Detailed Estimation of Worklife Expectancy for the Measurement of Human Capital: Accounting for Marriage and Children

Daniel L. Millimet
Southern Methodist University

Michael Nieswiadomy
University of North Texas

Daniel Slottje
Southern Methodist University

Abstract

Measuring an individual's human capital at a point in time as the present actuarial value of expected net lifetime earnings has a lengthy history. Calculating such measures requires accurate estimates of worklife expectancy. Here, worklife estimates for men and women in the United States categorized by educational attainment, race, marital status, parental status, and current labor force status are presented. Race has a much larger impact on the worklife expectancy of men than women. Education is associated with larger worklife differentials for women. The association between marriage and worklife expectancy is significant, but of opposite sign, for men and women: married women (men) have a lower (higher) worklife expectancy than single women (men). Parenthood is associated with a reduction in the worklife expectancy of women; the association is smaller and varies from positive for some education/marital status groups to negative for others for men.

JEL Classification: J17, J22, J26, J64
Key Words: Human Capital, Worklife Expectancy, Markov Processes, Children, Marriage

1. Introduction

Dating back at least until the work of Farr (1853), one measure of an individual's human capital is the present actuarial value of his or her expected net lifetime earnings. Dublin and Lotka (1930) extended the work of Farr (1853) by allowing for the possibility of non-employment, and Jorgenson and Fraumeni (1989, 1992) incorporate an individual's gender, age, and education into the analysis. Le et al. (2003, 2006) and Oxley et al. (2008) provide an extensive review and apply the Jorgenson and Fraumeni approach to the measurement of human capital in New Zealand. See Folloni and Vittadini (2010) for an excellent survey, as well as Wößmann (2003) for a discussion of human capital measurement in the context of the empirical growth literature. Finally, see Tchernis (2010) for a detailed analysis of the role of prior employment spells on human capital accumulation and lifetime wage paths.

Under this approach to human capital measurement, accurate data on worklife expectancies are required.[1] In the United States, the U.S. Bureau of Labor Statistics (BLS) first published worklife tables in 1950. However, the most recent BLS revision occurred in 1986 (BLS, 1986) and thus do not reflect recent changes in labor force participation rates. For instance, labor force participation rates (LFPR) for men (aged 16 and above) declined from 77.4% in 1980 to 74.9% in 1998, while the rate increased for women from 51.5% in 1980 to 59.8% in 1998 (Fullerton, 1999). In the past decade, the male LFPR has continued to decline to 72.3% in December 2008, while the female LFPR has remained around 59% (59.5% in December 2008) (BLS, 2009).

Millimet et al. (2003) recently updated the BLS worklife tables and provided a new econometric method for estimating worklife expectancies. The authors estimate worklife expectancies by gender, age, education, and race. In this paper, we utilize the Millimet et al. (2003) methodology to construct even more detailed worklife tables based on gender, age, education, race, marital status and parental status, conditional on current labor force status. These detailed estimates enable one to obtain much more precise measures of the stock of human capital.

Accounting for individual attributes such as marital status and parental status in worklife expectancies is likely to be very important. First, as articulated in Wößmann

[1] Worklife expectancies are also extensively utilized to project lost earnings (see, e.g., Nieswiadomy and Slottje, 1988; Nieswiadomy and Silberberg, 1988; Skoog and Ciecka, 2001, 2002).

(2003), Lovaglio (2010), and elsewhere, human capital is a complex concept, incorporating and depending on many underlying factors. Family background is one such factor. As such, observable attributes reflecting family background should play a salient role in any measure of human capital.

Second, beginning with the seminal work of Mincer and Polachek (1974), the impact of children and marriage on labor supply, housework, and earnings has received considerable attention in the human resources literature (e.g., Klerman and Leibowitz, 1999; Waldfogel, 1998; and Polachek and Siebert 1996). For example, Lundberg (1988) finds that labor supply of husbands and wives is jointly determined only when young children are present. Hersch and Stratton (1994) analyze the division of housework for employed spouses, finding significant interactions between the time allocation of husbands and wives, their education, and the presence of children under age 12. Angrist and Evans (1998), Millimet (2000), Ebenstein (2008) and others document a significant, negative effect of children on the labor supply of married women, but no corresponding effect on the time allocation of men. Millimet (2000) also finds a positive impact of children on the wages of men, while Waldfogel (1998), Hersch and Stratton (1997) and others discuss the "motherhood wage penalty." Finally, a number of studies provide empirical evidence in support of the marriage premium: the fact that married men earn more on average than single men even after controlling for common measures of human capital (e.g., Chun and Lee, 2001; Loh, 1996; Maasoumi et al. 2009).

Despite this rich literature examining the interplay between family life and labor market interactions, no study (to our knowledge) has analyzed the association between these attributes and worklife expectancy.[2] In addition, our worklife estimates differ from the earlier BLS approach in three other respects. First, we use the 1992 through 2001 March Annual Demographic Surveys in the Current Population Survey (CPS) reports. We use multiple years to account for a wide variety of economic impacts on labor supply, as well as to ensure that our results are not influenced by any particular business cycle. Second, we use an econometric model, rather than a simple relative frequency model. This econometric approach enables us to compile detailed worklife tables by race, education, marital status, and parental status, in addition to age and gender. Third, our

[2] In related work, Booth et al. (1999) examine work propensities over a several year period using the British Household Panel Survey for men and women and decompose these propensities into observable and unobservable components.

approach uses three labor force states: employed, unemployed and inactive (out of the labor force), in contrast to the BLS method that categorized persons as active (in the labor force) or inactive (out of the labor force). In the BLS two-state model, an individual's worklife at a particular age represents the number of years the individual is expected to remain in the labor force, not necessarily the number of years spent working. Our model differs because it defines worklife as the expected number of employed years remaining in a person's life.

The remainder of the paper is organized as follows. The next section discusses the decisions households must make and their possible impact on worklife expectancy. Section three presents the methodology. Section four discusses the data. Section five presents our econometrically estimated worklife tables. Section six concludes.

2. Household Decisions

Every household faces the decision of how to efficiently allocate its time between labor market and non-market pursuits. In the standard neo-classical model of time allocation, when there is only one adult (and no children) in the household, that decision depends on the market wage, non-labor income, and individual preferences for consumption versus leisure. For single-adult households with children, the decision becomes more complex. Now, the allocation of time between the labor market and the home will also depend on preferences for time spent with children, the relative quality of market-based versus own-provided child care, the cost of market-provided child care, and any social insurance benefits to which the household is entitled (e.g., Aid to Families with Dependent Children (AFDC), replaced by Temporary Aid to Needy Families (TANF) in 1996).[3] The host of factors involved in the time allocation decisions of single parents – relative to single individuals without children – makes the effect of children on worklife expectancy ambiguous.

