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# HEALTH AND LABOR FORCE PARTICIPATION OF OLDER MEN, 1900-1991

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## ABSTRACT

I investigate how the relationship between health status and retirement among older men has changed since 1900 using weight adjusted for height or Body Mass Index (BMI) as a proxy for health. I find that both in 1900 and in 1985-1991 the relative risk of labor force non-participation increases for the excessively lean and obese and that the BMI level that minimizes the relative risk of labor force non-participation remains unchanged. However, in 1900 both the relative risk of non-participation among men at low and high BMI levels and the elasticity of non-participation with respect to BMI were greater than today, suggesting that health is now less important to the retirement decision than in the past. The difference in the relative risk of non-participation is especially pronounced at high BMI levels. Declining physical job demands and improved control of chronic conditions may explain the difference.

The findings suggest that the impact of improvements in health on participation rates is increasingly more likely to be outweighed by the impact of other factors. Greater efforts made to increase the incorporation of the old and disabled into the labor force may therefore have a minimal impact on retirement rates. The findings also imply that in the past the economic costs of poor health were substantial.

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# 1 Health and Labor Supply

Health is a critical component of labor supply. Chronic conditions such as heart disease, arthritis and other musculoskeletal conditions, and respiratory disorders substantially reduce hours worked and the probability of participation and this reduction in labor supply accounts for up to 45% of the decline in earnings observed among middle-aged men (Bartel and Taubman 1979; Yelin and Katz 1991; Pincus *et al.* 1989; Burkhauser *et al.* 1986). Because the prevalence of chronic disabilities rises with age, the effects of health at older ages are especially pronounced. Older men commonly cite poor health as their main reason for withdrawing from the labor force and cross-sectional evidence indicates that health plays an important role in the retirement decision.[1]

Despite the importance of health in labor supply functions, we are still ignorant of the complex relation between health and participation. How do health, jobs skills, work environment, income, and societal standards of what constitutes incapacity to work interact to produce participation? Work place flexibility, such as that afforded by self-employment, can disproportionately affect the participation rates of the less healthy because the severity of chronic conditions often follows a cyclical pattern either within the day or over a longer period of time. For example, arthritics suffer morning flareups as well as periods of remission. Asthmatic conditions are more severe in the spring and summer (Yelin 1992: 116). On the other hand, participation depends upon physical job requirements such as the extent of lifting, climbing, stooping, reaching, or exposure to humidity, fumes, or dust. Lastly, those with low income and few non-labor market opportunities may not be able to afford to consider themselves disabled. Withdrawal from the labor force requires resources to sustain consumption. If these resources are available then poor health will lead to retirement.

Has health become more or less important in the retirement decision over the course

[1] See Quinn and Burkhauser (1990) for a review of the retirement literature.

of the century? The shift from the manufacturing and agriculture sectors to the service sectors, increasing mechanization, and the shortening of the work day has lessened the expenditure of physical energy required for jobs, thus easing the incorporation of the disabled into the labor force. Improved control or alleviation of chronic conditions provided by innovations such as antihypertensive drug therapy and anti-inflammatory drugs used in the care of arthritis have also eased the incorporation of the disabled into the labor force. But, job flexibility, as measured by self-employment, was more common in the past. In 1910, 27% of all men in the labor force were self-employed compared to only 12% in 1980. Among men 65 years of age or older, 79% were self-employed in 1910. In contrast, 33% were in 1980. Futhermore, men in 1900 may not have had the resources to be able to afford to consider themselves disabled. Retirement incomes in the past were low and private pensions were rare. In fact, men who by today's standards would be clearly disabled were participating in the labor force circa 1900 (Fogel, Costa, and Kim 1993).

Labor force participation rates for men 65 years of age or older were 65% in 1900. Participation rates fell steadily to reach 16% in 1990 despite improvements in average health (Moen 1987; Table No. 622 in U.S. Bureau of the Census 1993). Ill health among older men has been declining at a rate of about 7% per decade from 1910 to 1985-1988 (Fogel, Costa, and Kim 1993; Manton, Corder, and Stallard 1993; cf. Yelin 1992: 25-32; Verbrugge 1984).

All cross-sectional evidence indicates that health is an important factor in the retirement decision. One possible explanation that would reconcile the time series and cross-sectional evidence is that the effects of other factors, such as increased income, could be swamping improvements in health. Health may also indirectly affect labor supply through the wage. Evidence from the U.S. and from developing countries suggests that there is a relation between wages and health (Bartel and Taubman 1979). Changes in wages have both income and substitution effects. If the income effect dominates then men will chose earlier retirement. If an increase in health improves productivity and therefore wages then, at a high enough wage, labor supply may be backward bending with respect to wages and therefore health. Thus a large fraction of the increase in health may have been spent on increased leisure. Although health is an important determinant of participation within cross-sectional data, the impact of health on labor force participation rates may be declining. Therefore the potential increase in participation rates from improvements in health may now be smaller, increasing the likelihood that the effects of improved health on participation rates are outweighted by other factors.

This research examines the relationship between health status and labor force participation in both the National Health Interview Survey and a sample of Civil War veterans and assesses whether this relationship has changed over time. The focus in this paper is on how *relative* participation rates have changed. Participation rates were higher in the past, but were they uniformly higher or were the disabled more or less likely to work relative to the healthy? Responsiveness to changes in income may well vary by health status. If so, then the secular increase in income, the income effect of wage gains from improved health, and the growth of programs specifically aimed at the disabled will have changed the composition of the labor force. If we are to increase the solvency of the Social Security system by raising the retirement age or tightening eligibility standards for disability benefits, then we must learn more about the relation between health and labor force participation. Are increases in non-labor market income resulting from retirement and disability programs affecting mainly their target population?

The findings also have implications for economic growth. Policy makers have long recognized that the health of a population is crucial to economic development (e.g. United Nations 1968). The estimated relationship between health and labor force participation will allow us to determine the impact of improved health on participation rates and therefore on national income. To calculate the benefits of the improvements in health that have occured since the beginning of the century we must know the functional form between health and labor force participation for that time period. We must also know how the functional form has changed and what factors underly that change.

## 2 Measuring Health

Health comparisons between modern and past populations have been drawn using life expectancy, time lost from work (Riley 1991), prevalence of chronic conditions, and anthropometric measures such as height and weight adjusted for height (Fogel, Costa, and Kim 1993). The difficulty in using life expectancy as a health measure is that life expectancies can be high and health poor if advances in medical technology have led to an increased burden from chronic conditions (Verbrugge 1984; Gruenberg 1977). Using time lost from work avoids these problems, but time lost from work may be determined by cultural factors or economic incentives. An individual from a poor household may be more likely to work than an individual from a rich household. Prevalence rates for past populations are rare.[2] Furthermore, advances in diagnostic techniques make comparisons difficult. Both height and weight adjusted for height have been used as health proxies by development economists (e.g. Deolaliker 1988; Strauss and Thomas 1993; Behrman and Deolaliker 1989), as well as by economic historians. The measure that is used here is weight adjusted by height as measured by Body Mass Index (BMI) and defined as body weight in kilograms divided by the square of body height in meters.

