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## The expanding workweek? Understanding trends in long work hours among US men, 1979 - 2004

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January 2006

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## **ABSTRACT**

### **The Expanding Workweek? Understanding Trends in Long Work Hours Among U.S. Men, 1979-2004\***

After declining for most of the century, the share of employed American men regularly working more than 50 hours per week began to increase around 1970. This trend has been especially pronounced among highly educated, high-wage, salaried, and older men. Using two decades of CPS data, we rule out a number of factors, including business cycles, changes in observed labor force characteristics, and changes in the level of men's real hourly earnings as primary explanations of this trend. Instead we argue that increases in salaried men's marginal incentives to supply hours beyond 40 accounted for the recent rise. Since these increases were accompanied by a rough constancy in real earnings at 40 hours, they can be interpreted as a compensated wage increase.

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## **I. Introduction**

A significantly larger share of employed U.S. men is now regularly putting in long work weeks than two decades ago. For example, the share of employed, 25-64-year-old men who usually work more than 50 or more hours per week on their main job rose from 14.7 to 18.5 percent between 1980 and 2001. This trend has been especially pronounced among highly skilled men, with the long-hours share rising from 22.2 to 30.5 percent among college-educated men employed full-time (30 hours or more).

The recent rise in the prevalence of long work hours is puzzling for at least two reasons. One is the fact that it reverses, for essentially the first time<sup>1</sup>, a trend towards shorter work weeks that goes back at least a century-- a trend that is widely attributed by labor economists to rising prosperity. Second, the rise in hours occurred during a period when a second key dimension of male labor supply --the labor force participation rate-- moved in the opposite direction.<sup>2</sup> What factors might have caused this century-long trend towards shorter work weeks to change direction, especially among the most highly skilled workers in the U.S. economy? And how can the trend towards longer work weeks be reconciled with declining male labor force participation? The goal of this paper is to answer these questions. Our primary data source is the Current Population Survey Outgoing Rotation Groups, from 1979 to 2004.

We begin our analysis by documenting the increase in long work hours, and by identifying the parts of the male labor force where this increase was the strongest: highly

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<sup>1</sup> The only other reversal occurred during World War Two.

<sup>2</sup> These two puzzling features of the lengthening work week explain our focus on men rather than women. While long work hours increased even more dramatically among women over this period, this continues a long-term trend towards greater labor supply in all dimensions (including both participation and hours), whose causes have been debated by labor economists for decades.

educated, high-wage, salaried men. Next, we rule out a number of simple explanations for the increase, including changes in the demographic composition of the labor force and changes in the mix of occupations or industries. Third, we consider the suitability of simple labor supply models (such as Juhn, Murphy and Topel 1991, or Pencavel 2002) in which long term changes in the level of real average hourly earnings have causal effects on hours worked. Regardless of whether men's uncompensated labor supply elasticity with respect to average hourly earnings is assumed to be positive or negative, we show that such models cannot avoid some key inconsistencies with the the main trends in our data.

Our preferred explanation for the recent increase in men's work hours is based on an increase in salaried workers' marginal financial incentives for working long (versus "standard", i.e. 40) hours. Since this increase was not accompanied by an increase in real earnings at standard hours, it can be interpreted as a *compensated* wage increase. In particular, we study the evolution over time of two alternative proxies for the marginal work incentives facing full-time men: the *long-hours premium* and the within-detailed-occupation dispersion of earnings at fixed hours. The first of these gives the total earnings differential within a labor force subgroup between those who usually work (say) 55 hours per week versus 40 hours. The second, as suggested by Bell and Freeman (2001a,b), is used to proxy the distance between the "rungs" of promotion ladders within detailed occupations or industries in a tournament-type model of labor supply. Both of these measures show substantial increases over our sample period. Only the latter, however, does a good job of explaining *which* labor force subgroups experienced the

largest increase in long work hours; this is true whether subgroups are defined by 2- or 3- digit industry, 2- or 3-digit occupation, or age-education cell.

## **II. Data and Descriptive Statistics**

The phenomenon we hope to explain is illustrated in Figure 1, which is calculated from seven decades of U.S. Census microdata. 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 Census week, followed by a rise between 1970 and 1990. The Census did not collect information about actual weekly hours in 2000, but information on usual hours (while employed) in the calendar year preceding the Census is available from 1980 to 2000. This shows a similar increase. Part (b) of the Figure, 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. The increase in work hours after 1970 however remains just as strong from about 16 to 26 percent between 1970 and 1990.

Our primary data sets for this paper are the Outgoing Rotation Groups (ORGs) of the CPS from 1979 to 2004. Clearly, these focus on the period of increasing weekly hours identified from the Census data in Figure 1. 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 method of pay (salaried vs. hourly) which plays an important role in

some of 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.<sup>3</sup>

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 will be 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.<sup>4</sup> 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. Third, weekly hours are of interest in their own right, with implications for the quality and rhythm of family life that are distinct from other margins of labor supply variation.

In addition to focusing on weekly hours, we place most of our emphasis 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.<sup>5</sup> One reason is that the incidence of long work weeks strikes us as of particular importance for worker welfare; another is that 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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<sup>3</sup> 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.

<sup>4</sup> Related, 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).

<sup>5</sup> 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 2004.<sup>6</sup> 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 make it clear that (a) the incidence of long work hours fell in the recessions of 1983, 1992 and 2002; but (b) the secular trend in long work hours was upward and (perhaps surprisingly) stronger in the 1980's than the 1990's.<sup>7</sup> Further, note that the increase in long work hours coincides with a secular *decline* in men's labor force participation. This decline plays an important role in distinguishing among alternative explanations of the rise in hours later in this paper.

Of course, 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, it is however clear that 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 twenty-three-year period.<sup>8</sup> Second, as Figure 2 clearly 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

<sup>6</sup> The CPS defines the main job as the one with the highest weekly hours.

<sup>7</sup> This phenomenon is partially obscured in Figure 1's Census data because the year 2000 was at a higher point in the business cycle than 2001.

<sup>8</sup> 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.

change in men's long work hours are also observed 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.<sup>9</sup>

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: 1980, 1990 and 2001. As Figure 2b shows, each of these was one year after a business cycle peak.<sup>10</sup> Overall, according to CPS data the incidence of long work hours increased relatively modestly over this period, from 14.7 percent in 1980 to 18.5 percent in 2001, with essentially all the increase during the 1980's. Of course, time trends in the fraction of men working long hours may also have been affected by changes in the fraction of men who work part time. To eliminate the effect of part time workers on the data, row 2 of Table 1 restricts attention to the population of men who work at least 30 hours per week. Clearly this has little effect on the results, primarily because the fraction of employed men aged 25-64 who usually work part time is very low. Since our main interest in this paper is on the fraction of full-time men who choose to work long work

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<sup>9</sup> 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.