In households with two adults, determining the time allocation of each individual becomes more challenging, even without children present. In the standard economic framework, the household must decide on the utility-maximizing combination of home-produced and market-produced goods, subject to the household's budget constraint and production functions of the husband and wife for home-produced goods. Once this optimal bundle is known, the household must allocate the time of each spouse to market

[3] See Blundell and MaCurdy (1999) for an in-depth analysis of labor supply models.

and non-market work. Even assuming identical productivity for men and women in the production of home-produced goods (e.g., a clean house or a home cooked dinner), women often have a comparative advantage in these endeavors because of their lower market wage.[4] According to this logic, one might expect married men to work more than their single counterparts, but for marriage to have the opposite effect on women. However, the effect on worklife expectancy is distinct from the predictions of this static framework since marriage may also influence retirement decisions in a more complex, dynamic model. Thus, the overall effect of marriage on worklife expectancy is ambiguous. Finally, it is fairly straightforward to extend the static analysis of a two-adult household to one with children by assuming that one of the home-produced goods being "produced" represents children. Consequently, the presence of children adds more variables into the decision-making process of married couples, but does not alter the fact that marriage has an ambiguous effect on worklife expectancy. Furthermore, comparing married individuals, by gender, with and without children is also ambiguous (both for static time allocation, as well as worklife expectancy) since the relative quality and cost of home-produced child care now become relevant considerations.

Prior to continuing, a final comment is warranted. Our discussion heretofore of both children and marriage has presumed both decisions to be exogenous. Since this is not likely to be the case, the *association* between children, marriage, and worklife expectancy also reflects any unobserved attributes correlated with these decisions as well as decisions concerning labor supply. In our empirical analysis, our differences in worklife estimates across individuals differentiated on the basis of children or marital status reflect these associations and, hence, should not be interpreted as *causal* effects. That said, since our motivation stems from improved measurement of human capital, and

[4] The likelihood that married men have a comparative advantage in market labor are increased by the marriage premium. Chun and Lee (2001) investigate one possible reason for the marriage premium: married men earn more because they are able to specialize in labor market activities, whereas their wives specialize in home production. Chun and Lee (2001) verify that the marriage premium decreases the more one's wife works. Loh (1996) and Maasoumi et al. (2009) examine the household specialization hypothesis, as well as two other theories behind the marriage premium: (i) positive selection (i.e., unobservables associated with higher earnings are valued in the marriage market), and (ii) employer discrimination (i.e., employers favor married men, viewing them as more responsible or less mobile). While these, and other such, analyses provides insight into the origin of the earnings differential of married and single men, it does not help determine the differences in worklife expectancy across married and single men since the standard income and substitution effects are ambiguous. On the one hand, married men may work more years than single men due to their higher wage (the substitution effect). On the other hand, married men may choose to "buy" more time with their family (the income effect).

not estimation of the effect of exogenous changes in children or marital status, this "weaker" interpretation of our findings is not problematic.

3. The Econometrically Estimated Worklife Model

The BLS (1986) worklife tables categorize individuals into two types: active and inactive. The active category includes all individuals in the labor force (employed and unemployed). The BLS's inclusion of the unemployed among the active is based on the assumption that the unemployed are more similar to the employed than those out of the labor force. The BLS (1982, p. 2) states, "It has long been recognized that persons who are already in the labor force are more likely to work in the future than are those not currently active." However, the result of this pooling is that the BLS definition of "worklife" is not comparable to the number of remaining working years due to the inclusion of the unemployed in the active category. Thus, the BLS (1986) "worklife" expectancies overestimate the *actual* worklife (i.e., the remaining years of employment) of the currently active, and may exaggerate the worklife of the unemployed if their labor force behavior is more like that of the inactive. For example, Hall (1970), Clark and Summers (1979, 1982), Tano (1991), Flinn and Heckman (1983), Gönül (1992), and Jones and Riddell (1999) find that the distinction between the unemployed and inactive is fuzzy. Many individuals bounce between periods of unemployment and labor force inactivity within a given nonemployment spell. However, rather than simply re-defining the inactive group to include the unemployed, we allow for three distinct groups by estimating a multinomial logit for the three labor market states (employed, unemployed and inactive).[5] We estimate the model on three distinct subsets of the data: individuals initially employed, initially unemployed, and initially inactive. We then use the estimates to estimate the nine relevant age-specific transition probabilities described below.

Formally, we use the model in Millimet et al. (2003) that classifies individuals into three labor force states. Let q_x represent the probability of death in the year following exact age x, let l_x represent the number of survivors at age x, and let ${}^{E}l_x$ ${}^{U}l_x$, and ${}^{I}l_x$ represent the number of employed, unemployed and inactive survivors at age x,

[5] Given the ambiguity surrounding the similarity between the unemployed and the employed and those out of the labor force, we focus on a multinomial logit model as opposed to a nested logit model. We do conduct Hausman tests (Greene, 1993, p. 671) of the Independence of Irrelevant Alternatives (IIA) assumption.

respectively. In any given year, there are nine relevant conditional probabilities of work force transition:

- ${}^{E}p_x^{E}$ = the probability that someone who is employed at age *x* will be employed at age *x+1*;
- ${}^{E}p_x^{U}$ = the probability that someone who is employed at age *x* will be unemployed at age *x+1*;
- ${}^{E}p_x^{I}$ = the probability that someone who is employed at age *x* will be inactive at age *x*+1;
- ${}^{U}p_x^{E}$ = the probability that someone who is unemployed at age x will be employed at age *x*+1;
- ${}^{U}p_x^{U}$ = the probability that someone who is unemployed at age *x* will be unemployed at age *x*+1;
- ${}^{U}p_x^{I}$ = the probability that someone who is unemployed at age *x* will be inactive at age *x*+1;
- ${}^{I}p_x^{E}$ = the probability that someone who is inactive at age *x* will be employed at age *x*+1;
- ${}^{I}p_x^{U}$ = the probability that someone who is inactive at age *x* will be unemployed at age *x*+1;
- ${}^{I}p_x^{I}$ = the probability that someone who is inactive at age *x* will be inactive at age *x*+1.

As the above transitional probabilities are conditional on survival from age *x* to age *x*+1, ${}^{E}p_x^{E} + {}^{E}p_x^{U} + {}^{E}p_x^{I} = 1$, ${}^{U}p_x^{E} + {}^{U}p_x^{U} + {}^{U}p_x^{I} = 1$, and ${}^{I}p_x^{E} + {}^{I}p_x^{U} + {}^{I}p_x^{I} = 1$.

