Accounting for Trends in the Labor Force Participation Rate of Older Men in the United States

Preliminary

David Blau and Ryan Goodstein

University of North Carolina at Chapel Hill

May 16, 2006

We are grateful for financial support from the National Institute on Aging (Grant P30 AG024376)

Corresponding author: Ryan Goodstein
University of North Carolina at Chapel Hill
Department of Economics, Gardner Hall
Chapel Hill, NC 27599-3305
ryan_goodstein@unc.edu

Abstract

After nearly a full century of decline, the Labor Force Participation Rate (LFPR) of older men in the U.S. leveled off in the 1980s, and began to increase in the late 1990s. We use synthetic panel data from 1962 to 2005 to model the LFPR of older men, with the aim of accounting for these trends. We find that the decline in the LFPR of older men cannot be explained by demographic trends, changes to Social Security and private pension rules, health, or availability of employer provided retiree health insurance. We attribute the increase in the LFPR in recent years mainly to changes in the distribution of education among older men. The share of high school dropouts among older men declined from 65% in 1970 to 19% in 2005, and dropouts have a much lower LFPR than more educated men. Our results also suggest that increases in the Social Security delayed retirement credit, increased employment of the wives of older men, and the shift from Defined Benefit to Defined Contribution pension plans all may have contributed to the recent LFPR gains.

1. Introduction

The Labor Force Participation Rate (LFPR) of older men in the United States trended downward for much of the twentieth century. The magnitude and duration of this trend is remarkable. Among men aged 65 and older, the LFPR fell from 68% in 1900 to 19% in 1980 (Moen, 1987). However, by the end of the 1980s this long-run trend appeared to have ended. More recently, the LFPR of some age groups began to rise. After falling to a 20th century low of 24% in 1985, the LFPR of men aged 65 to 69 increased to over 33% in 2005. The participation rate for men aged 60 to 64 increased from 55% in 1985 to 58% in 2005 (Figure 1). With the population aging rapidly in the next two decades, it is important to understand why the downward trend in the LFPR of older men ended, and whether we can expect the recent increases in the LFPR of older men to persist.

The goal of this paper is to quantitatively assess alternative explanations for the trends described above. The main explanations considered include changes in (1) Social Security benefits; (2) coverage and type of employer-provided pensions; (3) the availability of employer provided retiree health insurance (EPRHI); (4) health; (5) employment of older married women; and (6) the composition of the older male population. We incorporate each of these candidate explanations in a unified framework. There is no single data source with which all of these alternatives can be examined, so we assemble data from several sources to generate a synthetic panel spanning the period 1962 to 2005. This preliminary draft merges data from the Current Population Survey (CPS), the Survey of Program Participation (SIPP), and the Social Security Administration (SSA). In ongoing work, we are assembling additional data from the National Health Interview Survey (NHIS) and the 1940, 1950, and 1960 Censuses of Population. Data from the various sources are merged by age and cohort, where cohorts are defined by single year of birth year and four categories of educational attainment.

Some of the proposed explanations that we analyze are not new; for example, many studies have analyzed the impact of increased generosity of Social Security retirement and disability benefits on the decline in the older male LFPR in the 1960s and 1970s (e.g. Hurd and Boskin, 1984; Parsons, 1980, Moffitt, 1987; Krueger and Pischke, 1992). The connection between trends in pension coverage and type and employment trends of older men has also been analyzed (e.g., Anderson, Gustman and Steinmeier, 1999; Friedberg and Webb, 2005), as has the impact of EPRHI (Madrian, 1994; Blau and Gilleskie, 2001a). A novel contribution of our study is consideration of increasing LFP of married women as a potential explanation for trends in LFP among older men. The increase in LFP of married women has been widely documented and analyzed (Costa, 2000), but the impact of this trend on the LFPR of older men has not been evaluated. This is a promising avenue to explore because the proportion of older married women with significant work experience and substantial Social Security and pension benefits has increased. Wives are generally younger than their husbands, so if LFP choices are coordinated within a household, then husbands may delay retirement until their wives become eligible for retirement benefits.

An important contribution of this study is to assess alternative explanations in a unified framework and during a period in which there was a major reversal of the long run downward trend in the LFPR of older men. This setting provides a challenge to any mono-causal explanation: such an explanation will have to account for many years of decline, a roughly 15 year period of no change, and the recent increase.

Our preliminary findings indicate that the most likely "explanation" for the downward trend observed over the period 1962-1985 is unobserved changes in preferences, constraints, and institutions that are not incorporated in our specification. In some specifications, we find that changes in Social Security benefits were important. Social Security benefits became increasingly generous from 1962 through 1976. Subsequently, Social Security reforms reduced retirement benefits. This pattern is consistent with declining LFP followed by a leveling off. However, the results for Social Security are quite sensitive to specification: allowing for either a linear birth year trend or an unrestricted set of birth year fixed effects eliminates almost all of the impact of Social Security. The recent increase in labor force participation turns out to be mainly a compositional effect: low-participating high school drop outs have been rapidly replaced in the labor force by higher-participating high school graduates, college attendees, and college graduates. We find that the increasing generosity of the Social Security delayed retirement credit can account for some of the increase in LFPR. Since 1987 lifetime Social Security income has been increasing for men who claim benefits at an older age relative to those who claim at younger ages, other things equal. We also find that the shift from Defined Benefit to Defined Contribution pensions and the increase in employment of older married women may have played a role in recent increase in older male LFP.

2. Background

The long-run trend of declining labor force participation among older men is not unique to the U.S. Similar patterns are found in other industrialized countries, suggesting that the principal explanations for the movement towards earlier retirement are common to all developed nations. Analysts generally attribute the long-run downward trend to rising lifetime income as a result of growing real wages (Costa, 1998; Burtless and Quinn, 2000). Other things equal, wealthier men would have a higher demand for leisure, and could more readily "afford" to retire.

However, the increase in LFP rates of older men since the mid 1990s has occurred during a period when real earnings have continued to increase in the U.S. This suggests that the wealth effect may have diminished in importance or that other forces are now dominating the wealth effect. Costa (1998) cites a number of studies which suggest that the effect of retirement income on retirement behavior is not as strong in recent years, in part because retirement has become more attractive due to changing social norms and the development of leisure technologies that have made retirement more affordable and enjoyable. Kopecky (2005) calibrates a model that uses increases in real wages and declines in the price of goods that are complementary with leisure to explain labor force participation rates since 1850. Her model captures the general pattern of a declining LFPR, but fails to predict the leveling and reversal of the trend since the mid 1980s and does not capture differences in the rate of decline by age group.

Circumstantial evidence suggests that changes to the generosity and structure of Social Security may have had some effect on LFP rates of older men (Ippollito, 1990; Stewart, 1995). Benefits steadily increased from the inception of Social Security in 1935 through the early 1970s, coinciding with declines in the LFPR of older men. The end of the downward trend in the 1980s coincides with several changes to Social Security policy that have encouraged men to work at older ages. Amendments in 1977 reduced benefits for men who turned 65 beginning in 1982. The 1983 amendments increased the Delayed Retirement Credit (DRC), which is an adjustment to benefits for entitlement after the normal retirement age. The adjustments took effect in 1987 through 2005. The result is that by 2005 the expected present discounted value of lifetime benefits is on average independent of the age of entitlement. The 1983 amendments also increased the normal retirement age from 65 in 1999 to 66 in 2005, effectively reducing lifetime social security benefits[1]. Finally amendments in 1983 (effective in 1990) and in 2000 modified the Social Security Earnings Test (SSET), first reducing and then eliminating the implicit tax on earnings for men at and above the normal retirement age.

Economists have not reached a consensus on the quantitative effect of changes to Social Security benefits on LFP rates of older men. Moffitt (1987) uses time-series data to assess the impact of increases in benefits from the 1950s through the 1970s on the LFPR of older men. He concludes that unexpected Social Security policy changes can explain no more than 20% of the observed decline in the 1970s. However, in a similar analysis using a longer time-series, Stewart (1995) finds that up to 40% of the change in the LFPR of older men between 1965 and 1990 can be attributed to changes in Social Security benefits. Researchers have also used panel data to study the effects of Social Security, taking advantage of richer individual-level data to assess the impact of particular SS amendments. Hurd and Boskin (1984) find that increases in SS benefits between 1970 and 1972 account for nearly the entire decline in the LFPR of older men between 1969 and 1973. In contrast, Kreuger and Pischke (1992) use synthetic panel data and find that the 1977 amendments had almost no impact on LFP rates of older men in the 1970s and 1980s. There is also disagreement over the role of Social Security Disability Insurance in explaining the decline in participation of older men at ages before eligibility for retirement benefits (Parsons, 1980; Bound, 1989; Chen and van der Klaauw, in press).

Changes in the availability and structure of private pension plans may also have had some impact on LFP rates of men at older ages. Traditionally, firms have offered

[1] A person who retires at the normal retirement age of 66 in 2005 collects Social Security benefits for a full year less than an equivalent individual who retired at the normal retirement age of 65 in 1999, holding constant life expectancy. The reduction in lifetime benefits is also reflected in an increased penalty associated with claiming benefits before the normal retirement age.

their employees Defined Benefit (DB) pension plans, in which benefits are determined as a function of age at retirement, tenure, and average earnings. DB plans generate incentives for workers to leave their jobs after becoming eligible, as the expected present value of lifetime benefits typically decreases for each additional year worked after becoming eligible (Lazear, 1986). However, in recent years employers have increasingly offered Defined Contribution (DC) plans in place of DB plans. Participation in DB plans fell from 84% in 1980 to 33% in 2003 among full-time employees of medium and large private firms, with a corresponding increase in DC plan participation (Employee Benefit Research Institute, 2005). In DC plans pension wealth accumulates as a function of employer and employee contributions and the returns on those contributions. DC plans do not impose disincentives for working at older ages, because the pension value depends only on the account balance rather than age or job tenure. As participation in DB plans has declined, disincentives for working at older ages associated with DB have become less important. However, these pension plan changes appear at older ages only with a significant lag, since the changes often affect only new employees.