<sup>10</sup> Very similar results occur when we compare business cycle peaks or troughs over time.

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 greater among salaried workers (from 22.7 to 32.2 percent), than among the hourly-paid (from 7.3 to 9.4 percent). Perhaps surprisingly, the recent increase in long hours is smallest in our youngest age group (25-34) and largest among older men: while older men were less likely to work long hours than young men in 1980, this differential essentially vanished by 2001. 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 much greater among the college-educated. In fact, our data show no increase in long work hours among high school dropouts at all.

The bottom panel of Table 1 examines the correlation of work hours trends with the hourly level of pay. It does so by ranking workers in each of the three selected years according to their average hourly earnings.<sup>11</sup> Clearly, the recent increase in long work hours has been concentrated among the highest wage earners: between 1979 and 2002, the frequency of long work hours increased by 14.4 percentage points among the top quintile of wage earners, while *falling* by 6.7 percentage points in the lowest quintile.

This panel also casts a new light on the relative constancy of overall long hours during

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<sup>11</sup> 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 24. 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 compelling 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.

the 1990's: during this period, the incidence of long hours continued to increase among the highest wage earners while falling among the less skilled. The continuity of these disparate trends by skill level over our entire sample period is clearly illustrated in Figure 3, which presents long hours rates by wage quintile for every year in our data (with each quintile normalized to one in 1979). Clearly, the differential trends by skill shown in Table 1 are not artifacts of the particular survey years shown there.

Perhaps the most striking feature of Table 1 is the *reversal* in the cross-sectional relationship between hourly wages and long work hours since the early 1980's: in 1983, the worst-paid 20 percent of workers were more likely to put in long work hours than the top 20 percent; by 2002 the top 20 percent were *twice* as likely to work long hours than the bottom 20. This reversal of men's wage-hours relationship is surely one of the most intriguing developments in labor supply during the last quarter-century, and forms part of the puzzle we attempt to answer in this paper.<sup>12</sup>

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 hours for other aspects of labor supply.

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<sup>12</sup> 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.

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 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 labor force participation. As noted, men's aggregate employment-population ratio fell over this period; this raises the possibility that what happened to men 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.<sup>13</sup> 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 recent decline in men's employment rates is *smallest* among precisely those men (the better-educated) who experienced the largest *increases* in

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<sup>13</sup> This scenario could in fact emerge quite naturally from a life-cycle labor supply framework (e.g. Heckman and MaCurdy 1980) subject to at 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.

long work hours. This 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. Instead, highly educated men raised their hours while reducing their participation rates only slightly, while high school dropouts did essentially the opposite. Since our interest here is in explaining the recent increase in long work hours, and since this increase was largely confined to skilled men, our primary challenge in this paper will be to explain an increase in skilled men's hours that is consistent with at most a modest decline, or a rough constancy, in their employment rates.

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. We have already noted (in the discussion of Table 1) that there has been very little change over our period in the share of men working part time; further inspection of the data also shows that the very 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.<sup>14</sup>

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<sup>14</sup> See U.S. Department of Labor (2004). We provide a summary table at <http://www.econ.ucsb.edu/~pjkuhn/Data/DataIndex.html>.

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 overall employment rate fell over this period, it remained roughly constant among the group of workers—the college educated—where long work hours increased the most. Thus the challenge we face in the remainder of the paper is explaining an increase in weekly work hours among skilled men, accompanied by what is at most a small decline in their employment rates.

### **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. Of course, one well-known way to adjust for composition effects of this kind is Oaxaca's (1973) method, which we use in what follows to decompose the change in the proportion of men working long hours between the beginning and end of our sample period. To represent the beginning of our sample period we now pool observations for 1983, 1984 and 1985; our end-of-period sample comprises 2000, 2001 and 2002. We chose 1983 as a starting date because the CPS

switched from 1970 to 1980 Census occupation and industry codes beginning in 1983, and the 1970 codes are considerably different from the later ones.

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 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—clearly 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*

Of course 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.<sup>15</sup> 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 2000/02.<sup>16</sup> 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.

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<sup>15</sup> David Autor kindly supplied 3-digit occupation codes that are consistent over the entire 1983-2002 period.

<sup>16</sup> 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.

#### IV. Explanations Based on Real Wage Levels

Having eliminated composition effects as primary explanations of the increase in long work hours, we now turn to the effects of changes in financial work incentives. We begin by examining the explanatory potential of a class of very simple static labor supply models in which (a) workers treat the level of real average hourly earnings (i.e. the real “wage rate”) as parametric in their choice of hours, and (b) there is a stable relationship, which we can think of as an uncompensated labor supply elasticity, between the level of real wages and hours worked. The models used by Juhn, Murphy and Topel (JMT) (1991) and by Juhn (1992) to understand trends in men’s employment rates are examples of this approach; we shall refer to these models as “wage-level-based” in what follows.

To address this question, we examine the covariation between real wage changes and hours changes across subgroups of men over this period. Since the biggest divergence in hours and wage trends across subgroups of men in Table 1 was across wage quintiles, we focus mainly on this disaggregation of the male labor force in Table 6 and in Figure 4.<sup>17</sup> Table 6 and Figure 4 show trends in real average hourly earnings for the five wage quintiles shown 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) there was essentially no aggregate real wage growth among men during the two decades under study. The top quintile of earners did experience a gain, most which occurred in the 1990’s. The second quintile experienced a slight decline in the 1980’s, followed by a larger rise in the 1990’s. The remaining three quintiles all experienced real wage losses over the period as a whole.

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<sup>17</sup> In the following section we shall present a quantitative assessment of wage-level-based models using a variety of other disaggregations.

At first glance, the wage trends in Table 6 can be seen as supportive of a wage-level-based model. Consistent with the notion of a positive, uncompensated labor supply elasticity, there is positive covariation across subgroups of men between real wage changes and changes in the incidence of long work hours over the last two decades. However, closer examination of the Table reveals some problems involving both the timing and sign of the changes. In particular, the largest increases in long work hours occurred during the 1980's when overall real wage growth was negative. For example, comparing Tables 1 and 6, between 1980 and 1990 the middle quintile of wage earners increased their incidence of long work hours by 5.9 percentage points (or, viewed another way,  $(.180-.121)/.121 = 49$  percent) while experiencing a real wage *decline* of 6.4 percent. At a minimum, some other feature must therefore be added to a wage-level-based labor supply model to make it consistent with the trends documented here.