Assuming the probability of death and the probability of transition between work force states are independent, the number of employed survivors at age *x+1* (${}^{E}l_{x+1}$), the number of unemployed survivors at age *x+1* (${}^{U}l_{x+1}$), and the number of inactive survivors at age *x*+1 (${}^{I}l_{x+1}$) can be defined as:

- ${}^{E}l_{x+1} = {}^{E}l_x\,{}^{E}p_x^{E} + {}^{U}l_x\,{}^{U}p_x^{E} + {}^{I}l_x\,{}^{I}p_x^{E}$
- ${}^{U}l_{x+1} = {}^{E}l_x\,{}^{E}p_x^{U} + {}^{U}l_x\,{}^{U}p_x^{U} + {}^{I}l_x\,{}^{I}p_x^{U}$
- ${}^{I}l_{x+1} = {}^{E}l_x\,{}^{E}p_x^{I} + {}^{U}l_x\,{}^{U}p_x^{I} + {}^{I}l_x\,{}^{I}p_x^{I}$

where $l_x = {}^{E}l_x + {}^{U}l_x + {}^{I}l_x$ and $l_{x+1} = l_x(1 - q_x)$.

It is assumed that persons who die, become employed, become unemployed, or become inactive in a given year do so at mid-year. According to this Markov process with one-period memory, the transition probabilities do not depend on previous transition probabilities.

The BLS calculated its 1986 increment-decrement tables by using relative frequency process to matched CPS data from the 1979 - 1980 survey. The BLS's worklife tables are sub-divided based on education or race, but not both at the same time, due to insufficient sample size. For similar reasons, marital status and parental status could not be used in constructing worklife tables. Yet, as stated previously, it is well known that labor supply decisions depend on these factors (see also Killingsworth, 1983). To account for these factors we estimate the transition probabilities as a function of marital status, number of children present in the home, age, race, gender, and education.

For the purpose of estimating worklife expectancies, we are interested in transitional (or conditional) probabilities. Thus, we model the transition probabilities using a multinomial logit framework conditional on initial labor force status. Specifically, using matched individuals from the CPS, we estimate three multinomial logit models by standard maximum likelihood methods using mutually exclusive sub-samples: individuals employed in year t, individuals unemployed in year t, and individuals inactive in year t. Using the notation $Y = 0$ for employed in year $t+1$, $Y = 1$ for unemployed in year $t+1$ and $Y = 2$ for inactive in year $t+1$, we first estimate the following probabilities for individuals employed in year t:

$$\text{Prob}\ (Y = 2) = e^{\beta_2 x_i} / (1 + \sum_{k=1}^{2} e^{\beta_k x_i}) \tag{1a}$$

$$\text{Prob}\ (Y = 1) = e^{\beta_1 x_i} / (1 + \sum_{k=1}^{2} e^{\beta_k x_i}) \tag{1b}$$

$$\text{Prob}\ (Y = 0) = 1 / (1 + \sum_{k=1}^{2} e^{\beta_k x_i}) \tag{1c}$$

We then estimate Eqn. (7.1a-c) using only individuals who were initially unemployed in year t, as well as using only individuals who were initially inactive ($Y = 2$) when first interviewed in year t. These models are estimated separately for men and women, thereby allowing the coefficients to differ by gender and initial labor force status.

The x vector in Eqn. (7.1a-c) includes controls for: age, age squared, a race dummy, race interacted with age and age squared, marriage dummy, marriage dummy times age, number of children under six, number of children under 18, each child variable interacted with age, occupation dummies, occupation dummies interacted with age, and time dummies. We use four occupation categories: managerial and professional; technical, sales and administrative; service; and operators and laborers.[6]

Upon completing the estimation, we obtain the average predicted transition probabilities at each age:

- ${}^{E}p_x^{E} = \frac{1}{N_x^E}\sum_{i=1}^{N_x^E} {}^{E}p_i^{E}$; ${}^{E}p_x^{U} = \frac{1}{N_x^E}\sum_{i=1}^{N_x^E} {}^{E}p_i^{U}$; ${}^{E}p_x^{I} = \frac{1}{N_x^E}\sum_{i=1}^{N_x^E} {}^{E}p_i^{I}$
- ${}^{U}p_x^{E} = \frac{1}{N_x^U}\sum_{i=1}^{N_x^U} {}^{U}p_i^{E}$; ${}^{U}p_x^{U} = \frac{1}{N_x^U}\sum_{i=1}^{N_x^U} {}^{U}p_i^{U}$; ${}^{U}p_x^{I} = \frac{1}{N_x^U}\sum_{i=1}^{N_x^U} {}^{U}p_i^{I}$
- ${}^{I}p_x^{E} = \frac{1}{N_x^I}\sum_{i=1}^{N_x^I} {}^{I}p_i^{E}$; ${}^{I}p_x^{U} = \frac{1}{N_x^I}\sum_{i=1}^{N_x^I} {}^{I}p_i^{U}$; ${}^{I}p_x^{I} = \frac{1}{N_x^I}\sum_{i=1}^{N_x^I} {}^{I}p_i^{I}$

where N_x^E, N_x^U, and N_x^I are the number of employed, unemployed and inactive individuals, respectively, in the sample of age x.[7] Finally, in accordance with other studies, we substitute the age-specific transition probabilities with a nine-year moving average to lessen variation due to small sample size, particularly at the tails of the age distribution (Schoen and Woodrow, 1980).

After the smoothed transition probabilities are obtained, the expected worklife expectancy for a currently active or inactive individual can be estimated recursively. Specifically, given a terminal year $T+1$, after which no one is assumed to be employed, the worklife for an employed individual of age T is given by the probability that an employed T year-old remains employed at $T+1$ (call this E_T^E). For an unemployed individual of age T, the worklife is given by one-half times the probability that an unemployed T year-old becomes employed at age $T+1$ (call this E_T^U). For an inactive individual of age T, the worklife is given by one-half times the probability that an

[6] The occupation dummies only enter the models used predict the probability of exiting or remaining in the state of employment.
[7] Note, in the actual estimation, weighted averages are obtained, where the weights are the sample weights.

inactive *T* year-old becomes employed at age *T+1* (call this E_T^I).[8] The worklife expectancy for an employed individual of age *T-1* is given by

$$E_{T-1}^{E} = {}^{E}p_{T-1}^{E}\left(1 + E_T^E\right) + {}^{E}p_{T-1}^{U}\left(0.5 + E_T^U\right) + {}^{E}p_{T-1}^{I}\left(0.5 + E_T^I\right) \tag{2}$$

and the worklife expectancy for an unemployed individual of age *T-1* is given by

$$E_{T-1}^{U} = {}^{U}p_{T-1}^{E}\left(0.5 + E_T^E\right) + {}^{U}p_{T-1}^{U}\left(E_T^U\right) + {}^{U}p_{T-1}^{I}\left(E_T^I\right) \tag{3}$$

and the worklife expectancy for an inactive individual of age *T-1* is given by

$$E_{T-1}^{I} = {}^{I}p_{T-1}^{E}\left(0.5 + E_T^E\right) + {}^{I}p_{T-1}^{U}\left(E_T^U\right) + {}^{I}p_{T-1}^{I}\left(E_T^I\right) \tag{4}$$

This process is repeated down to age *T*-(*T-16*).