There is good reason to think that increased LFP among married women may have contributed to the reversal in LFP rate trends of older men since the mid 1980s. The LFP rate of married women has nearly tripled since 1950 (Costa, 2000). The presence of a working spouse in the household could have opposing effects on the husband's decision to work. Men with career working wives are likely to have higher lifetime household income, and therefore may be able to afford earlier retirement (Gustman and Steinmeier, 2000). However, if husbands and their working wives coordinate their retirement behavior, men with working spouses may be more likely to work at older ages, other things equal. Hurd (1990), Blau (1998), Gustman and Steinmeier (2000), and others find that working husbands and wives tend to retire at the same time. As noted earlier, husbands may wait until their younger wives become eligible for Social Security or pension benefits before leaving the labor force. In addition, husbands may simply value leisure more highly when it is shared with a spouse. Coile (2004) finds evidence that husbands are less likely to retire as their wives' retirement benefits (from Social Security and pensions) increase, and that men strongly prefer leisure shared with their spouse.

The effect of employer provided retiree health insurance (EPRHI) on LFP may also have contributed to observed trends. Eligibility for public health insurance for the elderly (Medicare) begins at age 65 in the U.S. Men under the age of 65 who choose to retire without EPRHI must bear the cost of purchasing health coverage from another source or go uninsured, bearing the full brunt of medical expenditure risk. Blau and Gilleskie (2001a) estimate that roughly 13% of the decline in the LFP rate of men aged 55 to 59 between 1965 and 1984 can be attributed to increases in the availability of EPRHI. If the availability of EPRHI has declined in recent years then LFP rates of older men may be increasing as a result (see also Madrian, 1994).

Trends in health of older men have been discounted as a potential explanation for the observed trends in LFP rates of older men. Although health has a major impact on the LFPR of older men (Bound, 1991; Peracchi and Welch, 1994; Blau and Gilleskie, 2001b), recent trends in average health have been positive rather than negative (Burtless and Quinn, 2000). Similarly, changes in the occupational composition of the labor force are unlikely to have caused much of the changes in LFP rates of older men. Costa (1998) finds that the decline of the farming sector did not contribute to declines in LFP rates in the early 20th century. However, Quinn (1999) speculates that shifts in the U.S. economy from manufacturing to service may be contributing to recent increases in LFP rates of older men, as the physical demands of working may have declined. We control for trends in health and sectoral shifts in employment in our model. We also examine the role of the changing educational composition of the labor force.

3. Empirical Model

Economic models of employment decisions at older ages are generally based on the life cycle framework, and usually incorporate Social Security, pensions, earnings, wealth, and health. Life cycle models may also account for a liquidity constraint, savings decisions, uncertainty, and many other relevant constraints and features of behavior. Such models can be solved numerically and estimated with longitudinal data (e.g., Blau and Gilleskie, in press; French, 2005; van der Klaauw and Wolpin, 2005). Our analysis of trends over a period of more than 40 years requires use of synthetic panel data, which makes structural estimation much more difficult. Hence we specify a simple regression model that can be rationalized by a life cycle framework, but we do not specify the underlying life cycle model. Our empirical specification can be interpreted as a linear approximation to the employment decision rule implied by a life cycle model with a perfect capital market and imperfect information (see Moffitt, 1987). The model is:

$$L_{it} = \gamma_0 + \gamma_1 B_i + \gamma_2 E_i + \gamma_3 t_i + \gamma_4 X_{it} + \gamma_5 L_{wit} + \gamma_6 Y_{wit} + \gamma_7 H_{it} + \gamma_8 P_i + \gamma_9 SSW_{i65} + \gamma_{10} SSW_{i70} + \gamma_{11} AIME_i + u_{it}$$

where

L_{it} = 1 if man *i* is employed in year *t*, and 0 otherwise;

B_i is a vector of birth year fixed effects, or a polynomial in birth year

E_i is a vector of education fixed effects;

t_i is vector of age fixed effects;

X_{it} is a vector of demographic characteristics, including race and marital status;

L_{wit} =1 if the man is married and is wife is in the labor force, and zero otherwise;

Y_{wit} is the wife's earnings in the previous year (= 0 if not married)

H_{it} is an indicator for poor health;

P_i is a vector of pension and health insurance indicators;

SSW_{ia} is a measure of Social Security wealth conditional on retirement at age *a*, *a*=65, 70;

$AIME_i$ is Average Indexed Monthly Earnings (the earnings base on which Social Security benefits are calculated), conditional on full time work until 65.

In the remainder of this section, we describe the specification of the key explanatory variables.

Social Security Wealth:

SSW_{i65} is an approximation to the expected present discounted value of lifetime Social Security retirement benefits a man would receive conditional on claiming his benefit at age 65, assuming employment at the mean age-specific earnings for a man of his cohort in every year of his working life through age 64, followed by permanent exit from the labor force at age 65. SSW_{i70} is defined similarly. Following Krueger and Pischke (1992), Social Security wealth is computed as $SSW_{ia} = \sum_{j=a}^{T} SSB_{ia}(1+r)^{a-j}$, *a* = 65, 70; where *SSB* is the Social Security benefit given the above assumptions about earnings and retirement. Social Security wealth is a function of lifetime earnings (*AIME*), the

Social Security benefit rules in effect for the man's birth cohort, life expectancy *T*, and the interest rate *r*, here set at 2%. We assume the individual survives with certainty to his expected age at death, *T*, in order to simplify the calculations. SSB_{td} is the monthly Social Security disability benefit the individual would receive at age *t* conditional on full time work through period *t-2* conditional on good health, then bad health in periods *t-1* and *t* and zero hours worked in periods *t* and *t-1*. SSB_{td} is set to zero for *t* greater than or equal to the normal retirement age, because at this age disability benefits are converted to retirement benefits.

In this specification, SSW_{i65} captures the wealth effect of Social Security, and SSW_{i70} captures the incentive for later retirement induced by the Delayed Retirement Credit (DRC). In a simple life cycle model with an actuarially fair Social Security program and a perfect capital market, the Social Security retirement program would only affect retirement behavior through a wealth effect (Moffitt, 1987). Conditioning the SSW measure on retirement at age 65 is arbitrary, but the results are not very sensitive to other assumptions since SSW conditional on other ages of retirement is very highly correlated with SSW_{i65}. If the capital market is not perfect, Social Security can affect retirement behavior through other channels as well, such as the early retirement penalty, the delayed retirement credit, and the earnings test. We account for the delayed retirement credit because it increased from 1% to 8% per year from the 1980s to the early 2000s. We experimented with many other specifications designed to capture the effects of other channels. The main findings are very similar in all cases, so we focus on a simple and easily interpretable specification. These SSW measures do not explicitly account for expectations about future changes to SS rules. Therefore the current rules as captured by our measures may operate in part by helping to forecast future rule changes.

The SSW measures are computed based on mean earnings at each age for men of a given birth year and educational attainment. There are many nonlinearities in the Social Security benefit formula, so ideally we would compute SSW for each man using his individual earning history, and then aggregate to the cohort level. However, we do not have true panel data, so this is impossible. To ensure that variation in Social Security benefits across cohorts is driven by program rules and not by differences in earnings across birth years and education groups, we include in the model the measure of lifetime earnings, *AIME*, used to compute the benefits.

Pensions and Health Insurance: Unlike Social Security benefits, current and future rules governing private pension benefits vary significantly across individuals, depending on their employer. We lack data on pension plan features, so we include only indicator variables for Defined Benefit and (D_b) and Defined Contribution (D_c) coverage. We assume that the only variation in health insurance coverage that might affect the LFP decision is whether EPRHI is available at ages less then 65. For $t < 65$ we define D_{ht} = 1 if the man has EPRHI and D_{ht} = 0 if the man does not have EPRHI. For $t \geq 65$, D_{ht} = 0.

Other Variables: If the man is married, his wife's employment status, and her earnings are included to capture the potential impact of increased employment and earnings of older married women on the behavior of their husbands. It is likely that these variables are jointly determined with the husband's employment decisions, but we do not account for this in estimation. Health is included as a control variable. In some specifications, we included measures of self-employment, occupation (manual versus non-manual), and the lagged hourly wage rate. However, these variables can only be measured for workers, and including them results in a significant reduction in sample size. The effects of the main explanatory variables of interest were robust to their inclusion, so we omit them from the results reported here. The birth year, education, and age fixed effects control for any unobserved differences in behavior across cohorts and age that might be correlated with the explanatory variables of interest.

4. Data

We estimate the empirical model on a synthetic panel data set constructed from a variety of sources, including (to date) the Current Population Survey (CPS), the Social Security Administration (SSA), and the Survey of Income and Program Participation (SIPP). Individual records from the CPS and SIPP are aggregated into cohorts defined by single year of birth and education (high school dropout; high school graduate; some college; college graduate). The SSA data are available only in aggregate form. Disaggregation by education is important because of substantial differences in earnings trends by level of education and major changes in the educational composition of the older population over time. The aggregated data from each source are matched by cohort and year. The result is a synthetic panel data set covering 232 cohorts (4 education groups * 58 birth years, 1892 to 1949) between 1962 and 2005, although no cohort has data for all of these years, and some cohorts are dropped due to small sample sizes. Data from 1963 are dropped because there is no information on education in the 1963 CPS. Because we focus on LFP behavior at older ages, we include only cohorts that can be observed at ages 55 to 69 in our sample. The estimation sample contains 2,453 observations[2].