Weight adjusted for height is a measure of current nutritional status and as such reflects all demands made upon the body, including those of disease, labor, and even climate. It is an excellent predictor of subsequent mortality. Waaler (1984) found a marked relationship between BMI and risk of death in a sample of 1.8 million individuals. Mortality risk would decline rapidly at low weights as BMI increased, would stay relatively flat over BMI levels from the low to high

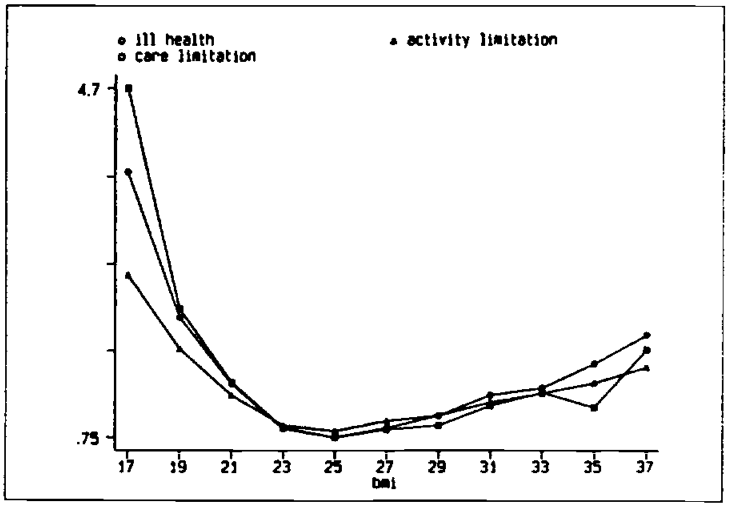
[2]See Fogel, Costa, and Kim (1993) for the use of prevalence rates in a past population.

twenties (28), then would start rising again, but less steeply than at very low BMIs. A similar U-shaped relation is evident in a sample of Union Army veterans, in which mortality risk first decreased with rising BMI, reached an optimum BMI in terms of mortality risk between 21 and 28, and then increased (Costa 1993), thus suggesting that BMI standards derived from recent populations can be applied to past populations. When height and weight are simultaneously related to mortality through a mortality surface, it is obvious that the optimal weight varies with height. To minimize their mortality risk, the shorter should be heavier and the taller leaner (Fogel, Costa, and Kim 1993).

When BMI curves are plotted by cause of death, there is a strong U-shape for obstructive lung disease, stomach cancer, and cerebrovascular disease, a very slight U-shape for cardiovascular disease and diabetes, and none at all for colon cancer, tuberculosis, and lung cancer (Waaler 1984). However, the relationship between BMI and subsequent all cause mortality does not appear to be determined by the moribund. Waaler (1984) found that when he deleted men who died within 5 years after measurement, the U-shape was even more pronounced. A 32-year follow-up study of a cohort measured at age 18 found that the most obese men had higher mortality rates from coronary heart disease while the leanest men had higher mortality rates from cancer (Hoffmans, Kromhout, and Coulander 1989).

BMI and ill health are correlated. The BMI-risk curve for the relative risk of being in fair or poor health (self-reported) shows the same U-shaped pattern for males aged 50-64 as seen for mortality risk. When height and weight are simultaneously related to ill health through a surface, the resulting surface is similar to that for mortality (Fogel, Costa, and Kim 1993). Figure 1 shows the relation between BMI and the presence of a self-reported activity or care limitation or of ill health in the National Health Interview Survey (described later) among men age 50-64. BMI is also related to more objective measures. The prevalence of hypertension, heart disease, and other circulatory disorders and of diabetes and chronic neck pain rises with

Figure 1: BMI and Relative Risk of Activity or Care Limitation and Ill Health Among White Men 50-64

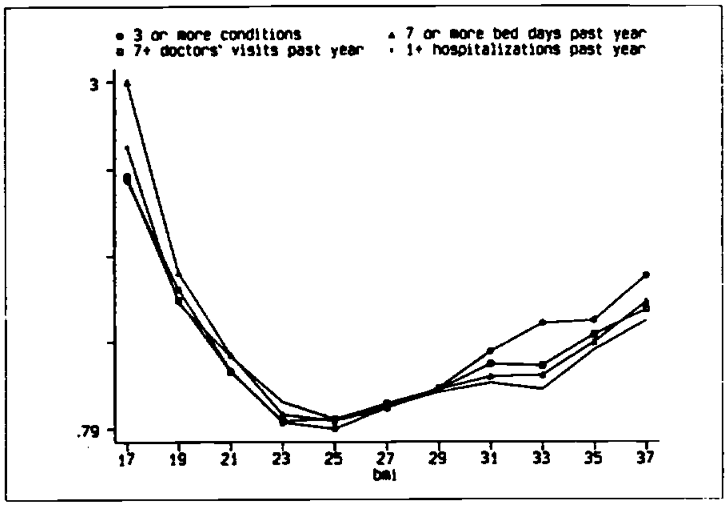



Source: National Health Interview Survey, 1985-1991.
Note: All observations are centered at the marks. 39,651 observations.

increasing obesity. Chronic respiratory disorders show a U-shaped relation to BMI (Roman Diaz 1992; Makela *et al.* 1991; Negri *et al.* 1988). The incidence of pulmonary tuberculosis is greater at lower BMI levels (Tverdal 1988). None too surprisingly BMI is also related to the number of chronic conditions, of bed days, of hospitalizations, and of doctors' visits (see Figure 2).[3]

[3] When height and weight are jointly related to the presence of an activity, work, or care limitation or to the number of chronic conditions, bed days, hospitalizations and doctors' visits, once again the resulting surface is similar to the mortality surface. However, in case of the chronic conditions, bed days, hospitalizations, and doctors' visits the relation with height is greatly diminished.

Figure 2: BMI and Relative Risk of Chronic Conditions, Bed Days, Hospitalizations, and Doctors' Visits Among White Males 50-64




Source: National Health Interview Survey, 1985-1991
Note: All observations are centered at the marks. 39,651 observations.

# 3 Health Data

## 3.1 Union Army Records

Although no random health surveys are available for either the nineteenth or the early twentieth century, the pension program run for the benefit of Union veterans and their dependent children provides a rich source of information on a large fraction of the population. By 1900 this was a program of enormous scope. Close to 85% of all Union Army veterans were collecting a pension.[4] Among all white males, 35% of those aged 55-59 were on the rolls, 21 percent of those aged 65-69, 14 percent of those aged 65-69, and 9 percent of those 70 or older. The annual value of the average veteran pension was $135 or 36% of the annual income of a laborer.

[4]Calculated from Table Y 957-970 in U.S. Bureau of the Census (1975: 1145) and Glasson (1918).

The basic system of pension laws, known as the General Law pension system, was established in 1862 by the United States Congress and provided pensions to both regular and volunteer recruits who were disabled as a direct result of military service.[5] The dollar amount received depended upon degree of disability, where disability was determined by the applicant's capacity to perform manual labor. Under later reinterpretations the total disability standard soon meant incapacity to performeven lighter types of manual labor. In fact, men judged disabled continued to labor in physically demanding, manual occupations. Inability to perform manual labor remained the standard in this and all subsequent laws, regardless of the wealth of the individual or his ability to earn a living by other than manual means. Withdrawal from the labor force was not a necessary prerequisite for the receipt of a pension. If the claimant had lesser disabilities then he received an amount proportionate to his disabilities. Application was through a pension attorney and the degree of disability was determined by a board of three local doctors employed by the Pension Bureau and following guidelines established by the Bureau.

