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THE EXPANDING WORKWEEK? UNDERSTANDING TRENDS IN LONG WORK HOURS AMONG U.S. MEN, 1979-2004

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ABSTRACT

According to Census and CPS data, the share of employed American men regularly working more than 48 hours per week is higher today than it was 25 years ago. Using CPS data from 1979 to 2006, we show that this increase was greatest among highly educated, highly-paid, and older men, was concentrated in the 1980s, and was largely confined to workers paid on a salaried basis. We rule out a number of possible explanations of these changes, including changes in measurement, composition effects, and internet-facilitated work from home. Among salaried men, increases in long work hours were greatest in detailed occupations and industries with larger increases in residual wage inequality and slowly-growing real compensation at 'standard' (40) hours.

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I. Introduction

Between 1970 and 1990, the share of employed-for-pay U.S. men who worked more than 48 hours in the Census reference week rose from 15.4 to 23.3 percent. Between the 1980 Census and 2005 American Community Survey, the share usually working more than 48 hours per week rose from 16.6 to 24.3 percent. Seen in the context of the long-term increase in leisure during the bulk of the 20th century (see for example Costa 2000, Aguiar and Hurst 2007), this recent increase seems surprising, and of likely interest to labor economists wishing to understand long-term trends in work behavior.

Why are employed American men are more likely to put in long work weeks today than a quarter century ago? We attempt to answer this question using Current Population Survey Outgoing Rotation Group data from 1979 to 2006. We begin by identifying both the time periods and the parts of the male labor force where the recent increase was the strongest: The increase in long hours was strongest before 1990, slowed during the 1990s, and actually reversed somewhat after 2000. It was concentrated among highly educated, high-wage, and older men, and largely confined to workers paid on a salaried basis. Next, we assess a number of simple explanations for trends in long hours over this period, including changes in the demographic composition of the labor force, changes in the mix of occupations or industries, changes in survey techniques, and technological changes like internet-facilitated work from home. Finally, we examine the determinants of (roughly) decadal changes in long hours at the detailed industry and occupation level among salaried men. We find that increases in long work hours were greatest in detailed occupations (and industries) with (a) rising residual wage inequality and (b) slowly-growing real compensation at ‘standard’ (40) hours.

II. Data and Descriptive Statistics

Figure 1 illustrates long-term trends in the prevalence of employed men’s long work hours, taken from seven decades of U.S. Census microdata and the 2005 American Community Survey (ACS). Part (a) of the Figure shows a decline between 1940 and 1970 in the share of employed 25-64-year-old men who worked more than 48 hours in the reference week, followed by a rise between 1970 and 1990. The Census did not collect information about reference week hours in 2000, but information on usual hours (while employed) in the calendar year preceding the Census is available from 1980 to 2000, and in the 2005 ACS. This shows an increase from 1980 to 2000, followed by a decline, but not to 1980 or 1990 levels. Part (b) of Figure 1, which removes self-employed men from the sample, shows that most of the decrease in men’s hours between 1940 and 1970 was associated with the decline in self employment (much of it agricultural) that occurred over that period. After 1970, it shows similar trends to part (a), with an even larger increase between 1970 and 2000.[1]

Our primary data sets for this paper are the Outgoing Rotation Groups (ORGs) of the CPS from 1979 to 2006. Aside from better wage information, these data have the advantages of consistent hours measures across many years, a large sample size that is representative of hours worked during the entire calendar year, and information on the

[1] In parallel with all other indicators of women’s labor force attachment, the incidence of long work hours increased among women over the past three decades as well. We focus on men in this paper in part because they are less affected by long-term changes in culture and gender role expectations than women, and in part because of the surprising contrast between men’s lengthening work week and their reduced labor force attachment on other dimensions.

method of pay (salaried vs. hourly) which plays an important role in our analysis. Throughout the paper we restrict attention to men aged 25 to 64, who are employed in the survey week, but not self employed on their main job.[2]

In contrast to some labor supply studies which focus on annual hours of work (e.g. Coleman and Pencavel 1993), our focus throughout this paper is on weekly hours of work among those with positive hours. In part this is because long-term trends in annual work hours are affected by technical innovations affecting labor market matching efficiency, such as the growth of the temporary help industry (Katz and Krueger 1999) and internet job matching services (Kuhn and Skuterud 2004), which are not of interest to us here.[3] Another motivation is that CPS annual work weeks information suffers from a serious measurement problem: it does not subtract vacations and other forms of leave from measured work time. In addition, adults' weekly hours may be of interest in their own right, for example if their effect on child development differs from those of weeks worked per year. In addition to focusing on weekly hours, we place most of our emphasis in this paper on a particular feature of the weekly hours distribution, namely the fraction of full-time (30 or more usual hours) workers usually working 50 hours or more.[4] In part this is because this measure is less affected by high-hours outliers than the mean; that said, we have replicated most of the results in this paper for mean hours among full time workers, and they are very similar to those reported.

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[2] On average, self-employed men work longer hours than other men. However since 1979 there has been no upward trend in either men's self employment rate, or in the fraction of the self employed who work long hours.

[3] Relatedly, our interest here is in labor supply choices made voluntarily by workers, and it unclear whether periods of unemployment reflected in annual hours are best modeled in a simple voluntary labor supply framework (e.g. Ham 1986).

[4] The Census data shown in Figure 1 focuses on men working over 48 hours because of the categorical nature of older Census data. Our CPS results are not sensitive to the cutoff point.

Turning to our CPS data, the fraction of men in our sample who report that they usually work 50 or more hours per week on their main job is plotted in panel (a) of Figure 2 for every year between 1979 and 2006.[5] For context, panel (b) plots the employment-population ratios of men aged 25-54, taken from published BLS data over a slightly longer period. Together, these figures show that (a) the incidence of long work hours fell in the recessions of 1983, 1992 and 2002; and (b) that long work hours rose sharply in the 1980s, more slowly in the 1990s, and --as in the Census data-- declined somewhat between 2000 and 2006.[6] Further, note that the increase in long work hours coincides with a secular *decline* in men's employment-population ratio.

It is well known that the CPS underwent a major redesign in 1994, and that this redesign included substantial changes to some of the work hours questions. For three main reasons, however, this redesign can explain neither the increase in long reported work hours nor its distribution across population subgroups. First, unlike the survey-week hours question which did change, the wording of the main-job hours question used in this paper --"How many hours does … USUALLY work at this job?"--, is unchanged over this entire period.[7] Second, as Figure 2 indicates, most of the increase in long hours predates the CPS redesign, and the series exhibits no detectable "jumps" between 1993 and 1994. Third, similar patterns of change in men's long work hours are also observed

[5] The CPS defines the main job as the one with the highest weekly hours.
[6] The Census data does appear to show a more robust increase in the 1990s than the CPS. This may be attributable to a number of differences between the Census and CPS hours questions, including the Census's inclusion of all jobs held (not just the main one).
[7] There is some difference in placement of the questions: after 1994 this question became part of the basic monthly survey rather than the ORG earner supplement. Unlike some other hours questions, however, in neither case is this question immediately preceded by other questions about work hours, which might frame responses differently in different years. Finally, after 1994, the CPS allowed workers to answer "hours vary" instead of reporting a usual level of hours. Interestingly, mean survey week hours for workers who chose this response were almost identical to those who did not (43.8 versus 43.3 in 1994). Further, the share of workers choosing "hours vary" remained virtually unchanged at about 6 percent throughout the period 1994-2002.

in other nationally-representative data sets. For example, in addition to the Census data reported in Figure 1, our own analysis of employed, non-self-employed men aged 25-64 in the General Social Survey (GSS) survey shows a higher overall incidence of long hours but exactly the same time trends: a substantial increase in long work hours, most of it concentrated in the 1980's rather than the 1990's.[8]

More detail on the size and distribution of the increase in men's long work hours is provided in Table 1. To abstract from business cycle effects, Table 1 (and a number of subsequent Tables) focuses on three years at similar points in the business cycle: 1979, 1989, 2000 and 2006. As Figure 2b shows, each of these constituted a peak in men's employment rates, with the possible exception of 2006.[9] Overall, according to CPS data the incidence of long work hours increased relatively modestly between 1979 and 2006, from 16.1 percent in 1979 to 17.8 percent among all men and 16.4 to 19.5 percent among men usually working more than 30 hours per week ("full time"). Since our main interest in this paper is on the fraction of full-time men who choose to work long work weeks, most of the remaining calculations in this paper restrict attention to full time workers only.

In the remainder of Table 1 we subdivide the population of full-time men in various ways in order to identify subgroups where the increase in long work hours has been the largest. This shows that salaried men are much more likely to work long hours than hourly-paid men, and also that the increase in long hours has been substantially

[8] Results available at http://www.econ.ucsb.edu/~pjkuhn/Data/DataIndex.html. To our knowledge, the only data that fail to confirm the trend of an increase in long work hours are Robinson and Godbey's (1997) time-diary studies. We conjecture that this may be explained by the low, and secularly declining response rates to their surveys. If those men who work long hours are less likely than others to participate in the arduous process of filling out a time diary, it seems quite likely that their survey technique could fail to detect an increasing incidence of long hours in the population.

[9] Very similar results occur when we compare business cycle troughs over time.

greater among salaried workers (from 24.4 to 30.1 percent), than among the hourly-paid (from 8.6 to 9.6 percent). Perhaps surprisingly, the incidence of long hours actually fell between 1979 and 2006 in our youngest age group (25-34), and was greatest among men aged 55-64: while older men were less likely to work long hours than young men in 1979, this differential had reversed by 2006. In all the years for which we have data, long work hours were much more common among college graduates than among workers with less education; the increase in long hours was also greatest among the college-educated. In fact, our data show a decline long work hours among high school dropouts over our entire period.