The foundation of our data is the March supplement to the CPS from 1962 to 2005. These data are used to construct measures of demographic characteristics, labor force participation, and earnings of older men and their spouses. Figure 2 shows the trend in male LFP averaged over all education groups and ages for the period 1962-2005. The LFPR was flat in the 1960s, and then fell from 70% in 1970 to 52% in the mid 1980s. The LFPR rose by about five percentage points beginning in the late 1990s. Figure 3 shows trends in the education distribution during this period, illustrating the remarkably rapid shift from an older labor force consisting mainly of high school dropouts in 1962 to one containing mainly high school graduates and college attendees today. Figure 4 shows that the LFPR is on average about 10 points lower for high school dropouts than for high school graduates, so educational composition effects may be important. Figure 5 shows the trend in LFP among wives of the men in the sample (with zeros for men who are not married). The LFP of wives began to increase in the late 1980s, followed several years later by the increase in male LFP. Figure 6 shows the trend in health. In this preliminary version of the paper, our measure of health uses data from the CPS. We follow Peracchi and Welch (1994) in defining a man to be in bad health if he did not work full time in the previous week or in the previous year and he attributes that choice to disability. Because this measure depends on labor force status in previous periods it is surely endogenous with respect to LFP choice in the current period. In future versions of this paper we will use data from the National Health and Interview Survey (NHIS) to measure trends in health status. The CPS measure shows a decline in the

[2] Only 1,458 observations are available for specifications that include the lagged wage and lagged job characteristics.

incidence of poor health from 18-20% in the early 1970s to around 12% in the 1990s.

We use data from various issues of the *Annual Statistical Supplement* to the *Social Security Bulletin* published by the SSA to measure Social Security benefits by cohort. These data are combined with CPS earnings data to form earnings histories that are input into the ANYPIA Social Security Benefit Calculator available on the SSA website[3]. The ANYPIA program calculates benefits using the appropriate benefit rules by birth year, so variation in benefit rules resulting from changes to the average indexed monthly earnings (AIME) and primary insurance amount (PIA) formulas, the normal retirement age, the SSET, and delayed retirement credit are all incorporated. Details on the construction of the benefit measures are provided in the data appendix. We then use measures of life expectancy from annual life tables published by the National Center for Health Statistics to convert benefits to wealth.

Figure 7 illustrates trends in real SSW for entitlement at age 65. SSW follows an upward trend during the entire period, with slower growth in the 1980s than in other periods. The trends are a result of changes in SS rules and changes in the lifetime average earnings on which benefits are based. Note that SSW is measured gross of payroll taxes, but payroll taxes are captured implicitly by the AIME measure. Figure 8 shows the trend in the SSDI benefit, averaged over ages 55-65. The trend in the SSDI benefit is generally upward, but is more irregular than the retirement benefit trend because benefits are age-specific, and the rules used to compute benefits are the same for all awardees in each year regardless of birth year. Finally, we include lifetime average real monthly earnings as a control variable in order to avoid attributing the effects of earnings growth to Social Security. Figure 9 shows trends in lifetime average monthly earnings by education group, and highlights the increasing earnings disparity between low-wage and high-wage workers.

We use data from topical modules of various SIPP panels to measure participation in DB and DC pensions, and availability of EPRHI. Respondents are asked detailed questions about pension benefits provided by the current and past employers. Respondents are asked if they are covered by EPRHI only if they are receiving income from a private pension at the time of the survey. To deal with small sample sizes for

[3] The ANYPIA program is available at http://www.ssa.gov/OACT/ANYPIA/anypia.html

early birth cohorts, our measures of participation in DB and DC and availability of EPRHI are averaged across birth years 1900 to 1910 separately by education group. Data for the earliest birth years likely suffer from mortality bias. There are additional biases for our measure of EPRHI, as individuals covered by EPRHI are more likely to be retired and receiving retirement income than those not covered by EPRHI. Details on how DB, DC, and EPRHI indicators are constructed are included in the data appendix. Figure 10 shows that DB pension coverage trended upward until the 1990s and only began to decline in the late 1990s. DC pension coverage increased slowly but steadily during the entire period. EPRHI coverage was roughly constant during the entire period. These trends are for men aged 55-69, and therefore do not reflect economy-wide trends fully until the latest years.

5. Estimation Results

Regression results for three specifications are shown in Table 1. The first has no controls for birth year, the second includes a linear birth year trend, and the third includes a full set of birth year fixed effects. The test statistics in the last row show that the specification without any control for birth year is strongly rejected against the specification with a linear birth year trend, and the latter is rejected against the unrestricted birth year specification.

Our discussion above led us to expect a negative effect of SSW_{i65} and a positive effect of SSW_{i70}. The results in column 1 show negative effects of both *SSW* measures, but the anticipated pattern does appear in columns 2 and 3. Focusing on column 3, the estimates imply that a $100,000 increase in Social Security wealth at ages 65 and 70 is to reduce LFP by 5 percentage points (-.1739 + .1212). Clearly, the estimated effect is sensitive to the specification of cohort effects. We evaluate the implied effect of the actual changes in *SSW* below. Higher monthly SSDI benefits are estimated to reduce LFP, and the magnitude of the coefficient estimate is fairly robust across the alternative specifications at -.03 to -.06 per thousand dollars. Average monthly earnings are estimated to have a positive impact on LFP.

DB pension coverage is estimated to reduce LFPR of older men, but the magnitude of the effect is very sensitive to specification, and the estimate is small and insignificantly different from zero in the birth year fixed effects specification. DC pension coverage is estimated to increase LFP by 10 points in column 3, but the estimates are quite different in the other columns. EPRHI coverage reduces LFP by 8-9 points in columns 2-3. The effect of the wife's labor force participation, which is interacted with the man's age, is close to zero at age 55 and increases by two points per year of age. The effect at age 65 is estimated to be +.26. Bad health has a large negative impact on LFP.

Figure 11 shows the actual and fitted trend in LFP for men aged 55-69 from 1962 through 2005, based on specification 3 (unrestricted birth year fixed effects). The model fits quite well.

The main issue of interest is how the results can be used to account for the LFP trends described above. We use the results to simulate several counterfactual experiments, in order to determine which, if any, of the explanatory variables can account for the trends. Figure 12 shows the result of an experiment in which Social Security retirement rules are fixed at their 1978 values while other variables take on their actual values[4]. We picked the 1978 rules because these were among the most generous rules in the history of Social Security for men claiming benefits at the normal retirement age or earlier. Benefit amounts were increasing prior to 1978, and subsequent reforms all reduced the overall generosity of Social Security benefits. If changes to Social Security benefits are an important contributor to the downward LFPR trend, then fixing benefits at their 1978 level should result in a much flatter LFPR trajectory. Figure 12a shows the results of a simulation based on the specification in column 1 of Table 1, which has no birth year controls. The simulated counterfactual trajectory based on the 1978 rules is in fact substantially flatter than the actual or predicted trend, suggesting that changes in Social Security retirement rules can account for a substantial portion of the downward trend from 1962 through the mid 1980s if cohort effects are omitted. However, Figure 12b shows that this finding is not robust: using the column 3 specification with birth year fixed effects, the simulated counterfactual trend is very similar to the observed trend. According to these results, the decline would have occurred even if there had been no changes in Social Security retirement rules. Figure 13 shows that the same is true for the

[4] Benefits are computed for each cohort as if they turn 62 in 1978 (birth year 1916), but using their actual earnings history. This allows us to capture the effect of rule changes while holding earnings constant.

rules that determine Social Security Disability benefits. This is not surprising, since the rules are very similar for disability and retirement. Figure 14 shows that changes to average lifetime earnings since 1970 also cannot explain the downward trend: the simulation indicates that if average lifetime earnings had remained constant at their 1970 levels, the downward trend would have been nearly identical. Figure 15 shows that birth year effects can "explain" the downward trend in the LFPR. In this simulation, birth year is fixed at the 1970 level by age (i.e. birth year is set to 1901 for 69 year olds, 1902 for 68 year olds, etc.). The results indicate that if birth year effects are held constant, the LFPR trajectory would have been much flatter, and the LFPR in 1990 would have been slightly *higher* than in 1970.

Thus, these results imply that changes in Social Security benefits are not a major cause of the decline in LFP of older men. This finding is consistent with the results of Moffitt (1987) and Krueger and Pischke (1992), who also use synthetic panel or time series data, but is inconsistent with the results of Hurd and Boskin (1984) using longitudinal data on individuals. Blau (1994) used longitudinal data on individuals and found that Social Security is important in accounting for variation across individuals in the timing of labor force exit, but that trends over time in Social Security benefits could not explain the secular trend in the exit rate from the labor force over the period 1961-1979. Peracchi and Welch (1994) reached a similar conclusion.

Table 2 presents an accounting exercise that quantifies the effect of selected factors on decline in LFPR from the period 1965-1970 to 1988-1993. The actual LFPR declined by 19.1 percentage points over this time, while simulated LFPR based on our model declined by 17.3 points. Essentially the entire decline can be attributed to birth year effects; when birth year is fixed at the 1970 level simulated LFPR increases slightly. Changes to SS disability rules, pensions, spouse's LFPR, and the educational distribution have little impact over this time period. Fixing the SS benefit rules at 1978 levels results in a simulated decline of 16.9 percentage points over this time period, almost identical to the observed decline.

We now describe the extent to which our estimates can explain the increase in the older male LFPR in recent years. We find that the increase can be explained by changes in the education distribution. More educated men participate in the labor force at higher rates, and the proportion of older men with high school degrees or more has increased significantly since 1970. Figure 16 shows the implications of our estimates: if the education distribution of the older male labor force was fixed at its 1985 level, LFP would have continued to decline in the 1990s and 2000s rather than increase.

Changes to Social Security benefit rules in the 1980s and 1990s appear to have had a modest positive impact on LFP behavior of older men. Figure 17 shows the results of an experiment in which benefit rules are fixed at their 1984 levels while other variables take on their actual values[5]. The figure shows a small impact of this hypothetical scenario.

Increases in the LFPR of the wives of older men since 1985 have had a modest impact on the older male LFPR. Figure 18 shows that if the fraction of men who have working spouses remained constant at the 1985 level, the LFPR of older men would have increased a bit more slowly than in the baseline case.

