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Differential Effects of Graduating during a Recession across Race and Gender

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DIFFERENTIAL EFFECTS OF GRADUATING DURING A RECESSION ACROSS RACE AND GENDER[1]

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Abstract

This paper examines the differential effects of the unemployment rate at entry to the labor market, defined as completion of education, on subsequent wages across race and gender. Economic theories about search frictions, human capital accumulation and the internal labor markets all predict less persistence for low skilled or disadvantaged workers and weaker effects on those with weak attachment to the labor force. Consistent with these predictions, I find that the effect fades faster for blacks, although the initial impact of a recession at entry is stronger for them. I also find weaker effects for women.

Keywords: unemployment rate; labor market entry; race; gender.
JEL classifications: E24, J15, J16, J31.

| 1 | INTRODUCTION

Recent studies have established that entering the labor market during a recession has long-term effects on the person's subsequent wages and career path (Oreopoulos, von Wachter and Heisz 2006, Kahn 2006). However, how the long-term consequences of such exogenous shocks at entry vary across different demographic groups has not been explored. Since the effect of contemporaneous labor market conditions differs considerably across demographic groups, the long-term effect of a recession at entry into the labor market may well differ. Given that the existing studies have mainly targeted high-skilled workers with strong attachment to the labor force, namely white male college graduates, focusing on African Americans and women may uncover different patterns for workers who are relatively less-skilled and with weaker attachment to the labor force.

This paper examines the effect of the unemployment rate at entry to the labor market, defined by completion of education, on subsequent wages by race and gender using data from the National Longitudinal Survey of Youth. To control for nation-wide cohort effects, I exploit variations in unemployment rates across states, with controls for time invariant state characteristics and regional differences in black-white wage gaps. I also address the endogenous determination of time and place of entry conditional on the decision of college degree attainment using an instrumental variable and provide evidence that the correlation between attainment of a college degree and the unemployment rate at age 18 is negligible.

I find a weaker effect of a recession at entry for women than for men; this contrast is especially stark for whites, for the effect for white women is almost zero. Also, although the initial impact is stronger for blacks than for whites, the effect fades faster for blacks, too. These findings are consistent with the predictions by economic theory: less persistence for low skilled workers and weaker effects on those with weak attachment to the labor force.

The rest of the paper is structured as follows. First, Section 2 contains a brief overview the previous literature. Section 3 reviews relevant theories and their predictions. Then, Section 4 describes data and Section 5 describes empirical strategy. Estimated effects of a recession at entry are presented in Section 6. Section 7 concludes.

| 2 | BACKGROUND

There is a good deal of evidence that entering labor market during a recession lowers subsequent wages for relatively high skilled men. Oreopoulos *et al.* (2006) have established these key observations using a large panel dataset of Canadian male college graduates to estimate the effect of graduating colleges during a recession. They find a significant negative effect on earnings that gradually fades away and a positive effect on turnover and mobility across industries and regions. My empirical specification follows their work. Kahn (2006) reports similar results using the sample of white male college graduates in the National Longitudinal Survey of Youth, the same data source used in this paper, with a slightly different empirical specification. Nevertheless, to my knowledge, the long-term effect of a recession at entry for less educated men, African Americans or women in North America have not been explored.

On the other hand, there is substantial evidence for differential affects of the contemporaneous labor market conditions across different demographic groups. Although it is controversial whether an economic upturn benefits economically disadvantaged people as much as the more advantaged,[1] there is little doubt that the costs of a downturn are borne disproportionately by the disadvantaged (Cutler and Katz 1991, Hines, Hoynes and Krueger 2001, Clark and Summers 1981). The stronger effect of the contemporaneous labor market condition suggests, on one hand, a stronger initial impact for the relatively disadvantaged group; on the other hand, the vulnerability to aggregate shocks after entry to the labor market may reduce the persistence of a negative shock at entry.