The Act of June 27, 1890 marked the beginning of a universal disability and old-age pension program for Union veterans. The new law, according to the veterans' lobby, would "place upon the rolls all survivors of the war whose conditions of health are not practically perfect."[6] In fact, within a year of the act's passage, the number of pensioners on the rolls more than doubled. Any disability now entitled a veteran to a pension. However, an applicant who could trace his disability to wartime service received substantially more for the same disability than his counterpart who could not. Even though old age was not recognized by statute law as sufficient cause to qualify for a pension until 1907, the Pension Bureau instructed the examining surgeons in 1890 to grant a minimum pension to all men at least 65 years of age, unless they were unusually

[5] Details on the pension program are provided in Costa (1994).

[6] Quoted in Glasson (1918).

vigorous.

Copious records were generated by the Union Army pension program. Examining surgeons recorded height, weight, pulse rate, and general appearance. They also described each condition in detail and rated its severity. Pension applications included rejections, retrospective information, and all applications for increases. Applications for increases could be filed at any time. In the sample studied here, each veteran filed an average of twelve prior to 1900. For every twelve, about two were rejected, most commonly on the grounds the alleged disabilities were unrelated to the war.

The Union Army records used in this research represent a 6 percent sample of a larger project that is still underway (see Fogel *et al.* 1991 for details) and consist of men in 20 companies who have been linked from their army records to their pension records, including the successive medical reports of the examining surgeons of the Pension Bureau, and to the 1900 census. The twenty companies are predominately from Ohio and New York state, but the Civil War cohort in these states did not differ in observable characteristics, such as home ownership, marital status, literacy, occupational distribution, foreign-birth, and age from other Northern states.[7] Seven hundred and twelve men were linked to the 1900 census out of 1036 men at risk to be found. Virtually all men found in the 1900 census were on the pension rolls by 1900.[8] An examining surgeons' report is available for 88% of these men. Wages, incomes, and wealth are not explicitly reported.

The analysis presented in this paper is based on the sample of non-institutionalized men

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[7]Calculated from the Public Use sample of the 1900 census (Preston and Higgs 1983).

[8]Men who were rejected would have a pension record. Men not yet on the rolls would frequently provide retrospective information. Pension applications included not only the original application, but also applications for increases, which could be filed at any time. Not all claims for pensions or for pension increases were accepted. Out of an average of 12 complaints filed, about 2 were rejected. Causes of rejection are known for 195 out of 557 rejections. Twenty-four percent of all men for whom causes of rejection are known were rejected because their disabilities were judged to be unrelated to the war.

found in the 1900 census. The sample appears to be representative of the northern population in terms of mortality and wealth.[9] Because the BMI that minimizes mortality risk varies with age (Andres 1985), BMI at ages 50-64 is used and men without this information were excluded from the sample.[10] Total sample size is 597. Ages ranged from 50 to 81. Men were classified as 1) farmers, 2) professionals or proprietors, 3) artisans, and 4) semi-skilled or unskilled laborers, including farm laborers, on the basis of their 1900 occupation, if in the labor force, or, if retired, on the basis of their previous occupation as given in the pension records.[11] Retired men whose occupation was unknown were assigned to an occupational class on the basis of their occupation at enlistment and their probability of switching occupational category given their individual characteristics. A man is considered to be retired in this analysis if the census enumerator specifically stated that he was "retired" or had "no occupation," or if he left the occupational field empty. Sixteen percent of veterans were retired.

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[9]Life tables were constructed for men found in the 1900 Census and compared with mortality schedules constructed for states that kept death registration records. The two life tables are similar. Also, the distribution of causes of death of veterans who died circa 1910 and were in the pension records is not significantly different from the distribution of causes of death reported by the death registration states. Recruits' households were neither disproportionately rich nor poor in 1860 (Fogel 1991; Fogel *et al.* 1991) and using height as a proxy for wealth, I find that wealth does not predict war survivorship.

[10]Men without a surgeon's record consisted of two types of men – those who were so severely injured during the war that they did not need a surgeons' exam to prove their disabilities and those who were exceptionally healthy and applied on the basis of old age. By 1900, on average, men without a surgeon's record were more likely to be collecting a higher pension, to have been discharged for disability, to have entered the pension rolls earlier, and to be out of the labor force. The sample of men with a surgeon's record used in this analysis may therefore be slightly healthier than those without. This is unlikely to affect the findings. When the unhealthy with a surgeon's exam are deleted from the sample, the findings remain unaffected.

[11]The census enumerators were asked to record an individual's primary occupation. In the few cases where two occupations were given, the first occupation was taken. Neither past nor current occupation is known for 51 men in the sample.

### 3.2 National Health Interview Survey

The National Health Interview Survey (NHIS) is a continuing nationwide sample survey on personal and demographic characteristics, illnesses, injuries, impairments, chronic conditions, utilization of health resources, and other health topics, obtained through personal household interviews. Seven years of data from the NHIS, from 1985 to 1991, were combined. To facilitate comparisions with the Union Army veteran sample, the sample was restricted to white males aged 50-64. Twenty-three percent of all men aged 50-64 were out of the labor force. Since each NHIS sample is representative of the noninstitutionalized U.S. civilian population, no adjustments were made to account for merging seven years of data. Although average BMI rose slightly, but significantly, over time, the results are robust to adjustments for cohort effects or year to year variations.

Unlike the weight and height data in the Union Army sample, that in the NHIS is self-reported. However the mean and standard distributions of height, weight, and BMI in the combined sample for men aged 18-74 match the published results of clinician-reported measurements from the National Health and Nutrition Examination Survey (NHANES II) very well. Height is greater in the NHIS sample in part because of the secular increase in heights and in part because of increased variance in self-reported measures.

Among the socio-economic characteristics in the NHIS is family income coded by income group. Each individual was therefore assigned the mean income within each group.

## 4 Health and Labor Force Participation

The improvement in health since 1900 has been substantial. Differences in BMI between men in the NHIS and the Union Army sample are striking (see Figure 3). Mean BMI in the NHIS is 26.4 compared to 23.0 in the Union Army sample. Most of the increase in BMI is due to rising weights. Average weight increased from 69kg to 83kg and height from 1.73m to 1.78m. The low weights of the Union Army veterans can partially explain why mortality for their cohort was higher than for cohorts today. Had it been possible to shift the Union Army BMI distribution so that the mean in the Union Army sample was equivalent to that in recent data, the increase in BMI would explain roughly 20% of the total decline in mortality above age 50 from 1900 to 1896 (Costa 1993).