Table 3 quantifies the effect of selected factors on the increase in LFPR from the 1988-93 period to 2000-05. Actual LFPR increased by 4.7 percentage points over this time, while our model predicts an increase of 3.2 points. Changes in the education distribution can "explain" about 200% of the observed LFPR increase. Fixing Social Security benefit rules at their 1984 levels suggests that benefit changes since 1984 can account for one third of the increase in LFPR. Increasing employment of wives can explain 28% of the observed increase. The switch from DB to DC pensions can also account for about one quarter of the observed increase. Changes in other factors shown in the table go in the wrong direction.

Other results not shown here indicate that changes in health and the availability of EPRHI cannot account for the observed changes in the LFPR of men. We also estimated specifications that included lagged wages and job characteristics, and found that these variables also cannot account for trends in LFPR. As noted above, we estimated many different specifications with alternative approaches to measuring Social Security benefits. In every case, the results imply that controlling for birth year fixed effects, Social Security cannot explain the decline in labor force participation.

[5] Benefits are computed for each cohort as if they turn 62 in 1984 (birth year 1922) using their observed earnings history.

6. Discussion

We are unable to attribute the long-run decline in LFP of older men to any of the main hypotheses investigated in this paper. Our results suggest that the long run trend has occurred as a result of unobserved changes in preferences, constraints, or institutions across cohorts. There are at least two omitted variables in our model that in principle could potentially help explain the downward trends. We do not include a measure of net worth in our model; a trend toward increased real wealth could be responsible for some part of the decline in participation. However, we do include a measure of lifetime average monthly earnings which is surely correlated with trends in wealth. Furthermore, empirical evidence on the impact of wealth on retirement suggests that the effect is very small (Blau 1994; Diamond and Hauman, 1984). We also omit from our model the price of goods that are complementary with leisure. As noted earlier, Kopecky (2005) demonstrates that declines in the price of such goods can help explain the long-run downward trend. However, she does not control for a time trend.

We find that the recent increase in the LFPR of older men is due to several contributing factors. The most important is a compositional effect: low-participating high school drop outs have been rapidly replaced in the labor force by higher-participating high school graduates, college attendees, and college graduates. This result is robust to all specifications that we estimated, including those that control for wages. Education is likely correlated with unobserved factors that influence preference for work, such as motivation. Men with higher levels of education also enter the labor force at later ages, so it is possible that these men work longer to reach retirement incentive milestones (e.g. tenure rules associated with DB plans) or may simply prefer to exit the labor force at later ages. Changes to Social Security rules after 1984 also appear to have had some modest impact on the increase in LFP. Rule changes after 1984 increased incentives for working at older ages relative to younger ages. Our results thus suggest that shifts in the relative value of benefits at different ages may be important in determining when older men choose to exit the labor force. Increasing employment of older married woman and the shift from DB to DC pensions may have contributed to the recent increase in LFP.

Appendix

A1. Data – Social Security Benefits

Our analysis requires measures of mean Social Security benefits by cohort. Cohorts are defined by sex (male), birth year (1900 to 1945), and education group (less than high school; high school graduate; some college; college graduate). Cohort Social Security benefits are a function of Social Security regulations (which vary by birth year) and mean earnings history of each cohort (which varies by birth year and education group). Section A.1.1 details the methods used to construct earnings histories, and Section A.1.2 describes how these earnings histories are used to compute cohort specific measures of Social Security benefits.

A1.1 Cohort Specific Earnings Histories

We construct mean earnings for each cohort at ages 27 through 70 using data from Current Population Surveys (CPS) between 1962 and 2005 and from editions of the *Annual Statistical Supplement* published by the Social Security Administration (SSA) between 1973 and 2005. The SSA data contain median earnings of male workers by age group and earnings year, and can be transformed to median earnings by age and birth year. For instance, median earnings of workers born in 1945 are $7,405 at age 27 (from the cell labeled age group 25 – 29 in earnings year 1972), are $12,762 at age 32 (from the cell labeled 30 – 34 in earnings year 1977), and so on through age 57. We ignore earnings data for ages below 27 and above 57 to limit biases due to non-participation. We assume that earnings at age 58 to 64 remain constant in real terms at the age 57 level (as in Englehardt and Gruber, 2004)[6]. Because only 35 years of earnings are incorporated in the Social Security benefit computation, earnings at ages 26 and younger do not affect Social Security benefits of workers who work in every year from age 27 until retirement at age 65, assuming that earnings before age 26 are less than at older ages.

The SSA data are problematic for our purposes in at least three ways. First, median earnings are reported instead of mean earnings. To resolve this issue, we

[6] We use the Average Wage Indexing series provided by the SSA to adjust for inflation. We assume no inflation in years 2006 and later. For instance, for cohorts born in 1945, we use observed inflation rates in 2003, 2004, and 2005 to create an earnings "history" for ages 58, 59, and 60. Earnings at ages 61 to 70 are assumed to be nominally equal to earnings at age 60. Calculation of social security benefits (including adjustments to the PIA bend points and the Maximum Family Benefit formula) assume 0.0% inflation in years after 2006, accordingly.

calculated the ratio of mean earnings to median earnings in the CPS. These ratios were used to convert the median earnings reported in the *Annual Statistical Supplement* to means. In the March supplements of the CPS from 1962 through the present, respondents report their earnings from the previous year. These data are used to calculate mean earnings, median earnings, and their ratio by age group for earnings years 1961 through 2004[7,8]. We then use OLS to estimate models of the form $MM_{ay} = \alpha_{a0} + \alpha_{a1} y + \varepsilon_{ay}$, where MM_{ay} is the mean-median ratio for age group *a* in birth year *y*. The OLS estimates are used to generate a predicted value of MM_{ay} for each age birth-year cell. Each value for median earnings reported in the *Annual Statistical Supplement* was then multiplied by the predicted value of the mean-median ratio for the corresponding age cell to create measures of mean earnings by age group and earnings year.

A second issue with our data is that values for some birth year and age cells are not available in the *Annual Statistical Supplement*. Earnings at age 27 are missing for most years between 1910 and 1932; earnings at age 32 are missing for most years between 1905 and 1927; earnings at age 37 are missing for birth years between 1900 and 1922; and so on. Further, we have no data on mean earnings at ages 28 to 31, 33 to 36, and so on for every cohort in our sample. We "fill in" the missing values using a two step procedure. First, we regress mean earnings on birth year *y* separately by age *a* according to the model $\ln(E_{ay}) = \sum_{i=0}^{6} b_i y^i + \varepsilon_{ay}$. OLS estimates of this model allow us to generate predicted values of earnings at ages 27, 32, 37, 42, 47, 52, and 57 for all birth years. The results show that the predicted values from our estimates closely match the actual values, and that the models generate reasonable predicted values for the earlier ages and birth years not included in the SSA tables. Next, we regress predicted mean earnings (from the previous step) on age separately for each birth year according to the model $\hat{E}_{ay} = \sum_{j=0}^{4} \beta_{jy} a^j + \tau_{ay}$. OLS estimates of this model are used to generate predicted

[7] We capped reported earnings in the CPS at the Maximum Taxable earnings for that year based on appropriate Social Security regulations before computing means and medians by cell.

[8] To more closely match the aggregation used by the SSA, we aggregated to the age groups used in the *Annual Statistical Supplement* (25 – 29, 30 – 34, and so on) instead of the individual ages (27, 32, and so on).

values of mean earnings for at all ages and birth years in our data, including ages 28 to 31, 33 to 36, and so on. Again, the results show that OLS estimates of these models generate reasonable values of mean earnings at all ages and birth years. As a result of these two steps, we now have predicted values of mean earnings for each age and birth year cell in our sample.

The third issue with our data is that earnings are not reported separately by education level in the SSA. We need to disaggregate earnings histories by education group to compute cohort-specific measures of Social Security benefits. We use data from the CPS to accomplish this task. We compute the ratio of mean earnings for each education group to mean population earnings (the "earnings-ratio") in the CPS separately by birth year[9], denoted $\hat{ER}_{ye}$. Using this ratio, we compute estimates of mean earnings by birth year, age, and education group according to the formula $\hat{E}_{aye} = \hat{E}_{ay} * \hat{ER}_{ye}$, where E is mean earnings, , a is age, y is birth year, and e is education group. To compute the earnings ratio, we calculate mean earnings between ages 27 to 57 separately by birth year for each education group and for all education groups, then divide the first by the second to create the birth year, education group specific earnings ratio[10]. We then computed predicted values of the earnings ratio from OLS estimates of the model $ER_{ye} = \sum_{k=0}^{3} a_{ek} y^k + \mu_{ey}$. Because we don't observe earnings prior to age 57 for birth years before 1906, we assume the earnings ratios for those birth years are constant at the 1906 level. The results verify that OLS estimates of our model do a reasonable job of predicting earnings ratios for each education group and birth year.

Because we do not compute earnings ratios for each birth-year-education-group cohort separately by age, our measure of the earnings ratio "averages out" life-cycle earnings patterns for each birth year. This creates biases for at least two reasons. First, the returns to schooling are higher at older ages, so for higher levels of education we

---

[9] Ideally we could compute earnings-ratios separately for each education group by birth-year and age. However, because the CPS only goes back to 1962 we lack data on earnings at younger ages for earlier birth years. This makes it quite difficult to construct reasonable estimates of earnings by education group for these cells. However, when we include data from the 1940, 1950, and 1960 U.S. Census we will have data for some of these cells. This may enable us to construct birth year and age specific earnings ratios.
[10] We omit cells with sample size less than 30

overstate mean earnings at younger ages and understate mean earnings at older ages. The opposite is true for lower levels of education. Second, because we do not observe younger ages for earlier birth years in the CPS, and earnings for better educated men are relatively higher at later ages, we are overstating (understating) the ratio of mean earnings to population earnings for higher (lower) levels of education at earlier birth years.

We should also note some limitations of using CPS data to calculate Social Security earnings histories. The CPS includes some respondents who worked but may not have been included in the social security program, particularly in earlier years. In addition, the CPS data (generated from surveys) is likely subject to a higher degree of measurement error then the SSA data (generated from administrative records). To address this, we removed observations with suspect earnings data from our sample. We dropped all records where the real weekly wage (total income last year / number of weeks worked last year) was below $50 and above $40,000. This reduced the number of observations with positive earnings in our CPS sample by 2.8% (18,881 records).