One element that might “rescue” a wage-level-based explanation of recent trends in long work hours is CPI bias. If this bias was large during the 1980s, it might be sufficient to 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.<sup>18</sup> However, this resolution would immediately raise three other problems. First, unless CPI bias *decelerated* dramatically during the 1990's, 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, large real wages increases for the median man make it very hard to explain the aggregate

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<sup>18</sup> Estimates of the amount of CPI bias in recent real wage indexes vary widely (see for example Moulton (1996, Table 1). The vast majority, however, are bracketed by the well-known Boskin Commission's range of plausible estimates of from 0.8 to 1.6 percent per year (see for example Boskin et al., 1998). Cumulated over 20 years, these two estimates amount to about 17 and 37 percent respectively, certainly sufficient to convert most or all of the estimated wage declines in Table 6 into real wage gains.

declines in men's employment rates documented earlier in the paper. Third, any model based on the assumption of positive uncompensated labor supply responses faces the problem of explaining why real wage increases were associated with declining, not increasing hours of work during the century before 1970. We conclude that CPI bias cannot convincingly reconcile simple labor supply models (of the type that treat average hourly earnings as parametric to workers) with the most basic trends concerning men's labor supply in the 20<sup>th</sup> century.<sup>19</sup>

## V. Changes in Marginal Work Incentives

In this section, we consider a slightly more sophisticated labor supply model. In particular, we propose the following hypothesis: during the latter third of the twentieth century, the *marginal* financial incentives for supplying hours of work beyond 40 per week increased substantially for skilled U.S. men. Further, this marginal change occurred without a substantial increase in the *level* of earnings for an individual who kept his work hours fixed at 40. In other words, and in contrast to the earlier part of the 20<sup>th</sup> century, we argue that skilled, salaried US men experienced a *compensated* real wage increase between 1979 and 2002. The compensated nature of this wage increase explains both (a) why labor supply responses were positive and (b) why there was no associated

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<sup>19</sup> A second element that, at first glance, might "rescue" a wage-level-based model is the Reagan tax cuts of the 1980's. Aside from the implausible labor supply elasticities required for this argument (see the discussion in Alesina *et al* 2005), this faces some other problems. For example, if the tax cuts are viewed as a simple wage increase that was not accompanied by a decline in the provision of government services, it is hard to understand why men *left* the labor market at the same time they increased their weekly hours. On the other hand, if we view the tax cut as accompanied by a decline in public services (as does Prescott 2004 for example), the tax hypothesis has trouble explaining the *distribution* of the hours increase. In particular, given the highly redistributive nature of most government spending, any reductions in transfers or public services should disproportionately affect *unskilled* workers. If leisure is a normal good, the resulting income effects would *raise* work hours among the less-skilled and reduce it among the skilled, which is the opposite of what happened.

increase in labor force participation (since the overall attractiveness of work versus inactivity did not rise substantially).

The simple intuition behind our hypothesis is illustrated in Figure 5. Consider first an unskilled “representative” hourly paid worker who, over the period from 1900 to 1970, experienced a substantial increase in the real wage earned for every hour worked. This familiar scenario is illustrated in part (a) of the Figure; as shown, utility-maximizing weekly hours of work fell over this period due to income effects. Now consider a skilled man who in 1970 was paid on a salaried basis for a 40-hour work week. Presumably, if this worker regularly failed to supply the minimum contracted number of hours (40), he would lose his job. But what incentives does this worker have to supply more than forty hours of work per week? In contrast to the hourly paid, salaried workers receive no such reward in the short run. Such a situation would be illustrated by the darkly shaded rectangular budget constraint in Figure 5, with optimal work hours exactly at  $168 - 128 = 40$ .

Over a longer time period, of course, a substantial set of rewards could accrue to salaried workers who put in “extra” hours. These include the possibility of 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, or contacts) that may be rewarded in either the current firm or another one, and perhaps an enhanced prospect of keeping one’s current job if the firm is forced to lay off workers in the future. Finally, suppose for the sake of argument that these rewards to “extra” hours were essentially zero in 1970, but substantial (when future rewards are converted to expected present values) in 2000. This

would add the ray  $ab$  to the budget constraint, and is unambiguously predicted to raise work hours. Because the original labor-leisure bundle is assumed to remain (just) affordable after this change, the change can be thought of as an (income-) compensated wage increase.

Is there any evidence that a compensated wage increase of the form sketched in Figure 5b affected salaried American men over the past two decades? In the remainder of this section we propose and implement two proxies for the strength of these marginal incentives: the *long-hours premium*, and the within-group dispersion of earnings.

### 1. The Long-Hours Premium

If a full-time worker regularly puts in “extra” work hours, then according to the hypothesis sketched above, this will eventually manifest itself in a higher weekly salary. If tastes for work are a relatively fixed personal characteristic (ensuring that current hours are indicative of past hours), the cumulated effects of past work hours decisions on an individual’s rate of pay will be detectable in cross-sectional salary comparisons.<sup>20</sup> In what follows we shall refer to cross-sectional estimates of the slope of  $ab$  as estimates of the *long-hours premium*. Since (under the above interpretation) this estimated premium

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<sup>20</sup> Somewhat more formally, if a given type of workers (for example a detailed occupation) faces a common budget constraint, then cross-sectional within-group comparisons of total weekly earnings at different hours levels will identify exactly the slope of segment  $ab$  in Figure 5a. 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. An increase in the within-group slope over time could thus indicate an increased sorting of abler workers into higher levels of hours. We readily admit that we cannot distinguish 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.

reflects the effects of *past* hours on current compensation, it has a retrospective aspect that we discuss further below.<sup>21</sup>

Pursuing the above logic, Table 7 presents estimates from the following set of cross-section regressions for the same two periods, 1983-85 and 2000-02, used in Tables 3 through 5. Separately for each period, panel (a) regresses total log real weekly earnings on usual weekly hours, education indicators, a quartic in age, and (where noted) a full set of three-digit occupation fixed effects. To focus on the incentives facing full-time workers, and to avoid any possible disproportionate impact of very-high-hours outliers on the estimated shape of the hours-wage relationship in the range of hours relatively near 40, the sample throughout Table 7 is restricted to men with usual hours between 40 and 65. As shown, the apparent marginal reward to putting in extra work hours within an occupation increased substantially between 1983/85 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.<sup>22</sup>

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. Again, the differences increase over time, and do so much more dramatically for salaried workers. In the early 1980's, a

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<sup>21</sup> An alternative interpretation of the long-hours premium is as a within-group market locus: if a relatively dense array of jobs with different wage-hours bundles is available to workers, then our estimated premium simply reflects the market's equilibrium reward for marginal hours. We are indebted to John Pencavel for this suggestion.