The bottom panel of Table 1 examines the correlation of work hours trends with a worker's rank in the wage distribution. It does so by ranking workers in each of the four selected years according to their average hourly earnings.[10] The results show that the recent increase in long work hours has been concentrated among the highest wage earners: between 1979 and 2006, the frequency of long work hours increased by 11.7 percentage points among the top quintile of wage earners, while falling by 8.4 percentage points in the lowest quintile. Figures 3 and 4 show the long-hours share and the mean level of hourly earnings for all five wage quintiles for every year in our sample; they show (a) that the increase in long hours was strongest among the *only* quintile to experience significant real wage growth between 1979 and 2006 (the highest earners), but (b) that this increase was strongest during the period (before 1997) when that group

[10] Table 1 makes no attempt to adjust for the top coding of earnings information in many years of the CPS. Later in the paper we use a Tobit procedure to make this adjustment: see footnote 22. Division bias driven by measurement error in hours (e.g. Borjas 1980) could also affect the bottom panel of Table 1, counteracting any positive cross-sectional association between true hours and wages. Since there is no obvious reason to expect the amount of division bias to change over time, it still seems likely that Table 1 provides useful information about time trends.

experienced no wage growth. We study these trends in greater depth in Section IV of the paper.

Perhaps the most striking feature of Table 1 is the reversal in the cross-sectional relationship between hourly wages and work hours since the early 1980's: in 1979, the worst-paid 20 percent of workers were more likely to put in long work hours than the top 20 percent; by 2006 the top 20 percent were twice as likely to work long hours than the bottom 20. This reversal of men's wage-hours relationship forms part of the recent change in men's labor supply behavior that we attempt to understand in this paper.[11]

We have already noted that, overall, men's employment-to-population ratio fell over the period we are studying. This raises the possibility that the increase in long work weeks documented above is "illusory" in the sense that, for a randomly selected worker, it was offset by decreases in other dimensions of labor supply. In the remainder of this section, we briefly examine weekly hours changes in the context of other dimensions of labor supply variation; the story that emerges is considerably more complex than a simple substitution of main-job hours for other aspects of labor supply.

We turn first to moonlighting. As noted, the main indicator of work hours used in this paper refers only to the respondent's "main" job. Is it possible that the increases in long hours documented in Table 1 were offset by a decline in multiple job holding, or by fewer hours worked in "second" or higher-order jobs? To address this issue, we examined evidence from May CPS surveys in 1979, 1991 and 2001. Unlike the regular CPS before 1994, the 1979 and 1991 Supplements contain information on multiple

[11] Some of the trends described in Table 1 have been noted by other authors. For example, Rones, Ilg and Gardner (1997) report an increase in the share of persons working more than 48 hours in CPS data from 1976 to 1993. Coleman and Pencavel (1993) document an increase between 1940 and 1988 in mean *annual* hours among well-educated workers. Finally, Costa (2000) documents a reversal between 1973 and 1991 in the relative *daily* hours worked by high-wage versus low-wage men, with the high wage earners working longer days by 1991.

jobholding, as well as on total usual hours worked in all jobs. As suggested by the Census data in Figure 1 (which includes hours on all jobs), this analysis shows a similar increase in work hours when all jobs are taken into account. The reason is simple: there was little change in either the rate of multiple jobholding or in the incidence of long work hours among multiple jobholders. Instead, essentially the entire increase in long *total* work hours was in the usual hours of workers who held only one job.

We next turn to employment rates. As noted, men's aggregate employment-population ratio fell over this period; this raises the possibility that what happened to men in the 1980s was a "concentration" of a relatively constant total amount of labor into periods of more intense work activity separated by more bouts of inactivity, low activity, and/or earlier retirement. [12] To address this hypothesis, Table 2 provides more detail on the nature of employment changes over this period. Looking first at all age groups combined, Table 2 shows that the decline in men's employment rates between 1979 and 2006 was greatest in the two middle education categories, not the highest where the increase in long hours was concentrated. Focusing on the period from 1979 to 1989, when most of the rise in long hours occurred, the decline in men's employment rates was much greater among the education group --high school dropouts-- among whom weekly long hours became less common, and very small among college graduates, where long hours rose considerably.

[12] This scenario could in fact emerge quite naturally from a life-cycle labor supply framework (e.g. Heckman and MaCurdy 1980) subject to two types of exogenous shocks. One is simply an increase in the year-to-year variation in labor productivity, or hourly wages, facing each worker. Another would be an increased nonconvexity in the production function relating (say) weekly hours worked by an individual to weekly output, of the sort modeled by Rogerson (1988) and Mulligan (1999), among others. For example, it is sometimes argued that the production of computer code is most efficiently accomplished in bouts of long hours, or that managerial jobs are (increasingly) optimally very intense but followed by early retirement.

The above pattern is illustrated much more dramatically if we restrict attention to men aged 45-54, among whom the overall decline in men's employment rates was particularly steep, but who are unlikely to be affected by changes in Social Security policy over this period. Clearly, these trends are not consistent with a scenario in which a "representative" man's total labor supply remained roughly constant over this period, with declines in employment probabilities roughly offsetting increases in hours when employed.

Two remaining dimensions along which compensatory declines in labor supply might have occurred during this period are via an increase in the incidence of part-time work, or longer annual vacations. Examination of the data underlying Table 1 reveals that there has been relatively little change over our period in the share of men working part time, and that the small increase that did occur –like the increase in inactivity-- was greater among *less*-skilled men. Finally, although we know of no consistent microdata over a long period on paid vacations, published statistics from the BLS's Employee Benefits Survey show very little trend between 1980 and 1997 in essentially all dimensions of paid leisure, including annual vacation days and holidays, paid lunch minutes and rest time, and paid sick leave.[13]

In sum, an examination of men's long hours changes in the context of other dimensions of labor supply variation does not support a simple "concentration" hypothesis, in which those groups of workers whose weekly hours increased the most when working also took more breaks between bouts of intense work, in the form of nonemployment, vacations, part-time work, or earlier retirement. Instead, while men's

[13] See U.S. Department of Labor (2004). We provide a summary table at http://www.econ.ucsb.edu/~pjkuhn/Data/DataIndex.html .

overall employment rate fell over this period, it fell the least among those group of workers –the most skilled-- where long work hours increased the most.

III. Composition Effects

1. Demographic Shifts

Both the marginal productivity and the marginal disutility of an extra work hour are likely to vary with worker characteristics (such as age and education) and with job characteristics (such as industry and occupation). As a result, optimal work hours will vary across types of workers and types of jobs, and the secular increase in long hours documented in the previous section might be simply explained by long-term changes in the mix of workers and jobs in the labor force. In this section we decompose trends in long work hours to assess the importance of such factors.

Because we wish to control for the most detailed occupations and/or industries possible, in this section --and in all the subsequent analysis that requires these controls--, we restrict our attention to the period from 1983 to 2002.[14] Also, to generate adequate sample sizes for a large number of detailed industry and occupation groups, we now pool observations for 1983, 1984 and 1985 to represent the beginning of this period; our end-of-period sample comprises 2000, 2001 and 2002. Recalling Figure 2(a), this period contains essentially all of the long-term increase in long hours observed in our data.[15]

Coefficients from a linear probability model for working long hours, and means of the regressors are reported in Table 3 for both of these sample periods. Holding other

[14] Three-digit occupation and industry codes changed dramatically, both in 1983 and near the end of our sample period.
[15] Later in the paper, we distinguish the 1983/85 - 1991/93 and 1991/93-2000/02 subperiods. Of course, ending the sample period at 2002 unfortunately rules out analysis of the dip in long hours observed after about 2000; we briefly speculate on possible reasons for this decline near the end of the paper.

characteristics (including 48 industry and 46 occupation groups) constant, Table 3 shows that educated workers were more likely to work long hours than less-educated workers in both periods. Mirroring the unadjusted trends shown in Table 1, this education differential was considerably larger in 2000-02 than in 1983-85. Somewhat differently from Table 1, older men were less likely to work long hours in both periods; this "pure" age effect was obscured in Table 1 because older men are more likely to be salaried and married (characteristics which contribute to higher hours). Consistent with Table 1, however, the negative impact of age on the propensity to work long hours does weaken (slightly) between the two sample periods. Similarly, the positive partial correlation between marriage and hours, as well as between salaried status and hours, strengthened over time. Black and Hispanic men were less likely to work long hours than other workers in both periods; these coefficients did not change much over time. Interestingly, the union coefficient changed sign from negative to positive, perhaps reflecting a secular decline in union power.

Table 3 also shows that the population of working, American men became better educated, less married, and much less unionized during our sample period. The population of working men did not become unambiguously older over this period; in fact the fraction aged over 55 fell. Finally, as noted by Hamermesh (2002) the share of men paid on a salaried basis did not increase over this period; despite the secular decline in blue-collar employment it actually fell.

The results of a standard Oaxaca decomposition of changes in long work hours are reported in Table 4, which shows a total increase in the fraction of (full-time) men working long hours of 19.6 – 16.6 = 3.0 percentage points between 1983/85 and 2000/02.