Despite the large differential in health between Union Army veterans and recent populations, the relationship between BMI and labor force participation is similar in both the NHIS and the Union Army sample. The relationship between BMI and the relative risk of labor force non-participation in the NHIS is illustrated in Figure 4.[12] The probability of being out the labor force falls precipitously from a BMI level of 17 to a level that maximizes the probability of participation at 22 to 28 and then rises gradually. A similar BMI level that maximixes the probability of participation is evident in the Union Army sample (see Figure 5). However, there the probability of being out of the labor force increases sharply at both high and low levels of BMI, although, as indicated by the confidence intervals, this may be a small sample phenomenon. Also, because of correlation between income and BMI, health may not be the driving factor in determining the relationship between BMI and the relative risk of labor force non-participation. The impact of BMI on labor force participation rates can only be determined controlling for socioeconomic and demographic characteristics.

The choice between retirement and labor force participation is modelled as depending on the well-being experienced in each option as determined by the income flows associated with

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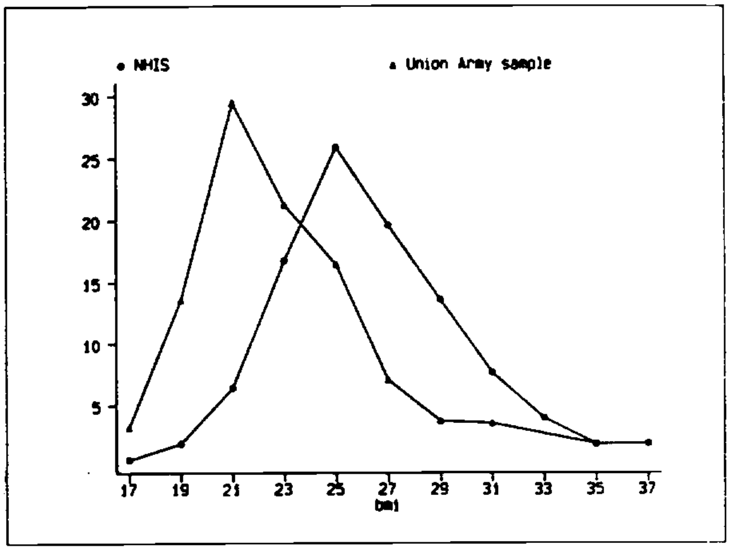
[12]The relative risk of labor force non-participation, $R_i$, at a given level of BMI is defined as

$$R_i = M_i/(\sum N_i M_i/N) = M_i/\bar{M}$$

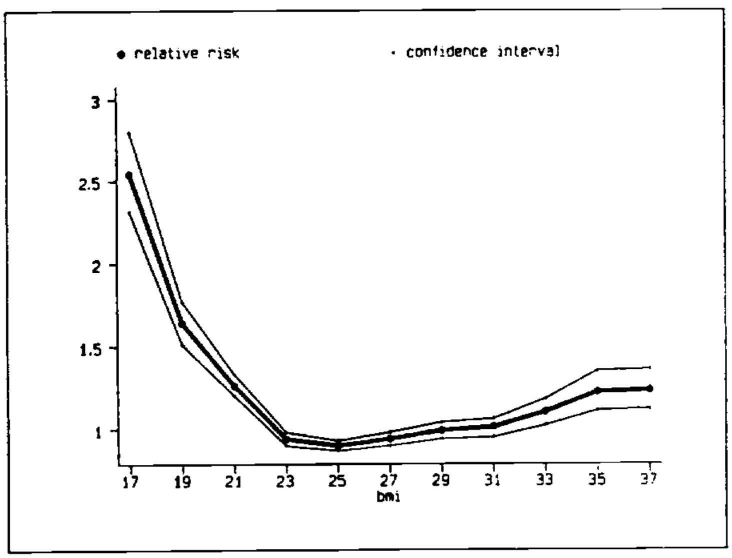
where $M_i$ is the non-participation rate at BMI level $i$, $N_i$ is the number of individuals at BMI level $i$, $N = \sum N_i$, and $\bar{M}$ is the mean non-participation rate over all levels of BMI.

Figure 3: BMI Distribution in the Union Army Sample and the NHIS, White Males 50-64



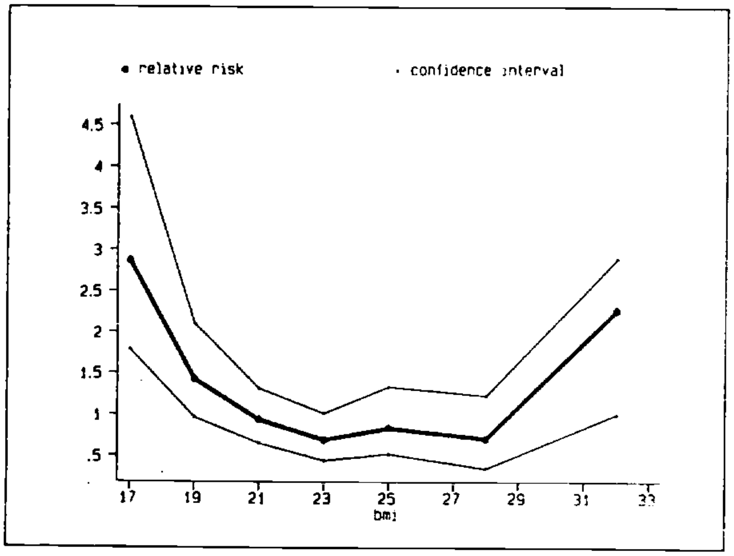

Note: All observations are centered at the marks. The NHIS contains 39,651 observations and the Union Army sample 597.

Figure 4: Relationship BMI and Relative Risk of Non-participation, White NHIS Men, 50-64, 1985-1991




Note: 39,651 observations. All observations are centered at the marks.

Figure 5: Relationship BMI and Relative Risk of Non-participation, Union Army Sample, 1900




Note: 597 observations. All observations are centered at the marks.

each option, the utility of time spent in leisure, the stigma costs of not working, and other sources of utility. Utility when not working can be written as $U_w(Y_w, \bar{H}; Z)$ and utility when not working as $U_l(Y_l, 0; Z)$ where $Y_w$ is income received when working, $Y_l$ is income received when not working, $\bar{H}$ is hours of work, and $Z$ is a vector of demographic variables and other utility shifters. Then, if the utility functions are assumed to be linear in their arguments, the utility maximizing individual evaluates the decision function,

$$
\begin{aligned}
I^* &= U_l(Y_l, 0; Z) - U_w(Y_w, \bar{H}; Z) \\
&= -X'\beta + \epsilon
\end{aligned}
$$

where $X$ is a vector containing $Y_l$, $Y_w$, $\bar{h}$, and $Z$, $\beta$ is a parameter vector, and $\epsilon$ is a standard normal error term. Although the value of $I^*$ is not observed, a discrete retirement indicator is observed given by

$$I = \begin{cases} 0 & \text{if } I^* < 0 \\ 1 & \text{otherwise} \end{cases}$$

where 1 represents retirement and 0 labor force participation. The probability of labor force non-participation can then be estimated by means of a probit,

$$P(I = 1) = Prob(\epsilon < X'\beta) = \Phi(X'\beta)$$

where $\Phi()$ is a standard normal cumulative distribution function.