A1.2 Computing Cohort Specific Social Security Benefits

We use the ANYPIA benefit calculator provided by the SSA to calculate the monthly social security benefit amount. Based on birth year and retirement age, the ANYPIA program computes the appropriate Primary Insurance Amount (PIA) and monthly benefit for a given earnings history[11] [12].

B. Pensions and Employer-Provided Retiree health Insurance (EPRHI)

Pension measures were derived from SIPP topical modules in the 1984 panel (wave 4), 1986 panel (wave 4), 1990 panel (wave 4), 1991 panel (wave 7), 1992 panel (wave 4), 1996 panel (wave 7), and 2001 panel (wave 7). Other panels were excluded due to incompleteness of data or changes in questionnaire design. These data have small sample sizes for earlier birth years: those who were born in 1900 are 84 at the time of the first survey so there is likely to be significant problems with mortality bias.

[11] Earnings in years prior to 1937 are not incorporated, as these years predate the formation of the Social Security program and are not included in the calculation of Social Security benefits for any cohorts.

[12] We used the Average Wage Indexing (AWI) series published by the Social Security Administration to convert nominal earnings to real (2004) earnings. As noted earlier we assume inflation in future years is 0%.

Different questions on pensions are asked depending on whether the respondent is currently working, has had a job in the past, has received a lump sum payment from a retirement plan, or is currently receiving retirement benefits (other than Social Security). Three (binary) pension measures are derived from these questions:

-- PENSION: Respondent has a pension plan if (a) his current job has a pension or (b) he expects to receive pension benefits from a previous job or (c) he is currently receiving pension benefits or (d) he received a lump sum payment from a retirement plan in the past or (e) he owns a business that has a pension plan he participates in; else no pension

-- DB: Respondent has a defined benefit pension plan if (a) the pension from his current job is a defined benefit plan or (b) he expects to receive pension benefits from a past job or (c) the retirement benefits he is currently receiving are from a defined benefit type plan.

-- DC: Respondent has a defined contribution pension plan if (a) the pension from his current job is a defined contribution plan, or (b) he owns a business that has a pension plan he participates in, or (c) he is receiving retirement benefits from a defined contribution plan or (d) he received a lump sum payment from a pension plan in the past.

If the respondent is currently receiving retirement benefits then he is asked if he is covered by EPRHI. The sample sizes for later birth years at older ages are again quite small.

References

Anderson, Patricia M., Gustman, Alan L., and Steinmeier, Thomas L. "Trends in Male Labor Force Participation and Retirement: Some Evidence on the Role of Pensions and Social Security in the 1970s and 1980s." *Journal of Labor Economics* 17, no. 4 (Oct. 1999): 757-783.

Blau, David M. "Labor Force Dynamics of Older Men." *Econometrica* 62, No. 1 (Jan. 1994): 117-56.

Blau, David M. "Labor Force Dynamics of Older Married Couples." *Journal of Labor Economics* 16, no. 3 (Jul. 1998): 595-629.

Blau, David M. and Gilleskie, Donna B. "Retiree Health Insurance and the Labor Force Behavior of Older Men in the 1990s." *The Review of Economics and Statistics* 83 no. 1 (Feb. 2001): 64 – 80.

Blau, David M. and Gilleskie, Donna B. "The Effect of health on Employment Transitions of Older Men," in Solomon W. Polachek (ed.) <u>Worker Wellbeing in a Changing Labor Market</u>, Research in Labor Economics, Vol. 20, Greenwich CT, JAI Press, 2001: 35-66.

Blau, David M. and Gilleskie, Donna B. "Health Insurance and Retirement of Older Married Couples." *Journal of Applied Econometrics* (in press).

Burtless, Gary and Quinn, Joseph F. "Retirement Trends and Policies to Encourage Work Among Older Americans." Boston College Working Paper 436 (January 2000).

Bound, John. "The Health and Earnings of Rejected Disability Insurance Applicants," <u>American Economic Review</u> 89 no. 3 (1989): 482-503.

Bound, John. "Self-reported and Objective Measures of Health in Retirement Models." *Journal of Human Resources* 26, No. 1 (Summer 1991): 106-38.

Chen, Susan, and Wilbert van der Klaauw. "The Effect of Disability Insurance on Labor Supply of Older Individuals in the 1990s," *Journal of Econometrics* (in press)

Coile, Courtney. "Retirement Incentives and Couples' Retirement Decisions." *Topics in Economic Analysis and Policy* 4 no. 1 (2004), article17.www.bepress.com/bejeap/topics/vol4/iss1/art17/

Costa, Dora L. *The Evolution of Retirement: An American Economic History, 1880 – 1990*. The University of Chicago Press (1998), Chicago, IL.

Costa, Dora L. "From Mill Town to Board Room: The Rise of Women's Paid Labor." *The Journal of Economic Perspectives* 14, No. 4 (Autumn, 2000): 101 – 122.

Diamond, Peter, and Hausman, Jerry. “The Retirement and Unemployment Behavior of Older Men.” In Henry Aaron and Gary Burtless (eds.) Retirement and Economic Behavior Washington D.C.: The Brookings Institution, 1984.

Englehardt, Gary, and Gruber, Jonathan. “Social Security and the Evolution of Elderly Poverty,” NBER Working Paper 10466, May 2004

Employee Benefit Research Institute Website: Databook on Employee Benefits, Chapter 4: Participation in Employee Benefit Programs. Updated November 2005. http://www.ebri.org/publications/books/index.cfm?fa=databook

French, Eric. “The Effects of Health, Wealth, and Wages on Labour Supply and Retirement Behavior,” *Review of Economic Studies* 72 (2005): 395-427.

Friedberg, Leora, and Webb, Anthony. “Retirement and the Evolution of Pension Structure.” *Journal of Human Resources* 40, No. 2 (Spring 2005): 281-308.

Gustman, Alan L., and Steinmeier, Thomas L. “Retirement in Dual-Career Families: A Structural Model.” *Journal of Labor Economics* 18, No. 3 (Jul. 2000): 503 – 545.

Hurd, Michael D. “The Joint Retirement Decision of Husbands and Wives.” In David Wise (ed.) Issues in the Economics of Aging, Chicago: University of Chicago Press for the NBER, 1990: 231-258.

Hurd, Michael D., and Boskin, Michael J. “The Effect of Social Security on Retirement in the Early 1970s.” *The Quarterly Journal of Economics* 99, No. 4 (Nov. 1984): 767 – 790.

Ippolito, Richard A. “Toward Explaining Earlier Retirement after 1970.” *Industrial & Labor Relations Review* 43, no. 5 (Jul. 1990): 556 – 569.

Kopecky, Karen. “The Trend in Retirement”, Economie d’avant garde, Research Report No. 12 (July 2005)

Krueger, Alan B., and Pischke, Jorn-Steffen. “The Effect of Social Security on Labor Supply: A Cohort Analysis of the Notch Generation.” *Journal of Labor Economics* 10, No. 4 (Oct. 1992): 412 – 437.

Lazear, Edward P., “Retirement from the Labor Force.” *Handbook of Labor Economics*, Vol 1. Ashenfelter, O. and Layard, R. eds., Elsevier Science, London (1986): pp. 305 – 355.

Madrian, Brigitte C. “The Effect of Health Insurance on Retirement,” *Brookings Papers on Economic Activity, Microeconomics*, 1 (1994): 181-232.

Moen, Jon Roger, “Rural Nonfarm Households: Leaving the Farm and the Retirement of

Older Men, 1860 – 1980." *Social Science History* 18, no. 1 (1994): 55 - 76

Moffit, Robert A. "Life Cycle Labor Supply and Social Security: A Time Series Analysis." *Work, Health, and Income among the Elderly*, Burtless, Gary, ed., Brookings: Washington, DC (1987).

Quinn, Joseph F. "Has the Early Retirement Trend Reversed?" Working Paper 424 (May 1999), Department of Economics, Boston College.

Parsons, D.O. "The Decline in Male Labor Force Participation," Journal of Political Economy 88 (Feb. 1980): 117-134.

Peracchi, Franco, and Welch, Finis. "Trends in Labor Force Transitions of Older Men and Women," *Journal of Labor Economics* 12, No. 2 (Apr. 1994): 210 – 242.

Stewart, Jay, "Do Older Workers Respond to Changes in Social Security Benefits? A Look at the Time Series Evidence," Bureau of Labor Statistics Working Paper 271, 1995.

Van der Klaauw, Wilbert, and Kenneth Wolpin. "Social Security, Pensions, and the Savings and Retirement Behavior of Households," manuscript, University of North Carolina at Chapel Hill (March 2005).