According to Table 4, using the 1983/85 regression coefficients, changing observed characteristics accounted for (17.3 – 16.6 =) 0.7 of the 3.0 percentage point increase in long work hours over this period. Using the 2000/02 regressions, only 0.2 points out of 3.0 percentage point difference are thus explained. Although the overall increase in long hours was greater (4.5 percentage points) among salaried workers, changes over time in their observed characteristics are also largely unsuccessful in explaining changes in their work hours: the explained portion is 1.8 or 1.6 percentage points, or about 40 percent of the total, depending on the baseline regression used. Thus, while observed characteristics --including rising education levels, an aging workforce, declining unionization, and a shifting mix of 48 industry and 46 occupation categories—certainly play a role, the majority of the recent increase in long work hours cannot be accounted for by these factors.

2. Detailed industry and occupation mix

It remains possible that the analysis in Tables 3 and 4 fails to capture the true effects of industry and occupational shifts because the categories are too broad. To address this possibility, Table 5 conducts a shift-share analysis of the change in long hours over the same time period using very detailed (three-digit) occupation and industry categories.[16] In all, when we restrict attention to cells with 50 or more observations in both the early (1983-85) and late (2000-02) sample periods, this leaves us with 315 occupations and 201 industries. According to columns 1 and 2 of Table 5, the fraction of men working long hours increased from 16.3 to 20.4 percent between 1983/85 and

[16] Sample sizes make it impractical to include a full set of three-digit industry and occupation fixed effects in the regression-based decomposition of Table 4. David Autor kindly supplied 3-digit occupation codes that are consistent over the entire 1983-2002 period.

2000/02.[17] According to column 3, if the within-occupation long-hours means remained at their 1983/85 levels, but the occupation mix changed to its 2000/02 level, the fraction working long hours would be 17.0 percent. Similarly, column 4 indicates that if we impose the 1983/85 occupation mix on the 2000/02 cell means, there is only a slight reduction in the predicted fraction working long hours for both samples. Thus, confirming the regression-based results in Table 4, detailed occupational shifts explain almost none of the increase in long hours.

The second row of Table 5 replicates the above analysis for industry rather than occupation cells. The results are essentially identical: detailed industry shifts cannot explain the trend, since the great bulk of the increase occurs within cells. Finally, the same conclusion applies when the sample is restricted to salaried workers only. In sum, Table 5 shows that the vast majority of the increase in men's long work hours over the past two decades occurred within very detailed occupation and industry groups. Changes in the mix of jobs performed (including for example the shift from blue collar manufacturing work to service sector jobs) thus cannot account for this increase in hours.

IV. Changes in Work Incentives: Aggregate Trends

Having considered the contribution of measurement techniques and composition effects to the increase in long work hours, we now turn to the effects of changes in financial incentives to supply those hours. In turn, we will consider three proxies for such incentives that can be calculated in the CPS data available to us. The first and most familiar of these is the level of real average hourly earnings. We also consider two

[17] The difference from Table 3 is because (as noted) Table 4 restricts attention to occupation and industry cells with more than 50 observations in each time period.

additional measures which might better capture a salaried worker's marginal incentives to work more than 40 hours per week: a measure we shall call the *long hours premium,* and the level of earnings inequality (at fixed hours) within groups such as 3-digit occupations. In this section, we motivate and describe these measures and document the aggregate trends in each of them over our sample period. The following section conducts a simple econometric analysis of the association between measured incentives and hours at a more disaggregated level.

1. Changes in Average Hourly Earnings

Some simple descriptive statistics on time trends in men's real wages over our sample period are presented in Figure 4 and Table 6.[18] (For convenience, Figure 4's wage quintiles match those in Figure 3; note that both figures normalize all groups to a constant base in 1979.) They show, as is well known, that when wages are deflated using the Consumer Price Index (CPI) men experienced essentially no aggregate real wage growth during the two and a half decades under study. Salaried workers did experience a real wage gain of about 4.5 percent between 1979 and 2006, most of which occurred in the 1990s. Hourly workers experienced an overall loss of 14.5 log points, consisting of a large decline in the 1980s followed by a rough constancy after that. Looking across quintiles of the wage distribution, all quintiles but the top one experienced a real wage decline in the 1980s, while all quintiles experienced gains in the 1990s. That said, as is well known, real wage increases in the 1990s were concentrated in the top of the wage

[18] As in all studies using CPS data, these hourly wages are directly reported by hourly paid workers, but calculated as the ratio of weekly earnings to hours for salaried workers.

distribution. Between 2000 and 2006 there were small wage gains in the top quintile and small declines in the bottom.

In well-known papers, Juhn, Murphy and Topel (JMT) (1991) and Juhn (1992) argue that a very simple model based on a positive uncompensated labor supply response to real wages can explain recent trends in men's employment rates, both overall and across skill groups. Examination of Table 6 reveals that such a model is not a good candidate for explaining the recent increase in long work hours. While both the increase in long hours and any increases in real wages were concentrated among skilled and salaried workers, a simple story of this nature faces two problems. First, the bulk of the increase in long hours occurred in the 1980s, while almost all the real wage growth during this period was confined to the 1990s. Even more problematic, for some groups long hours rose in the 1980s despite a decline in average hourly earnings. For example, comparing Tables 1 and 6, between 1979 and 1989 the middle quintile of wage earners increased their incidence of long work hours by 3.4 percentage points (or, viewed another way, (.176-.132)/.132 = 33 percent) while experiencing a real wage *decline* of 7.7 percent.[19] Thus, a JMT-type model faces difficulties in explaining why long hours rose at the middle of men's wage distribution in the 1980s, and why the increase in long hours was more muted for all men in the 1990s than the 1980s.

For similar reasons, several variants of the JMT model that also try to explain hours changes by real wage changes alone face clear problems as well. One such variant might invoke CPI bias: If this bias was large during the 1980s, it might be sufficient to

[19] One consideration that might alleviate the latter problem is CPI bias. it would now be difficult to explain the much smaller gains in long work hours during that decade, when real wage growth was much stronger. Second, it would be difficult to reconcile a large real wage increase for the median man over the last two decades with the aggregate declines in men's employment rates documented earlier in this paper.

convert the estimated real wage declines of the middle three quintiles of men into real wage increases sufficient in magnitude to explain those groups' increased incidence of long work hours in that decade. However, unless CPI bias decelerated dramatically during the 1990's and 2000's, CPI bias makes it even harder to explain the deceleration and reversal of long hours in those periods. Another variant would be to assume, contrary to JMT, that men's uncompensated labor supply response is negative (as is often argued in discussions of the decline in men's workweek between 1900 and 1950). In this case, it is of course difficult to explain why the largest increases in long hours have occurred among the most skilled workers.[20] Given these potential difficulties, before evaluating models like JMT's –which treat average hourly earnings as parametric to the worker-- more formally, we first describe two alternative proxies for men's financial incentives to supply long hours.

2. The Long-Hours Premium

Given our interest in the prevalence of long hours among full-time men, it may be more appropriate to focus specifically on trends in the *marginal* incentives to supply hours beyond the standard full time amount. Of course, if workers face a linear budget constraint then the traditional hourly "wage" measure examined above will measure these marginal incentives adequately. But, especially for salaried workers --who account for the bulk of the rise in hours, and who are not explicitly rewarded in their current jobs for hours worked beyond the contracted minimum-- such a measure could be quite a poor

[20] Considering the likely effects of changes in nonlabor income available to men (primarily their wives' earnings) only adds to these difficulties because skilled men's wives had the largest earnings increases (see Juhn and Murphy 1997). Thus if leisure is a normal good, skilled men should have experienced the smallest, not the largest rise in weekly work hours.

proxy for the actual, long-term financial gains associated with supplying additional hours. For salaried workers, the primary financial rewards to putting in "extra" hours consists of factors like earning a bonus or raise within one's position, winning a promotion to a better one, signaling to the labor market that one is productive or ambitious and thus securing a better job in another firm, acquiring extra skills (or networks and contacts) that may be rewarded in either the current firm or another one, and (especially relevant to older workers, for whom displacement is much more costly) perhaps an enhanced prospect of keeping one's current job if the firm is forced to lay off workers in the future.

Measuring the above rewards for extra hours is of course a challenge, especially in cross-section data sets like the CPS. In the current paper we examine two crude proxies for the magnitude of these rewards, the first of which we call the *long hours premium*. This is simply the cross-sectional relationship, within a labor market subgroup such as a three-digit occupation, between usual weekly hours and total weekly earnings over the interval from 40 to 65 hours. To the extent that tastes for work are a relatively fixed personal characteristic, these regressions will identify the slope of the implicit long-term compensation schedule relating earnings to hours, at least within a group of persons who might plausibly face a common long-run pay schedule.[21] Of course, if individuals have earnings intercepts that are correlated (within groups) with their tastes for work, these cross-sectional comparisons will also contain an element of sorting; in that case an increase in the within-group slope over time will also indicate an increased sorting of abler workers into higher levels of hours. We readily admit that we cannot distinguish

[21] Somewhat more formally, if a given type of workers (for example a detailed occupation) contains individuals with heterogeneous tastes for leisure but faces a common budget constraint relating hours worked to total earnings, then cross-sectional within-group comparisons of total weekly earnings at different hours will identify the slope of the common budget constraint.

such a sorting story from an increase in the slope of the compensation function with CPS data. However, we find it hard to imagine what –aside from an increase in the payoff to long hours among skilled workers itself—might cause this sort of change in worker sorting by ability.

Our estimates of the cross-sectional long hours premium are presented in Table 7 for three periods, 1983-85, 1991-93 and 2000-02, for which comparable 3-digit occupation codes are available. Separately for each of these three periods, panel (a) regresses total log real weekly earnings on usual weekly hours, education indicators, a quartic in age, and a full set of three-digit occupation fixed effects. Because of our sample restrictions, the slope coefficient reported there represents an average slope over the interval from 40 to 65 hours, which could (in principle) be very different from other regions of the budget set. Panel (b) poses the same question as panel (a), substituting a quartic in hours for the linear hours term. Shown are predicted log earnings for an average sample member at 40 versus 55 hours, and the difference between the two.