Income flows under either the retirement or participation option are observed only if the respective choice was made. Therefore, in the NHIS, income flows in either labor force state are estimated (see Table 1). First, a participation probit is estimated where the explanatory variables are BMI, age, education, veteran and marital status, year of survey, extent of urbanization, geographic region, foreign birth, and the presence of a care limitation. Then income is estimated for both participants and non-participants using the same explanatory variables and the inverse Mill's ratio to account for selection (Heckman 1976; Lee 1979). Estimated non-participation income will therefore include income from disability programs, as well as from family members. In the final step, the probability of labor force non-participation is then estimated by means of a probit in which the dependent variable is a dummy equal to one if the individual is not participating and the independent variables are estimated family income when working, estimated family income when not working, BMI, marital status, age, and survery year (see Table 2).[13] The

[13]Provided that not all of the variables used in the two-step selection correction are included in the final probit. the model will be identified. When the log of height was included as an explanatory variable. the coefficient was insignificant. Height appears to be highly correlated with income.

Table 1: Participation Probit and Earnings Equations Used to Estimate Income in and out of the Labor Force, NHIS, Men 50-64

| | | participation probit | | participant income | | non-participant income | |
|---|---|---|---|---|---|---|---|
| variable | mean | para-meter | std err | para-meter | std err | para-meter | std err |
| probability participation | 0.23 | . | | . | | . | |
| log of participant income | 10.41 | . | | . | | . | |
| log of non-participant income | 9.83 | . | | . | | . | |
| intercept | | $22.713^{\ddagger}$ | 1.856 | $8.792^{\ddagger}$ | 0.949 | $3.804^{*}$ | 1.936 |
| BMI | 26.4 | $-0.894^{\ddagger}$ | 0.112 | $0.189^{\dagger}$ | 0.055 | 0.119 | 0.088 |
| BMI squared | 712.24 | $0.029^{\ddagger}$ | 0.004 | $-0.006^{\ddagger}$ | 0.002 | -0.003 | 0.003 |
| BMI cubed/10 | 1959.45 | $0.003^{\ddagger}$ | 0.000 | $0.001^{\dagger}$ | 0.000 | 0.000 | 0.000 |
| age | 56.86 | $-0.628^{\ddagger}$ | 0.054 | -0.036 | 0.027 | $0.107^{*}$ | 0.060 |
| age squared | 3233.06 | $0.007^{\ddagger}$ | 0.000 | 0.000 | 0.000 | -0.001 | 0.001 |
| dummy=1 if | | . | | . | | . | |
| care limitation | 0.03 | $1.780^{\ddagger}$ | 0.044 | . | | . | |
| less than high school | 0.39 | . | | . | | . | |
| high school graduate | 0.36 | $-0.121^{\ddagger}$ | 0.018 | $0.047^{\ddagger}$ | 0.008 | $0.177^{\ddagger}$ | 0.018 |
| some college | 0.14 | $-0.200^{\ddagger}$ | 0.024 | $0.143^{\ddagger}$ | 0.010 | $0.366^{\ddagger}$ | 0.026 |
| graduate school | 0.11 | $-0.468^{\ddagger}$ | 0.029 | $0.328^{\ddagger}$ | 0.011 | $0.651^{\ddagger}$ | 0.034 |
| married | 0.86 | $-0.397^{\ddagger}$ | 0.021 | $0.372^{\ddagger}$ | 0.011 | $0.587^{\ddagger}$ | 0.020 |
| veteran | 0.65 | 0.020 | 0.017 | $0.098^{\ddagger}$ | 0.007 | $0.135^{\ddagger}$ | 0.019 |
| foreign-born | 0.03 | -0.045 | 0.047 | $-0.104^{\ddagger}$ | 0.019 | $-0.306^{\ddagger}$ | 0.049 |
| non-msa | 0.73 | . | | . | | . | |
| msa 1,000,000 or more | 0.38 | $-0.097^{\ddagger}$ | 0.020 | $0.276^{\ddagger}$ | 0.008 | $0.242^{\ddagger}$ | 0.021 |
| msa <= 1,000,000 | 0.35 | $-0.033^{*}$ | 0.020 | $0.182^{\ddagger}$ | 0.008 | $0.151^{\ddagger}$ | 0.020 |
| northeast | 0.22 | . | | . | | . | |
| midwest | 0.27 | $-0.045^{\dagger}$ | 0.023 | $-0.060^{\ddagger}$ | 0.009 | -0.009 | 0.024 |
| south | 0.31 | $0.119^{\ddagger}$ | 0.022 | $-0.078^{\ddagger}$ | 0.009 | $-0.083^{\ddagger}$ | 0.023 |
| west | 0.20 | $0.100^{\ddagger}$ | 0.100 | $-0.028^{\dagger}$ | 0.010 | -0.020 | 0.025 |
| 1985 | 0.14 | . | | . | | . | |
| 1986 | 0.09 | 0.009 | 0.034 | 0.009 | 0.014 | $0.080^{\dagger}$ | 0.034 |
| 1987 | 0.16 | 0.031 | 0.029 | $0.070^{\ddagger}$ | 0.012 | $0.088^{\ddagger}$ | 0.029 |
| 1988 | 0.16 | 0.022 | 0.029 | $0.111^{\ddagger}$ | 0.012 | $0.112^{\ddagger}$ | 0.029 |
| 1989 | 0.15 | 0.047 | 0.029 | $0.156^{\ddagger}$ | 0.012 | $0.219^{\ddagger}$ | 0.030 |
| 1990 | 0.15 | 0.040 | 0.030 | $0.196^{\ddagger}$ | 0.012 | $0.249^{\ddagger}$ | 0.030 |
| 1991 | 0.15 | $0.065^{\dagger}$ | 0.029 | $0.214^{\ddagger}$ | 0.012 | $0.286^{\ddagger}$ | 0.030 |
| inverse Mills' ratio | | . | | $-0.221^{\ddagger}$ | 0.032 | $-0.121^{\ddagger}$ | 0.020 |

Note: The symbols $*$, $\dagger$, $\ddagger$ indicate significance at the 10%, 5%, and 1% level, respectively. The participation probit was run on 39,923 observations. The likelihood ratio test for all coefficients except the intercept being different from 0 is 9062. The participant income regression was run on 25,665 observations, adjusted $R^2 = 0.19$. The non-participant income regression was run on 7,517 observations, adjusted $R^2 = .25$. The standard errors in the income regressions are unadjusted, but since the equations are used only for prediction, this is unimportant to interpretation.

covariance matrix is approximated as if estimated family income were the exact exogenous variable.

Tests revealed that a quadratic specification for age and a cubic specification for BMI fit the data. The use of dummy variables for age and BMI indicated that interactions between age and BMI were small and insignificant.[14] The BMI level that maximizes the probability of participation is 25.4 – statistically indistinguishable from the BMI level that at the sample mean height of 1.78m minimizes mortality risk. When BMI is interacted with a dummy for household income out of the labor force of less than $20,000 per year, the BMI level that maximizes the probability of participation of men in households where non-labor force participation income is less than $20,000 per year is 25.9 compared to 24.7 among men in households where annual income is more than $20,000. This difference in BMI levels that maximize the probability of participation is accounted for by differences in height. When dummies are used for BMI levels, the risk of non-participation is minimized at BMI levels of 23-26 and the risk of non-participation is much greater at lower BMI levels (cf. Figure 4). When dummies for a low BMI ($< 22$) and a high BMI ($> 30$) are included in the regression and interacted with income both in and out of the labor force, the resulting coefficients are small and insignificant, suggesting that the impact of health on the probability of participation does not vary by income.