4. Data

The most recent Current Population Survey (CPS) March annual surveys are used to create a series of two-period longitudinal data sets. The CPS is one of the most frequently cited data sources on individual income and employment in the U.S. In the CPS, households are interviewed for four consecutive months, left alone for eight months, and then surveyed for four additional consecutive months. New households are inserted each month into the survey as earlier ones finish their interviews.

In constructing a two-period panel data set, we match persons in rotation groups 1 - 4 (the entering group) with persons in rotation groups 5 - 8 (the outgoing group) in the following year. Next, we learn the change in labor force status from the prior year for all individuals between the ages of 17 and 71, not presently in school or disabled. We assume that no person works past age 71 or before age 16. Consequently, our worklife tables cover the ages 16 to 70.[9]

The 1992 to 2001 March annual surveys provide detailed individual information in addition to labor market status, yielding three benefits. First, the most current changes in labor force participation will be integrated into our worklife tables. Second, our data set reflects labor market activity over a ten-year period, and thus is not sensitive to any specific set of economic conditions. Finally, the availability of a host of other individual attributes likely to influence labor supply decisions is available.

---

[8] As in the BLS approach and advocated by Alter and Becker (1985), we assume that all transitions occur at mid-year.

[9] Note, the original BLS tables terminated at age 75. Despite the inclusion of several years of data, our tables terminate earlier due to the allowance of three unique labor market states. Were we to apply the method utilized herein and maintain the assumption of only two labor market states, we would be able to extend the tables out to age 85, as in Millimet et al. (2003).

Our matching algorithm is similar to Peracchi and Welch (1995). First, we match households in rotations 1 - 4 in one year with the same household (using the unique household identification number) the following year in rotations 5 - 8. Second, we require that an individual in a matched household must have the same sex and race, and must be one year older, when interviewed in rotations 5 - 8. We should note the CPS does not follow movers, which may lead to a nonrandom attrition problem. Fortunately, Peracchi and Welch (1995) discovered no systematic bias in transition estimates after controlling for sex, age, and labor force status at the time of the first survey, based on 13 years (1979 - 1991) of CPS data. Peracchi and Welch (1995) matched two-thirds of their March annual surveys from 1979 to 1991. Similarly, we are able to match approximately 63% of our 1992 - 2001 samples. (1993 cannot be matched with 1994, and 1995 cannot be matched with 1996 because of survey modifications.)

To allow comparison to previous worklife tables, we sort the data by sex and education, as represented in the BLS 1986 study. Our three educational categories are nearly identical to those of the BLS: less than a high school degree, high school degree to some college, and a college degree and above.[10] In addition, we utilize data on marital status, race, number of children under age six, number of children under age 18, and occupation. Finally, we include the appropriate sampling weights not only when estimating the multinomial logit models, but also when obtaining the mean age-specific transition probabilities.

5. Empirical Results

Tables 1 through 6 present the worklife expectancies obtained after permitting three unique labor force states (employed, unemployed and inactive).[11] We do not present the full tables, but they are available at http://faculty.smu.edu/millimet/pdf/worklifekids.pdf. In order to construct the worklife tables for women with children, we assume that a woman has the typical number of children based on the demographic statistics from the Census Bureau. Specifically, we assume that a married woman with less than a high school degree has three children at ages 22, 25, and 28. A married woman with a high school degree has two children at

---

[10] The BLS's second and third categories were somewhat different. The BLS second category was high school degree to 14 years of schooling and the third category was 15 or more years of schooling.
[11] According to the Hausman tests, we cannot reject the Independence of Irrelevant Alternative (IIA) assumption in the majority of the multinomial logit models estimated.

ages 22 and 25. A married woman with a college degree has two children at ages 22 and 25. We assume that a single woman with less than a high school degree has two children at ages 22 and 25. A single woman with a high school degree has one child at age 22. A single woman with a college degree has one child at age 22. These assumptions of the number of children and the age of the mother at the time of each birth are only meant to be approximate.[12] We now present the results, first for women, then for men. For both women and men, we first discuss the relationship between race and worklife expectancy, then education, marital status, and children.

5.1. *Women: race, education, and marriage*

For women, race has little association with worklife expectancy except for high school educated women. As shown in Table 1, for employed 30 year-olds, there is less than a one-year difference between single white and non-white employed women with college and less than high school education levels; the same is true for married white and non-white women.[13] However, the worklife expectancy of an employed 30 year-old white woman with a high school degree exceeds that of similar non-white woman by over two years.

In contrast, the association between educational attainment and worklife expectancy is more pronounced, especially for single women. As shown in Table 1, a 30 year-old college educated single white female has a worklife expectancy that is 13.9 years longer than a woman with less than a high school degree. A 30 year-old college educated married white woman has a worklife expectancy that exceeds that for a white woman with less than a high school degree by 9.1 years. For non-white women, the differences in worklife expectancies between college and less than high school educated

[12] According to a recently released report by the US Centers for Disease Control and Prevention (*National Vital Statistics Report*, Volume 51, Number 1, 11 December 2002, available at http://www.cdc.gov/nchs/data/nvsr/nvsr51/nvsr51_01.pdf), the average age of first birth in 1970 was 21.4 years of age. In 2000, this has increased to 24.9 years of age. The report does not differentiate age at first birth by education or marital status. As a result, the worklife expectancies presented in this study are more accurate for women in older cohorts.

[13] Throughout the discussion of the results, it is important to bear two caveats in mind. First, worklife estimates are estimated *conditional* on current labor force status (e.g., employed, unemployed, or inactive); the relationship between race or any other demographic characteristics of individuals and current labor force status is ignored. Thus, while race appears unrelated to, say, the worklife expectancy of single, *employed* females, race may play a pivotal role in the probability that a single female *is* employed. Second, in principle there are more than three labor force states into which individuals may be categorized. For example, an unemployed individual with a long history of employment may be fundamentally distinct from an individual who has spent the majority of his or her life unemployed or inactive. Unfortunately, such distinctions are not possible using CPS data.

women are slightly larger (15.3 years for singles and 10.3 years for married persons). Moreover, when comparing college to the high school level, race matters as well. For instance, the worklife expectancy of a 30 year-old college educated single white woman is 4.3 years longer than for a white woman with a high school degree; for a non-white woman the difference is 7.5 years. The worklife expectancy of a 30 year-old college educated married white woman exceeds the worklife expectancy of her high school counterpart by 1.6 years; for a non-white woman the difference is 4.6 years.

As expected, the association between marriage and the worklife expectancy of women is considerable. The relationship is strongest for women with college degrees. As shown in Table 1, for both white and non-white college educated 30 year-old women, marriage is associated with a decline in worklife expectancy of roughly 4.2 years. For white and non-white high school educated women, marriage is associated with about a 1.5 year reduction in worklife expectancy. At the less than high school level, there is almost no correlation between marital status and worklife expectancy. The negative association between marriage and worklife expectancy for women with at least a high school education is consistent with the model of joint labor supply discussed previously. Specifically, marriage typically results in one spouse specializing in the production of household goods. Since women tend to (i) be more productive at home, and/or (ii) have a lower opportunity cost (due to a lower market wage), the labor force participation of women suffers with marriage.