According to Panel (a) of Table 7, the apparent marginal reward to putting in ‘extra’ work hours within an occupation increased substantially between 1983/85 and 1991/93, and again between 1991/93 and 2000/02. Overall, an extra hour beyond 40 was associated with a 1.2 percent increase in earnings in 1983/85, and with about a two-percent increase by 2000/02. For obvious reasons, hourly workers’ total earnings are more strongly associated with current hours than salaried workers’, but the strength of this association grew over time among both hourly and salaried men. For hourly paid workers, the increase was confined to the 1980s; for salaried workers it continued through the 1990s as well, as did the increase in long hours (though much more modestly).[22]

Panel (b) shows essentially the same results from a more flexible specification. Again, the estimated slope of the hours-pay schedule rises over time, and does so much more dramatically (in percentage terms) for salaried workers. In the early 1980's, a randomly selected salaried man regularly putting in 55 hours per week earned a total weekly salary 10.5 percent higher than an observationally-equivalent man working usually working 40 hours in the same 3-digit occupation.[23] By the early 21st century, that gap had more than doubled, to 24.8 percent. The functional form of the predicted relationship in 1983/85 versus 2000/02 is shown in Figures 5a and b for hourly and salaried workers respectively. According to these figures, hourly-paid workers' weekly earnings have always been positively associated with hours worked. However a substantial positive association of this form emerged among *salaried* workers only after the early 1980's.

Panel (b) of Table 7 illustrates one remaining point of note: despite the large increases in our measures of marginal work incentives, real earnings at 40 hours remained essentially constant over our entire sample period (salaried workers experienced a slight increase that was confined to the 1990s). This contrast highlights the need to distinguish average and marginal rewards to work, and may help reconcile the

[22] To check whether time trends in the estimated long-hours premium might be driven by changes over time in top coding of nominal weekly earnings, we re-estimated the models reported in Table 7 as Tobits, with right-hand censoring at the top-coding point. The results were virtually identical.

[23] Put a different way, a (15/40=) 37.5 percent increase in hours corresponded to only a 10.5 percent gain in total earnings, implying an overtime *penalty* of (37.5 – 10.5)/37.5 = 72 percent of the straight-time wage for hours worked beyond 40. In this sense, marginal incentives to provide hours beyond 40 appear to have been very low indeed.

observed increase in long hours among full-time men with the fact that men's labor force participation did not increase during our sample period.

3. Within-Group Earnings Dispersion

In some recent papers, Bell and Freeman (2001a,b) have proposed an alternative proxy, in cross-sectional data, for the magnitude of the change in the lifetime present value of earnings that is likely to be associated with an extra current hour of work. Arguing that, within occupations, (or even in the labor market as a whole) compensation can be interpreted as a tournament scheme in which the cross-sectional earnings distribution measures the set of prizes available to workers, Bell and Freeman argue that an increase in the spread of this prize distribution should elicit more work hours (see for example Lazear and Rosen 1981).[24]

In addition to capturing a literal reading of Bell and Freeman's tournament-based hypothesis, we note that an increase in unexplained within-occupation earnings dispersion should also capture a wide variety of other departures from standard rates of pay, towards more individualized (and thus potentially performance-related) pay. These include individual and team bonuses, stock options, changes in job security, and the size of pay differentials *between* firms. Together, all of these pay components should provide firms (and the market) more ways to reward "extra" hours with higher pay, compared to a world where all salaried workers doing a similar job are paid the same. We thus hypothesize that increases in this measure of dispersion will be associated with increases in salaried workers' propensity to supply 'extra' hours.

[24] Bell and Freeman's main interest is in explaining cross-country differences in hours worked. To our knowledge, our paper is the first to apply their hypothesis to changes over time in work hours in any country.

Table 8 presents data on two measures of the within-occupation dispersion of earnings, both calculated from the regressions underlying part (b) of Table 7. It is noteworthy that, in contrast to the measures of the long-hours premium in Table 7, the statistics in Table 8 refer to earnings dispersion net of very detailed controls for current work hours (specifically, a quartic in the interval from 40 to 65 hours). Thus the Table 8 indicators are net of any variation in compensation that might be interpreted as within-survey-year compensation for high survey year work hours. As in Table 7, the indicators in Table 8 show a substantial increase in this proxy for marginal work incentives, and this increase is substantially greater among salaried than hourly paid workers. The increase continues throughout the 1980s and 1990s; the increase in salaried workers' 90/10 differential is strongly concentrated in the 1980s.[25] Thus, like the rise in the long-hours premium, the increase in within-occupation earnings dispersion also seems a promising *prima facie* candidate to explain the increase in men's long work hours.[26]

4. Summary: Aggregate Evidence on Changes in Work Incentives

In sum, aggregate trends in long work hours over the past two decades are not easy to reconcile with a model that treats average hourly earnings as parametric to an individual worker. Regardless of whether we assume a positive uncompensated labor supply elasticity (following JMT) *or* a negative one, it is hard to tell a consistent aggregate story based only on real wage changes. More promising in this regard are our

[25] As in Table 7, we were concerned that these increases in residual earnings variation could be an artifact of changing top coding of earnings across CPS years. However, replacing the OLS regressions underlying Table 8 by a Tobit yielded very similar results.

[26] We are of course hardly the first researchers to document an increase in residual wage inequality over this period (see for example Juhn, Murphy and Brooks 1993). That said, we are unaware of any previous evidence that such an increase took place within very detailed (three-digit) occupation groups with very flexible controls for work hours, nor of any such work that considers salaried versus hourly workers separately or updates this analysis into the 2000s.

two proxies for full-time workers' marginal incentives to supply weekly hours beyond 40: both these measures did increase throughout our sample period, and like the increase in long work hours were more pronounced among salaried workers. That said, the aggregate trends in these indicators still raise some concerns about timing. For example, why did long hours increase much more slowly in the 1990s despite the continuing rise in measured incentives? In part for this reason, we turn in the next section to a disaggregated analysis of changes in the incidence in long work weeks at the occupation/industry/decade level.

V. Changes in Long Work Hours among Salaried Men: Disaggregated Analysis

In this section we restrict our attention to men who are paid on a salaried basis – among whom the recent increase in long weekly hours was concentrated—and conduct a disaggregated analysis of long-term (roughly decadal) trends in long hours among this group. More precisely, we study the association between changes in measured work incentives and changes in the incidence of long hours, across subgroups of full-time, salaried men. Since, from a theoretical point of view, we prefer to work with groups who share a compensation schedule, the ideal disaggregation is probably the most detailed occupation group available (in our case, 3-digit codes). That said, to assess the robustness of our results we present estimates for 2-digit occupations, and for 2- and 3-digit industries in this section as well.

Our first set of results are presented in Table 9. Each entry in this Table is a coefficient from a univariate regression where the dependent variable is the within-cell change of the incidence of long work hours in one of two periods: 1983/85 - 1991/93, or

1991/93 - 2000/02. [27] Observations for all of these regressions are thus industry- or occupation/period cells. Each row of the table uses a different characteristic of the cell as the independent variable. All the regressions contain a fixed effect for 1991/93 - 2000/02. Finally, since the precision with which these cell characteristics are measured depends on the number of observations from which they are calculated, estimation is by weighted least squares with cell counts as weights. As in Table 5, cells were included in the analysis only if they contained at least 50 workers.

To see how Table 9 works, consider row 1. It shows that 2-digit industries with higher average hourly earnings at the beginning of a period (i.e. in either 1983/85 or 1991/93) experienced greater growth in long work hours during that period (i.e. between 1983/85 and 1991/93 or between 1991/93 and 2000/02 respectively). The same is true for the other disaggregations: 3-digit industries, and 2- and 3-digit occupations. According to row 2, the same is true when we substitute the beginning-of-period level of total weekly earnings for the level of average hourly earnings. Together, rows 1 and 2 of Table 9 confirm our earlier findings that the recent increase in long work hours was concentrated among highly paid men.

Row 3 of Table 9 asks whether the industries or occupations that experienced the largest increases in real hourly wage rates (as conventionally measured, i.e. earnings divided by wages) experienced the largest increases in long work hours. This is clearly not the case; in fact the three-digit occupation analysis shows a strong negative effect.[28]

[27] We also ran separate regressions for each of the two 'decadal' periods, with similar results (but lower significance levels) in both cases.

[28] A possible concern is that this strong negative relationship could be driven by "division bias", e.g. Borjas 1980) related to measurement error in hours. This seems unlikely to us given that the units of analysis are cell means, with minimum cell sizes of 50. We also note –see immediately below-- that even the estimates in Row 4, which should be biased in the opposite direction by measurement error in hours, remain negative in the 3-digit industry specification.

Row 4 of the Table repeats this analysis using total log weekly earnings changes instead of hourly wage changes. Importantly, this exercise will be biased towards finding a positive correlation if any of our observed increases in weekly earnings incorporate some immediate compensation for extra hours worked. But even in the presence of this positive bias, no significant positive correlation between total earnings changes and hours changes is found.