Because the receipt of a Union Army pension was not contingent upon labor force status, the only income difference between the retirement and participation option in the Union Army sample was the income received when working. Income when working is $Y_w = B + Y + N$ and income when not working $Y_l = B + N$ where $B$ is pension amount, $N$ is non-labor market income, and $Y$ is labor market income. Neither wages nor income are observed in the Union Army sample. Therefore I use past occupation as a proxy for the opportunity cost of not working.

[14] The use of a spline also indicated that retirement did not rise abruptly at age 62. The impact of Social Security benefits appears to be captured by income when out of the labor force.

Table 2: Probit of Determinants of Probability Non-participation with Non-partici:ation as the Dependent Variable, NHIS 1985-1991

| variable | mean | para-meter[a] | std error | mean derivative |
|---|---|---|---|---|
| non-participating | 0.23 | . | | |
| intercept | | 20.230‡ | 1.811 | |
| family income/10,000[b] | | . | | |
| if non-participant | 2.15 | 0.292‡ | 0.031 | 0.033 |
| if participant | 3.33 | −0.595‡ | 0.031 | -0.111 |
| age | 56.86 | −0.537‡ | 0.522 | -0.133 |
| age squared | 3251.86 | 0.005‡ | 0.000 | 0.001 |
| dummy=1 if married | 0.86 | −0.052† | -0.024 | -0.015 |
| dummy=1 if year is | | . | | |
| 1985 | 0.12 | . | | |
| 1986 | 0.09 | -0.019 | 0.033 | 0.000 |
| 1987 | 0.17 | 0.082‡ | 0.028 | 0.020 |
| 1988 | 0.16 | 0.133‡ | 0.029 | 0.030 |
| 1989 | 0.15 | 0.161‡ | 0.029 | 0.041 |
| 1990 | 0.15 | 0.209‡ | 0.030 | 0.050 |
| 1991 | 0.16 | 0.233‡ | 0.030 | 0.057 |
| BMI | 26.43 | −0.733‡ | 0.108 | -0.208 |
| BMI squared | 712.36 | 0.024‡ | 0.004 | 0.007 |
| BMI cubed/10 | 1959.90 | −0.003‡ | 0.000 | -0.001 |
| 39,923 obs | | . | | |

[a]The symbols *, †, and ‡ indicate 10%, 5%, and 1% significance levels, respectively. The likelihood ratio for the test that the coefficients on all variables except for the intercept are equal to 0 is 6773.

[b]Family income was estimated by means of the Heckman two-step selection correction. Additional variables that were included in the income and participation regressions were the extent of urbanization, education, geographic region, foreign birth, and the presence of a care limitation.

Past occupation may be a poor proxy for opportunity cost if the ill are no longer able to work in their usual occupation. For these men the opportunity cost of retirement is underestimated.

Several other proxies for earnings and wealth are available. Illiteracy and foreign birth may indicate lower than average earnings. Marital status may also be measuring earnings if employers favor married men or if married men were more skilled. In 1900, married males in manufacturing earned 17 percent more than unmarried males, controlling for the observable characteristics of workers and their jobs (Goldin 1990: 102). Among home owners in cities, letting rooms to boarders increases family income, but may be symptomatic of economic difficulties (Modell and Harevan 1973). The hire of a servant is an indicator of affluence. Homeownership meant that the person had wealth, because a substantial down payment, generally equal to half of the value of the purchased property, was required (Haines and Goodman 1992).[15] Higher unemployment in the veteran's current state may induce more retirement.[16]

In the probit that is estimated for the Union Army sample the dependent variable is a dummy equal to one if veteran was retired (see Table 3). The independent variables are pension amount, occupational dummies, BMI, property ownership, age, characteristics of region of residence, and other socio-economic variables. Information on pension amount is unavailable for 126 men even though these men were on the pension rolls. Because tests indicated that the unavailability of pension amount is determined by random factors, these men were excluded from the sample. Tests revealed that both pension amount and age should be entered linearly in the specification. In the past retirement did not rise abruptly with age. The impact of Union Army pensions on retirement is discussed in previous work (Costa 1994).

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[15]However, because property was one of primary modes of saving, men who had retired might already have liquidated their property.

[16]Margo (1993) finds that the long-term unemployed soon retired. The statewide unemployment numbers are from Table A.13 in Keyssar (1986: 340-341).

Table 3: Probit of Determinants of Probability of Non-participation with Non-participation as the Dependent Variable, Union Army Sample

| variable | mean | parameter[a] | std error | mean derivative |
|---|---|---|---|---|
| non-participating | 0.16 | . | | |
| intercept | | $-6.257^{*}$ | 3.337 | |
| BMI | 22.97 | $-0.551^{\ddagger}$ | 0.160 | -0.107 |
| BMI squared | 540.56 | $0.011^{\ddagger}$ | 0.003 | 0.002 |
| age | 61.32 | $0.054^{\ddagger}$ | 0.012 | 0.011 |
| monthly pension | 12.73 | $0.030^{\ddagger}$ | 0.010 | 0.006 |
| state unemployment[b] | 3.62 | $2.168^{\ddagger}$ | 0.708 | 0.419 |
| dummy=1 if farmer | 0.46 | . | | |
| professional or proprietor | 0.17 | $-0.552^{\dagger}$ | -0.107 | -0.119 |
| laborer | 0.22 | −0.083 | 0.227 | -0.016 |
| artisan | 0.15 | 0.135 | 0.234 | 0.026 |
| does not own residence | 0.31 | $0.319^{*}$ | 0.178 | 0.062 |
| atlantic seaboard | 0.20 | . | | |
| midwest | 0.74 | $0.583^{\dagger}$ | 0.279 | 0.113 |
| other region | 0.06 | −0.146 | 0.547 | -0.028 |
| urban county | 0.36 | $0.363^{*}$ | 0.177 | 0.070 |
| 4 or more dependents | 0.14 | −0.538 | 0.314 | -0.104 |
| servant present | 0.02 | −0.747 | 0.663 | -0.144 |
| boarder present | 0.04 | −0.034 | 0.421 | -0.007 |
| illiterate | 0.06 | 0.079 | 0.330 | 0.015 |
| foreign-born | 0.10 | −0.065 | 0.272 | -0.012 |
| 471 obs | | . | | |

[a]The symbols *, †, and ‡ indicate 10%, 5%, and 1% significance levels, respectively. The likelihood ratio for the test that all coefficients except for the intercept are equal to 0 is 96.11.

[b]Mean duration of unemployment in months for manufacturing workers (Table A.13 in Keyssar 1986: 340-341)

Because Figure 5 indicated that the risk of labor force non-participation was strongly U-shaped in BMI levels, a quadratic rather than a cubic specification was employed.[17] Note that the BMI level that maximizes labor force participation is 25.2 – well within the range that maximizes labor force participation in Figure 5. This BMI level is statistically indistinguishable from that which for the sample mean height of 1.73m minimizes mortality risk. Coefficients on the interactions of BMI dummies with pension amount were small and significant.