<sup>22</sup> 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.



randomly selected salaried man putting in 55 hours per week earned a weekly salary of 10.5 percent than an observationally-equivalent man in the same 3-digit occupation. By the early 21<sup>st</sup> century, that gap had more than doubled, to 24.5 percent. The functional form of the predicted relationship is shown in Figures 6a 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. As we shall argue below, one possible interpretation of this trend is as a change, towards greater "incentivization", in the compensation schedule facing a typical salaried man during this time period.

Panel (b) of Table 7 illustrates one remaining point of note: the increase in the apparent *marginal* return to extra work hours beyond 40 over this period was not associated with an increase in the *level* of real earnings for workers putting in 40 hours per week: real earnings at 40 hours remained essentially flat among hourly workers and increased only slightly among salaried workers. In this sense, the change between 1983/85 and 2000/02 in the estimated budget constraint facing salaried men matches Figure 5b almost exactly: the total earnings associated with a standard work week (point *a* in the Figure) remained essentially fixed over these two decades; a fact which eliminates the income effects that would normally reduce hours worked in the case of a more conventional wage increase.

## 2. *Within-Group Earnings Dispersion*

In some recent papers, Bell and Freeman (2001a,b) have proposed a second way to proxy the change in lifetime present value of earnings associated with an extra current hour of work in cross-sectional data. Arguing that, within occupations, (or even in the labor market as a whole) compensation can be understood 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. Because the current within-occupation distribution of wages can be seen as the set of jobs a worker can aspire to, in this sense Bell and Freeman's measure is more "prospective" than the long-hours premium measure. In what follows, we ask whether Bell and Freeman's argument might help us understand changes over time in the incidence of long work hours in the U.S.<sup>23</sup>

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 marginal work incentives presented in Table 7, the statistics in Table 8 refer to earnings dispersion at a fixed level of work 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 our proxy for marginal work incentives, and this increase is substantially greater among salaried than hourly paid

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<sup>23</sup> 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.

workers. Thus, the increase in within-occupation earnings dispersion also seems a promising *prima facie* candidate to explain the increase in men's long work hours.<sup>24</sup>

We realize, of course, that we are not the first to document an increase in residual wage inequality over this period (see for example Juhn, Murphy and Topel 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. A more important question, however, is whether increases in within-group earnings dispersion can predict *which* detailed labor force subgroups experienced the largest increase in long work hours over the past two decades. We turn to this question in the following section.

### *3. Explaining the Distribution of Changes in Long Work Hours*

The previous two subsections have shown that our two proxies for the marginal financial incentives for working long hours have indeed increased over the last two decades. They also show that these increases were larger among salaried workers-- the group who experienced the larger increase in long hours. That said, our argument that changes in marginal financial incentives have played a causal role would be strengthened if the covariance between the change in long hours and the change in measured incentives was positive across other disaggregations of the labor force as well.

This question is addressed in Table 9, which disaggregates the population of full-time men in five distinct ways: by 2- and 3-digit industry, by 2- and 3-digit occupation,

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<sup>24</sup> 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. For example, the estimated value of the standard deviation of the earnings residual rose from .353 to .398 for hourly paid workers, and from .417 to .501 for salaried workers.

and according to sixteen age-education cells (these are simply the cross product of the four age and education categories used in Table 1). Since (as noted) the bulk of the increase in long work hours occurred among salaried men, and since the hypotheses sketched in the preceding subsections are theoretically most appropriate for salaried workers, Table 9 restricts attention to that group. Each entry in this Table is a  $t$ -statistic from a univariate regression of the within-cell change (between 1983/85 and 2000/02) of the incidence of long work hours on some other characteristic of that cell. 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 *in 1983/85* experienced greater *growth* in long work hours between 1983/85 and 2000/02. The same is true for 3-digit industries, 2- and 3-digit occupations, and age/education cells. According to row 2, the same is true for when we substitute 1983/85 level of *total weekly earnings* for the level of average hourly earnings. Together, rows 1 and 2 of Table 9 provide further evidence that the recent increase in long work hours was highly concentrated among highly skilled men.

Row 3 of Table 9 returns to the question of whether simple wage-level-based models like those described in Section IV (models that treat average hourly earnings as parametric to the worker) to explain the recent increase in long hours. According to such models (assuming the uncompensated labor supply elasticity is positive) then those labor force subgroups that experienced the largest increases in hourly wage rates should have

seen the largest increases in long work hours. While this is (marginally) true when the labor force is broken down by age-education cells, it is decidedly not true for any of the other disaggregations in the Table. 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 those groups that raised their hours the most were compensated for those extra hours in the form of higher earnings. But even in the presence of this positive bias, only a zero or weak positive correlation between earnings changes and hours changes is found.

The next four rows of Table 9 test the ability of various indicators of the long-hours premium to explain the recent increases in long work hours. Rows 5 and 6 use very simple measures of the long-hours premium: the within-cell difference in average hourly earnings (or 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, a separate regression like that in Table 7 was run for each cell in the data; in row 7 using the linear specification of Table 7(a); in row 8 using the quartic specification in Table 7(b).) 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 labor force subgroups 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 the within-cell dispersion of earnings at fixed hours caused increases in the incidence of long work hours, using four alternative indicators of earnings dispersion. The first two of these (in

rows 9 and 10) are simple descriptive statistics: the within-cell standard deviation of log earnings, or of log wages. The remaining two, computed from the cell-specific regressions used in Rows 7 and 8, are earnings residuals of the types reported in Table 8. Clearly, these results are much more robust, with positive coefficients in all 20 cases, and statistically significant ones in 18. Thus, no matter how the labor force is disaggregated, those subgroups of men that experienced an increase in residual earnings variation (whether overall *or* at a fixed level of hours) were more likely than other subgroups to experience increases in long work hours.

Could the strong results in rows 9 through 12 of Table 9 somehow be a spurious consequence of our statistical and/or computational techniques, including how those techniques interact with possible measurement error in hours? To check for these possibilities, we replicated rows 11 and 12 for a sixth disaggregation of the labor force, one we refer to as “pseudo-cells”. In this exercise, each cell consisted of 50 workers, each of whom was selected at random from a *different* state. Thus, nothing would lead us to expect that the workers in each cell would share a common compensation schedule against which they optimize; further, no common characteristic links the persons in a particular cell across the two time periods under study. Any estimated association between cell-level changes in work hours and residual earnings variation should therefore be driven purely by artifacts of our statistical procedures. When we did this, no relationship was found (specifically, the *t*-statistics corresponding to rows 11 and 12 of Table 9 were -0.5 and -1.0 respectively).