The next three rows of Table 9 test the ability of various estimates of the long-hours premium to explain decadal changes in long work hours. Row 5 uses a very crude measure of the long-hours premium: the within-cell difference in total weekly earnings between men working 40-49 hours and those working 50-65 hours. Rows 7 and 8 use regression-adjusted measures of the form examined in Table 7. To be clear, for row 7 of Table 9, we first ran a separate regression of the form presented in Table 7a for *every* industry or occupation in each of our three periods (1983/85, 1991/93 and 2000/02). Between-period changes in the estimated cross-sectional salary slope coefficients from these regressions are then used as regressors in row 7 of Table 9. Row 8 of Table 9 repeats this exercise using the quartic specification in Table 7(b) to estimate the within-cell long-hours premium. Overall, these results are disappointing, showing either no association or a negative one between long-hours changes and changes in the long-hours premium. Put another way, contrary to our expectations, those industries or occupations (and decades) which experienced the largest increases in the long-hours premium did not experience the largest increase in the incidence of long work hours.

The last four rows of Table 9 examine the hypothesis that increases in within-group residual earnings inequality were associated with increases in the incidence of long work hours, using four alternative indicators of earnings inequality. Rows 9 and 10 use very simple measures of earnings dispersion: the within-cell standard deviation of log weekly earnings, or of log hourly wages. The remaining two indicators of within-cell earnings dispersion are the cell-specific equivalents of the earnings residuals reported in Table 8. Specifically, we first ran a separate regression of the form presented in Table 7(b) for every industry or occupation in each of our three periods (1983/85, 1991/93 and 2000/02). Between-period changes in the standard deviation of the residuals in these regressions are then used as regressors in row 10 of Table 9. Changes in the difference between the 90th and 10th percentile of the residuals are used in row 11.

These results are considerably more robust, with positive coefficients in all 20 cases, the majority of them statistically significant. Thus, no matter whether we disaggregate the population of salaried men by industry or occupation, or at the 2- or 3-digit level, we find that those subgroups of men that experienced larger increases in within-group earnings inequality (either overall or net of controls, including a very flexible control for hours worked) in a particular decade were more likely to experience increases in long work hours in that decade. The results also appear to be strongest for three-digit occupations, which is theoretically the most appropriate specification.

In sum, Table 9 provides some evidence that increases in the incidence of long work hours over this period were more common in occupations and industries where (from row 3) increases in average hourly earnings were lowest, and (b) where within-group pay inequality increased the most. The former association is suggestive of a negative income effect of higher pay on labor supply, the latter with the notion that individualized pay setting provides greater marginal incentives for salaried workers to supply “extra” hours beyond the ‘normal’ 40-hour week.

To explore the two above relationships further, Table 10 runs three bivariate regressions on the same data as Table 9. Intuitively, we propose that the propensity of members of a detailed occupation group to work long hours depends on (a) the level of earnings they can expect to receive if they work ‘normal’ (40) hours, and (b) the marginal incentives to provide hours beyond 40. We expect the former relationship to be negative if leisure is a normal good, and the latter to be positive. In the results presented in Table 10, the measures of (a) and (b) are both derived from the same within-cell regressions, specifically those underlying rows 7, 10 and 11 of Table 9. In each of the three regression specifications shown in Table 10, one of the two independent variables is the change in the predicted level of log earnings at exactly 40 hours of work. The other variable in each regression is one of our three proxies for changes in the marginal incentive to provide hours beyond the ‘normal’ 40 per week: the long-hours premium, the standard deviation of the within-cell salary residual, or the 90/10 gap in salary residuals.

The vast majority of specifications in Table 10 (the main exception is two-digit occupations) show a negative partial relationship between the change in real log weekly earnings at 40 hours and the change in the share of workers supplying ‘extra’ hours beyond 40. As in Table 9, the long-hours premium is unsuccessful in predicting whichsubgroups of men will become more likely to work long hours.[29] Coefficients on both indicators of residual wage inequality remain positive and for the most part

[29] One possible reason for the better performance of our earnings-dispersion measure of marginal work incentives may involve lags: an increase in wage inequality both immediately raises workers’ incentives *and* is immediately reflected in our data. In contrast, if (as we argue) the cross-sectional long-hours premium largely reflects the impact of past hours on current pay, this could take considerable time to be detected in our data.

significant, however. Thus, a simple characterization of the increases in long hours among salaried men between 1983 and 2002 is that these were more common when a group of men experienced an increase in residual earnings inequality, and when they experienced a decline in the level of real earnings at 'standard' (40) hours.

Together, the evidence in Table 10 suggests a possible explanation of the deceleration in the increase in long work hours in the 1990s (and perhaps its reversal after 2000, though we do not address this formally): Throughout the entire period from 1983 to 2002, an ongoing rise in within-occupation earnings inequality created incentives for salaried men to put in long work weeks. In the second half of this period, however, this effect was counteracted by income effects related to the rise in the level of real earnings at 'normal' hours for a typical salaried man.[30] While this interpretation of our data is hardly conclusive, we hope it will provide a useful basis for further study of hours trends, especially studies that are based on more explicit measures of work incentives than those available to us here.

Throughout the past two sections we have argued that the salaried workers' marginal incentives to supply 'extra' hours rose throughout our sample period; the only evidence offered for this so far has consisted of simple proxies calculated from the CPS. In the remainder of this section we examine other sources of information for evidence of changes in pay policy away from a standardized salary and towards more individualized (and potentially "incentivized") systems among skilled or salaried men. One way in

[30] To assess this notion more formally, we calculated the share of the actual changes in salaried men's long hours between 1983/85 and 2000/02 that be explained by the relationships estimated in Table 10. Focusing on three-digit occupations, the share of salaried men working long hours increased by .065 between 1983/85 and 1991/93, and fell by .001 between 1991/93 and 2000/02. Using regression 2 of Table 10, we predict changes in long hours in these two periods of .001 and .-.001 respectively; using regression 3 these numbers are .004 and -.001. Thus, while regression 2 is not very successful in predicting the 1980s rise in long hours, regression 3 is reasonably successful in doing so. Further, both specifications match the change in long hours after 1991 very well, and both predict the rapid deceleration after 1991.

which salaried jobs might have become more incentivized would involve a greater willingness by firms to fire or lay off underperforming workers. Interestingly, in an exhaustive study of time trends in the incidence and consequences of job loss, Farber (1997) finds little overall change, but does find substantial increases in displacement rates among skilled workers over the past two decades. Likewise, while Schmidt (2000) finds no overall trend in perceptions of job insecurity over the past two decades, she does find that perceived job insecurity rose substantially among highly educated workers in GSS data over the same period.

Are there any direct measures of firms' compensation policies that might substantiate the claim of increased incentivization in salaried jobs? Perhaps surprisingly, very few surveys of firms' pay policies that yield consistent measures over time exist. One such source is the BLS's Employee Compensation Survey; since 1983 this survey has collected information on the prevalence of "nonproduction bonuses" as part of its database on employee benefits. By definition, nonproduction bonuses cannot be tied directly to employee productivity; that said, they can be used to reward things like attendance, safety, suggestions for productivity improvement, and may be more relevant to salaried workers than production bonuses. In addition, these bonuses may proxy for the existence of other forms of "variable pay". According to the Employee Compensation Survey, the incidence of nonproduction bonuses expanded during the past two decades, from 17 percent of employees at large and medium-sized enterprises in 1983 to 42 percent in 1997.

A second source of information on changes in firms' pay practices is a periodic survey of pay practices in Fortune 1000 firms (Lawler, Mohrman and Benson 2001).[31] Summary tabulations from this survey are presented in Table 11; unfortunately the earliest information available is from 1987. As is apparent, *all* forms of incentive pay on which the survey collects information became more prevalent between 1987 and 1999. Particularly striking is the increased use of individual incentives, work group or team incentives, and "gainsharing" –a form of plant-level incentives-- with the latter two more than doubling in popularity over this period. In contrast, and supporting the notion that increased fear of layoffs might contribute to work incentives, the survey shows that corporate policies designed to enhance employees' job security became much less common.

In sum, consistent with our CPS-based measures, available direct evidence of compensation practices does show increases in the 'incentivization' of pay during both the 1980s and 1990s, with some indication that these increases were more important for those groups of men (in particular the more skilled) where long hours increased the most over this period.

VI. Discussion

In this paper we have shown that employed American men are more likely to work 50 or more hours per week today than a quarter century ago. We think that this

[31] Some striking, but less-well-documented evidence is reported by Cappelli (1999, pp. 150-151). Based on unpublished data, Cappelli reports that the share of fixed compensation (salary and benefits) in managers' compensation fell from over 60 to under 40 percent between 1984 and 1995. Also according to Cappelli, data collected by Hay Associates shows that, in 1989, the average salary increase associated with the highest level of performance among its clients' employees was 2.5 times larger than for the lowest performance level; by 1993 that factor had risen to 4.

trend is likely to be of interest to labor economists because (a) it reversed (perhaps temporarily) a secular decline in the work week that dates back over a century; and (b) it occurred at a time when other dimensions of men's labor supply, such as employment rates, were in decline.