Table 1: Weighted OLS Regressions for LFPR of Older Men

	No Birth Year Effects	Linear Birth Year Effects	Birth Year Fixed Effects
Birth Year		**-0.0080**	
		(0.0003)	
High School Degree	**0.0711**	**0.0641**	**0.0595**
	(0.0070)	(0.0061)	(0.0072)
Some College	**0.1161**	**0.0883**	**0.0864**
	(0.0083)	(0.0073)	(0.0090)
College Degree	**0.1912**	**0.1534**	**0.1472**
	(0.0103)	(0.0092)	(0.0105)
Black	**-0.3027**	-0.0531	-0.0501
	(0.0467)	(0.0420)	(0.0423)
Married	**0.1997**	**0.1121**	**0.1798**
	(0.0505)	(0.0445)	(0.0445)
Divorced	-0.0242	**0.1499**	**0.1606**
	(0.0566)	(0.0501)	(0.0500)
Bad Health	**-0.3363**	**-0.2671**	**-0.3064**
	(0.0340)	(0.0300)	(0.0303)
Spouse LFPR	-0.3946	**-0.9752**	**-1.1412**
	(0.2818)	(0.2485)	(0.2598)
Age * Spouse LFPR	**0.0094**	**0.0194**	**0.0216**
	(0.0046)	(0.0041)	(0.0043)
Married * Sp. Income *	0.0006	**0.0007**	0.0005
	(0.0004)	(0.0004)	(0.0004)
DB Pension	**-0.3052**	**-0.1034**	-0.0310
	(0.0269)	(0.0248)	(0.0311)
DC Pension	**-0.1279**	0.0073	0.1005
	(0.0485)	(0.0429)	(0.0516)
EPRHI	-0.0141	**-0.0914**	**-0.0804**
	(0.0253)	(0.0224)	(0.0237)
Average Monthly Earnings*	0.0193	**0.0363**	0.0335
	(0.0111)	(0.0098)	(0.0199)
SS Disability Benefit*	**-0.0580**	**-0.0374**	**-0.0343**
	(0.0087)	(0.0077)	(0.0082)
SS Wealth, Claim at Age 65**	**-0.0708**	**-0.1107**	**-0.1739**
	(0.0152)	(0.0134)	(0.0523)
SS Wealth, Claim at Age 70**	**-0.0509**	**0.1112**	**0.1212**
	(0.0182)	(0.0171)	(0.0465)
Intercept	**0.9681**	**16.1104**	**0.9634**
	(0.0479)	(0.5665)	(0.0682)
R Squared	0.9309	0.9467	0.9507
Sample Size	2453	2453	2453
p-value, test of SS Benefits	0.0000	0.0000	0.0038
p-value, test of Birth Year FE	0.0000	0.0000	

Notes: Measures marked with * are divided by 1,000; Measures marked with ** are divided by 100,000. All Specificaitions include a full set of age fixed effects. Observations are weighted by the square root of the number of observations per cell, using March CPS weights to compute the number of observations.

Table 2: Accounting for the Decline in LFPR of Older Men between 1965 and 1993

Period	Actual LFP	Predicted LFP	SS Benefit Rules	SS Disability	Pensions	Spouse LFPR	Birth Year Effects	Education Effects
1965 to 1970	0.709	0.695	0.653	0.686	0.695	0.697	0.678	0.667
1988 to 1993	0.519	0.524	0.484	0.523	0.519	0.523	0.667	0.471
Decrease	0.191	0.171	0.169	0.163	0.176	0.174	0.011	0.196
Difference	-	-	0.001	0.007	-0.005	-0.003	0.160	-0.025
% of Decrease	-	-	1%	4%	-3%	-2%	93%	-15%

Table 3: Accounting for the Increase in LFPR of Older Men between 1988 and 2005

Period	Actual LFP	Predicted LFP	SS Benefit Rules	SS Disability	Pensions	Spouse LFPR	Birth Year Effects	Education Effects
1988 to 1993	0.519	0.524	0.528	0.527	0.522	0.520	0.555	0.502
2000 to 2005	0.565	0.556	0.547	0.565	0.547	0.544	0.648	0.474
Increase	0.047	0.032	0.019	0.038	0.024	0.024	0.093	-0.027
Difference	-	-	0.012	-0.006	0.007	0.007	-0.061	0.059
% of Increase	-	-	39%	-19%	23%	23%	-193%	186%

Figure 1: LFPR Trends by Age Group (Source: BLS Website)