In sum, Table 9 does provide some evidence that increases in the incidence of long work hours over this period were more common in labor force subgroups where

measured marginal work incentives increased the most. However the only measure of marginal incentives which consistently has a significant effect on work hours is within-group pay inequality. This suggests that a tournament-type framework might be a better way to model the work hours decisions of salaried workers than some other alternatives, including the “traditional” labor supply model where the ratio of earnings to hours is taken as parametric, or the long-hours premium model discussed above. Perhaps the better performance of the earnings-dispersion measure of marginal work incentives relates to its more “prospective” nature: an increase in wage inequality both immediately raises workers’ incentives *and* is immediately reflected in our data. An occupation-specific increase in the effects of past hours on current wages could take considerable time to be detected in our data.

#### *4. Corroborating Evidence*

In this subsection we review other studies for evidence that incentives to supply work hours beyond 40 have increased among skilled or salaried men. Clearly, one way in 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 –precisely those groups where long hours increased the most in our data-- 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).<sup>25</sup> Summary tabulations from this survey are presented in Table 10. Clearly, *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

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<sup>25</sup> Some impressive --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.



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.

## **VI. Discussion**

In this paper we have shown that American men are more likely to work over 50 hours per week now than a quarter-century ago. This trend is of particular interest because (a) it reverses a secular decline in the work week that dates back over a century; and (b) it occurred in the face of a decline in men’s employment rates and in their median real hourly earnings.

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 phenomenon associated with 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. 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, 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.

In our search for explanations, we also reject simple “wage-level-based” models in which average hourly earnings are interpreted as an hourly wage rate against which workers optimize. Whether a positive *or* negative uncompensated labor supply curve is assumed, models of this type simply cannot account for the most basic facts of men's hours and work participation during the 20th century. Such models are also—at least superficially-- inconsistent with the contractual features of salaried jobs, where the incentives to supply extra work hours are harder to measure, and where the bulk of the recent increase in long work hours took place.<sup>26</sup>

In our view, the most likely explanation for the recent increase in men's long hours is a rise in the marginal incentives to supply hours beyond 40 among skilled, salaried workers. Since these added incentives only applied to hours beyond 40 per week, they can be thought of as a *compensated* wage increase, which explains why they were not counteracted by income effects. Unlike wage-level-based hypotheses, this explanation is consistent with broad trends in men's hours *and* labor force participation throughout the twentieth century. Our explanation is also supported by the fact that both measures of marginal work incentives in this paper—the long-hours premium and the

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<sup>26</sup> 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. If 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.

within-occupation residual dispersion of earnings—increased markedly between 1983 and 2002, especially among salaried men. Only the latter measure, however, does a good job of explaining *which* occupations, industries and age-education groups experienced the largest increases in long work hours during our sample period. Perhaps the better performance of this “inequality-based” measure relates to its prospective, rather than retrospective, means of capturing marginal work incentives.

We conclude by noting that this paper does not attempt to explain *why* U.S. firms might have changed their methods of compensation for skilled, salaried workers over the past quarter century; in that sense we have identified only a proximate cause for the rise in long hours. Our conjectures regarding more fundamental causes include the possibility that contracts with stronger incentives are a socially efficient response to an increase in the demand for skilled workers, or to 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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**Table 1: Fraction of Men Usually Working Long ( $\geq 50$ ) Hours**

	1980	1990	2001	Change	Change	Change
				1980- 1990	1990- 2001	1980- 2001
All Men	0.147	0.192	0.185	0.045	-0.007	0.038
Full Time Men ( $\geq 30$ hours)	0.150	0.198	0.203	0.048	0.005	0.053
<b>Among Full Time Men:</b>						
Salaried	0.227	0.310	0.322	0.083	0.012	0.095
Hourly	0.073	0.093	0.094	0.020	0.001	0.021
Age 25-34	0.153	0.195	0.186	0.042	-0.009	0.033
Age 35-44	0.171	0.215	0.218	0.044	0.003	0.047
Age 45-54	0.144	0.197	0.210	0.053	0.013	0.066
Age 55-64	0.116	0.159	0.190	0.043	0.031	0.074
Less than High School	0.109	0.124	0.109	0.015	-0.015	0.000
High School Graduates	0.123	0.154	0.148	0.031	-0.006	0.025
Some College	0.152	0.188	0.185	0.036	-0.003	0.033
College Graduates	0.222	0.298	0.305	0.076	0.007	0.083
Average Hourly earnings quintile:						
1 (highest wage)	0.146	0.235	0.290	0.089	0.055	0.144
2	0.121	0.201	0.216	0.080	0.015	0.095
3	0.121	0.180	0.195	0.059	0.015	0.074
4	0.155	0.192	0.179	0.037	-0.013	0.024
5 (lowest wage)	0.210	0.189	0.143	-0.021	-0.046	-0.067

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

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

	1980	1990	2001	Change	Change	Change
				1990-1980	2001-1990	2001-1980
<b>AGES 25-64:</b>						
<b>Share of Men Employed<sup>1</sup></b>						
Less than High School	0.734	0.701	0.689	-0.033	-0.012	-0.045
High School Graduates	0.869	0.853	0.820	-.0016	-0.033	-0.049
Some College	0.890	0.885	0.864	-0.005	-0.021	-0.026
College Graduates	0.939	0.925	0.911	-0.014	-0.014	-0.028
<b>AGES 45-54 ONLY:</b>						
<b>Share of Men Employed<sup>1</sup>:</b>						
Less than High School	0.789	0.739	0.688	-0.050	-0.051	-0.101
High School Graduates	0.902	0.879	0.826	-0.023	-0.053	-0.076
Some College	0.916	0.901	0.870	-0.015	-0.031	-0.046
College Graduates	0.958	0.950	0.931	-0.008	-0.019	-0.027
<b>Share of Employed working Long Hours<sup>2</sup>:</b>						
Less than High School	0.111	0.123	0.132	0.012	0.009	0.021
High School Graduates	0.116	0.150	0.180	0.034	0.030	0.064
Some College	0.141	0.183	0.233	0.042	0.050	0.092
College Graduates	0.239	0.310	0.346	0.071	0.036	0.107

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

**Notes:** Robust t-statistics errors in parentheses: \* significant at 5%; \*\* significant at 1%

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: usual weekly hours 50 or more. 1983-1985 long hours mean is 0.163 and 2000-2002 long hours mean is 0.203.