5.2. *Women: children*

The worklife expectancy of a single woman with children (as described in our model) is between one and two years lower than a comparable childless woman. As shown in Table 2, for a 30 year-old employed single white college educated female, children are associated with a reduction in worklife expectancy of 2.2 years; for a non-white woman the reduction is 1.4 years. At the high school level the reduction is 1.1 to 1.2 years for single white and non-white women, respectively. Finally, for those with less than a high school education, worklife expectancy is 1.5 to 1.7 years lower for single whites and non-whites, respectively.

The worklife expectancy of a married woman with children (as described in our model) is between one and three years lower than comparable childless woman. As shown in Table 3, for a 30 year-old employed married white college educated female, the worklife expectancy for women with children is 2.9 years shorter than for similar childless women; for a non-white woman the reduction is 1.5 years. At the high school level the reduction associated with children is 1.0 to 1.1 years for married whites and non-whites, respectively. Lastly, for women without a high school diploma, the reduction in worklife expectancy is 1.2 to 1.3 years for married whites and non-whites, respectively.

The negative relationship between children and female worklife expectancy are consonant with the vast empirical studies documenting the deleterious (causal) effects of children on female labor force participation (e.g., Millimet, 2000; Angrist and Evans, 1998). Of course, studies of the effects of children on female labor force participation typically focus on just the impact of young children on labor market behavior (while the children still reside at home). While differences in the worklife expectancies of women with and without children reflect these effects, children may also influence worklife expectancy even when they no longer reside with the parents. For example, adult children may influence retirement decisions due to income transfers from adult children to their parents (e.g., Jellal and Wolff, 2000), or due to the desire of the parents to relocate closer to their adult children to be near grandchildren.

5.3. *Men: race, education, and marriage*

In contrast to women, the association between race and male worklife expectancy is particularly acute, with the largest correlation occurring at the high school level. As shown in Table 4, for employed 30 year-olds with less than a high school education, the worklife expectancy of single white men exceeds that of similar non-white men by one year. The same holds for married men. For men with a high school education, the worklife expectancy of single white men is 4.4 years longer than of similar non-whites; for married men, the difference is 3.5 years. For men with a college degree, worklife expectancy is 2.9 years greater for single white men; for married men, the difference is 2.3 years.

Worklife expectancies also vary considerably by educational attainment, particularly for 30 year-old single white males. However, the association is not as strong as for women. As shown in Table 4, an employed 30 year-old single white male with a college education will work, on average, an additional 11.2 years relative to a white male with a less than high school education. (The comparable differential for women is 13.9

years; see Table 1). The difference is of lower magnitude for single non-white males, with the difference being only 9.2 years (15.3 for comparable women; see Table 1). For 30 year-old married white males, the worklife expectancy of the college educated exceeds that of similar men with less than high school educated by 9.0 years. For non-white males, this difference is only 7.8 years.

The differences in worklife expectancy are not as great when comparing college versus high school educated men. The worklife expectancy of a 30 year-old single white male with a college education is 3.5 years greater compared with his high school educated counterpart. The comparable difference for a 30 year-old single non-white male is 5.0 years. The worklife expectancy of a 30 year-old married white male exceeds that of a comparable high school educated individual by 2.7 years; for similar non-white males, the difference is 3.9 years.

Marriage is associated with large differences in worklife expectancy for men of both races, especially at the high school and less than high school education levels. In contrast to women, marriage increases worklife expectancies for men. As shown in Table 4, a 30 year-old married white male with a college-level education works an additional 1.9 years relative to his single counterpart; an additional 2.5 years for non-white males. These numbers increase at the high school level. The worklife expectancy of a 30 year-old married white male exceeds that of a comparable single male by 2.7 years. For a 30 year-old married non-white male, this difference increases to 3.7 years. Finally, regardless of race, the worklife expectancy of a 30 year-old married male with less than a high school education exceeds that of a comparable single male by roughly 4.0 years. The positive association between marriage and worklife expectancy is consistent with marriage premium increasing male labor force participation due to the substitution effect dominating the income effect.

5.4. *Men: children*

While the relationship between children and female worklife expectancy is clear, the association for males varies, mainly across education levels, but also by race. Moreover, while children are always associated with a reduction in a woman's worklife, regardless of marital status, this does not hold for men.

For single men with children relative to childless men, worklife expectancy varies from 1.1 years lower to 2.0 years greater. As shown in Table 5, for a 30 year-old employed single white college educated male, children are associated with a reduction in worklife expectancy of 1.1 years; for a non-white male, worklife expectancy is 0.2 years greater. At the high school level, children are associated with a 0.4 to 0.8 year increase in worklife expectancy for single white and non-white men, respectively. Finally, for males without a high school education, children are associated with a 2.0 year increase in worklife expectancy for single white men; a 0.7 year decrease for single non-white men.

The relationship between children and the worklife expectancy of married males is much weaker than for single males. As shown in Table 6, the worklife expectancy of a 30 year-old employed married college educated male with children is 0.3 to 0.1 years lower relative to comparable childless men for white and non-white men, respectively. At the high school level, children are associated with an increase in worklife expectancy by 0.3 to 0.5 years for married whites and non-whites, respectively. Finally, the worklife expectancy of married males without a high school diploma with children is 0.4 and 0.3 years lower relative to comparable single males for whites and non-whites, respectively.

For males, married males in particular, the association between children and worklife expectancy is complex. As documented in Millimet (2000), Angrist and Evans (1998), and others, the presence of young children in the household has a statistically insignificant effect on male labor supply. However, children may influence the retirement behavior of males as discussed above for women. In addition, Millimet (2000) finds a statistically significant, positive effect of children on the wages of married men, thus providing an incentive to delay retirement if the substitution effect dominates the income effect.

6. Conclusion

The correlations between gender, race, education, marital status and children and labor market behavior have been the focus of much empirical research. In addition, recent research has investigated the effect of these attributes on the propensity to work over periods of time (Booth et al. 1999). This paper extends this research by estimating separate worklife expectancies for individuals differentiated by these attributes. Such estimates are a vital component of one frequently used measure of human capital: the present actuarial value of an individual's expected net lifetime earnings.