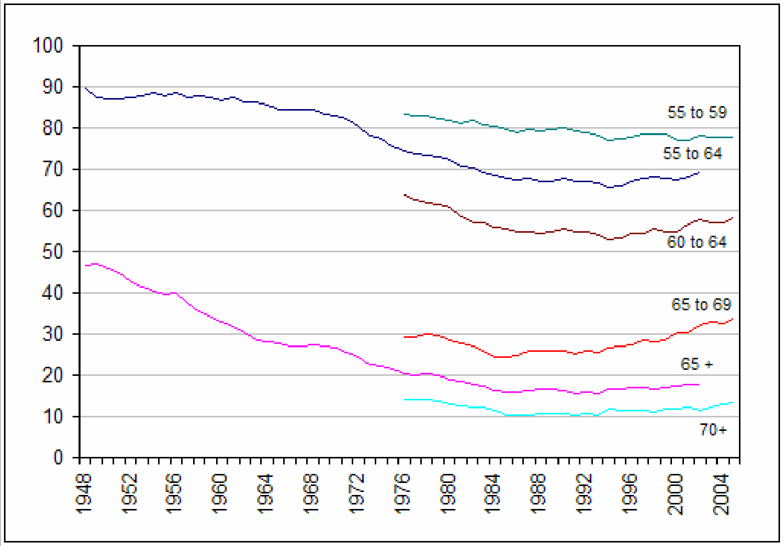

Figure 2: LFPR of Men Ages 55 to 69

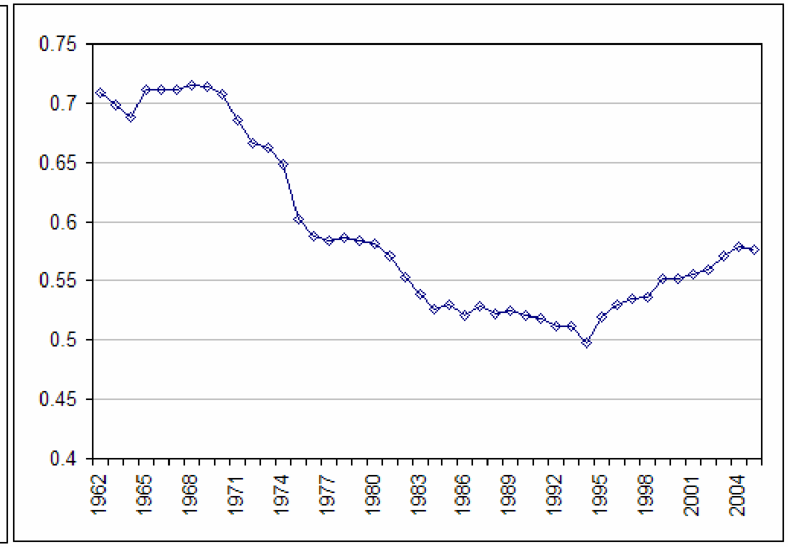

Figure 3: Trends in the Education Distribution of Older Men

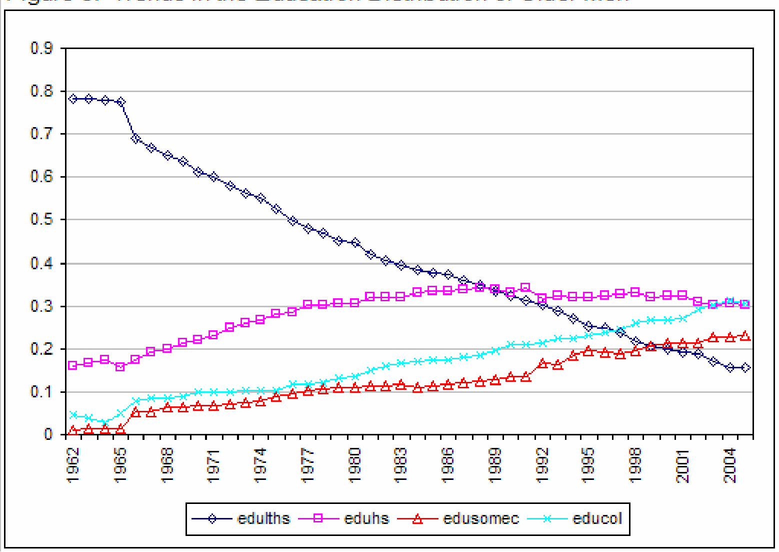

Figure 4: Trends in LFP of Older Men by Education

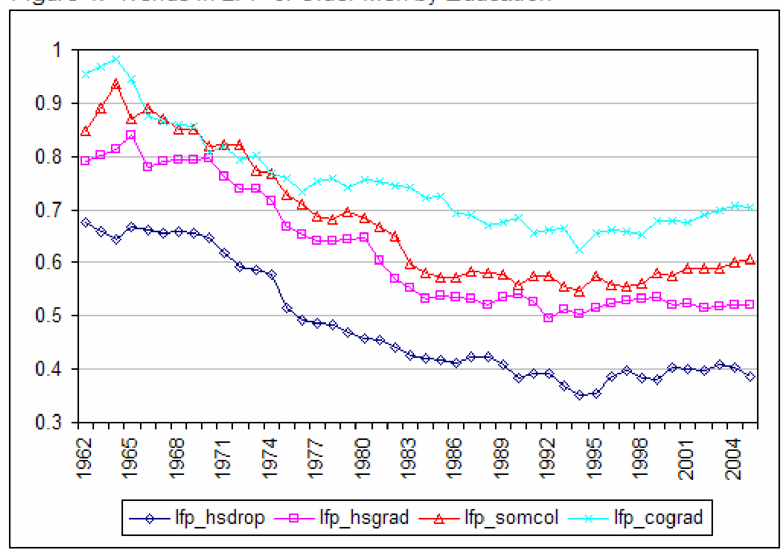

Figure 5: Trend in the LFPR of Wives of Older Men

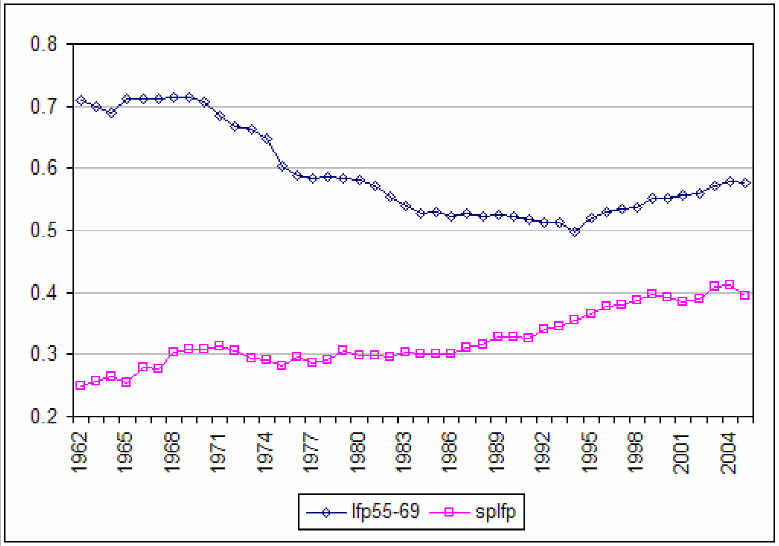

Figure 6: Trend in Poor Health

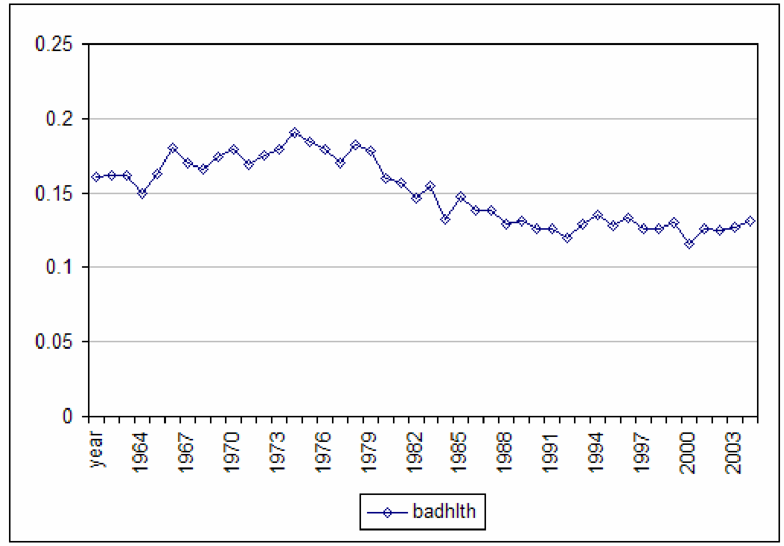

Figure 7: Trend in Lifetime SS Wealth by Claim Age

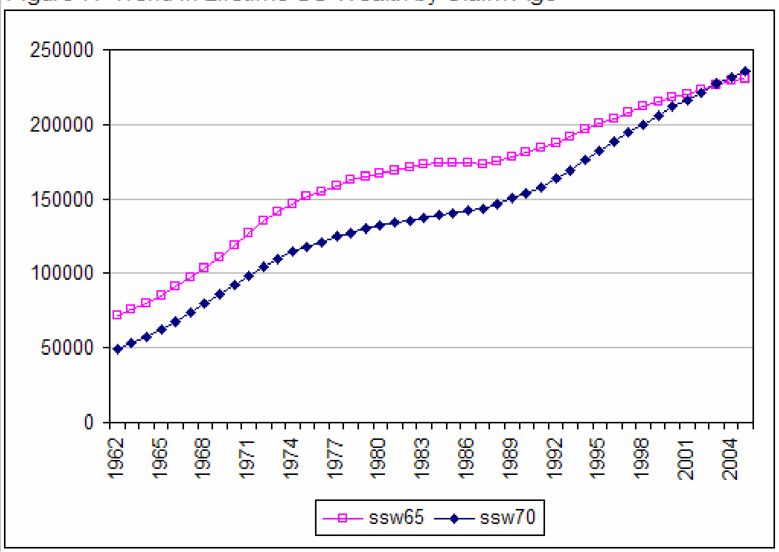

Figure 8: Social Security Disability Benefits

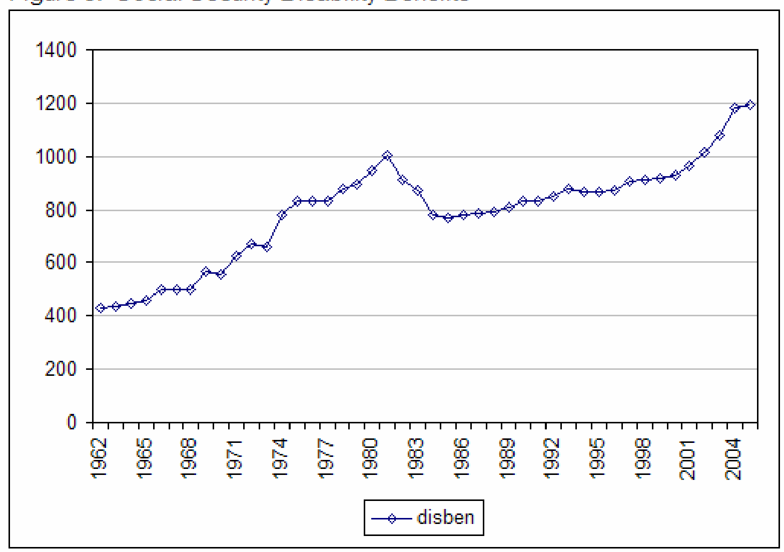

Figure 9: Average Monthly Earnings by Education Group

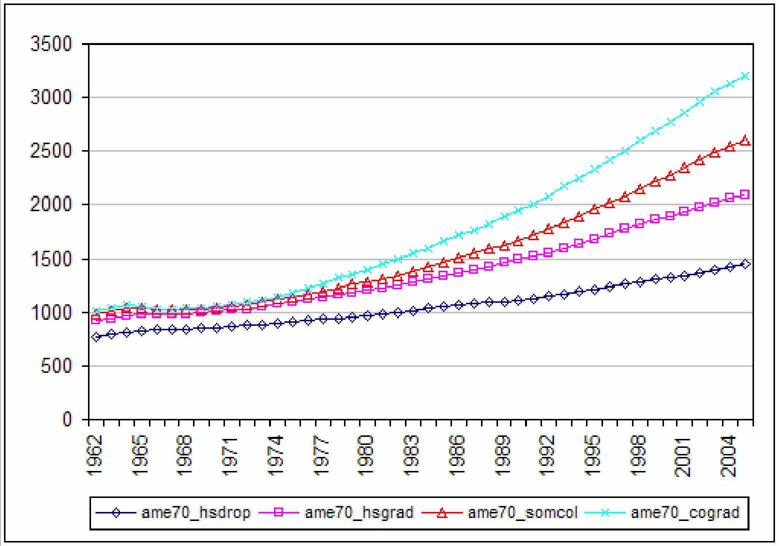

Figure 10: Trends in Pension and EPRHI Coverage

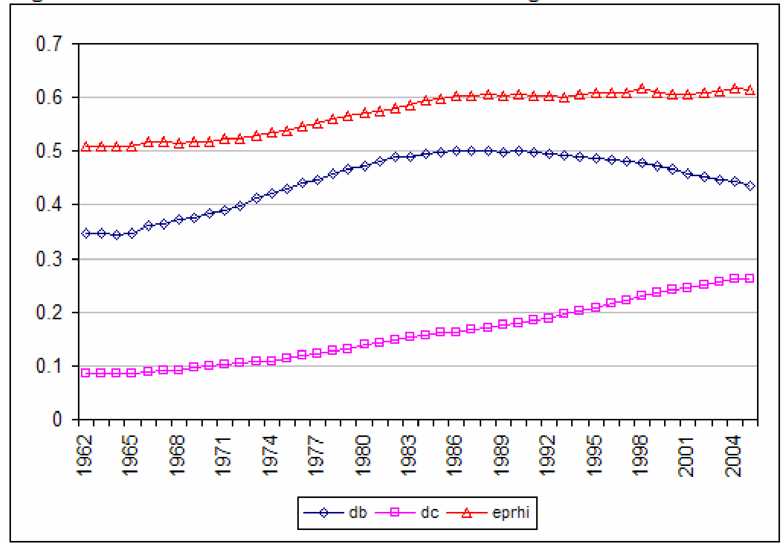