**Table 4: Oaxaca Decompositions of the Increase in Long Hours, 1983/85 to 2000/02**

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

**Table 5: Shift-Share Decompositions by 3-digit Occupation and Industry**

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

Note: 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),  
by Wage Quintile, 1979-2002**

	1980	1990	2001	Change	Change	Change
				1980- 1990	1990- 2001	1980- 2001
<b>A. Unadjusted wage, by hourly wage quintile:</b>						
1 (highest )	2.875	2.933	3.043	0.058	0.110	0.168
2	2.507	2.493	2.551	-0.014	0.058	0.044
3	2.290	2.226	2.256	-0.064	0.030	-0.034
4	2.039	1.946	1.959	-0.093	0.013	-0.080
5 (lowest)	1.630	1.518	1.526	-0.112	0.008	-0.104

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

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

**a. Linear hours effect model: coefficient on hours<sup>1</sup>**

Sample:	Controls for:			
	Age, Education		Age, Education and 3-Digit Occupation	
	1983/85	2000/02	1983/85	2000/02
<b>All workers</b>	.0123 (.0002)	.0232 (.0002)	.0119 (.0002)	.0197 (.0002)
<b>Hourly</b>	.0186 (.0003)	.0268 (.0003)	.0199 (.0003)	.0253 (.0003)
<b>Salaried</b>	.0064 (.0002)	.0169 (.0003)	.0068 (.0002)	.0147 (.0003)

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

Sample:	Controls for:			
	Age, Education		Age, Education and 3-Digit Occupation	
	1983/85	2000/02	1983/85	2000/02
<b>All workers:</b>				
<b>40</b>	5.950	5.943	5.952	5.956
<b>55</b>	6.140	6.309	6.135	6.271
<b>difference</b>	0.190	0.366	0.183	0.315
<b>Hourly:</b>				
<b>40</b>	5.796	5.760	5.793	5.764
<b>55</b>	6.084	6.150	6.094	6.134
<b>difference</b>	0.288	0.390	0.301	0.370
<b>Salaried:</b>				
<b>40</b>	6.117	6.157	6.119	6.173
<b>55</b>	6.218	6.437	6.224	6.418
<b>difference</b>	0.101	0.280	0.105	0.245

<sup>1</sup>Coefficients represent the effect of working one more hour per week on the log of total weekly earnings. Standard errors in parentheses.

Sample: Men ages 25-64, currently employed and not self-employed, working between 40 and 65 hours per week. Age is measured as a quartic, hours (in part b) as a quartic, education groups as shown in Table 1.

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

Sample:	Root Mean Square Error, log earnings regression		90-10 differential in log earnings residual	
	1983/85	2000/02	1983/85	2000/02
<b>All workers</b>	.376	.450	.944	1.045
<b>Hourly</b>	.351	.398	.901	.957
<b>Salaried</b>	.385	.488	.948	1.110

Sample: Men ages 25-64, currently employed and not self-employed, working between 40 and 65 hours per week. Log earnings regressions control for a quartic in age, quartic in hours, education categories, and 3-digit occupation fixed effects.

**Table 9: Univariate Regressions between Selected Industry/Occupation/Cell Characteristics and Long-Run (1983/85-2000/02) Changes in the Incidence of Long Work Hours, Salaried Full-Time Men (*t*-statistics)**

Ind/Occ/Cell Characteristic:	Unit of Analysis				
	2D Ind.	3D Ind.	2D Occ.	3D Occ.	Age/Ed Cell
<b>Initial Skill Levels:</b>					
1. Initial log Hourly Wage	4.4	8.9	7.2	9.7	9.9
2. Initial log Weekly Earnings	4.5	8.2	6.8	8.8	9.1
<b>Wage or Earnings Changes:</b>					
3. Change in log Hourly Wage	-2.1	-2.6	0.6	-1.2	1.9
4. Change in log Weekly Earnings	-0.8	0.6	1.9	1.7	2.5
<b>Change in Long-Hours Premium:</b>					
5. Change in LH Earnings Premium	-2.2	-0.7	0.1	0.6	1.8
6. Change in LH Hourly Wage Premium	-2.3	-1.3	-0.3	0.3	1.5
7. Change in Salary Slope (linear specification)	-1.8	-0.8	-0.4	-0.0	1.5
8. Change in Salary Slope (quartic specification)	-1.5	-0.3	0.1	-0.3	1.1
<b>Change in Earnings Dispersion:</b>					
9. Change in Standard Deviation of log Earnings	3.1	2.5	2.0	2.5	2.6
10. Change in Standard Deviation of log Wage	3.6	2.9	2.1	2.4	2.0
11. Change in Standard Deviation of Salary Residual	1.6	1.3	2.1	3.1	3.2
12. Change in 90-10 Salary Residual Gap	2.9	3.0	3.6	4.1	5.7
<b>N</b>	<b>44</b>	<b>180</b>	<b>43</b>	<b>197</b>	<b>16</b>

All regressions restrict attention to cell sizes of at least 50 in both periods, and are weighted by the average number of observations across the two periods.

**Table 10: Percent of Fortune 1000 Firms Surveyed in which over 20 percent of employees are covered by selected reward practices**

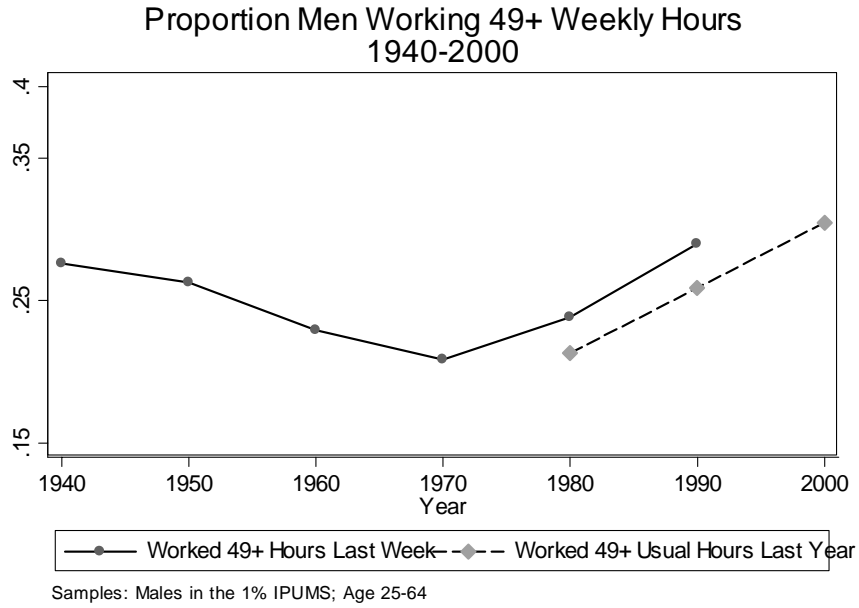
	<b>1987</b>	<b>1990</b>	<b>1993</b>	<b>1996</b>	<b>1999</b>
Individual Incentives	38	45	50	59	67
Work Group or Team Incentives	22	31	n.a.	41	48
Gainsharing	7	11	16	20	24
Profit Sharing	45	44	44	52	55
Employee Stock Ownership Plan	52	56	63	59	63
Stock Option Plan	n.a.	n.a.	30	41	49
Nonmonetary Recognition 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 performance 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 Men Working Long Hours**

