we focus on how the effects of these personal events, such as becoming ill or widowed, on the probability of employment differ according to race, class, and gender.

Specifically, in this research, we focus on three mechanisms of differentiation in retirement timing. Following Han and Moen (1999), we discuss institutional context, particularly eligibility for Social Security, and social heterogeneity, particularly race, gender, and education. We also discuss a third mechanism of intracohort differentiation in retirement timing, which we term patterned vulnerability. Patterned vulnerability refers to the ways in which the effects of chance events are structured by aspects of stratification, such as race, gender, and education.

Institutional Context

Institutional context refers to a wide range of economic, social, and policy factors that guide the typical life course. Historical changes in the institutional content help to explain changes in the average age of retirement and the variability in retirement age. Depending on the institutional context, some life course transitions such as retirement may be more or less tightly keyed to chronological age.

Until the 1960's, the institutional context surrounding work led to decreased variability in the average age of retirement (Kohli and Rein 1991). One major reason that retirement became more tightly keyed to chronological age was the institutionalization of a "normal" retirement age through the Social Security Act of 1935 (Guillemard and Rein 1993; George 1993; Mayer and Schoepflin 1989). An increasing number of workers, particularly men from the middle and upper middle classes, chose to retire at the age that they could begin to collect Social Security benefits. Rules surrounding private pensions, and eventually social norms, also encouraged workers to leave the labor force at a particular age. Theorists such as Mayer and Schoepflin (1989) argue

that the increasing regularity in the retirement age was part of a more general trend toward an institutional life course, in which life course transitions became increasingly dependent on welfare state policies.

Since the 1960's, there is evidence that the life course is becoming less rather than more institutionalized. The high degree of temporal regularity in age of retirement has begun to break down due to policy, labor market, and demographic factors. Policy changes include the introduction of provisions for taking early retirement benefits and the elimination of most mandatory retirement provisions. These policy changes allow workers to retire earlier or choose to remain in the labor force beyond the typical age of retirement. The changing labor market conditions have also led to a decrease in the regularity of the retirement age. The disintegration of the implicit contract between employers and employees and the shift away from manufacturing occupations feeds two countervailing trends. First, to the extent that workers have fewer economic resources due to unstable employment histories, they may need to remain in the labor force past the average age of retirement. Second, to the extent that factory closings and job loss prompt early retirement, workers may exit the labor force earlier than expected. Demographic changes, including increases in life expectancy and the resulting softening of age stereotypes, also allow workers to work past the "normal" age of retirement (Han and Moen 1999). The changing historical context has thus led to increasing variability in the age of retirement.

In this paper, we focus on Social Security eligibility as an example of the effect of institutional context. Some evidence suggests that public policies, particularly those surrounding Social Security, are particularly influential in determining the age of retirement. To the extent that factors other than Social Security eligibility are salient, it would suggest that the life course

is becoming less institutionalized, allowing other factors such as social heterogeneity and patterned vulnerability, to play a larger role. The de-institutionalization of the retirement age also raises questions about other factors that become more salient in determining the retirement age when the influence of welfare state policies decrease.

Social Heterogeneity

A second set of factors guiding retirement timing, particularly as the variability surrounding the retirement age increases, is social heterogeneity. Aspects of social heterogeneity include financial resources, religious and cultural background, and marital status. In the present analysis, we focus on five aspects of social heterogeneity: financial resources (including wage rate, pension wealth, nonhousing equity, and housing equity), occupation, race, gender, and education.

First, financial resources tend to structure the effect of institutional context on retirement decisions. For instance, Peracchi and Welch (1994) argue that increases in early retirement in recent years have been primarily among workers with low wage rates, for whom Social Security benefits replace a substantial portion of their wages. Fields and Mitchell (1984a, 1984b, 1984c) similarly argue that financial resources at the time of retirement, particularly for older married men, explain some but not all of the variation in the age of retirement. Postretirement employee benefits, such as pensions and health insurance, also tend to affect the probability that workers expect to retire at age 62 or 65 (Fronstin 1999). Thus, financial resources are one aspect of social heterogeneity

Second, occupation, particularly the complexity and physical demandingness of a job, can directly affect the retirement decision (Hayward, Grady, Hardy and Sommers 1989; Chirikos and Nestal 1989b). Occupation can also affect retirement timing in more subtle ways. The effects

of various determinants of retirement vary across occupational work contexts. For instance, workers in manufacturing occupations may be displaced due to new technology and lack the necessary skills to find alternative employment (Hayward and Hardy 1985; O'Rand and Landerman 1984). In addition, workers in some occupations are exposed to hazardous conditions or chemicals that can affect their health and disability status, and thus indirectly affect their retirement decision (Burtless 1987; Hayward, Grady, Hardy and Sommers 1989). Thus, occupation is a source of social heterogeneity that can indirectly and directly affect retirement decisions.

We also consider race as an aspect of social heterogeneity, focusing on the gap between the retirement timing of black and non-black workers. For most of the post-World War II period, the labor force participation rates of black men have been substantially lower than those of white men. While the gap between black women and white women is substantially smaller (Bound, Schoenbaum, and Waidmann 1996), the difference in the labor force participation rates of blacks has led some researchers to question the relevance of retirement for African Americans (Hayward, Friedman, and Chen 1996; Jackson and Gibson 1985). Unstable labor market histories and lack of financial resources make it necessary for many black men and women to return to the workforce past the typical age of retirement.

Gender is a fourth aspect of social heterogeneity that can affect the retirement decision. Childbearing history tends to have a stronger effect on the pension wealth and the labor force participation of women as compared to men (Quadagno 1988; O'Rand and Landerman 1984), as child care costs often preclude the mothers of young children from working full-time (Connelly 1992). In addition, Perkins (1993) argues that a worklife in sex segregated occupations and gender discrimination leave many working class women with an economically insecure

retirement. Thus, as a substantial number of studies have shown (Moen 1996a; Henretta, O'Rand and Chan 1993; Kohli and Rein 1991; DeViney and O'Rand 1988), retirement expectations and the average age of retirement differ for men and women.

The final mechanism of social heterogeneity included in this analysis is education. Education is a human capital variable that affects wage inequality. For instance, differences in job tenure and education largely account for the racial gap in wages (Tomaskovic-Devey 1993). In summary, various components of social heterogeneity, such as gender, race, and education, become increasingly important as the institutional context allows for more variation in retirement timing.

Patterned Vulnerability

As social regulation weakened, there was an increase in individuation, making the personal aspects of the life course, such as career history (see Han and Moen 1999) and unplanned changes, more important. In this paper we argue that, while these events are integral parts of the individual life course, race, class, and gender pattern the effects of these subjective changes.

Unplanned changes may lead to increased vulnerability by causing individuals to exit the labor force earlier than they otherwise would have. Access to economic and social resources, such as pension wealth and employment options, can cushion the potentially negative effects of unplanned changes. Because access to these resources differs according to race, class, and gender, unplanned changes have the potential to increase inequality. That is, while racial and ethnic minorities, individuals with lower education, and women enter the retirement planning years with relatively few resources, unplanned changes have the potential to exacerbate relative disadvantage by more strongly affecting those who are already disadvantaged.

There are a variety of individual events, including changes in living arrangements and lending large sums of money to parents or children, that affect labor force trajectories. However, in this paper, we focus on two types of unplanned changes,[2] changes in marital status and changes in health status, as illustrative examples of patterned vulnerability. We selected these events because previous literature indicates that, not only are marital status and health key determinants of retirement, but their effects also differ by race, gender, or education. Thus, we might also expect the effects of unplanned changes in marital status and health to differ by race, gender, and education.

First, previous literature indicates that married persons tend to be better off financially than unmarried persons. Old age pension schemes tend to assume that most married beneficiaries will remain married, and to make few provisions for widowhood and divorce (Meyer 1990). For instance, a large part of the gap between the retirement income of married and unmarried households stems from the structure of Social Security benefits (Hogan and Perrucci 1998). Due to the financial and social benefits of being married, unplanned changes in marital status can be expected to have adverse effects.

In addition, we can expect, based on previous literature, that the effect of changes in marital status will be patterned by gender and by social class. Historically, married women tend to retire earlier than their husbands. For many married women, the spousal benefits under Social Security exceed the retirement benefits that they would receive based on their own work history (O'Rand and Henretta 1999). Consequently, divorce and widowhood have lasting effects on the economic status of women, even among those who remarry (Holden and Kuo 1996; Holden and Smock 1991). Moreover, the effects of marriage, widowhood, and divorce for women differ depending on social class. Spousal benefits through Social Security heavily benefit upper middle

class women, whose average earnings are most likely to be less than half those of their husbands (Meyer 1996). Conversely, as Holden and Smock (1991) show, the steepest declines in financial resources for women who divorce or are widowed are those in families with the highest family income.

Changes in health status are also an example of patterned vulnerability. Due to environment and characteristics in early life, such as diet, smoking, and exercise, good health is unequally distributed in later life (O'Rand and Henretta 1999). Moreover, a large body of research (see Chirikos and Nestel 1989a; Hayward and Grady 1990) indicates that poor health predicts early exit from the labor force and that some involuntary retirements are directly due to illness.

The prevalence and effects of poor health differ across race and social class lines. The health of middle-aged blacks is worse than that of their white counterparts, and that difference widens with age. The gap in the health of blacks and whites is related to the availability of socioeconomic resources (Hayward, Crimmins, Miles, and Yang 2000). Other research (Hayward, Friedman and Chen 1996; Santiago and Muschkin 1996) has shown that the poorer health of African-Americans contributes to their higher rates of labor force withdrawal.

In this paper, we focus on patterned vulnerability as a determinant of retirement timing. Specifically, we evaluate the size and direction of the effects of unplanned changes in marital status and health. Based on previous literature and life course theory, we focus on two general hypotheses regarding the effect of unplanned changes on labor force participation. First, we can expect unplanned changes in disability and marital status to affect labor force participation, as unplanned changes represent unexpected deviations in life course trajectories. Second, we expect the magnitude of the effect of unexpected events to differ by race, gender, and education, with

the adverse effects of unplanned changes stronger for those with the fewest economic and social resources.

METHODS

This analysis is based on data from the first four waves of the Health and Retirement Study. The Health and Retirement Study follows a representative sample of non-institutionalized individuals born between 1931 and 1941 and their spouses. Beginning in 1992, this sample of respondents was interviewed every two years. This analysis draws on public release data for waves 1 through 3 (1992, 1994, and 1996) and preliminary release data for wave 4 (1998). We focus on the 5,942 respondents born between 1931 and 1941 who have valid data for all measures at all four time points.[3] Because the respondents are tracked over time, the majority of the respondents in the sample are ages 51 to 61 in 1992 but ages 57 to 67 in 1998.