Several earlier studies have explored the effects of a high unemployment rates in the beginning of job tenure on current wages (Beaudry and DiNardo 1991, Baker, Gibbs and Holmstrom 1994b) and quitting rates (Bowlus 1995). Although the effects of unemployment rates in the beginning of job tenure are conceptually different from those of the unemployment rates at entry to the labor market, these studies provide several clues to disentangle the confounded effects. Beaudry and DiNardo (1991) find that current wages depends on the minimum unemployment rates during the job tenure, rather than the unemployment rates at the beginning of the job, and regard it as evidence that worker can move to better jobs during booms. In contrast, Baker *et al.* (1994b) find a persistent effect of the labor market condition at entry to a large firm. Although, at first glance, the two results contradict to each other, both can be true if the only way for a worker to improve her situation is moving to a better employer, provided that the firm size serves as a measure of job quality. Bowlus (1995) finds that a job started during a recession ends earlier and interprets it as evidence of deterioration of matching quality.

| 3 | Relevant Theories and Predictions

Several interrelated factors produce a persistent negative effect of a recession at entry on subsequent wages. First, search frictions prolong the process to resolve the initial bad matches. The logic is straightforward: it takes time and costs to find a good job under incomplete information about available job opportunities. In a typical job search model, a new job opportunity drawn from a given distribution arrives randomly and a worker chooses to take the job if she finds it is better than remaining in her present job. Since a worker voluntarily changes her job only if her wage increases, workers who obtained high-wage jobs at entry remain in the high-wage jobs while workers who had low-wage jobs have to move up gradually through job search.

The advantage of obtaining a high-wage job at entry vanishes when the worker is dismissed or quits due to exogenous reasons such as marriage and childbirth. Therefore, the effect of entering the labor market during a recession will be more persistent for people who are

[1]While Cutler and Katz (1991) argue that the explanation in the mid-late 80s did not benefit the disadvantaged, Hines et al (2001) view the 1980s as an exception and conclude that the benefits from strong economic growth for the disadvantaged are at least as great as they are for the more advantaged using data spanning from the 1970s to the 2000s.

more likely to stay in the labor force without being dismissed. Since unemployment rates for blacks are more variable in response to the business cycle conditions (Altonji and Blank 1999), it is natural to think that blacks are more likely to be dismissed in a recession after entry. Studies on cyclical up-grading of the labor force (e.g. Okun 1973, Devereux 2004) also suggest that people who lose their jobs in recessions are less skilled while those who move up to higher paying jobs in booms are more skilled, that is, the story based on search theory is more relevant to more skilled workers. Likewise, a turnover due to marriage or childbirth will make the effect of recessions at entry less persistent for women than for men.

Note that, in this sort of search model, the initial impact of a recession is stronger for groups more sensitive to temporary fluctuations of labor demand in the entry-level labor market. Therefore, the impact of a recession on wages in the first few years will be greater for blacks. On the other hand, the initial impact on women will be weaker because they are less sensitive to aggregate labor market conditions (Clark and Summers 1981, Altonji and Blank 1999).

Further, missing opportunities for on-the-job training due to a bad match can aggravate the negative effect of entering the labor market during a recession.[2] It is ambiguous whether the loss of training opportunities matters more to African Americans, who on average have lower education and other pre-market human capital,[3] and fewer training opportunities on the job. If pre-market human capital increases returns to on-the-job training, missing a training opportunity will matter more to white workers; if the marginal return to training is diminishing, it will matter more to black workers. In any case, this loss of human capital would not occur if the worker could move to a job with better training opportunities without frictions; thus the amount of the lost training opportunity is likely to be greater for whites. Also, weak labor force attachment implies less incentive to invest in human capital, thus it predicts a weaker effect of a recession at entry on subsequent wages for women than for men.

The discussion drawing on search frictions implicitly assumes the existence of long-term employment contracts. Going a little further, if wages are also determined in long-term contracts, the initial conditions of labor market directly affect the subsequent wages. Although the evidence from existing studies is mixed,[4] at least for some workers wages are not very sensitive to the external labor market conditions. Moreover, access to internal labor markets of high-paying firms through the entry-level jobs may be limited to new graduates. Baker, Gibbs and Holmstrom (1994a) report suggestive evidence for this: majority of the entrants to a large firm are young workers who enter the lowest level job (ports of entry) and many of them stay in this firm for many years and are promoted within this firm. Although

[2]The idea of viewing an early job as a training opportunity goes back to Rosen (1972). Welch (1979) draws on this idea to explain why cohort size affects earnings.

[3]For example, Table 2 shows that African Americans are not only less educated but also their scores of an ability-test are lower. See Section 5.1 of Altonji and Blank (1999) for the literature on pre-market human capital.