a) all employed men



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

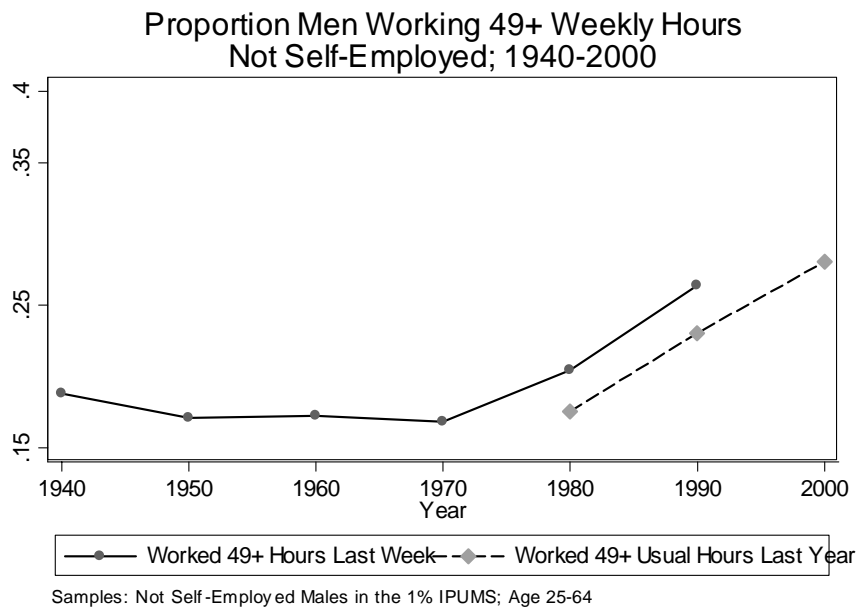
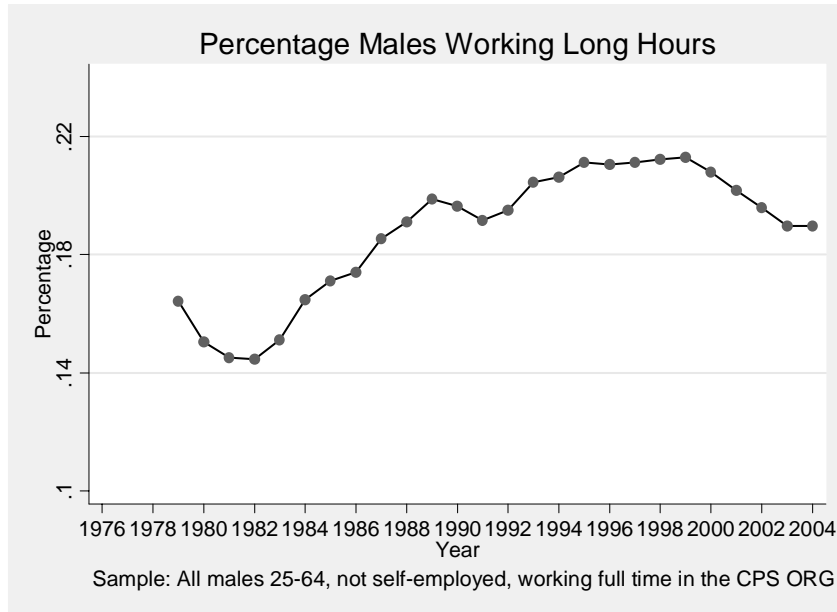
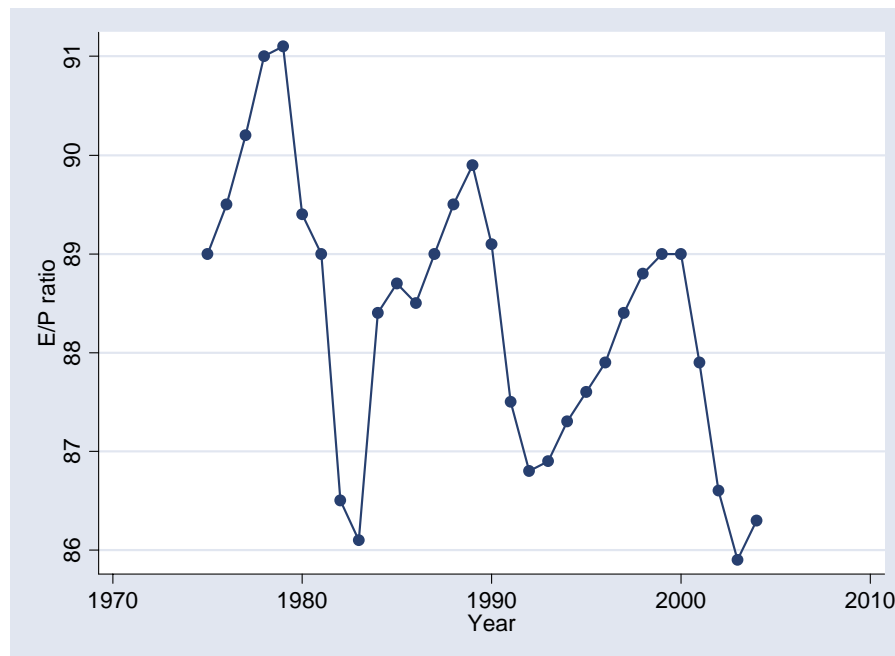


Figure 2: Long Hours and Employment Trends, Men, CPS Data.

a) Long Hours



b) Employment/Population Ratio



Sample: Men, ages 25-54. BLS Website, Public Data Query, Series Id: LNS12300061Q (quarterly, converted to annual by the authors)