Measures

In addition to measures of unplanned events, we have included a series of control variables drawn from previous studies. These include pension wealth (Blank 1999; Johnson, Sambamoorthi, and Crystal 1999), race and ethnicity (Bound, Schoenbaum, and Waidmann 1996; Burr, Massagli, Mutchler, and Pienta 1996; Gohmann 1990; Wray 1996), gender, occupation (Chirikos and Nestel 1989; Hayward and Hardy 1985; Hayward, Grady, Hardy and Sommers 1989; Stanford, Hapersett, Morton, Molgaard, and Peddecord 1991), disability status (Santiago and Muschkin 1996),[4] and Social Security eligibility.[5] Our analysis makes use of two groups of predictors.

The first group of predictors can vary for each respondent for each time point. These variables are sometimes referred to as time variant. They are treated as level 1 variables in the

discussion of HLM.[6] These variables include whether employed, age, whether 62, whether

disabled, marital status, whether disability status changed, and whether marital status changed.

Whether employed, the dependent variable, is a dichotomous variable referring to whether the

respondent is currently working for pay.[7] Age refers to age in years (and months) from age 48.

Whether eligible for Social Security is a dichotomous variable indicating whether the respondent

is at least 62 and 2 months old. Due to administrative issues, most people who intend to take

early benefits at age 62 do not begin receiving these benefits until age 62 and 2 months (Olson

1992). Also included in the first group of predictors are status variables, interruption variables,

and interaction terms. Status variables, including whether disabled and marital status, refer to the

present status of a respondent. While some researchers suggest that currently retired workers

may use poor health or disability as an excuse for not working, most research finds that self-rated

measures are strongly correlated with more objective measures of health (Dwyer and Mitchell

1999). Interruption variables, including marital change and disability change, refer to whether a

respondent has experienced a change in a status variable since the last observation. Finally,

interaction terms indicate which type change a person has undergone. We use the interruption

effect and the interaction effect to compute a net effect of a change in marital status or disability

status on the log odds of employment. We treat the interruption effect (X_{1ij}) as the independent

variable and the status effect (X_{2ij}) as the moderator variable. The interaction term is $X_{1ij}X_{2ij}$, the

product-term interaction between X_{1ij} and X_{2ij}.[8]

A second group of variables is time invariant. These will be treated as level 2 predictors.

These variables include whether female, whether black, and years of education. Whether female

and whether black are dichotomous variables. Education in years is an interval level variable,

centered at the mean. These three variables are assumed to be constant for each respondent

across the different time points. An additional six variables use only 1992 data.[9] These are logarithm of pension wealth, logarithm of wage rate, logarithm of nonhousing equity, logarithm of housing equity, whether manual occupation, and whether service occupation. The logarithms of all variables, except for the dichotomous variables, are used to reduce skewness and are centered at the mean.

Statistical Model. In this analysis, we estimate a two-level binomial hierarchical model using HLM 5.02. HLM 5.02, developed by Bryk and Raudenbusch (1992), allows for the estimation of models with dichotomous outcomes. HLM uses Bayesian estimation to compute coefficients and standard errors for each individual, using both their information and information from other cases. In addition, the HLM program uses an iterative algorithm to allow us to estimate the size and variance of error terms. In this analysis, the variables referred to above as time variant are level 1 predictors, while those referred to as time invariant are level 2 predictors.

The level 1 model includes observations for all respondents at all available time points. The dependent variable is "Whether employed," a dichotomous variable. The basic level 1 model for a dichotomous dependent variable is a logistic regression. The time-invariant variables are included as predictors of T_{1ij} (time, as measured using age) and X_{1ij} through X_{8ij}. Accordingly, the level 1 model is:

$$\ln(\frac{\text{pr } Y = 1}{\text{pr } Y \neq 1}) = \boldsymbol{b}_{0j} + \boldsymbol{b}_{1j}T_{ij} + \boldsymbol{b}_{2j}X_{1ij}....\boldsymbol{b}_{9j}X_{8ij} + v_{ij} \qquad \text{[Eq. 1]}$$

The predicted log of odds is equal to a linear combination of the independent variables, indexed by both respondent (*j*) and time (*i*). Each coefficient represents the predicted increase in the log odds of working for a one unit increase in that predictor. Taking the exponent of both sides of equation (1), we find that:

$$\frac{\text{pr } Y = 1}{\text{pr } Y \neq 1} = e^{b_{0j} + b_{1j}T_{ij} + b_{2j}X_{1ij} \dots b_{9j}X_{8ij} + v_{ij}}$$

$$= e^{b_{0j}} e^{b_{1j}T_{ij}} e^{b_{2j}X_{1ij}} \dots e^{b_{9j}X_{8ij}} e^{v_{ij}}$$

[Eq. 2]

Thus, the exponent of the level 1 coefficients is the factor by which the odds of working increase for a one unit increase in that predictor (Raudenbush, Bryk, Cheong, and Congdon, 2000).

The level 2 model is a set of regression equations predicting the level 1 coefficients. Our level 2 model has three groups of equations: a random effects equation for the constant, a random effects equation for time, and a fixed effects equation for predictors. The random effects model in equation 3 predicting the constant includes all nine time invariant predictors.

$$b_{0j} = g_{00} + g_{01}W_{1j} + g_{02}W_{2j} \dots + g_{09}W_{9j} + u_{0j}$$

[Eq. 3]

The random effects model predicting time, in equation 4, includes only race, education and gender as predictors.

$$b_{1j} = g_{10} + g_{11}W_{1j} + g_{12}W_{2j} + g_{13}W_{3j} + u_{1j}$$

[Eq. 4]

In addition, we explicitly model the error terms u_{0j} and u_{1j}. Treating the constant and time (age) as random effects models the diversity in individual trajectories through the retirement planning years. While gender, race, and education predict some of this diversity, a large proportion of the diversity in individual pathways is due to unique life events and attitudes. Thus, the error terms for the equations predicting age and the constant are explicitly modeled. Our model assumes that the error of age and the constant are meaningful, rather than due to sampling error alone.

Finally, as shown in equation 5, the coefficients for level 1 predictors X_{1j} through X_{9j} are predicted by race (W_{1j}), education (W_{2j}), and gender (W_{3j}). These equations explain in part the variation in the effects of marital status, disability status, and unplanned events.

$$b_{2j} = g_{20} + g_{21}W_{1j} + g_{22}W_{2j} + g_{23}W_{3j}$$

[Eq. 5]

For equations 4 and 5, we first test the coefficients W_{1j} through W_{3j} for significance. If the coefficient is not significant at p<.05, we omit it from the final model.

Because the level 1 coefficients (the βs) can be calculated, with the exception of random effects, based on the level 2 coefficients (the γs), we present only level 2 coefficients. However, we frequently present net coefficients based on the gamma values. For dichotomous level 2 predictors, such as men and women, we present net coefficients for each group. For interval level predictors, the level 2 constants (γ_{00}, γ_{10}, and γ_{20}) give the net effect of the coefficients (β_{oj}, β_{ij}, and β_{2j}) at the mean. In the analysis, we frequently present net coefficients (nc) at high (one standard deviation above the mean) and low (one standard deviation below the mean) values of the level 2 interval level predictors.[10] When discussing interaction terms, we also include net coefficients designated β(nc). We treat the interruption effect (X_{1ij}) as the independent variable and the status effect (X_{2ij}) as the moderator variable. The interaction term is $X_{1ij}X_{2ij}$, the product-term interaction between X_{1ij} and X_{2ij}.[11] In the analysis, we do not provide the statistical significance of net coefficients. However, all net coefficients reported are based on statistically significant gamma values.

RESULTS

Literature on retirement from a life course perspective often assumes that the majority of retirement transitions are orderly. Individuals clearly have varying trajectories through the retirement planning years. However, an approach that views most retirement transitions as orderly or continuous downplays the extent to which those trajectories can change during the retirement planning years.

The focus on established differences, such as pension wealth and wage rate, tends to de-emphasize the proportion of the retirement-age population that experiences an unplanned event.

As shown in Table 1, approximately one third of our sample experienced an unplanned marital or disability status event at some time from 1992 to 1998. The proportion of respondents reporting a disability change includes the respondents who experienced at least one disability change (recovers from a disability or gets a disability). Similarly, the proportion of respondents reporting a marital status change includes respondents who experienced at least one marital status change (gets married, gets divorced, and is widowed).[12] While a majority of respondents were not disabled at any time from 1992 to 1998, a substantial minority reported being disabled in at least one wave. While 36.7 percent of respondents were disabled in at least one wave, only 9.6 percent of the total sample was disabled in all four waves. The remainder of these respondents either recovered from a disability (15.5 percent) or got a disability (11.6 percent) during this time. Similarly, a substantial proportion of respondents experienced an unplanned change in marital status during this time. The majority of respondents (92.6 percent) had the same marital status during this time period, but substantial minorities got married (1.4 percent), divorced (1.8 percent) or were widowed (4.3 percent) at some point from 1992 to 1998. Thus, at least regarding disability and marital status, about a third of the sample have trajectories marked by unplanned events during the retirement planning years. This analysis thus distinguishes between two types of models: models of uninterrupted trajectories and models of interrupted trajectories marked by unplanned events.

Uninterrupted Trajectories through the Retirement Planning Years

While the existing retirement literature may be inadequate to model the shape of interrupted trajectories, it does provide a reasonable picture of the uninterrupted trajectories that characterize the majority of the sample. Model 1 in Table 2 assumes that all trajectories are uninterrupted. For all regression models in this analysis, education in years, whether female and

whether black were included in an initial model. Except for the constant, these level 2 predictors were retained only in cases where they were significant.

While a wide range of variables, including education, race, gender, occupation, and financial resources, guide the probability of a person working at any given time, this model assumes that unanticipated events are not a major factor. Education in years (γ=.123, p<.001) and pension wealth (γ=.025, p<.001) are positively associated with the probability of working. This suggests that people with more socioeconomic resources, in terms of education and pension wealth, are less likely to exit the labor force. Probably, this is due to a set of factors, such as higher job satisfaction and more access to stable employment. However, controlling for these resources, wage rate (γ=-.185, p<.001) is negatively associated with the probability of working. Consistent with previous research, blacks (γ=-.193, p<.01) and women (γ=-.857, p<.01) are less likely to be working than whites and men respectively. Due in part to family responsibilities, women have lower labor force participation rates throughout the life course. A combination of low education and lack of suitable employment opportunities also reduces the odds of blacks working. Finally, people in service occupations (γ=-.187, p<.01) and manual occupations (γ= -.399, p<.001) are less likely to be working than their counterparts in professional occupations. This suggests that people are more likely to remain employed if in less physically demanding and more rewarding jobs.