What might explain this change in work behavior? In this paper we are able to rule out a number of possible causes, some more fundamental than others. For example, we know that the recent increase in the prevalence of long work weeks is not an artifact of changing CPS survey techniques, not a purely cyclical phenomenon, and not easily attributable to the changing mix of occupations and industries in the male labor force. Because the change is strongly concentrated among skilled, salaried men, we know that it is not a direct consequence of the declining economic fortunes of unskilled American men over the past two decades. For a number of reasons including the fact that we measure usual hours, we know that this phenomenon is not an artifact of increased month-to-month variability in hours worked. We know that it is not a consequence of increased self employment (nor of higher hours among the self employed). Nor is it related to an increase in multiple jobholding, or hours worked on second- and higher-order jobs. Because the bulk of the increase occurred during the 1980s, it is not likely related to advances in communication technology (such as the internet) that facilitate additional work from home. Further, because it was concentrated in the subgroups of men who experienced the smallest declines in labor supply on other margins, the increase in long hours cannot be easily interpreted as a simple reallocation of a fixed amount of labor within a representative worker's life cycle.[32]

[32] A final alternative explanation that also seems unlikely to us is the notion that increased fixed costs of employment have led to an increase in the optimal level of hours per worker (see for example Cutler and Madrian, 1998). If

Our analysis also includes an examination of long-hours changes at the decadal level for salaried men (among whom the bulk of the recent increase in hours occurred), across detailed occupation and industry groups. We find that two group characteristics -- a rising level of within-group earnings inequality (at fixed hours) and a falling (or more slowly growing) level of mean earnings at 'standard' (40) hours-- are associated with increases in the share of workers usually supplying 50 or more hours per week. This suggests that an increased dependence of pay on unobservable factors (including job performance) over this period may have had some effect on salaried men's choices to work long hours. Negative income effects from rising compensation at 40 hours might help explain why the increase in long hours slowed down after 1993, and reversed after 1999.

We conclude by noting that this paper does not attempt to explain *why* the idiosyncratic component of salaried men's compensation has risen. Possible causes of this change include the possibility that employment contracts with stronger incentives are a socially efficient response to an increase in the demand for skilled workers, or to (within occupation and industry) changes in market structure that raise the relative productivity of long (versus normal) work weeks. Increased incentives to produce the industry's best product in "winner-take-all"-type markets for information goods come readily to mind. Clearly, however, these are questions for another paper.

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anything, changes in fixed costs should have a higher impact on the hours of low-wage workers than high-wage workers; yet we find that hours rose much more among the highly skilled.

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Table 1: Fraction of Men Usually Working Long (>=50) Hours

	1979	**1989**	**2000**	**2006**
All Men	0.161	0.193	0.190	0.178
Full Time Men (>=30 hours)	0.164	0.199	0.207	0.195
Among Full Time Men:				
Salaried	0.244	0.312	0.320	0.301
Hourly	0.086	0.094	0.105	0.096
Age 25-34	0.171	0.197	0.196	0.167
Age 35-44	0.185	0.221	0.222	0.208
Age 45-54	0.154	0.193	0.216	0.213
Age 55-64	0.128	0.154	0.178	0.191
Less than High School	0.124	0.121	0.116	0.099
High School Graduates	0.137	0.155	0.149	0.153
Some College	0.166	0.19	0.194	0.182
College Graduates	0.240	0.303	0.312	0.278
Average Hourly earnings quintile:				
1 (highest wage)	0.151	0.243	0.297	0.268
2	0.137	0.193	0.214	0.219
3	0.132	0.176	0.199	0.189
4	0.176	0.202	0.184	0.172
5 (lowest wage)	0.217	0.186	0.151	0.133

Notes: Sample is Employed, non-self-employed, Ages 25-64.

Table 2: Men's Labor Supply Indicators, by Education

	1979	1989	2000	2006
AGES 25-64:				
Share of Men Employed[1]:				
Less than High School	0.763	0.709	0.724	0.726
High School Graduates	0.892	0.859	0.831	0.807
Some College	0.904	0.892	0.871	0.843
College Graduates	0.940	0.928	0.914	0.890
AGES 45-54 ONLY:				
Share of Men Employed[1]:				
Less than High School	0.814	0.760	0.700	0.710
High School Graduates	0.910	0.886	0.837	0.825
Some College	0.920	0.911	0.874	0.863
College Graduates	0.961	0.943	0.935	0.929
Share of Employed working Long Hours[2]:				
Less than High School	0.111	0.126	0.111	0.110
High School Graduates	0.133	0.139	0.138	0.166
Some College	0.159	0.200	0.194	0.195
College Graduates	0.248	0.301	0.324	0.302

1. Sample: All Men

2. Sample: Men working full time (30 or more hours), not self-employed.

Table 3: Linear Probability Model Coefficients for Working Long Hours

	Regression Coefficients		Sample Means	
	1983-1985	2000-2002	1983-1985	2000-2002
High school grad	.006* *(.002)*	.007* *(.003)*	.344	.310
Some college	.014** *(.003)*	.025** *(.003)*	.215	.265
College graduate	.040** *(.003)*	.065** *(.004)*	.268	.314
Age 35-44	.001 *(.002)*	.008** *(.002)*	.285	.332
Age 45-54	-.020** *(.002)*	-.007** *(.002)*	.195	.261
Age 55+	-.039** *(.002)*	-.030** *(.003)*	.131	.115
Salaried	.113** *(.002)*	.130** *(.002)*	.501	.483
Married	.018** *(.002)*	.027** *(.002)*	.761	.672
Union	-.029** *(.002)*	.016** *(.002)*	.275	.172
Black	-.056** *(.003)*	-.054** *(.003)*	.095	.107
Hispanic	-.046** *(.003)*	-.052** *(.003)*	.058	.114
Observations	213,062	210,640		
R-squared	.134	.127		

Notes: Robust t-statistics errors in parentheses.

* and ** indicate significance at 5% and 1% respectively.

Both regressions include 47 industry controls, 45 occupation controls and 49 state dummies.

Sample: Non-self-employed, salaried and hourly paid men, working 30 usual hours or more

Dependent Variable: Indicator for whether usual weekly hours are 50 or more. The mean of the dependent variable is 0.163 in 1983-1985, and 0.203 in 2000-2002.

Table 4: Predicted Share of Men Working Long Hours, 1983/85 to 2000/02, from Linear Probability Model

	83/85 Coefficients, 83-85 Means	00/02 Coefficients, 00-02 Means	83/85 Coefficients, 00-02 Means	00/02 Coefficients, 83-85 Means
	(1)	(2)	(3)	(4)
All Full-Time Men	.166	.196	.173	.194
Salaried Full-Time Men	.253	.298	.271	.282

Table 5: Predicted Share of Men Working Long Hours, 1983/85 to 2000/02, from Shift-Share Decompositions

	83-85 Cell Means,	00-02 Cell Means,	83-85 Cell Means	00-02 Cell Means
	83-85 Mix	00-02 Mix	00-02 Mix	83-85 Mix
	(1)	(2)	(3)	(4)
ALL FULL-TIME MEN:				
a) by Occupation	.163	.204	.170	.198
b) by Industry	.163	.203	.170	.200
SALARIED FULL-TIME MEN:				
a) by Occupation	.252	.320	.258	.315
b) by Industry	.257	.325	.262	.318

Note: Cells are three-digit industries or occupations. Differences from Table 3, and slight variations in the column 1 and 2 means are explained by changes in sample composition generated by dropping industry or occupation cells with fewer than 50 observations. Number of cells are: 315 occupations and 201 industries for All Men; 199 occupations and 179 industries for Salaried Men.

Table 6: Average Log Real Hourly Wages ($1993), 1979-2006

	1979	**1989**	**2000**	**2006**
A. Total	2.297	2.221	2.242	2.240
B. By Pay Method				
Salaried	2.404	2.391	2.440	2.449
Hourly	2.192	2.063	2.062	2.047
C. By hourly wage quintile				
1 (highest)	2.910	2.939	3.018	3.037
2	2.537	2.504	2.535	2.540
3	2.321	2.244	2.245	2.229
4	2.090	1.961	1.951	1.924
5 (lowest)	1.687	1.485	1.503	1.479

Sample: employed, not self-employed men aged 25-64.

Table 7: Estimates of the Total Earnings Premium for Working Long Hours

a. Linear hours effect model: coefficient on hours[1]

Sample:	1983/85	1991/93	2000/02	Change, 1983/85 – 1991/93	Change, 1983/85 – 1991/93
All workers	.0118 (.0002)	.0164 (.0002)	.0199 (.0002)	.046	.035
Hourly	.0199 (.0003)	.0257 (.0003)	.0253 (.0003)	.058	-.004
Salaried	.0067 (.0002)	.0100 (.0002)	.0148 (.0003)	.033	.048

b. Polynomial in hours: Predicted log earnings at 40 versus 55 hours

Sample:	1983/85	1991/93	2000/02	Change, 1983/85 – 1991/93	Change, 1983/85 – 1991/93
All workers:					
40	5.953	5.903	5.955		
55	6.136	6.159	6.272		
difference	0.183	0.256	0.317	.073	.061
Hourly:					
40	5.797	5.720	5.768		
55	6.097	6.100	6.139		
difference	0.300	0.379	0.371	.079	-.008
Salaried:					
40	6.122	6.123	6.176		
55	6.227	6.284	6.424		
difference	0.105	0.162	0.248	.057	.086

[1]Coefficients represent the effect of working one more hour per week on the log of total weekly earnings. Standard errors in parentheses.

Dependent variable in all regressions is total weekly earnings.

Sample in all regressions is men ages 25-64, currently employed and not self-employed, working between 40 and 65 hours per week. Age is measured as a quartic, education groups as shown in Table 1. Hours are entered linearly in part (a), as a quartic in part (b). All regressions contain a full set of 3-digit occupation fixed effects.

Table 8: Estimates of Within-3-digit-Occupation Residual Earnings Dispersion at Fixed Hours

Sample:	Root Mean Square Error, log earnings regression			90-10 differential in log earnings residual		
	1983/85	1991/93	2000/02	1980	1991/93	2000/02
All workers	.376	.397	.449	.940	.988	1.047
Hourly	.351	.358	.398	.897	.910	.961
Salaried	.385	.420	.487	.942	1.032	1.104

Sample: Men ages 25-64, currently employed and not self-employed, working between 40 and 65 hours per week.