Figure 3: Percentage of Men Working Long Hours, by Hourly Earnings Quintile

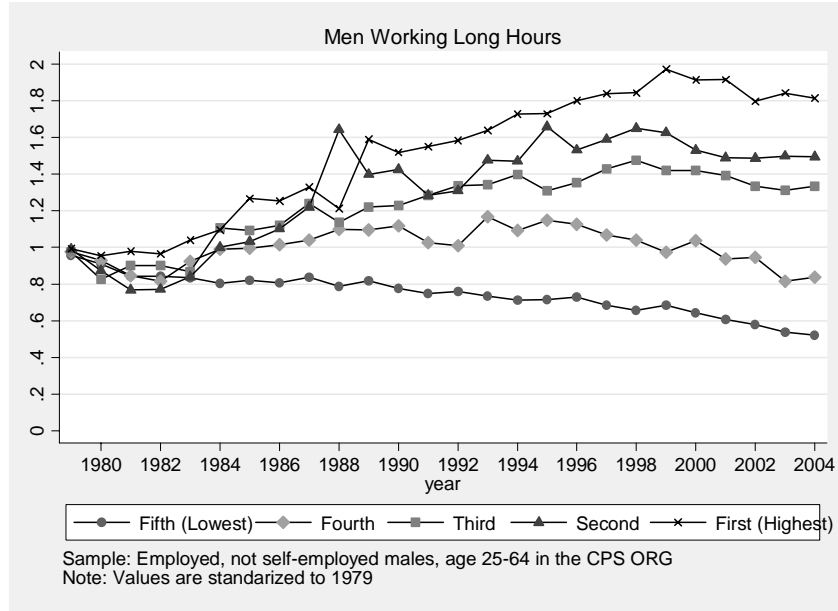


Figure 4 Average Hourly Earnings, by Quintile

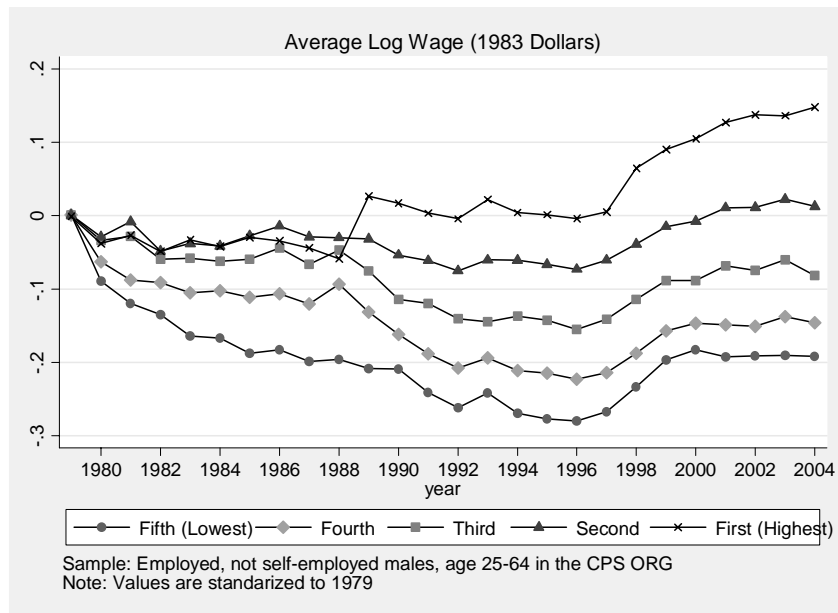
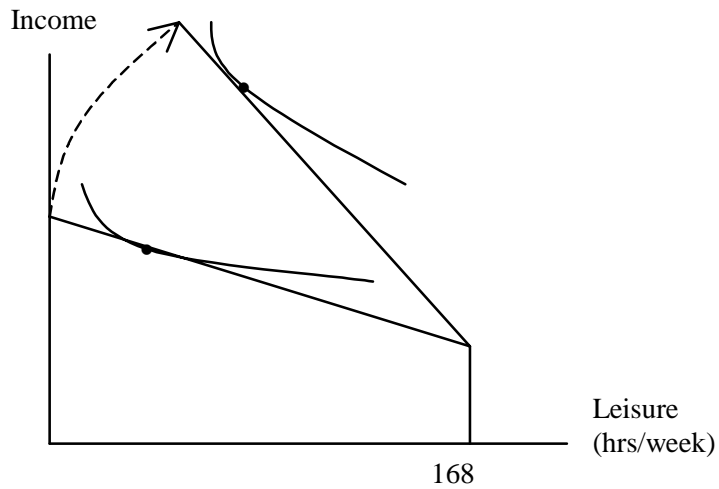
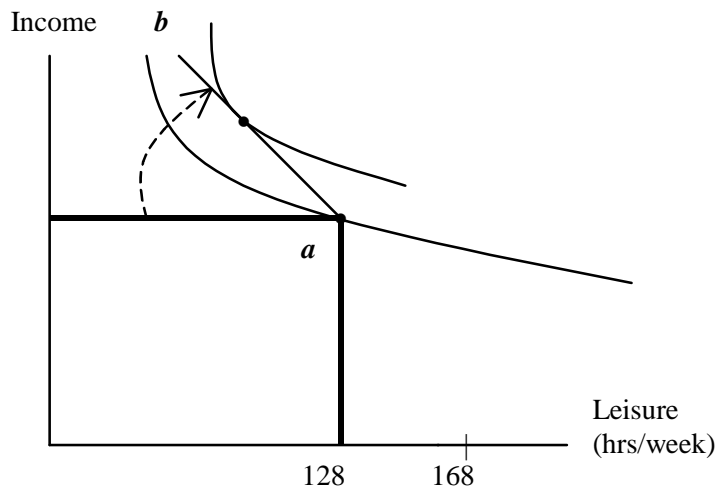


Figure 5: Hypothetical Wage Increases

a) Across-the-board real wage increase (posited for 1900-1970; negative labor supply response):

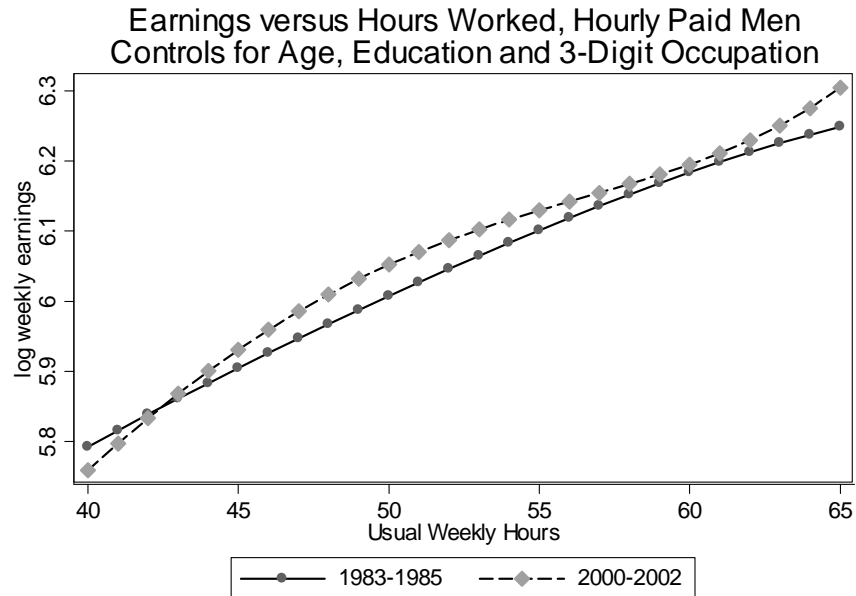


b) Increased marginal work incentives for salaried workers (posited for 1970-2000; positive labor supply response):



**Figure 6: Earnings versus hours worked, Salaried Men and Hourly-paid Men, 1983/85 and 2000/02**

**a) Hourly Paid Men**



**b) Salaried Men**