Note that in contrast to the NHIS the body build of Union Army veterans places them at high risk of relative non-participation. Mean BMI in the NHIS is 26.4 and the standard deviation is 3.7. Thus the mean is within the NHIS level that maximizes the probability of participation by less than half a standard deviation. However, because mean BMI in the Union Army sample is 23.0 and the standard deviation is 3.6, the mean is not within half a standard deviation of the Union Army level that maximizes the probability of participation.[18]

In both the Union Army sample and the NHIS, the elasticity of labor force non-participation with respect to BMI is greater at both lower and higher BMI levels, suggesting that in a population such as the Union Army sample in which mean BMI is low, health is more likely to have a larger impact on participation rates. But, even at the mean BMI prevailing in the NHIS, the elasticity in the Union Army sample is a substantial 0.88, while the elasticity in the NHIS is only 0.28. Of course, the elasticities are not strictly comparable. Different specifications were employed and characteristics differ across samples.[19] But, this finding is consistent with the

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[17]When either a cubic or a dummy variable specification is used the coefficients on BMI are no longer significant. Work with the NHIS indicates that significance of coefficients on BMI are more sensitive to sample size when a dummy variable rather than a continuous specification is used.

[18]The observed relation between BMI and labor force participation is not an artefact of distance of the mean from the level that maximizes the probability of participation. When the NHIS is randomly restricted to have the same fraction of men in BMI categories as the Union Army sample, the relation between health and participation remains unchanged.

[19]Even controlling for differences in characteristics, the magnitude of the difference in elasticities remains unchanged.

observed BMI-labor force participation relations in which the relative risk of non-participation among men at low and especially at high BMI levels was greater in 1900.

Those with higher Union Army pensions may be less healthy, but their poorer health may be unobservable. Futhermore, although pensions were awarded regardless of participation status, non-participation may have been viewed by employees of the Pension Bureau as evidence of inability to perform manual labor. Thus, because of dependence on unobservable retirement determinants, pension status is potentially endogenous and all coefficients are biased. But, as discussed in Costa (1994), the ability to establish whether a disability was traceable can be used as an instrument variable because it predicts pension amount and is arguably not related to unobserved retirement determinants conditional on measured health status. A Hausman test for exogeneity of pension amount suggests that endogeneity of pension amount is not a problem (Costa 1994). In fact, when an instrumented probit is estimated, the BMI level that maximizes the probability of participation remains unchanged at 25.4, as does the elasticity.

Another source of endogeneity may be that between BMI and labor force participation. In the case of Union Army veterans, retirement may have caused ill health if retirement incomes were insufficient. Contemporaries argued that unemployed men experienced a deterioration in health and skills that led to their becoming members of the casual laboring class. Long-term unemployment was frequently a prelude to retirement (Margo 1993). Unfortunately it is not possible to test whether BMI and labor force participation are endogenous. Food prices are often used as instruments for BMI, but county level food prices are not available. State price indices and the fraction of dwelling that were farms within a county are not correlated with BMI. But, because BMI did not differ among farmers, artisans, laborers, this suggests that insufficiency of income was probably not an important determinant of BMI.

## 5 Interpreting BMI

The meaning of BMI may have changed since 1900. Differences in muscle mass can skew optimal weight for height comparisons and the percentage of muscle mass may differ across the two samples because the occupational structure has shifted from one in which physically strenuous occupations are common to one in which sedentary occupations are. Cigarette smoking became a widespread practice after World War II. The higher risk of non-participation among the thin in the NHIS could be explained by the fact that those thinner are more likely to be smokers and thus more likely to be in ill health.[20] Differences in the disease environment might also skew optimal weight for height recommendations. In the Union Army sample the prevalence rate for diarrhea among men 65 years of age or older was 32% compared to less than 2% in the NHIS (Fogel, Costa, and Kim 1993). In an environment in which infectious diseases are common, optimal weight for height recommendations may be greater.

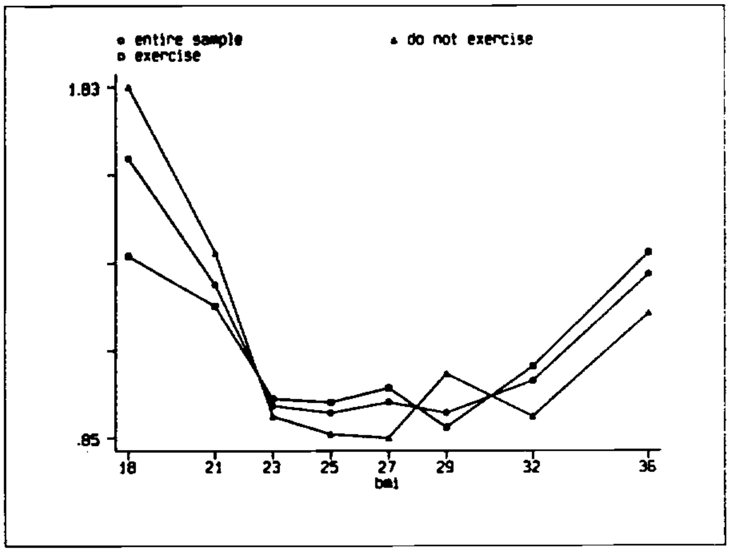
Although no information on muscle mass is available, information on excercise habits is available in the 1985, 1990, and 1991 NHIS. Figure 6 shows the relationship between labor force participation and BMI in the entire sample when the sample is divided into those who excercise regularly and those who do not. Note that the relationship between labor force participation and BMI does not change by that much, suggesting that the estimated relationship is not confounded by differences in muscle mass.[21]

Nor does the relationship between labor force participation and BMI change when a sample consisting of the years 1985, 1987, 1990, and 1991 is divided into smokers and non-smokers (see Figure 7). Furthermore, controlling for other characteristics, the addition of smoking

[20] Lee *et al.* (1993) argue that the influence of cigarette smoking can skew weight recommendations that are optimal in terms of subsequent mortality.

[21] In both Figure 6 and Figure 7, the relative risk of non-participation is lower than in Figure 4. Because of missing information, few men in the sample had a BMI of less than 18.

Figure 6: BMI and Relative Risk Labor Force Non-participation by Exercise Habits, 1985, 1990, and 1991 NHIS




Note: All observations are centered at the marks. There are 7704 men in the entire sample, 2149 of whom report absolutely no activity for exercise.

status as an explanatory variable in predictions of labor force participation does not change the coefficients on dummies for BMI levels or the coefficients on the cubic specification for BMI.[22]

The relationship between BMI and labor force participation remains similar in both the NHIS and the Union Army sample when the very sick are deleted from the sample, suggesting that the relationship between BMI and labor force participation is not the result of current illness alone (see Figure 8).[23] BMI can thus be thought of as a measure of the stock of health. Of course, chronic conditions will in part determine the relationship between labor force participation and BMI. When men who reported at least one condition are deleted from the NHIS, the obese are no