(Insert Table 2 about here)

In addition, men with average levels of education are less likely to work as they age (γ=-.151, p<.001). The negative effect of age is weaker for women (γ(nc)=-.126).[13] Keep in mind that throughout this article these net coefficients, indicated by the (nc) designation, are not presented in our tables. In addition, the negative effect of age is weaker for respondents with low

education (γ(nc)= -.136 for men and -.111 for women) and stronger for those with high education (γ(nc)= -.166 for men and -.141 for women).[14] Thus, while women and respondents with low education have lower probabilities of working at the beginning of the retirement planning years (age 48), their odds of working decrease less sharply with age.

The effect of whether 62 indicates that Social Security eligibility predicts a decrease in labor force participation rates (γ=-.509, p<.001). Education in years is positively associated with the net coefficient for "Whether 62" (γ=.041, p<.01); the effect of Social Security eligibility is weaker for those with high education and stronger for those with low education

Model 2 is an elaboration of model 1. It also includes whether disabled, whether married, whether divorced, and whether widowed. The results replicate what is already known about retirement in general. People with disabilities are disproportionately likely to exit the labor force at any age (γ=-1.468, p<.001), especially if they are black (γ(nc)= -1.898) or have low education (γ(nc)= -1.680).[15] Also, compared to those whose have never been married, married men are more likely to work (γ= .419, p<.01) while married women are less likely to work (γ(nc)= -.342).[16]

Models 1 and 2 are essentially models of uninterrupted trajectories. Based on a series of characteristics of an individual as they enter the retirement planning years, these models predict their probabilities of working over the retirement period. Some predictors, such as race and gender, are ascribed characteristics. Other predictors, such as wage rate and pension wealth, are relatively stable for each respondent over time.

Figure 1, based on the coefficients for model 1, shows the uninterrupted trajectories for white men, white women, black men, and black women. These trajectories are evaluated at the mean education, wage rate, pension wealth, nonhousing equity, and housing equity. In addition,

they are evaluated for the occupation reference group, professional. While there is variation in the probability of working, much of this variation is attributable to characteristics such as financial resources, education, race, and gender, which exist at the beginning of the retirement planning years. Thus, the models based on uninterrupted trajectories incorporate the roles of institutional context, such as Social Security eligibility, and aspects of social heterogeneity, such as race, education, gender, and financial resources. Additionally, although model 2 in Table 2 allows for the effect of disability and marital status, it does not model the unique effect of an unplanned event in either of these domains. According to this model, being disabled, for instance, has the same effect on labor force participation regardless of whether a person has been disabled for one year or for ten years.

Interrupted Trajectories through the Retirement Planning Years

Models 1 and 2 are consistent with previous research that assumes that the decision to retire is an orderly one for most people. However, we argue that unplanned events create interrupted trajectories and can drastically change a person's probability of retirement. Not only does an unplanned change mean that the new status (such as being disabled) affects the probability of retirement, but the stress associated with unplanned events can also affect retirement probabilities.

Model 3 in Table 3 tests the hypothesis that unplanned changes in marital status and disability status have effects on the probability of working over and above the effects of current disability or marital status. Whether marital status changed and whether disability status changed are dichotomous variables indicating whether the individual's present status differs from their status at the time of the last interview. A change in marital status does not have a significant effect on labor force participation among white males (γ= .194, p$\geq$.05). However, a change in

marital status is associated with a greater probability of working for black men ($\gamma(nc)$= .698), but a lower probability of working for white women ($\gamma(nc)$= -.547).[17] The effect is close to zero for black women ($\gamma(nc)$= -.043). A change in disability status is significantly associated with an increase in labor force participation (γ= .125, p<.01) for respondents with average levels of education. The effect is weak and negative for those with high education ($\gamma(nc)$= -.084) and stronger for those with low education ($\gamma(nc)$= .334).[18]

(Insert Table 3 about here)

While a change in disability status does affect the probability of working, the effect differs by the nature of the change itself. In model 4, for respondents with the mean level of education, the net effect of a getting a disability is weak and positive ($\gamma(nc)$= .456).[19] When this relatively weak positive effect is combined with the stronger negative effect of having a disability (γ= -1.593, p<.001), the result is a delay in the full effect of getting a disability. In the one or two years immediately following being disabled, a person is less likely to work overall. However, without the effect of the unplanned event, the negative effect of the disability status is still stronger three or four years after the person gets the disability. Possibly, many people may remain at work for a few years after getting a disability because they cannot yet afford to retire. However, even those disabled people who initially stayed at work are less likely to remain at work in the long run.

Figure 2, based on the coefficients in model 4, illustrates the effects of a change in disability status on existing trajectories. White men with the mean level of education are used as an example, as white males are the group most typically included in studies of retirement. The pattern is similar for white females, black males, and black females.[20] The solid lines represent interrupted trajectories. The dotted lines represent the uninterrupted trajectories. For men

disabled at age 60, the probability of working drops precipitously immediately after getting the disability. In the following years, the predicted rate drops when they reach the age for Social Security eligibility and again decreases more gradually from age 62 to age 68.

(Insert Figure 2 about here)

Conversely, for white men who recover from a disability at age 60, the predicted probability of working increases dramatically immediately after the recovery. Despite the small negative effect of the unplanned event itself (γ= -.168, p<.01), a person who has recovered from a disability is almost as likely to continue working as a person who has been healthy throughout the entire time period under consideration.[21] In part, this may be to compensate for lost income. Both scenarios illustrate ways that unplanned changes in disability status can derail existing trajectories and set people on very different pathways.

The delayed effect of unplanned changes in disability status suggests that institutional context plays a role in shaping the effects of unplanned changes, and that role is stronger for workers with fewer financial resources. Despite the increasing variability in the age of retirement, rules surrounding private pensions and Social Security continue to make it difficult for all but the most affluent workers to retire before their early sixties. Workers are unable to collect early Social Security benefits before age 62 or to collect full benefits before age 65. Similar minimum age requirements apply for many private pension plans. Accordingly, many workers who experience a disability change (particularly those who get a disability) may delay leaving the labor force until they can qualify for pension benefits. In addition, as the coefficients in model 3 indicate, a change in disability status is associated with a moderate positive increase in the probability of labor force participation among those with low education (γ(nc)= .334) but almost no effect among those with high education (γ(nc)= -.084). The "delayed" effect of a

disability change is more prominent among those with low education, who can least afford to

leave the labor force. Our interpretation indicates that, although retirement is becoming

deinstitionalized, institutional context is more important in guiding the retirement decisions of

those with fewer economic resources.

Unplanned changes in marital status do affect labor force participation, but the effect is

similar for the divorced and widowed. As shown in Table 3, model 4, the interaction terms

between whether marital status changed and current marital status are non-significant for both

the divorced (γ= -.133, p>.05) and the widowed (γ= .098, p>.05). However, the effect of a

marital status change is stronger for women than for men. For instance, while the effect of a

marital status change is nonsignificant for white men (γ= .230, p>.05), it is associated with a

lower probability of working for white women (γ(nc)= -.597).[22] In addition, the effects of a

change in marital status differ according to race.

Figures 3 and 4 illustrate the probability of working for widowed women, married

women, women widowed at age 55, and women married at age 58.[23] Figure 3 tracks the

pathways for white women, while Figure 4 tracks pathways for black women. First, the positive

effect of getting married during the retirement planning years appear to be substantially larger for

white women, as compared to black women. The probability of working for white women who

marry at age 58 drops substantially below that of women married throughout the retirement

planning years. The probability of labor force participation for black women who marry at age

58 drops much less steeply. This suggests that white women are more likely than black women

to marry spouses with substantial economic resources, making it feasible for them to retire. In a

parallel way, white women widowed at age 55 are approximately as likely to work as married

white women, but the probability of working for black women widowed at age 55 increases

substantially. Our interpretation is that widowed black women are less likely to have access

widow's benefits and survivor's benefits than are white women, and are correspondingly more

likely to need to return to work after the death of a spouse. Particularly for women, the adverse

effects of a change in marital status tend to be stronger for blacks, while the positive effects tend

to be stronger than whites. Unplanned changes in marital status can thus exacerbate existing

inequalities.

(Insert Figures 3 and 4 here)

Although the recentness of marital status change does not appear to play a role, it is

possible that the age of the respondent experiencing a marital status change does. Younger or

older women may be more likely to stop working when their marital status changes. Similarly, it

is possible that the timing of changes in disability status moderates the effect of the unplanned

event on the probability of working. That is, the effect of unplanned events can differ depending

on the point in the life course that they occur. Table 4 expands on model 4 by including the

interaction of age and status variables, change variables, and interactions between status and

change variable. These interaction terms evaluate whether the effect of unplanned events differ

by age.

(Insert Table 4 here)

The interaction between age and marital change variables are omitted from the model in

Table 4 because they were nonsignificant. However, we found that the interaction between

whether disability status changed and age is positive and significant (γ= .056, p<.001). The net

effect of getting a disability is stronger for older respondents. For instance, the net coefficient for

getting a disability at age 48 is .495, while the net coefficient for getting a disability at age 68 is

1.615. This indicates the delay in returning to work is a stronger factor for older workers. In a

parallel way, the net effect of recovering from a disability at age 48 is moderate and negative ($\gamma(nc)$= -.735) while at age 68 the effect is small and positive ($\gamma(nc)$= .385).[24] This indicates that those who recover from a disability at older ages experience a substantial upward jump in the probability of working after they recover.

The results for disability timing suggest that the point in the life course that a unplanned event occurs moderates the effect of that event. The time in the life course that the event occurs may be a factor for more than one reason. First, institutional factors may make returning to work less attractive at some ages than others. For instance, one reason that the delay in returning to work is stronger for older workers may be that these workers are more likely to have access to pension and Social Security income. Second, the financial and social resources that a person can acquire differ depending on when an event occurs. For instance, a person who has been disabled for a substantial amount of time (in the case of the above example, from age 48 to 68) may have fewer assets accumulated in the form of savings, pension plans, and Social Security wealth. These workers, as indicated above, may return to the labor force relatively quickly to make up for lost assets. In contrast, for workers disabled for a shorter period of time (e.g., those who recover at age 48) they have more years to make up lost assets.

These models of interrupted trajectories, in addition to incorporating aspects of social heterogeneity and institutional context, also tap into the role of patterned vulnerability. While personal events, such as type and timing of an unplanned event, serve to individuate life courses from each other, these effects are patterned by race, education, and gender.