To proceed, we utilize the econometric model developed in Millimet et al. (2003) and current data over a ten-year period (1992 – 2001) from CPS March annual surveys, thus allowing us to construct worklife tables using several demographic characteristics simultaneously. Relative to previous worklife tables, the estimates presented herein offer several advantages. First, the econometric approach permits the estimation of much more detailed worklife expectancy tables compared to the relative frequency approach, which is limited to using one variable at a time because of sample size problems. Second, our use of multiple rounds of up-to-date data reduces the sensitivity of the worklife expectancy estimates to business cycle conditions, as well as reflects the vast changes that have occurred in the U.S. labor market over the past few decades. Third, our use of a multinomial logit model to estimate transition probabilities allows us to categorize individuals into three states: employed, unemployed, and inactive. This avoids the arbitrary decision of pooling unemployed individuals with either employed individuals or those out of the labor force, as well as maintains the "true" definition of worklife.

The results indicate that worklife expectancy is sensitive to several factors. First, even controlling for differences in childbearing and marital status, race has a large impact, especially for men (up to 4.4 years longer worklife for single white men with high school degrees). Second, education also matters, particularly for women. The largest difference occurs between single non-white women with a college education and similar women with less than a high school degree (where worklife expectancy is 15.3 years greater for the former).

Third, marriage has opposite effects on the worklife expectancies of men and women. For women, marriage reduces worklife expectancy, where the largest differential is 4.2 years for married white and non-white women with college degrees compared to their single counterparts. The largest differential for men is the additional 4.0 years of worklife expectancy for married white males with less than a high school education relative to comparable single males. Failure to differentiate by marital status, as in the prior literature, yields estimated worklife expectancies that are weighted averages of those obtained here (see, e.g., Millimet et al. 2003). Moreover, use of such weighted averages to construct measures of human capital stock will be less precise and will not adequately capture changes in marriage trends on the stocks in the future. For example, the marriage rate declined in every state and Washington D.C. in the United States between 1990 and 2004 except Hawaii and West Virginia.[14]

[14] See http://www.cdc.gov/nchs/data/nvss/marriage90_04.pdf.

Finally, the association between children and worklife expectancy is large and negative for women; for men, the association varies and is of smaller magnitude. The largest relationship for women is the 2.9 year reduction in worklife for married white college educated woman relative to similar childless women. For men, the largest increase in worklife expectancy (2.0 years) is for single white men with less than a high school degree, while the largest decrease in worklife expectancy (1.1 years) is for single white college educated men. Again, failure to incorporate children into estimated worklife expectancies in the prior literature yields estimates that are weighted averages of those presented here. Moreover, as stated above, given the heterogeneity in worklife expectancy by child status, estimates of periodic human capital stocks based on previous estimates will not adequately reflect changes in fertility rates. For example, in the United States the total fertility rate rose from 1.84 in 1980 to 2.10 in 2006.[15]

The multinomial logit method of estimating worklife expectancy offers clear benefits. A variety of worklife tables, based simultaneously on race, education, sex, marital status and children can be constructed, as shown in this paper. This method could be used to construct additional tables based on other characteristics that are available in the CPS, as well as worklife tables and subsequent measures of human capital, for other countries for which similar data are available.

Acknowledgements

The authors would like to thank Jay Stewart and Bob Gaddie of the BLS for helpful comments.

References

Alter, G. and Becker, W. (1985). Estimating lost future earnings using the new worklife tables. *Monthly Labor Review* 108: 39-42.

Angrist, J. and Evans, W.N. (1998). Children and their parents' labor supply: evidence from exogenous variation in household size. *American Economic Review* 88: 450-477.

Blundell, R. and MaCurdy, T. (1999). Labor supply: a review of alternative approaches. In: Ashenfelter, O., Card, D. (Eds.), Handbook of Labor Economics, Volume 3, Elsevier Science.

[15] See http://www.census.gov/compendia/statab/tables/09s0082.pdf.

Booth, A.L., Jenkins, S.P., and Serrano, C.G. (1999). New men and new women? a comparison of paid work propensities from a panel data perspective. *Oxford Bulletin of Economics and Statistics* 61: 167-197.

Chun, H. and Lee, I. (2001). Why do married men earn more: productivity or marriage selection?" *Economic Inquiry* 39: 307-319.

Clark, K.B. and Summers, L.H. (1979). Labor market dynamics and unemployment: a reconsideration. *Brookings Papers on Economic Activity* 1979: 13-60.

Clark, K.B. and Summers, L.H. (1982). The dynamics of youth unemployment. In: Freeman, R., Wise, D. (Eds.), The Youth Labor Market Problem: Its Nature, Causes and Consequences. University of Chicago Press, Chicago, IL.

Dublin, L.I. and Lotka, A. (1930). The Money Value of Man. Ronald Press, New York.

Ebenstein, A. (2008). When is the local average treatment close to the average? evidence from fertility and labor supply. Journal of Human Resources, forthcoming.

Farr, W. (1853). The income and property tax. *Journal of the Royal Statistical Society of London* 16: 1-44.

Folloni, G. and Vittadini, G. (2010). Human capital measurement: a survey. *Journal of Economic Surveys* (this issue).

Fullerton, H. (1999). Labor force participation: 75 years of change, 1950 - 1998 and 1998 - 2025. *Monthly Labor Review* 122: 3-12.

Flinn, C.J. and Heckman, J.J. (1983). Are unemployment and out of the labor force behaviorally distinct labor force states?" *Journal of Labor Economics* 1: 28-42.

Gönül, F. (1992). New evidence on whether unemployment and out of the labor force are distinct states. *Journal of Human Resources* 27: 329-361.

Greene, W. (1993). Econometric Analysis, 2nd Edition. Prentice-Hall, Inc., Englewood Cliffs.

Hall, R.E. (1970). Why is the unemployment rate so high at full employment?" *Brookings Papers on Economic Activity* 1970: 369-402.

Hersch, J. and Stratton, L. (1994). Housework, wages, and the division of housework time for employed spouses. *American Economic Review* 84: 120-125.

Hersch, J. and Stratton, L. (1997). Housework, fixed effects, and wages of married workers. *Journal of Human Resources* 32: 285-307.

Jellal, M. and Wolff, F.-C. (2000). Shaping intergenerational relationships: the demonstration effect. *Economics Letters* 68: 255-261.

Jones, S.R.G. and Riddell, W.C. (1999). The measurement of unemployment: an empirical approach. *Econometrica* 67: 147-62.

Jorgenson, D.W. and Fraumeni, B.M. (1989). The accumulation of human and nonhuman capita, 1948-1984. In: Lipsey, R.E., Stone Tice, H. (Eds.), *The Measurement of Saving, Investment, and Wealth*. University of Chicago Press, Chicago, IL.

Jorgenson, D.W. and Fraumeni, B.M. (1992). The output of the educational sector. In: Griliches, Z. (Ed.), Output Measurement in the Service Sector, NBER Studies in Income and Wealth, Volume 56. University of Chicago Press, Chicago, IL.

Killingsworth, M. (1983). *Labor Supply*. Cambridge University Press, Cambridge.

Klerman, J.A. and Leibowitz, A. (1999). Job continuity among new mothers. *Demography* 36: 145-155.

Le, T., Gibson, J. and Oxley, L. (2006). A forward-looking measure of the stock of human capital in New Zealand. *Manchester School* 74: 593-609.