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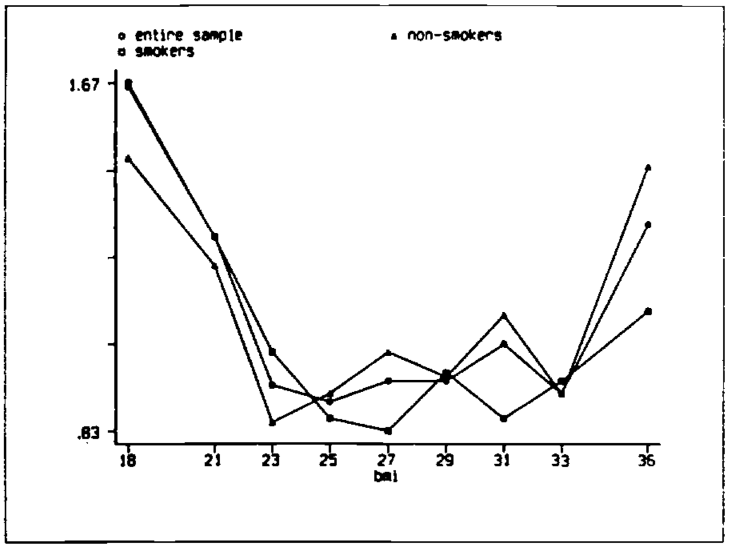
[22]When smoking is interacted with BMI the coefficient on the resulting variable is insignificant. Once again the coefficients on BMI remain unchanged.

[23]Among men who are in the labor force, the relation between BMI and sick days is U-shaped and similar to that between BMI and labor force participation.

Figure 7: BMI and Relative Risk Labor Force Non-Participation By Smoking Status, 1985, 1987, 1990, and 1991 NHIS



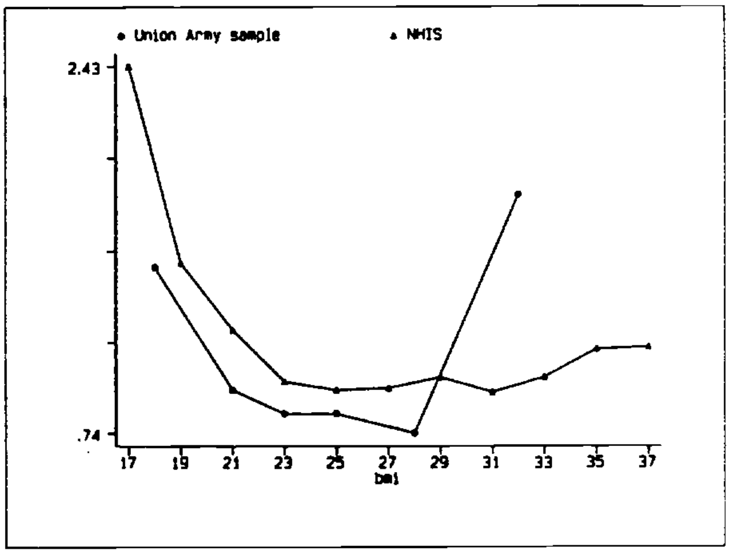

Note: All observations are centered at the marks. There are 9901 men in the entire sample, 2667 of whom are current smokers.

longer at greater risk of non-participation than those of average weight for height. However the thin still face a greater risk of non-participation.[24]

BMI is related to chronic conditions and the relative prevalence rates of chronic conditions has changed (Fogel, Costa, and Kim 1993). The impact of chronic conditions on labor force participation differs greatly by chronic condition (Bartel and Taubman 1979). If chronic conditions that are now correlated with a high BMI do not have as large an impact on labor force participation as conditions that in the past were correlated with high BMI then high BMI levels will no longer have as large an impact on participation rates. As more data is collected from the Union Army records it will be possible to investigate whether the relationship between BMI and

[24]The obese might be more likely to be diagnosed as having a chronic condition, because of the strong focus in the medical literature on the problems of obesity.

Figure 8: BMI and Relative Risk Labor Force Non-Participation, Union Army Sample and NHIS, Deleting the Very Ill




Note: All observations are centered at the marks. There are 445 men in the Union Army Sample and 34,336 men in the NHIS.

the prevalence of chronic disorders has changed.[25]

# 6 Implications

The relationship between BMI and the risk of labor force non-participation in 1900 was shown to be remarkably similar to that in 1985-1991 – the BMI level that maximizes the probability of participation was about 25 in both time periods and the relative risk of non-participation was greater at low and high BMI levels. There were, of course, differences. The observed BMI-participation relation and the elasticities of labor force non-participation with respect to BMI

[25]There is some evidence of a changing relationship between heart disease and BMI. In the Union Army sample prevalence decreased with increasing BMI, perhaps because much of the heart disease was valvular rather than arterio-sclerotic.

indicated that men in 1900 were much more responsive to changes in health than they are now. Several factors could account for this. One possibility is that the relationship between chronic conditions and BMI has changed. Another possible explanation for the change at high BMI levels is that the changing nature of jobs has allowed men with chronic conditions associated with high BMI levels to continue to work. The rise of the clerical sector, the increased mechanization of manufacturing, greater safety in the workplace, and the shortening of the work day have reduced the physical demands of most occupations. Heart disease, associated in recent populations with high levels of BMI, may not be a significant predictor of non-participation (Bartel and Taubman 1979).[26]

The findings do not necessarily imply that as the population grows increasingly healthier, jobs less physically demanding, control over chronic conditions improves, and greater efforts made to increase the incorporation of the old and disabled into the labor force, labor force participation rates among older men will rise. Secularly rising incomes may lower participation rates. In fact, the secular rise in income may explain up to 60% of the fall in labor force participation rates among men 65 years of age or older (Costa 1994). By affecting the wage, improved health may lower labor force participation rates. In the NHIS, conditioning on labor force participation, BMI is a strong predictor of family income. The elasticity of family income with respect to BMI ranges from 0.50 for the 10th income decile to 0.11 for the 50th, and to 0 for the 80th.[27] Recent cross-sectional evidence indicates that the male labor supply curve is gently backwards sloping (Killingsworth 1983: 102). In the nineteenth century, the labor supply curve for working men was strongly backwards bending (Whaples 1990).

Without these complications, the probit estimates from the Union Army sample suggest

[26]It is, however, a significant predictor of the number of hours worked each week.

[27]Calculated from quantile regressions for men 60-54 years of age.

that had the distribution of BMI in the Union Army sample been the same as that in the NHIS the probability of non-participation would have fallen by 6%. If participation rates among older men could be increased with out affecting the employment prospects of the young, then, abstracting from any complications caused by substitutability of healthy and unhealthy workers, the increase in participation among older men alone would have led to a 1.5% increase in GNP.[28] But how could such a shift in BMI have been achieved? Both early childhood and adult environmental conditions have improved due to advances in medical care, public health investments in sanitation, improved working conditions, and rising incomes. The mounting body of evidence of a relationship between chronic conditions at late adult ages and early childhood environmental factors (Barker 1992) suggests that these factors have improved adult health and therefore BMI. However, these improvements may have changed the relationship between BMI and the relative risk of non-participation. If the increase in BMI was necessarily accompanied by a shift in the relationship between BMI and the relative risk of labor force non-participation, then had the distribution of BMI in the Union Army sample been the same as that in the NHIS the probability of non-participation would have fallen by 10%. Ignoring whether the factors that also improved adult health affected the probability of labor force participation, GNP would have increased by 2.3%.

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