DISCUSSION

The goal of this study was to elaborate on the concept of interrupted trajectories through the retirement planning years. We found that approximately one-third of our sample, drawn from

the Health and Retirement study, experienced an unplanned change in disability status or marital status over this period. Moreover, we found that unplanned changes in disability and marital status had effects on the probability of working over and above the effect of the status itself. In addition, the effect of unplanned changes are shaped by race, gender, and education.

The results of this study contribute to the existing literature on retirement by highlighting the ways in which aspects of social heterogeneity, such as race, gender, and education, moderate the effects of institutional context and even unplanned events on retirement timing. In a historical context in which retirement timing is becoming more flexible, attention to individual factors, such as unplanned changes in disability and marital status, helps to explain how and why individual life course trajectories diverge from each other.

Methodological issues in this analysis centered on the measurement of time and the role of cohort. First, we used data based on whether a person was working at four times, at the interviews in 1992, 1994, 1996, and 1998. Because some populations, notably blacks and low-wage workers, are known to have irregular patterns of labor force participation, an appropriate way to address this population might have focused on whether they were working, disabled, or married in a given month. We replicated the analysis using the monthly data. However, because this approach involved more missing data and because the results were not substantially different, we presented the analysis based on two-year times. Second, one limitation of this study is that, in focusing on a fine grained analysis of the probabilities of working over six years, it foregoes a broader analysis of cohort. Studies using a life course perspective frequently focus on the unique historical events shaping successive birth cohorts. Thus, while our study focuses on modeling diversity within cohorts, this approach does not allow us to evaluate the role of diversity across cohorts.

This analysis also raises theoretical issues about the interaction of institutional context, social heterogeneity, and patterned vulnerability. First, previous literature suggests that life course transitions are becoming less strongly tied to chronological age, leaving room for biographical factors to become more important. While life course trajectories may become less governed by institutional factors, it is unclear how strong the constraints imposed by social heterogeneity (in the form of race, gender, and education) are compared to the effect of biographical factors. For instance, race, gender, and education guide the probability that a person will experience an unplanned change in disability or marital status. However, the event is biographically unique, as not all persons with a given class, gender, and racial background will experience the event. As discussed above, aspects of social heterogeneity shape the long-term effects of the unplanned event. As institutional factors become less important in determining retirement timing, the relative importance of unique biographical factors and social heterogeneity should be reevaluated.

Second, we found that the delayed effect of a disability change was stronger for those with lower education. Our interpretation was that, due to their lower levels of economic resources, these respondents were more strongly affected by age eligibility rules surrounding private pensions and Social Security. Although the retirement age is becoming less tightly keyed to chronological age, the institutional life course may retain a stronger hold over racial minorities, those with low education, and women. For these groups, welfare state policies dictate the amount and timing of bulk of their retirement income. Thus, this analysis raises questions about whether the deinstitionalization of the life course is limited to certain subgroups, particularly the financially well-off.

This analysis also suggests a general strategy for old age policy. Previous research indicates that people tend to plan for retirement (Anderson, Burkhauser, and Quinn 1986; Ekerdt, DeViney, and Kosloski 1996; Hall and Johnson 1980). Retirement is institutionalized, not only in programs such as Social Security and Medicare, but also in the subjective expectations of workers. Older workers have individual timetables for when they plan to retire. In addition, many workers retire within a year of their projected retirement date (Ekerdt, Kosloski, and DeViney 2000; Ekerdt, Vinick, and Bossé 1989). This research suggests that it is unfair to many workers to institute changes in policies, such as the Social Security eligibility rules, that are not phased in very gradually. While this is the case, our results indicate that for approximately one-third of the sample, an unplanned change occurs during their fifties and early sixties. Thus, while it may not be appropriate to make drastic changes, it is possible that smaller policy changes might be implemented to help cushion the potentially negative effects of unexpected events.

This analysis maps the ways in which interrupted trajectories, marked by unplanned events, differ from uninterrupted trajectories. Future research on interrupted trajectories might address a number of issues. First, how does the social and economic status of respondents who retire due to unplanned events differ from that of respondents with uninterrupted trajectories? Second, what institutional arrangements help people who experience unplanned events stay in the labor force? For instance, policies to help workers with disabilities and employer flexibility that allows for family leave might moderate the effects of unplanned changed in disability and marital status, respectively. Because many people leave the labor force due to unplanned events, policies that help to ease the stress associated with these changes might encourage labor force participation. Higher labor force participation, in turn, means that people contribute longer both through their work and by their contributions to Social Security and Medicare. In addition, by

preventing workers who experience unplanned events from exiting the labor force earlier than they otherwise would have, these policies might minimize the potentially adverse effect of unplanned events on subsequent socioeconomic status. Thus, greater attention to interrupted trajectories may help both to ease the transition for individuals experiencing unplanned events and to encourage them to remain at work when appropriate.

ENDNOTES

[1] In this article, the term "retirement planning years" refers to the years that a worker is between ages 48 and 68. Many workers are already retired during at least some of these years and other workers may not actively "plan" for retirement. However, previous research indicates that most workers in this age group are giving some thought to retirement if they are not already retired (Anderson, Burkhauser, and Quinn 1986; Ekerdt, DeViney, and Kosloski 1996; Hall and Johnson 1980).

[2] We refer to these events as "unplanned changes" because, in most cases, people do not "plan" to have a disability or a divorce. While they may anticipate changes in health or marital status, many of these respondents would not plan for changes in health or marital status.

[3] We use respondent level weights to calculate the descriptive statistics. The respondent level weights correspond to the number of individuals in U.S. population as measured by the March CPS for that year. While the weights for a given individual do not differ substantially from year to year, we use the weights from each year to capture the effect of changing population composition. We do not, however, use weights in our hierarchical models. As Lohr (1999) discusses, weighted regression models should produce parameter estimates that are consistent with unweighted regression models, if the model is properly specified.

[4] We use as a measure of disability the respondent's self-report of whether they have a disability or health problem that interferes with the amount or type of work they can do. While some research indicates that self-reports of health are unreliable, in general people's self-reported health tends to be highly correlated with more objective measures (Waidmann, Bound, and Schoenbaum 1995; Bound 1991; Anderson and Burkhauser 1985; Chirikos and Nestel 1984; Dwyer and Mitcher 1999).

[5] The dummy variable for Social Security eligibility refers to early eligibility at age 62. We omitted a dummy variable at age 65 because in all models it was nonsignificant.

[6] Hierarchical Linear Models (HLM), also referred to as multilevel or mixed models, refers to a class of regression models that takes the nesting of observations within groups into account. For instance, individuals are nested within families and, as in this paper, time points are nested within individuals.

[7] In this paper, we use the terms employment, labor force participation, and retirement interchangeably. While we recognize that "retirement" is a more complex concept than merely being not employed, the majority of retirement transitions are still a transition from full time employment to complete exit. Complex patterns of labor force exit and reentry are also less common after the age of Social Security eligibility (O'Rand and Henretta 1999).

[8] For instance, if a respondent is not disabled in time 1, gets a disability that lasts through times 2 and 3, and recovers in time 4, the variable pattern is:

	t_1	t_2	t_3	t_4
Status: Whether disability	0	1	1	0
Interruption: Disability change	0	1	0	1
Interaction: Whether disability * Disability change	0	1	0	0

Thus, the status variable captures the effect of having a disability. The interruption effect captures the effect of a change in disability status. The interaction effect distinguishes between the effect of getting a disability and the effect of recovering from a disability.

[9] While these variables are not time invariant, they have very little within-individual variation and are treated as invariant in this analysis. Due to the low level of within-individual variation, it is not possible to estimate a model treating all of these variables as time variant.

[10] These net coefficients are calculated in the straightforward way. The level 2 equation predicting a certain coefficient might be

$$b_{zj} = g_{z0} + g_{z1}W_{1j} + g_{z2}W_{2j} + g_{z3}W_{3j} \qquad \text{[Eq. 6]}$$

The net coefficient would be based on substituting the values for W_z. These net coefficients are based on the level 2 equation and designated as $\gamma(nc)$, or γ(net coefficient). We refer to these net coefficients as γ for the sake of consistency, although the net γ for a case is equivalent to the β for that case.

[11] The logistic regression equation is thus:

$$\ln(\frac{\text{pr } Y = 1}{\text{pr } Y \neq 1}) = b_{0j} + b_{1j}X_{1ij} + b_{2j}X_{2ij} + b_{3j}X_{1ij}X_{2ij} .... b_{kij}X_{kij} + r_{ij} \qquad \text{[Eq. 7]}$$

Algebraically, the net coefficient for a given value of X_{1ij} is equal to the base coefficient β_{1ij}, plus the product of β_{3ij} and X_2ij.

$$b_{1j}(nc) = b_{1j} + b_{3j}X_{2ij} \qquad \text{[Eq. 8]}$$

Like net coefficients based on Equation 6 above, net coefficients computed using Equation 8 are referred to as $\gamma(nc)$ rather than $\beta(nc)$.

[12] It is possible for a respondent to experience more than one unplanned event over the six year observation period. Of the 1958 respondents in our sample reporting at least one unplanned event, 1815 reported one unplanned change. Of the 143 respondents reporting more than one unplanned event, 120 reported one disability status change and one marital status change. The analysis takes these multiple unplanned events into account by calculating separate coefficients for marital and disability status changes.