Le, T., Oxley, L. and Gibson, J. (2003). Cost- and income-based measures of human capital. *Journal of Economic Surveys* 17: 271-308.

Loh, E.-S. (1996). Productivity differences and the marriage wage premium for white males. *Journal of Human Resources* 31: 566-589.

Lovaglio, P. (2010). The estimation of human capital by administrative archives in a static and longitudinal perspective: the case of Milan. *Journal of Economic Surveys* (this issue).

Lundberg, S. (1988). Labor supply of husband and wives: a simultaneous equations approach. *Review of Economics and Statistics* 70: 224-235.

Maasoumi, E., Millimet, D.L. and Sarkar, D. (2009). Who benefits from marriage? *Oxford Bulletin of Economics and Statistics* 71: 1-33

Millimet, D.L. (2000). The impact of children on wages, job tenure, and the division of household labour. *The Economic Journal* 110: C139-C158.

Millimet, D.L., Nieswiadomy, M., Ryu, H. and Slottje, D. (2003). Estimating worklife expectancy: an econometric approach. *Journal of Econometrics* 113: 83-113.

Mincer, J. and Polachek, S. (1974). Family investments in human capital: earnings of women. *Journal of Political Economy* 82: S76-S108

Nieswiadomy, M. and Silberberg, E. (1988). Calculating changes in worklife expectancies and lost earnings in personal injury cases. *Journal of Risk and Insurance* 55: 492-498.

Nieswiadomy, M. and Slottje, D. (1988). Estimating lost future earnings using the new worklife tables: a comment. *Journal of Risk and Insurance* 55: 539-544.

Oxley, L., Le., T., and Gibson, J. (2008). Measuring human capital: alternative methods and international evidence, *Korean Economic Review* 24: 283-344.

Peracchi, F. and Welch, F. (1995). How representative are matched cross-sections? evidence from the CPS. *Journal of Econometrics* 68: 153-179.

Polachek, S. and Siebert, S. (1996). Family and labor market incentives: men's and women's work behavior and earnings. In: Menchik, P. (Ed.), Household and Family Economics. Kluwer-Nijhoff, Boston, MA.

Schoen, R. and Woodrow, K. (1980). Labor force status life tables for the united states, 1972. *Demography* 17: 297-322.

Skoog, G. and Ciecka, J. (2001). A Markov (increment-decrement) model of labor force activity: extended tables of central tendency, variation, and probability intervals. *Journal of Legal Economics* 11: 23-87.

Skoog, G. and Ciecka, J. (2002). Probability mass functions for additional years of labor market activity induced by the markov (increment-decrement) model. *Economics Letters* 77: 425-431.

Tano, D.K. (1991). Are unemployment and out of the labor force behaviorally distinct labor force states? *Economics Letters* 36: 113-117.

Tchernis, R. (2010). Measuring human capital and wage growth. *Journal of Economics Surveys* (this issue).

U.S. Bureau of Labor Statistics (1982). Tables of working life: the increment-decrement model. *BLS Bulletin 2135*. U.S. Department of Labor, Washington, D.C.

U.S. Bureau of Labor Statistics (1986). Worklife estimates: effects of race and education. *BLS Bulletin 2254*. U.S. Department of Labor, Washington, D.C.

U.S. Bureau of Labor Statistics (2009). *The Employment Situation: January 2009,* USDL 09-0117, U.S. Department of Labor, Washington, D.C.

Waldfogel, J. (1998). The family gap for young women in the united states and britain: can maternity leave make a difference? *Journal of Labor Economics* 16: 505-545.

Wöβmann, L. (2003). Specifying human capital. *Journal of Economic Surveys* 17: 239-270.

Table 1
Worklife Expectancies of Women by Race and Marital Status

Single White Females

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	16.39	14.62	13.76	26.01	24.77	23.32	30.27	29.32	27.21
40	12.16	9.54	8.40	18.91	17.17	15.16	21.88	20.78	18.71
50	8.26	4.69	3.25	11.82	9.62	6.74	13.48	12.15	9.18
60	4.46	1.92	0.66	5.94	3.58	1.48	6.64	4.17	2.00

Single Non-White Females

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	15.56	13.47	12.80	23.42	21.85	20.71	30.89	30.24	27.50
40	11.70	8.60	7.77	16.98	14.85	13.21	22.76	20.90	19.62
50	7.85	3.85	2.88	10.34	8.22	5.81	14.36	11.89	10.47
60	4.09	1.23	0.55	4.94	3.43	1.37	7.26	3.24	2.71

Married White Females

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	16.90	15.60	14.49	24.39	23.16	21.92	26.02	24.93	22.87
40	12.34	10.23	8.97	17.19	15.51	13.79	18.39	17.23	14.86
50	8.19	5.02	3.54	10.23	8.24	5.87	10.75	9.63	6.25
60	4.36	1.98	0.75	4.90	2.91	1.27	5.11	3.53	1.13

Married Non-White Females

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	16.37	14.90	13.91	22.13	20.56	19.70	26.71	26.06	23.32
40	12.00	9.57	8.55	15.44	13.43	12.10	19.42	17.62	15.92
50	7.83	4.30	3.26	8.91	7.01	5.06	11.76	9.35	7.37
60	4.02	1.28	0.65	4.06	2.78	1.17	5.86	2.31	1.56

Table 2
Worklife Expectancies of Single Women by Race and Children

Single White Females without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	17.20	15.67	14.50	26.55	25.22	23.71	30.73	29.85	28.25
40	12.50	10.07	8.74	19.00	17.24	15.16	21.80	20.60	18.89
50	8.32	4.84	3.29	11.82	9.63	6.73	13.43	11.85	9.13
60	4.47	1.93	0.66	5.94	3.58	1.48	6.63	4.16	2.00

Single White Females with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	15.70	13.78	13.17	25.48	24.39	23.05	28.49	27.59	25.29
40	11.92	9.31	8.30	18.84	17.14	15.21	20.91	20.02	16.72
50	8.32	4.84	3.29	11.82	9.63	6.73	13.04	11.15	6.76
60	4.47	1.93	0.66	5.94	3.58	1.48	6.62	4.13	1.81

Single Non-White Females without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	16.51	14.72	13.71	24.01	22.29	21.13	31.37	30.70	28.50
40	12.08	9.21	8.18	17.08	14.90	13.21	22.64	20.57	19.72
50	7.92	4.03	2.94	10.34	8.22	5.80	14.29	11.52	10.40
60	4.10	1.23	0.55	4.94	3.43	1.37	7.26	3.24	2.71

Single Non-White Females with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	14.80	12.51	12.12	22.84	21.48	20.41	30.00	29.52	28.09
40	11.44	8.35	7.64	16.89	14.81	13.26	22.04	20.60	18.97
50	7.92	4.03	2.94	10.34	8.22	5.80	13.89	10.56	8.79
60	4.10	1.23	0.55	4.94	3.43	1.37	7.19	2.97	2.29