Earnings residuals are from the regressions in Table 7b; these regress total weekly earnings on a quartic in age, a quartic in hours, education categories, and 3-digit occupation fixed effects.

Table 9: Univariate Regression Coefficients of Long-Run Changes in the Incidence of Long Work Hours on Industry/Occupation Characteristics, Salaried Full-Time Men

Industry/Occupation Characteristic (Independent Variable):	Unit of Analysis			
	2 Digit Industry	3 Digit Industry	2 Digit Occupation	3 Digit Occupation
A. Initial Skill Levels				
1. Initial log Hourly Wage	0.060***	0.070***	0.065***	0.060***
	(0.015)	(0.009)	(0.010)	(0.008)
2. Initial log Weekly Earnings	0.068***	0.074***	0.062***	0.055***
	(0.016)	(0.010)	(0.010)	(0.009)
B. Wages or Earnings Changes				
3. Change in log Hourly Wage	-0.197***	-0.144***	-0.007	-0.219***
	(0.056)	(0.037)	(0.056)	(0.032)
4. Change in log Weekly	-0.092	0.034	0.083	-0.059*
Earnings	(0.063)	(0.038)	(0.055)	(0.036)
C. Changes in Long Hours Premium				
5. Change in Unadjusted Long-Hours Premium[1]	-0.135**	-0.002	0.005	0.023
	(0.062)	(0.026)	(0.057)	(0.022)
6. Change in Salary Slope	-1.567*	-0.469	0.118	0.205
(Linear Specification)	(0.847)	(0.396)	(0.939)	(0.320)
7. Change in Salary Slope	-0.137***	-0.017	0.003	-0.004
(Quartic Specification)	(0.050)	(0.015)	(0.009)	(0.005)
D. Changes in Earnings Dispersion				
8. Change in Standard	0.143	0.145**	0.128*	0.096**
Deviation of log Earnings	(0.088)	(0.047)	(0.068)	(0.039)
9. Change in the Standard	0.160*	0.151**	0.143**	0.105***
Deviation of log Wage	(0.088)	(0.048)	(0.070)	(0.039)
10. Change in the Standard	0.082	0.087*	0.170**	0.088**
Deviation of Salary Residual	(0.110)	(0.049)	(0.073)	(0.041)
11. Change in the 90-10 Salary	0.071	0.083***	0.098**	0.064***
Residual Gap	(0.062)	(0.030)	(0.040)	(0.023)
N	90	369	86	416

Observations are industry/period or occupation/period cells, where the two periods are 1983/85-1991/93 and 1991/93-2000/02. The dependent variable in all regressions is the within-period change in the share of workers in the industry or occupation working long hours. Changes calculated from cell sizes of fewer than 50 persons are excluded from the sample. All regressions are weighted by the average number of observations from which the change in long hours is calculated. All regressions include a fixed effect for 1991/93-2000/02. Standard errors in parentheses. ***, ** and * indicate significance at 1%, 5% and 10% respectively.

1. The unadjusted long-hours premium is just the ratio of average total weekly earnings among persons working 50-65 hours to those working 40-49 hours.

Table 10: Bivariate Regression Coefficients of Long-Run Changes in the Incidence of Long Work Hours on Industry/Occupation Characteristics, Salaried Full-Time Men

Independent Variables:	**Unit of Analysis**			
Bivariate Regression (1)	2 Digit Industry	3 Digit Industry	2 Digit Occupation	3 Digit Occupation
Change in log Weekly Earnings	-0.162**	-0.071**	0.019	-0.152***
at 40 Hours	(0.063)	(0.035)	(0.060)	(0.033)
Change in Long Hours Premium	-0.143***	-0.022	0.003	-0.004
(Quartic Specification)	(0.048)	(0.015)	(0.009)	(0.005)
Bivariate Regression (2)				
Change in log Weekly Earnings	-0.149***	-0.055	0.028	-0.145***
at 40 Hours	(0.066)	(0.035)	(0.058)	(0.033)
Change in Standard Deviation	0.050	0.078	0.172**	0.065
of Salary Residual	(0.109)	(0.049)	(0.074)	(0.040)
Bivariate Regression (3)				
Change in log Weekly Earnings	-0.162**	-0.060*	0.001	-0.150***
at 40 Hours	(0.066)	(0.035)	(0.058)	(0.032)
16. Change in the 90-10 Salary	0.086	0.082***	0.098**	0.062***
Residual Gap	(0.061)	(0.030)	(0.040)	(0.023)
N	88	369	86	416

Observations are industry/period or occupation/period cells, where the two periods are 1983/85-1991/93 and 1991/93-2000/02. The dependent variable in all regressions is the within-period change in the share of workers in the industry or occupation working long hours. Changes calculated from cell sizes of fewer than 50 persons are excluded from the sample. All regressions are weighted by the average number of observations from which the change in long hours is calculated. All regressions include a fixed effect for 1991/93-2000/02. Standard errors in parentheses.

***, ** and * indicate significance at 1%, 5% and 10% respectively.

Table 11: Percent of Fortune 1000 Firms Surveyed in which over 20 percent of employees are covered by selected reward practices

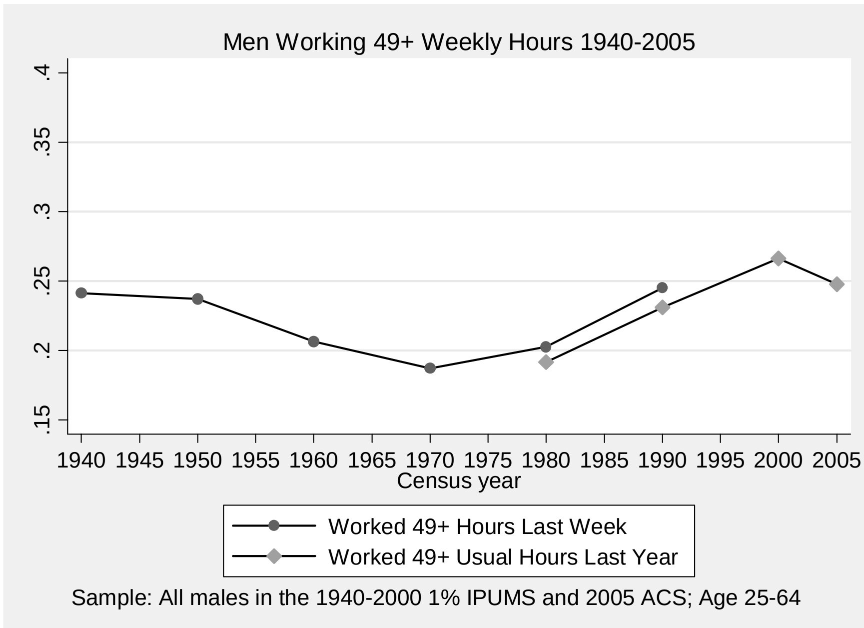
	1987	**1990**	**1993**	**1996**	**1999**
Individual Incentives	38	45	50	59	67
Work Group or Team Incentives	22	31	n.a.	41	48
Gainsharing	7	11	16	20	24
Profit Sharing	45	44	44	52	55
Employee Stock Ownership Plan	52	56	63	59	63
Stock Option Plan	n.a.	n.a	30	41	49
Nonmonetary Recognition Awards for Performance	n.a.	68	73	80	82
Employee Security	34	27	19	17	14

Source: Lawler, Mohrman and Benson (2001), Tables 5.1 and 5.3.

Notes: Gainsharing is a bonus based on improvements in productivity, cost-effectiveness, quality or other perfomance indicator at a large organizational level such as a plant. Employee security is defined as “corporation policy designed to prevent layoffs”. “n/a” denotes data not available.

Figure 1. Census Data: Proportion of Males Working Long Hours

a) All employed men



b) Men: paid workers only (not self employed)