[13] Based on the coefficients in Model 1, the net coefficient for age is equal to:

$$g* = -.151 - .005(education) + .025(female) \qquad \text{[Eq. 9]}$$

Education is centered at zero, so that the constant -.151 represents the net coefficient for men with an average level of education:

$$g* = -.151 - .005(0) + .025(0) = -.151 \qquad \text{[Eq. 10]}$$

For women with an average level of education, the net coefficient is:

$$g* = -.151 - .005(0) + .025(1) = -.126 \qquad \text{[Eq. 11]}$$

[14] The equation to compute these net coefficients is Equation 9 above. However, the net coefficients for high and low education are evaluated at one standard deviation above (3.03) and below (-3.03) the mean, respectively. Thus, for men with low education, the net coefficient for age is -.136.

$$g* = -.151 - .005(-3.03) + .025(0) = -.136 \qquad \text{[Eq. 12]}$$

For women with low education, the net coefficient is -.111.

$$g* = -.151 - .005(-3.03) + .025(1) = -.111 \qquad \text{[Eq. 13]}$$

For men with high education, the net coefficient is -.166.

$$g* = -.151 - .005(3.03) + .025(0) = -.166 \qquad \text{[Eq. 14]}$$

For women with high education, the net coefficient is -.141

$$g* = -.151 - .005(3.03) + .025(1) = -.141 \qquad \text{[Eq. 15]}$$

[15] The net coefficient for whether disabled, in Model 2, is equal to:

$$g* = -1.468 + .070(education) - .430(black) \qquad \text{[Eq. 16]}$$

For black respondents with average education, the value of the coefficient is:

$$g* = -1.468 + .070(0) - .430(1) = -1.898 \qquad \text{[Eq. 17]}$$

For white respondents with low education, the net coefficient is:

$$g* = -1.468 + .070(-3.03) - .430(0) = -1.680 \qquad \text{[Eq. 18]}$$

[16] The net coefficient for whether married is:

$$g* = .419 - .761(female) \qquad \text{[Eq. 19]}$$

The net coefficient for men is then:

$$g* = .419 - .761(0) = .419 \qquad \text{[Eq. 20]}$$

While the net coefficient for women is

$$g* = .419 - .761(1) = -.342 \qquad \text{[Eq. 21]}$$

[17] Based on the coefficients in Model 3, the net coefficient for whether marital status changed is equal to:

$$g* = .194 + .504(black) - .741(female) \qquad \text{[Eq. 22]}$$

The net coefficient for white men is:

$$g* = .194 + .504(0) - .741(0) = .194 \qquad \text{[Eq. 23]}$$

The net coefficient for white women is:

$$g* = .194 + .504(0) - .741(1) = -.547 \qquad \text{[Eq. 24]}$$

The net coefficient for black men is:

$$g* = .194 + .504(1) - .741(0) = .698 \qquad \text{[Eq. 25]}$$

And the net coefficient for black women is:

$$g* = .194 + .504(1) - .741(1) = .043 \qquad \text{[Eq. 26]}$$

[18] Based on the coefficients in Model 3, the net coefficient for a change in disability status is:

$$g* = .125 - .069(education) \qquad \text{[Eq. 27]}$$

For respondents with high education, the net coefficient is:

$$\boldsymbol{g}^* = .125 - .069(3.03) = -.084 \qquad \text{[Eq. 28]}$$

For respondents with low education, the net coefficient is:

$$\boldsymbol{g}^* = .125 - .069(-3.03) = .334 \qquad \text{[Eq. 29]}$$

[19] The equation for the net effect of getting a disability is:

$$\boldsymbol{g}^* = (-.168 - .062(education)) + .624 * whether\_disabled \qquad \text{[Eq. 30]}$$

This equation involves results from substituting the level 2 equation into the equation for

interaction (Equation 8). The level 2 equation for whether disability status changed is equivalent

to β_{1j}, .624 is equivalent is equivalent to β_{3j} and whether disabled is equivalent to X_{2ij}. A person

with average education who gets a disability has a score of 1 on "Whether disabled." Thus, the

net coefficient for this disability change is:

$$\boldsymbol{g}^* = (-.168 - .062(0)) + .624 * 1 = .456 \qquad \text{[Eq. 31]}$$

[20] The corresponding figures for all four groups are included in the appendix on pages 52 through

56.

[21] Parallel to equation 30, a respondent with average education who recovers from a disability

has a score of 0 on "Whether disabled." The net effect of the disability change is then:

$$\boldsymbol{g}^* = (-.168 - .062(0)) + .624 * 1 = -.168 \qquad \text{[Eq. 32]}$$

[22] The net effect of a marital status change is equal to:

$$\boldsymbol{g}^* = .230 + .501(black) - .827(female) \qquad \text{[Eq. 33]}$$

For white women this is:

$$\boldsymbol{g}^* = .230 + .501(0) - .827(1) = -.597 \qquad \text{[Eq. 34]}$$

While this is technically the coefficient for respondents who marry (the reference group), the interactions between whether marital status changed and marital status (whether divorced and whether widowed) are nonsignificant.

[23] The corresponding figures are not shown for men, as marital status has a smaller effect for men. The figures white men and black men corresponding to Figures 3 and 4 are in the appendix.

[24] As in equation 30 above, the net effect of a disability status change at different ages can be calculated by substituting the level 2 equations into the net coefficient equation (equation 8).

$$\textbf{g}* = (-.735 - .053(education)) + .056 * age \qquad \text{[Eq. 35]}$$

At the mean level of education, the net effect of a disability status change is:

$$\textbf{g}* = (-.735 - .053(0)) + .056 * 20 = .385 \qquad \text{[Eq. 36]}$$

For the sake of clarity, the interaction term between whether disability status changed and whether disabled and between whether disabled and age have been omitted. Consequently, the net coefficient .385 represents the effect of recovering from a disability.

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Preretirement-Age Americans." *The Gerontologist 36*: 287-298.

TABLE 1
Proportion of sample experiencing unplanned events [a]

	Proportion
Never disabled	.633
Always disabled	.096
Disability change	.271
Recovers from a disability	.155
Gets a disability	.116
Always married	.708
Always divorced	.126
Always widowed	.056
Never married	.036
Marital change	.074
Gets married	.014
Gets divorced	.018
Is widowed	.043

[a] All proportions are weighted with the average of respondent level weights

TABLE 2
Logistic Coefficients (γ) and Odds Ratios for Binomial
Hierarchical Linear Models for Uninterrupted Pathways[a]

	MODEL 1				MODEL 2			
	γ	se(γ)[b]	exp(γ)[c]	t	γ	se(γ)[b]	exp(γ)[c]	t
Intercept								
Constant	3.000	0.083	20.079	36.049 ***	2.976	0.168	19.608	17.653 ***
Years of education	0.123	0.017	1.131	7.101 ***	0.102	0.018	1.107	5.566 ***
Whether black	-0.193	0.057	0.824	-3.361 **	-0.091	0.061	0.913	-1.490
Whether female	-0.857	0.097	0.424	-8.859 ***	-0.344	0.212	0.709	-1.625
Log of wage rate	-0.185	0.007	0.831	-27.341 ***	-0.160	0.007	0.852	-24.014 ***
Log of pension	0.025	0.004	1.026	6.120 ***	0.013	0.004	1.013	3.320 **
Log of nonhousing equity	0.129	0.079	1.137	1.623	0.073	0.063	1.076	1.174
Log of housing equity	-0.109	0.037	0.896	-2.926 **	-0.123	0.025	0.884	-4.871 ***
Whether service occupation	-0.187	0.058	0.829	-3.219 **	-0.125	0.056	0.883	-2.236 *
Whether manual occupation	-0.399	0.068	0.671	-5.905 ***	-0.280	0.063	0.756	-4.426 ***
Age								
Constant	-0.151	0.006	0.860	-23.420 ***	-0.146	0.007	0.864	-21.860 ***
Education in years	-0.005	0.002	0.995	-3.094 **	-0.007	0.002	0.993	-4.017 ***
Whether female	0.025	0.008	1.026	3.122 **	0.023	0.008	1.023	2.731 **
Whether 62								
Constant	-0.509	0.040	0.601	-12.742 ***	-0.572	0.042	0.564	-13.732 ***
Education in years	0.041	0.013	1.041	3.114 **	0.057	0.014	1.059	4.164 ***
Whether disabled								
Constant					-1.468	0.043	0.230	-33.764 ***
Years of education					0.070	0.012	1.072	5.612 ***
Whether black					-0.430	0.103	0.651	-4.163 ***
Whether married								
Constant					0.419	0.153	1.520	2.735 **
Whether female					-0.761	0.197	0.467	-3.854 ***
Whether divorced								
Constant					0.120	0.170	1.127	0.703
Whether female					0.098	0.214	1.103	0.457
Whether widowed								
Constant					-0.117	0.210	0.889	-0.559
Whether female					0.229	0.249	1.257	0.917

* p<.05 **p<.01 ***p<.001
[a] There are 5,942 level 2 units and 23,768 level 1 units.
[b] se(γ) is the robust standard error of the coefficients
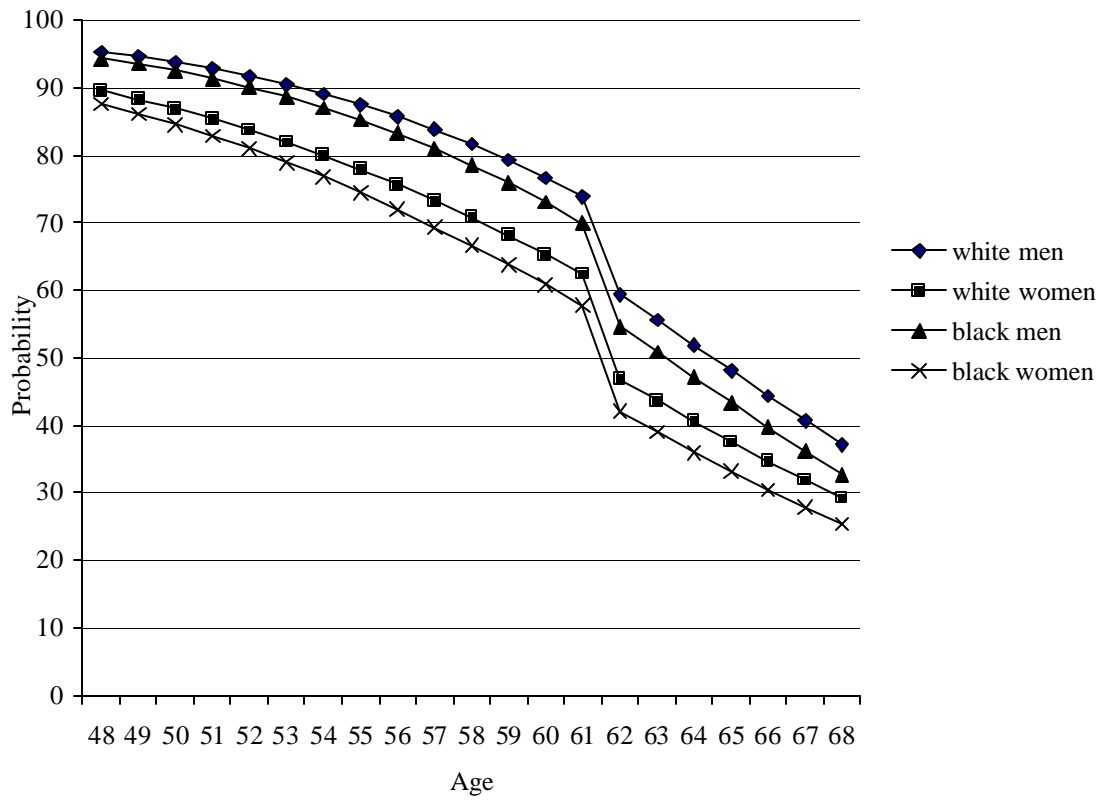[c] exp(γ) is the exponent of the coefficient. It is equivalent to the odds ratio

Figure 1: Uninterrupted Trajectories

Note: The probability for the four trajectories are calculated assuming that all predictors, except whether female and whether black, are 0.