Table 3
Worklife Expectancies of Married Women by Race and Children

Married White Females without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	17.50	16.33	14.97	24.92	23.59	22.31	26.85	25.82	24.16
40	12.61	10.62	9.17	17.26	15.56	13.79	18.35	17.05	15.04
50	8.24	5.14	3.56	10.23	8.24	5.86	10.68	9.32	6.21
60	4.36	1.99	0.75	4.90	2.91	1.27	5.10	3.53	1.13

Married White Females with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	16.35	14.96	14.06	23.89	22.81	21.65	23.92	22.97	20.92
40	12.11	10.02	8.89	17.12	15.47	13.84	17.38	16.49	13.02
50	8.24	5.14	3.56	10.23	8.24	5.86	10.39	8.85	4.59
60	4.36	1.99	0.75	4.90	2.91	1.27	5.14	3.57	1.16

Married Non-White Females without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	17.04	15.77	14.46	22.70	20.98	20.11	27.59	26.89	24.54
40	12.28	10.01	8.76	15.52	13.48	12.1	19.34	17.29	16.00
50	7.87	4.20	3.28	8.90	7.01	5.05	11.68	8.94	7.30
60	4.02	1.28	0.65	4.06	2.78	1.17	5.85	2.31	1.56

Married Non-White Females with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	15.74	14.13	13.39	21.60	20.24	19.4	26.14	25.69	24.39
40	11.74	9.31	8.44	15.36	13.40	12.14	18.89	17.58	15.67
50	7.87	4.20	3.28	8.90	7.01	5.05	11.47	8.44	6.38
60	4.02	1.28	0.65	4.06	2.78	1.17	5.85	2.30	1.51

Table 4
Worklife Expectancies of Men by Race and Marital Status

Single White Males

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	18.72	16.84	15.99	26.37	25.10	23.86	29.91	28.92	28.05
40	13.31	10.99	9.22	18.78	17.12	15.10	21.01	19.81	18.15
50	8.75	6.01	3.81	11.69	9.27	6.83	12.79	11.08	8.60
60	4.46	2.37	0.82	5.84	2.79	1.62	6.29	3.57	2.01

Single Non-White Males

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	17.75	16.11	15.01	21.96	20.57	19.82	26.99	26.12	25.29
40	12.75	11.00	8.85	15.45	13.80	12.28	18.40	17.63	15.65
50	8.44	6.26	3.85	9.43	7.49	5.47	10.67	9.07	6.89
60	4.36	2.44	0.93	4.61	2.63	1.35	5.12	1.92	1.53

Married White Males

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	22.72	21.04	19.35	29.07	28.08	26.54	31.76	30.93	30.15
40	15.54	13.45	10.68	20.29	19.11	16.77	22.54	21.65	20.06
50	9.45	6.93	4.13	12.16	10.46	7.61	13.81	12.73	9.92
60	4.41	2.57	0.84	5.57	3.28	1.85	6.53	4.90	2.47

Married Non-White Males

Age	Less than High School			High School			College		
	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	21.68	20.20	18.40	25.61	24.38	23.23	29.49	28.72	27.96
40	14.86	13.32	10.45	17.40	16.08	14.11	20.38	19.73	17.90
50	9.11	7.19	4.36	10.04	8.59	6.15	11.93	10.80	8.29
60	4.32	2.70	1.01	4.44	3.04	1.52	5.39	2.65	2.00

Table 5
Worklife Expectancies of Single Men by Race and Children

Single White Males without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	16.39	14.62	13.76	26.01	24.77	23.32	30.27	29.32	27.21
40	12.16	9.54	8.4	18.91	17.17	15.16	21.88	20.78	18.71
50	8.26	4.69	3.25	11.82	9.62	6.74	13.48	12.15	9.18
60	4.46	1.92	0.66	5.94	3.58	1.48	6.64	4.17	2.00

Single White Males with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	18.39	16.29	15.78	26.41	25.07	23.89	29.16	28.38	28.29
40	13.26	10.60	9.28	18.66	16.99	14.92	20.03	18.87	16.87
50	8.76	6.17	3.79	11.54	9.16	6.77	12.06	10.11	6.45
60	4.46	2.43	0.82	5.80	2.75	1.61	6.17	3.33	1.46

Single Non-White Males without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	18.07	16.57	15.28	21.28	19.94	19.18	26.09	25.19	24.30
40	12.81	11.28	8.89	14.95	13.40	11.95	17.75	16.96	15.06
50	8.44	6.39	3.85	9.26	7.37	5.41	10.45	8.84	6.69
60	4.35	2.49	0.93	4.57	2.59	1.35	5.08	1.86	1.50

Single Non-White Males with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	17.38	15.55	14.73	22.05	20.55	19.83	26.29	25.58	25.51
40	12.70	10.66	8.87	15.35	13.67	12.09	17.32	16.57	14.54
50	8.44	6.39	3.85	9.26	7.37	5.41	9.85	8.06	5.17
60	4.35	2.49	0.93	4.57	2.59	1.35	5.01	1.68	1.17

Table 6
Worklife Expectancies of Married Men by Race and Children

Married White Males without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	22.85	21.29	19.37	28.70	27.75	26.17	31.42	30.53	29.68
40	15.54	13.68	10.59	20.00	18.85	16.54	22.26	21.30	19.72
50	9.44	7.05	4.08	12.04	10.35	7.55	13.67	12.58	9.73
60	4.40	2.61	0.84	5.54	3.24	1.84	6.50	4.87	2.45

Married White Males with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	22.48	20.63	19.24	29.01	27.98	26.53	31.14	30.46	30.43
40	15.48	13.03	10.70	20.17	19.00	16.62	21.81	21.00	19.65
50	9.44	7.05	4.08	12.04	10.35	7.55	13.21	12.01	7.85
60	4.40	2.61	0.84	5.54	3.24	1.84	6.41	4.71	1.85

Married Non-White Males without Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	21.79	20.42	18.47	25.07	23.87	22.71	28.96	28.15	27.32
40	14.82	13.48	10.41	16.99	15.73	13.81	19.96	19.29	17.45
50	9.09	7.28	4.33	9.89	8.47	6.08	11.75	10.63	8.09
60	4.32	2.74	1.01	4.41	2.99	1.52	5.35	2.61	1.98

Married Non-White Males with Children

	Less than High School			High School			College		
Age	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive	Employed	Unemployed	Inactive
30	21.45	19.81	18.31	25.60	24.35	23.23	28.83	28.18	28.16
40	14.80	12.98	10.45	17.26	15.94	13.93	19.53	18.91	17.57
50	9.09	7.28	4.33	9.89	8.47	6.08	11.21	9.98	6.63
60	4.32	2.74	1.01	4.41	2.99	1.52	5.26	2.37	1.55