Figure 11: Actual vs Predicted LFPR of Older Men

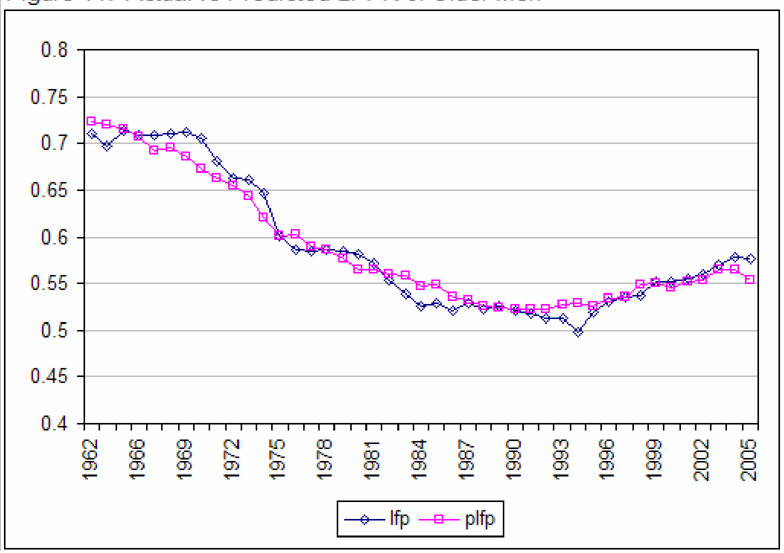

Figure 12(a): 1978 Social Security Counterfactual, Specification 1

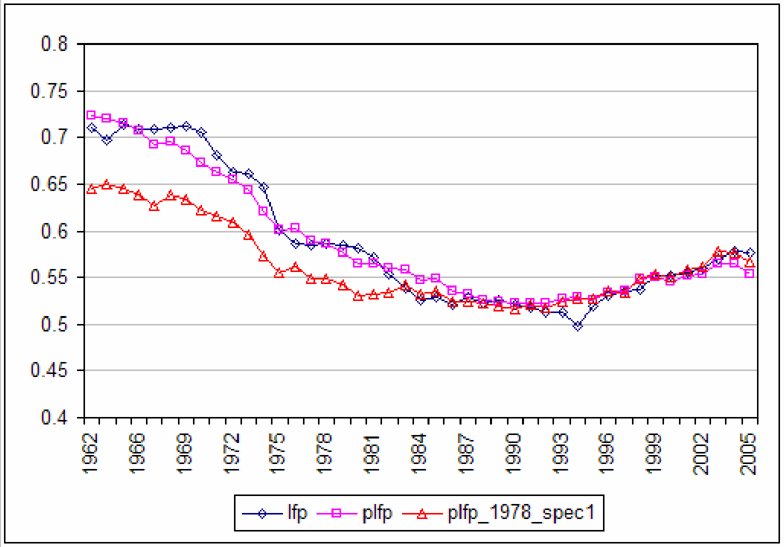

Figure 12(b): 1978 Social Security Counterfactual, Specification 3

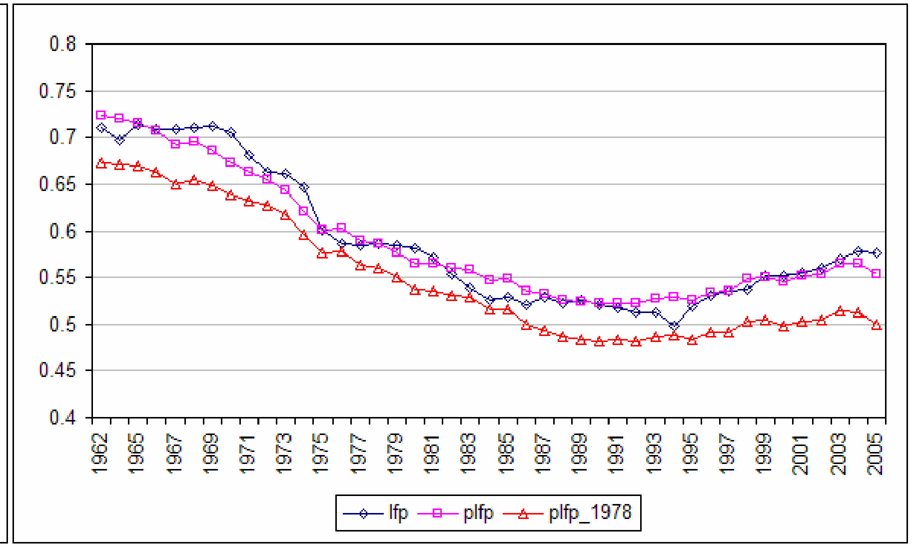

Figure 13: Counterfactual, 1978 SS Disability Benefit Rules

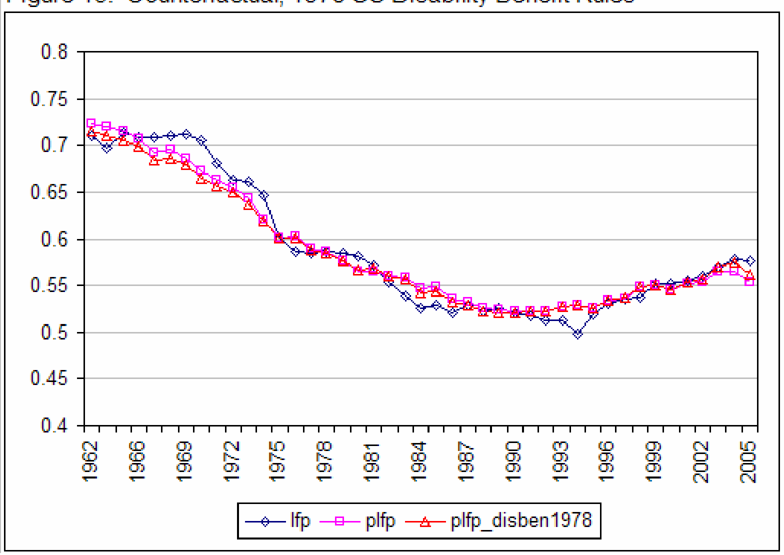

Figure 14: Counterfactual, Average Monthly Earnings fixed at 1970 Levels

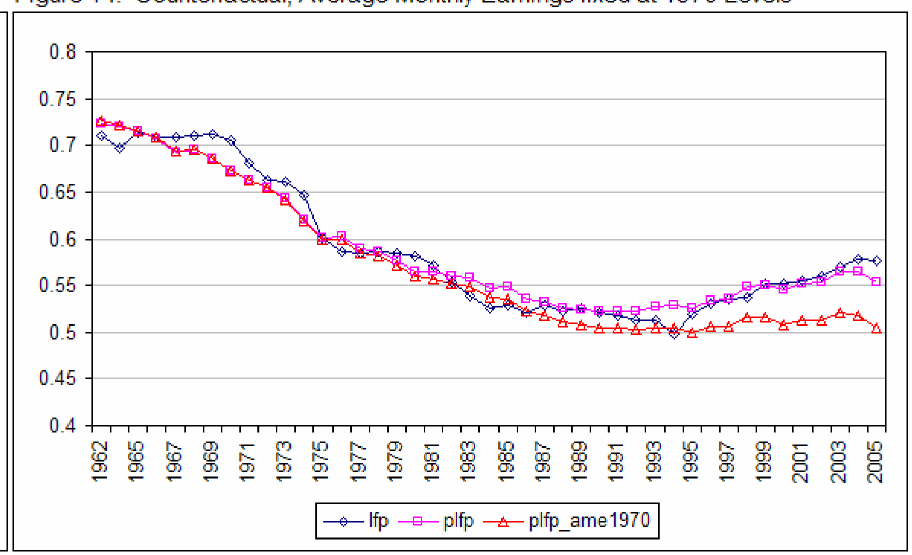

Figure 15: Counterfactual, Birthyear fixed at 1970 Level

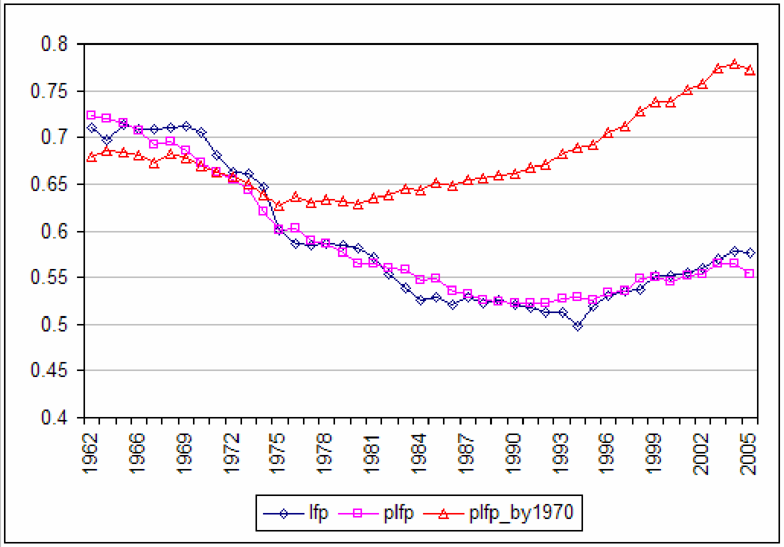

Figure 16: Counterfactual, Education fixed at 1970 Level

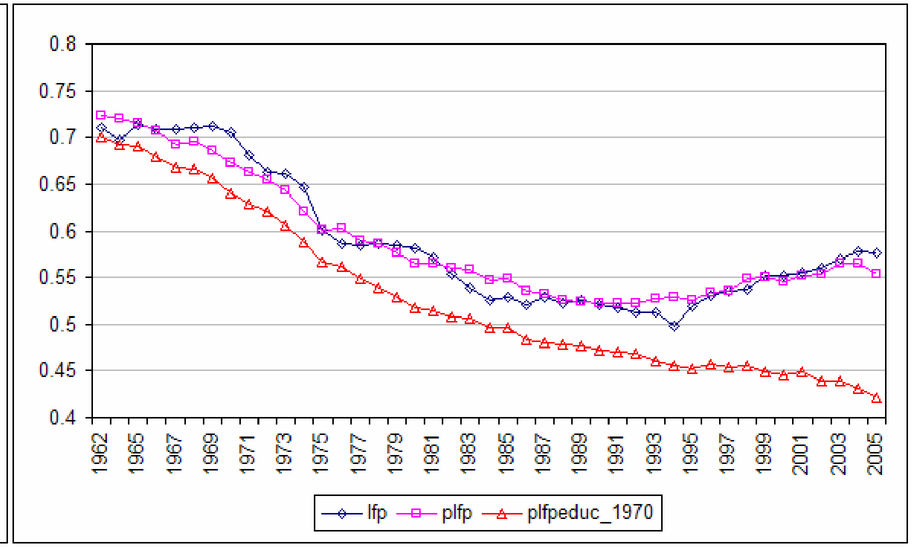

Figure 17: 1984 Social Security Counterfactual

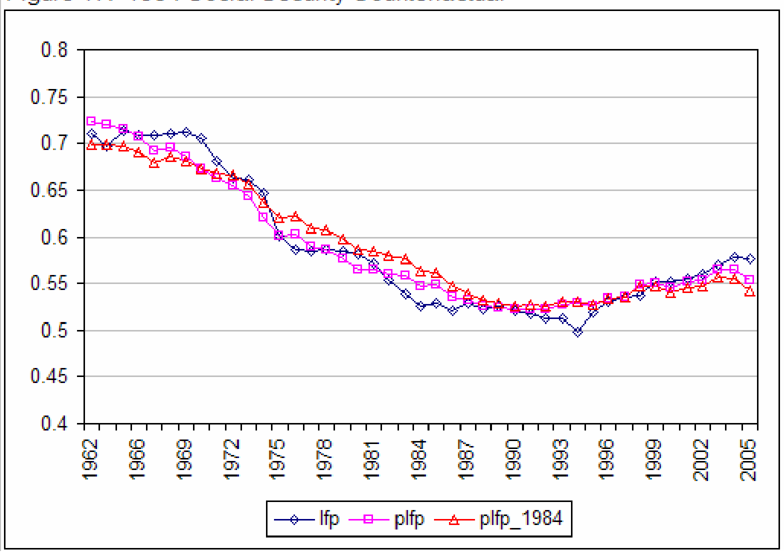

Figure 18: Counterfactual, Spouse LFPR fixed at 1985 Level