TABLE 3
Logistic Coefficients (γ) and Odds Ratios for
Binomial Hierarchical Linear Models for Interrupted Pathways[a]

	MODEL 3				MODEL 4			
	γ	se(γ)[b]	Exp(γ)[c]	t	γ	se(γ)[b]	exp(γ)[c]	t
Intercept								
Constant	2.927	0.167	18.663	17.573 ***	2.958	0.166	19.262	17.809 ***
Years of education	0.115	0.018	1.122	6.270 ***	0.125	0.018	1.134	6.880 ***
Whether black	-0.107	0.061	0.898	-1.746	-0.090	0.061	0.914	-1.486
Whether female	-0.316	0.212	0.729	-1.493	-0.318	0.210	0.728	-1.511
Log of wage rate	-0.160	0.007	0.852	-23.999 ***	-0.159	0.007	0.853	-23.904 ***
Log of pension	0.013	0.004	1.013	3.279 **	0.000	0.000	1.000	-6.370 ***
Log of nonhousing equity	0.072	0.062	1.075	1.156	0.064	0.062	1.066	1.026
Log of housing equity	-0.123	0.025	0.884	-4.949 ***	-0.109	0.026	0.897	-4.126 ***
Whether service occupation	-0.123	0.056	0.885	-2.198 *	-0.168	0.056	0.845	-3.012 **
Whether manual occupation	-0.275	0.063	0.759	-4.350 ***	-0.328	0.063	0.720	-5.181 ***
Age								
Constant	-0.145	0.007	0.865	-21.636 ***	-0.142	0.007	0.867	-21.414 ***
Education in years	-0.007	0.002	0.993	-4.452 ***	-0.007	0.002	0.993	-4.330 ***
Whether female	0.022	0.008	1.023	2.674 **	0.021	0.008	1.022	2.583 ***
Whether 62								
Constant	-0.574	0.042	0.563	-13.810 ***	-0.582	0.042	0.559	-13.973 ***
Education in years	0.058	0.014	1.059	4.242 ***	0.057	0.014	1.059	4.206 ***
Whether disabled								
Constant	-1.479	0.044	0.228	-33.867 ***	-1.593	0.047	0.203	-34.247 ***
Years of education	0.074	0.012	1.077	6.014 ***	0.071	0.012	1.074	5.782 ***
Whether black	-0.431	0.103	0.650	-4.175 ***	-0.441	0.103	0.644	-4.266 ***
Whether married								
Constant	0.424	0.152	1.528	2.785 **	0.478	0.151	1.613	3.162 **
Whether female	-0.788	0.197	0.455	-3.996 ***	-0.876	0.196	0.416	-4.480 ***
Whether divorced								
Constant	0.110	0.169	1.116	0.647 **	0.149	0.168	1.161	0.888
Whether female	0.104	0.214	1.109	0.485 ***	0.057	0.212	1.059	0.270
Whether widowed								
Constant	-0.206	0.214	0.814	-0.963	-0.224	0.217	0.799	-1.035
Whether female	0.395	0.253	1.484	1.559	0.386	0.252	1.471	1.531
Whether marital status changed								
Constant	0.194	0.136	1.214	1.430	0.230	0.183	1.259	1.256
Whether black	0.504	0.186	1.655	2.702 **	0.501	0.185	1.650	2.701 **
Whether female	-0.741	0.163	0.477	-4.536 ***	-0.827	0.164	0.437	-5.058 ***
Whether disability status changed								
Constant	0.125	0.045	1.133	2.781 **	-0.168	0.056	0.846	-2.994 **
Years of education	-0.069	0.014	0.933	-4.826 ***	-0.062	0.013	0.940	-4.570 ***
Whether marital status changed * Whether divorced								
Constant					-0.133	0.218	0.875	-0.610
Whether marital status changed * Whether widowed								
Constant					0.098	0.198	1.103	0.496
Whether disability status changed * Whether disabled								
Constant					0.624	0.091	1.866	6.849 ***

* p<.05 **p<.01 ***p<.001
[a] There are 5,942 level 2 units and 23,768 level 1 units.
[b] se(γ) is the robust standard error of the coefficients
[c] exp(γ) is the exponent of the coefficient. It is equivalent to the odds ratio
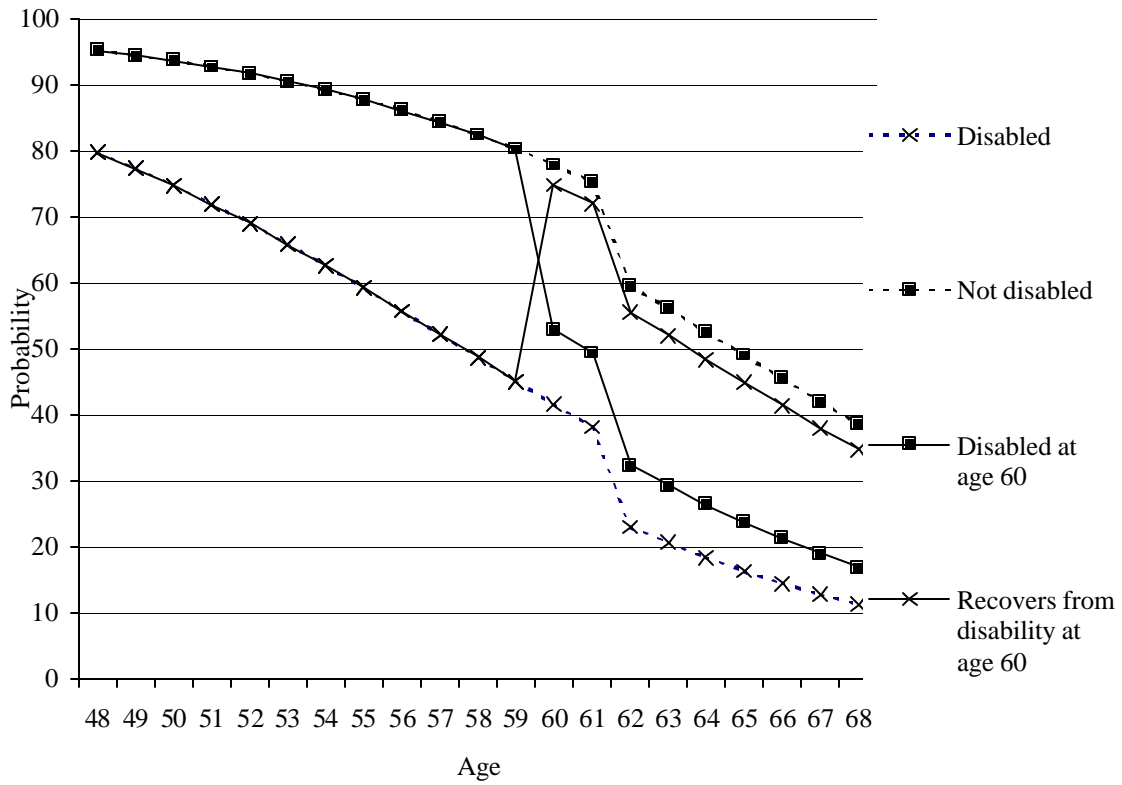
Figure 2: Probability of Working for Interrupted Trajectories in Disability Status

Note : The probability for the four trajectories are calculated for white men. All other predictors are assumed to be zero.
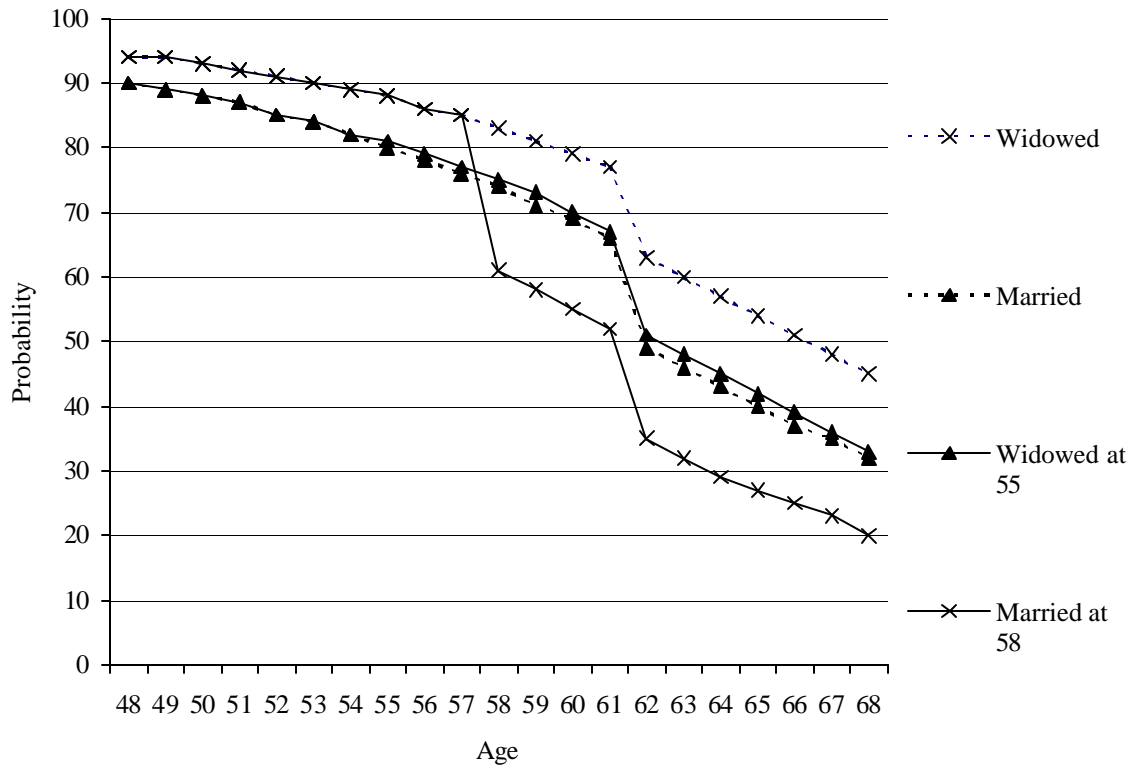
Figure 3: Probability of Working for Interrupted Trajectories in Marital Status for White Women