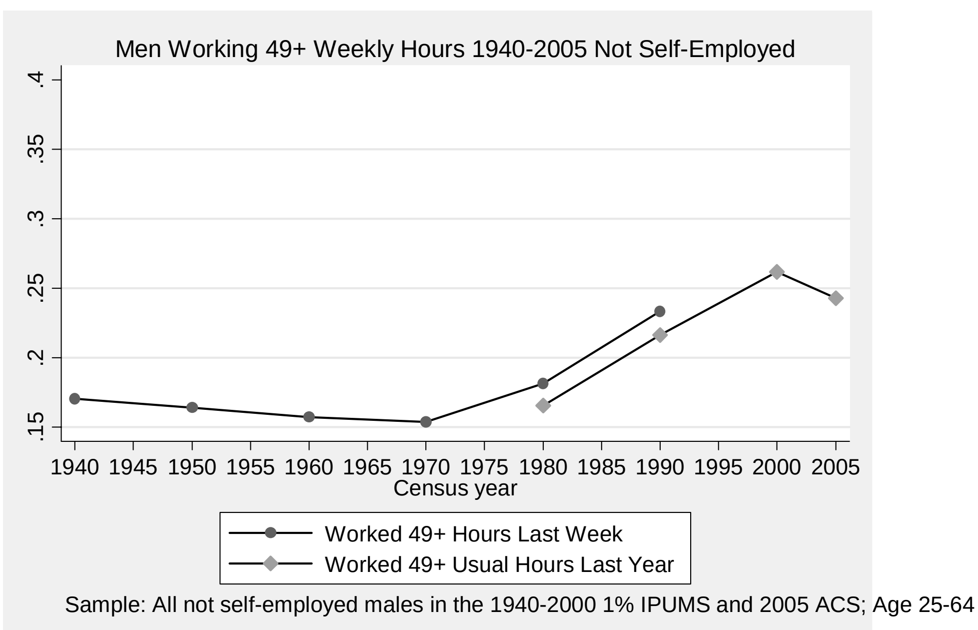
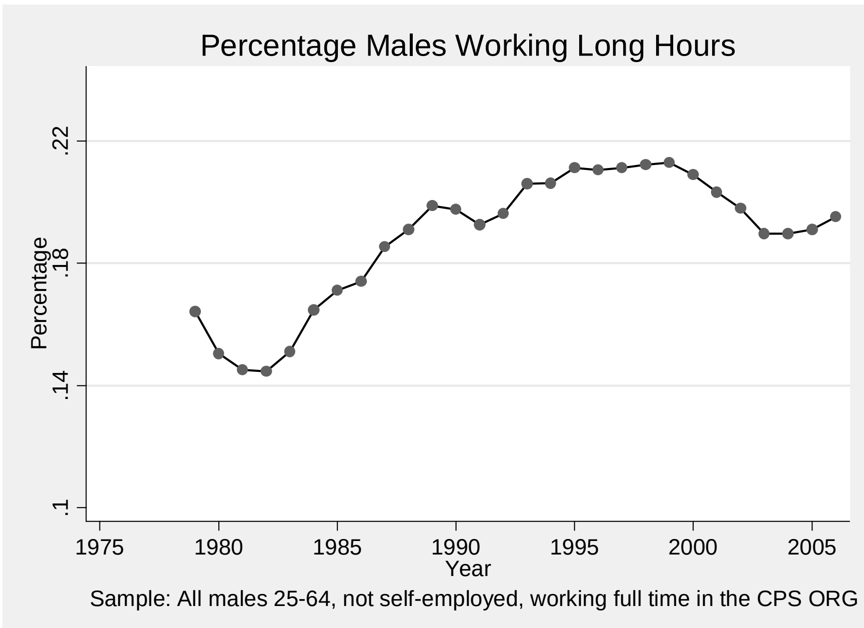


Figure 2: Trends in the Incidence of Long Hours versus Labor Force Participation Rates: CPS Data, Men, 25-64.

a) Long Hours



b) Employment Population Ratio

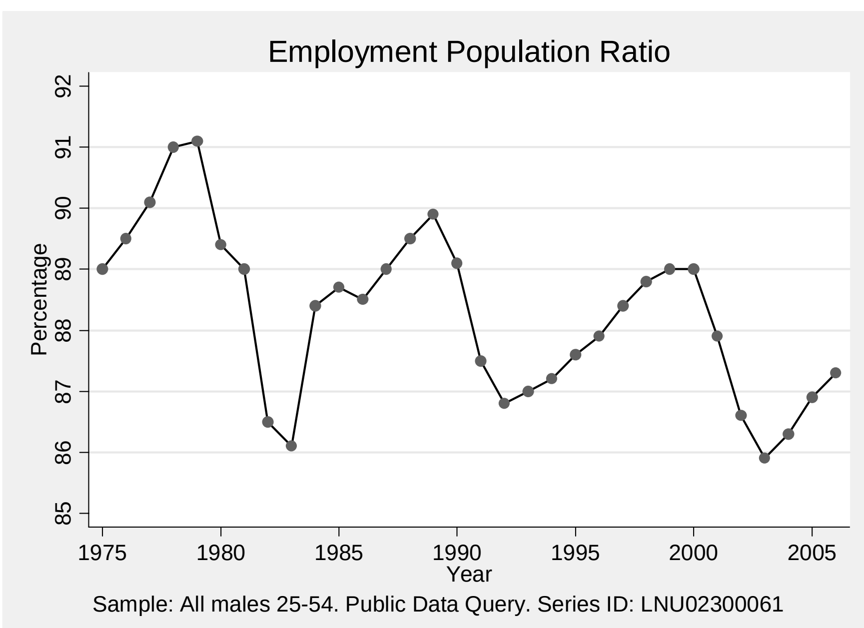


Figure 3: Percentage of Men Working Long Hours, by Hourly Earnings Quintile (1979=1)

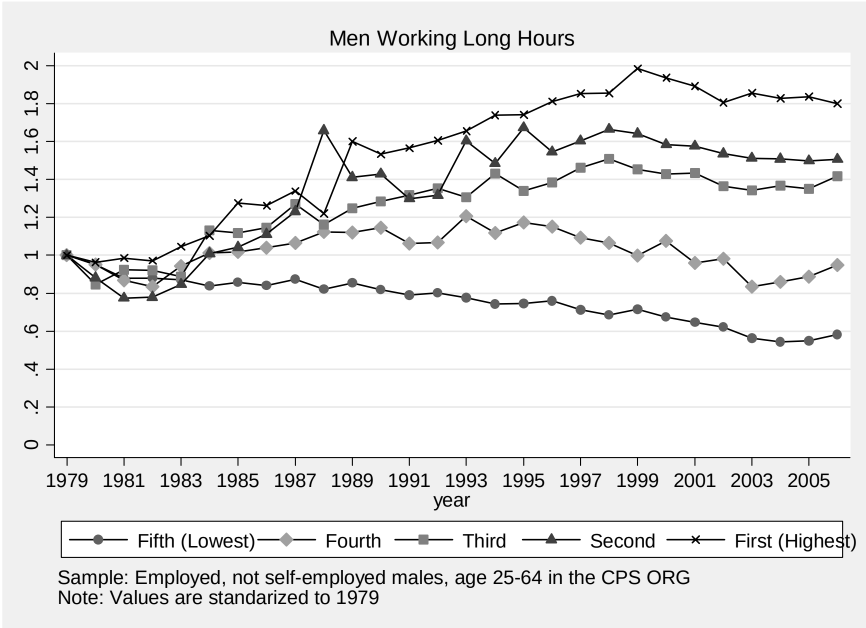


Figure 4 Average Hourly Earnings, by Hourly Earnings Quintile (In 1983 Dollars)

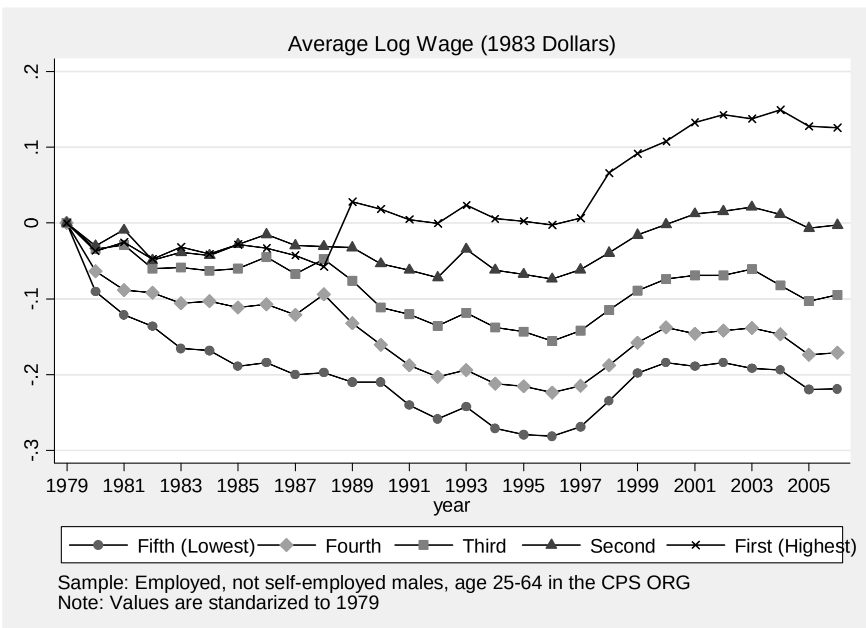
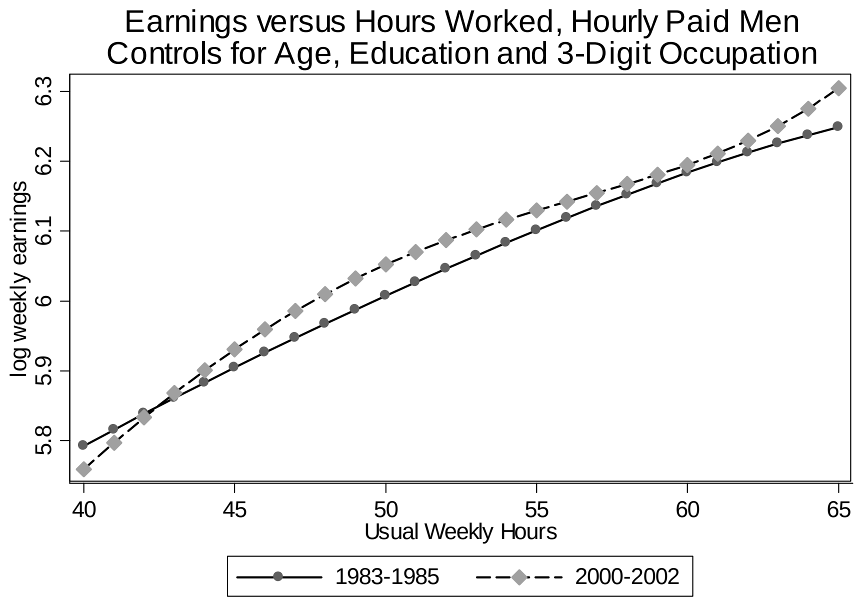


Figure 5: Earnings versus hours worked, Salaried Men and Hourly-paid Men, 1983/85 and 2000/02

a) Hourly Paid Men



b) Salaried Men