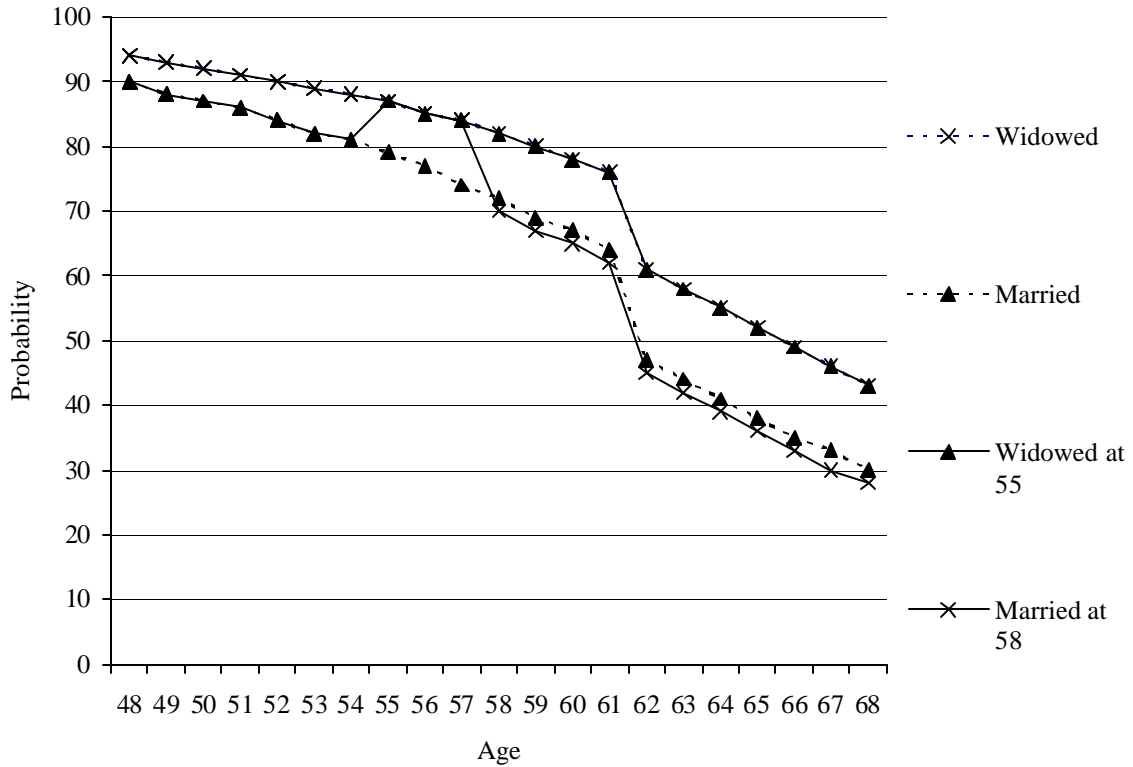*Note* : The probability for the four trajectories are calculated for white women. All other predictors, are assumed to be 0.

Figure 4: Probability of Working for Interrupted Trajectories in Marital Status for Black Women

Note : The probability for the four trajectories are calculated for black women. All predictors are assumed to be 0.

TABLE 4
Logistic Coefficients (γ) and Odds Ratios for
Binomial Hierarchical Linear Models for Uninterrupted Pathways with interactions with Age[a]

	MODEL 5			
	γ	se(γ)[b]	exp(γ)[c]	t
Intercept				
Constant	3.391	0.248	29.702	13.696 ***
Years of education	0.120	0.030	1.128	4.035 ***
Whether black	-0.082	0.083	0.921	-0.991
Whether female	-0.525	0.297	0.592	-1.764
Log of wage rate	-0.158	0.009	0.854	-18.098 ***
Log of pension	0.005	0.005	1.005	1.069
Log of nonhousing equity	0.067	0.082	1.069	0.819
Log of housing equity	-0.147	0.081	0.863	-1.815
Whether service occupation	-0.085	0.070	0.919	-1.211
Whether manual occupation	-0.231	0.081	0.793	-2.848 **
Age				
Constant	-0.172	0.011	0.842	-15.233 ***
Education in years	-0.008	0.003	0.992	-3.098 **
Whether female	0.031	0.012	1.032	2.527 *
Whether 62				
Constant	-0.605	0.058	0.546	-10.388 ***
Education in years	0.072	0.019	1.075	3.718 ***
Whether disabled				
Constant	-2.935	0.169	0.053	-17.404 ***
Years of education	0.088	0.017	1.092	5.054 ***
Whether black	-0.518	0.140	0.596	-3.702 ***
Whether married				
Constant	0.446	0.209	1.562	2.129 *
Whether female	-0.821	0.267	0.440	-3.073 **
Whether divorced				
Constant	0.095	0.237	1.100	0.402
Whether female	0.260	0.295	1.297	0.882
Whether widowed				
Constant	-0.492	0.304	0.612	-1.619
Whether female	0.764	0.351	2.146	2.174 *
Whether marital status changed				
Constant	0.420	0.294	1.522	1.427
Whether black	0.604	0.269	1.830	2.248 *
Whether female	-0.939	0.267	0.391	-3.513 **
Whether disability status changed				
Constant	-0.735	0.284	0.480	-2.589 *
Years of education	-0.053	0.019	0.949	-2.779 **
Whether marital status changed * Whether divorced				
Constant	-0.198	0.335	0.820	-0.592
Whether marital status changed * Whether widowed				
Constant	-0.018	0.289	0.982	-0.063
Whether disability status changed * Whether disabled				
Constant	1.230	0.430	3.422	2.864 **
Whether disability status changed*age				
Constant	0.056	0.014	1.058	4.019 ***

(continued)

TABLE 4 *continued*

	γ	se(γ)[b]	exp(γ)[c]	t
			MODEL 5	
Whether disability status changed*Whether disabled*age				
Constant	-0.042	0.040	0.959	-1.052
Whether disabled*age				
Constant	0.071	0.014	1.074	5.096 ***

* p<.05 **p<.01 ***p<.001
[a] There are 5,942 level 2 units and 23,768 level 1 units.
[b] se(γ) is the robust standard error of the coefficients
[c] exp(γ) is the exponent of the coefficient. It is equivalent to the odds ratio

APPENDIX

This appendix contains variations on Figures 2 and 3 for white men, white women, black men, and black women. In the article, only one figure is shown for each type of unplanned event in order to conserve space.

Figure 2, as presented in the article, tracks the probability of working for interrupted trajectories in disability status for white men, evaluated at the mean of all interval level predictors. While Figure 2a also shows the trajectories for white men, Figures 2b through 2d are corresponding figures for white women, black men, and black women respectively.

Figure 3, as presented in the article, tracks the probability of working for interrupted trajectories in marital status for black women, evaluated at the mean of all interval level predictors. While Figure 3d also shows the trajectories for black men, Figures 3a through 3c are corresponding figures for white men, white women, and black men respectively.

The figures included in this appendix are:

Figure 2a. Probability of Working for Interrupted Trajectories in Disability Status (White Men). *Shown in article as Figure 2.*
Figure 2b. Probability of Working for Interrupted Trajectories in Disability Status (White Women).
Figure 2c. Probability of Working for Interrupted Trajectories in Disability Status (Black Men).
Figure 2d. Probability of Working for Interrupted Trajectories in Disability Status (Black Women).

Figure 3a. Probability of Working for Interrupted Trajectories in Marital Status (White Men).
Figure 3b. Probability of Working for Interrupted Trajectories in Marital Status (White Women). *Shown in article as Figure 3.*
Figure 3c. Probability of Working for Interrupted Trajectories in Marital Status (Black Men).
Figure 3d. Probability of Working for Interrupted Trajectories in Marital Status (Black Women). *Shown in article as Figure 4.*
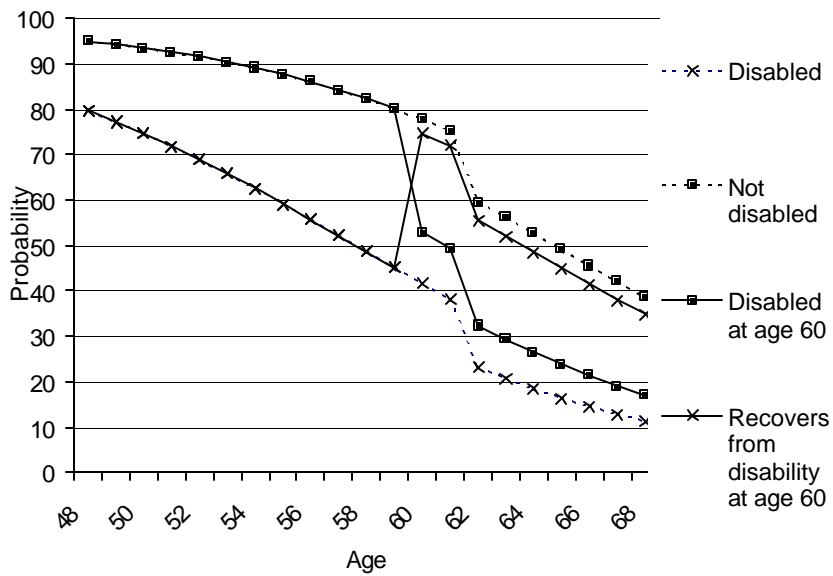
Figure 2a. Probability of Working for Interrupted Trajectories in
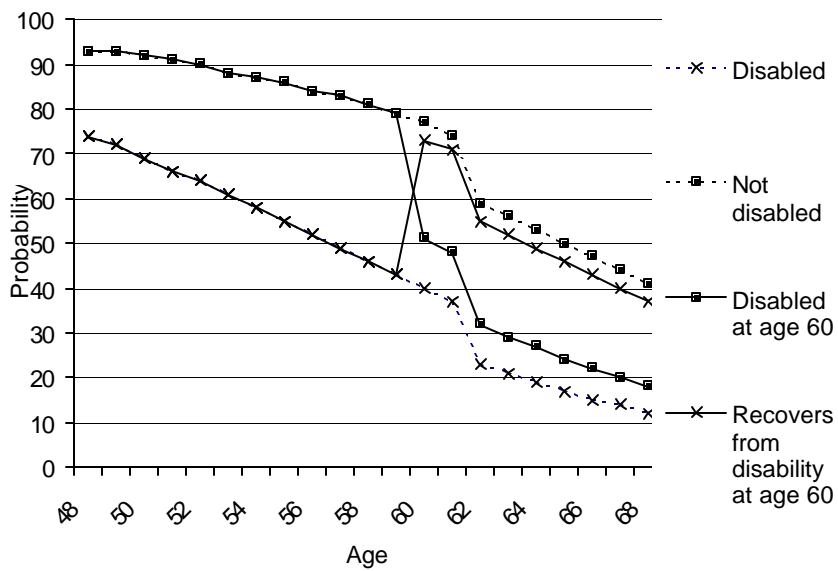Disability Status (White Men)

Figure 2b. Probability of Working for Interrupted Trajectories in
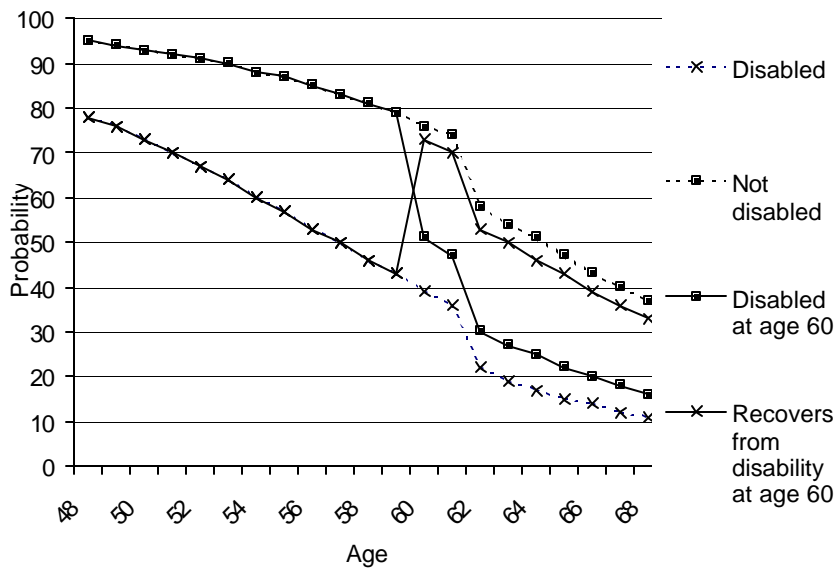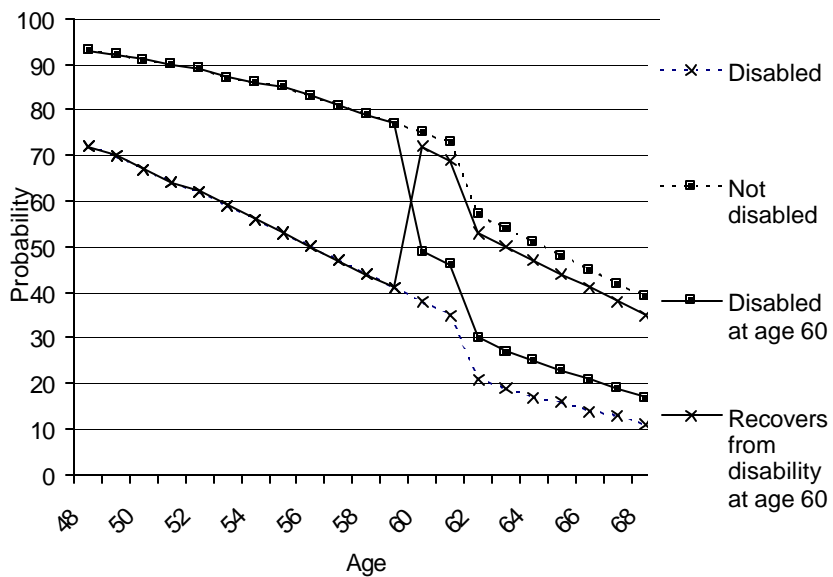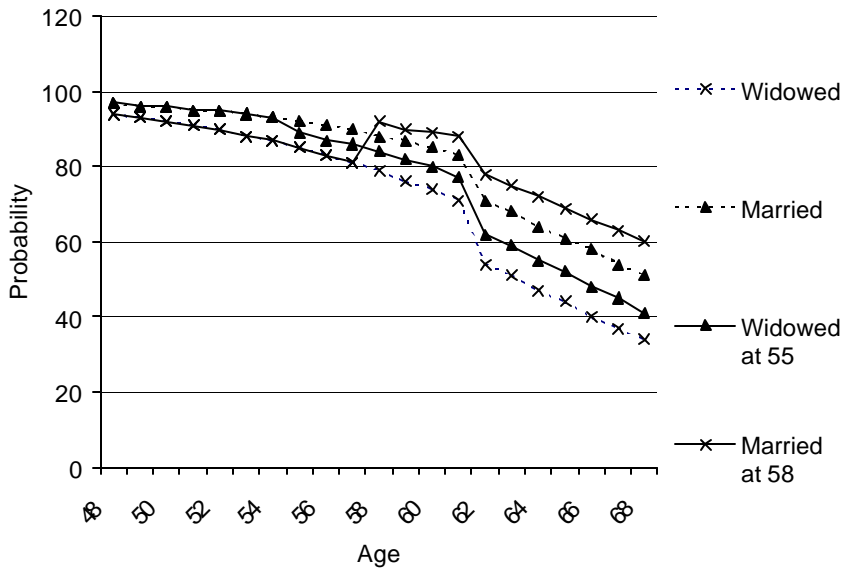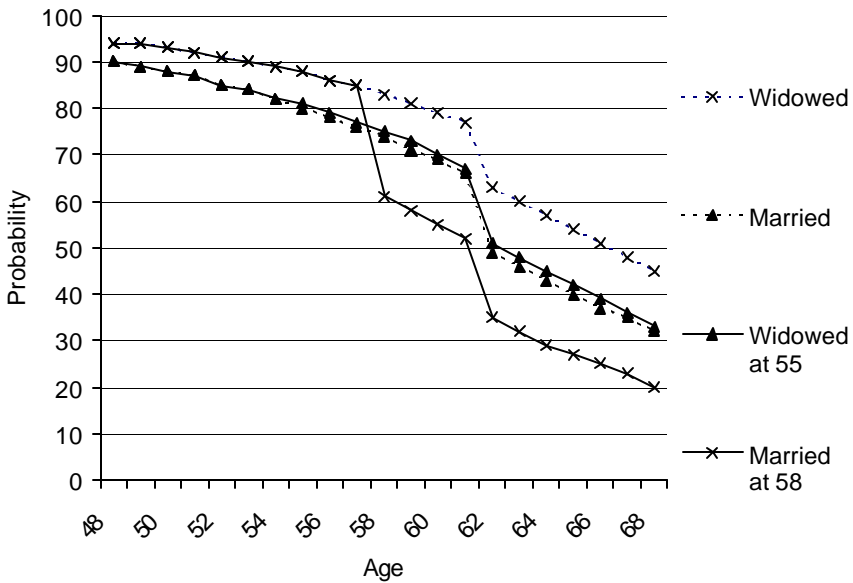Disability Status (White Women)

Figure 2c. Probability of Working for Interrupted Trajectories in
Disability Status (Black Men)

Figure 2d. Probability of Working for Interrupted Trajectories in
Disability Status (Black Women)

Figure 3a. Probability of Working for Interrupted Trajectories in
Marital Status (White Men)

Figure 3b. Probability of Working for Interrupted Trajectories in
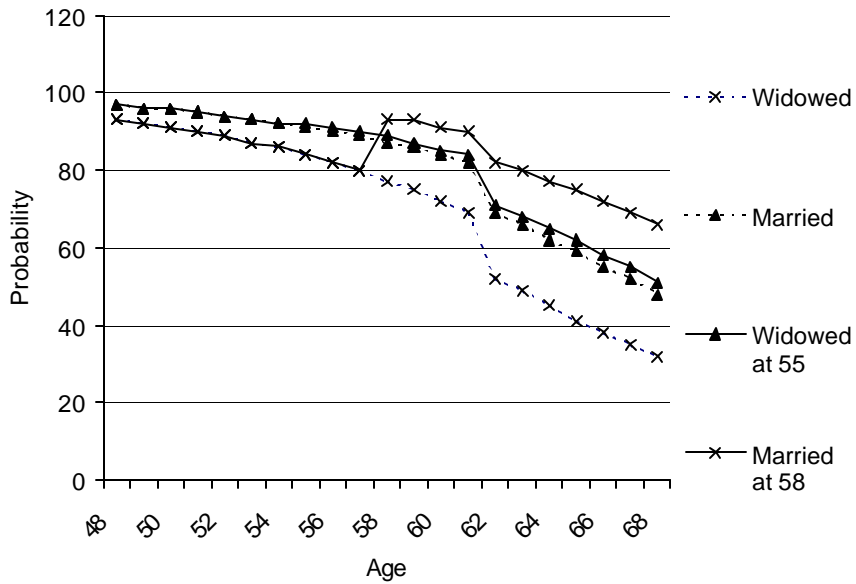Marital Status (White Women)

Figure 3c. Probability of Working for Interrupted Trajectories in
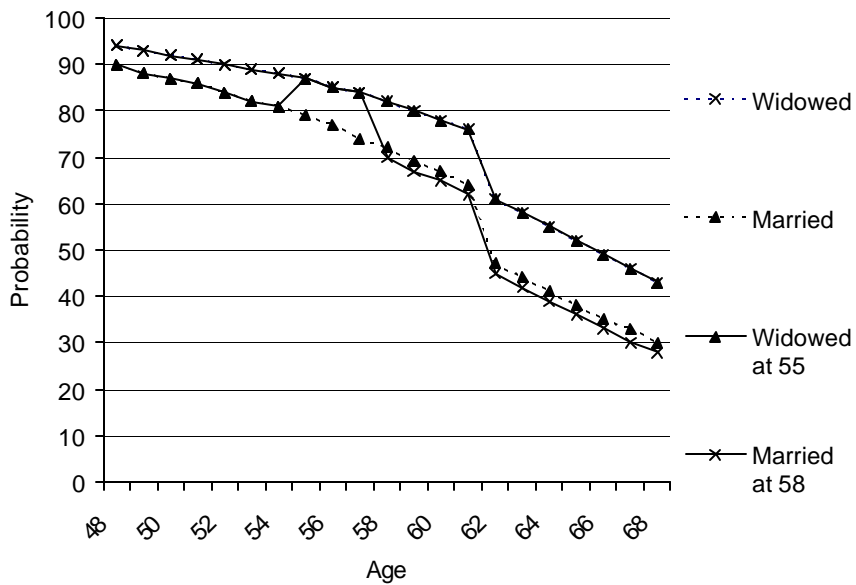Marital Status (Black Men)

Figure 3d. Probability of Working for Interrupted Trajectories in
Marital Status (Black Women)

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