[4]In addition to Baker et al (1994b) and Beaudry and DiNardo (1991), which are summarized in Introduction of this paper, Grant (2003) replicates and extends the Beaudry-DiNardo model to find that both the contemporaneous unemployment rate and minimum rates over job tenure affect wages, and Bertrand (2004) finds that import penetration increases the sensitivity of wages to the external labor market conditions.

they find a non-negligible number of entrants to higher level jobs, contrary to exclusive "ports of entry" described by Doeringer and Piore (1971), these entrants seem to have accumulated working experience elsewhere, which comes back to the issue of loss of training opportunity.

As pointed by Doeringer and Piore (1971), jobs that require higher skill and pay higher wages are more likely to be associated with internal labor markets because complicated tasks often need sophisticated incentive schemes. On the other hand, marginal workers with weak attachment to the labor force obtain jobs in the secondary sector, which works more like a spot market. Thus, long-term wage contracts and internal labor markets seem to be more relevant to more skilled workers with stronger labor force attachment. The higher share of white men in higher-paying occupations and industries (Altonji and Blank 1999) suggests that they are more relevant to white men. Grant (2003) also suggests that long-term wage contracts are actually slightly less relevant for women.

To recapitulate, economic theory reviewed above predicts a stronger effect of a recession at entry for people with stronger labor force attachment, and the effect will be more persistent for more skilled workers who are more likely to be on stable employment contracts. Combining with the existing evidence of a stronger effect of contemporaneous labor market conditions for African Americans, the effect of a recession at entry is expected to be initially stronger but fade away faster for African Americans than for whites. Also, given that more women withdraw from the labor force upon marriage and child births than men, the effect is expected to be weaker for women.

| 4 | DATA

4.1 NSLSY79

My main source of data is the National Longitudinal Survey of Youth 1979 (NLSY). The NLSY consists of a representative sample of the noninstitutionalized U.S. civilian segment of population aged 14-22 in 1979. In addition, a set of supplemental samples of Hispanics, blacks, and economically disadvantaged whites born in the same time period, and a military subsample are included. I drop the military subsample because it covers only cohorts born between 1957 and 1961, and the supplemental poor whites sample because they are dropped from the survey after 1991. I also drop Hispanics due to a fear of non-random selection by the sampling scheme (they had to be living in the US in 1979) and the small sample size.

The year of labor market entry is defined as the last year of enrollment before the respondent turned 30 years old. I do not count enrollment for one or two years in colleges after not being enrolled at least for a year or enrollment after the age of 30 because it is likely to be a part of vocational trainings in community college or master's courses for professional jobs.[5] I drop observations that are missing information necessary to define the year and the state of entry; the detailed sample restrictions are described in the appendix.

[5]This definition follows Neumark (2002), who also used the NLSY to identify the effect of early job mobility on the adult wages.

Finally, 5387 individuals remain in the sample. Table 1 summarizes the number of individuals of each race-gender group and state unemployment rates by the year of labor market entry. About 90% of the sample entered the labor market during the twelve years between 1975 and 1986, the period including the recessions in 1975 and 1982-83.

The variable for the years of schooling is "highest grade completed as of May 1," and that for enrollment is "enrollment status as of May 1." The highest degree/diploma attained, used to calculate the predicted year of entry as the instrumental variable for actual entry, is based on the following two questions "highest degree ever received" and "Does R have high school diploma or equivalent?" and *not* based on the highest grade completed.

The real wages are deflated by the consumer price index. The hourly wages in the NLSY are defined as "hourly rates of pay at current or most recent job," which is missing if the respondents did not work at all since last interview. At the same time, a small fraction of respondents are currently employed but lacking wages for some unknown reason. Table 2 shows the number of observations with valid (positive) hourly wages, missing wages due to non-employment and missing wages for other reasons, for each race-gender-potential experience cell. The proportion with missing wages due to non-employment varies across race and gender: African Americans and women are less likely to have valid wages. Moreover, white women's labor supply patterns over potential experience are opposite that of black women.

Since the expected effect of the unemployment rate at entry on subsequent wages is negative, there might be a negative effect on subsequent employment as well. If the likelihood of having a valid wage were affected by the unemployment rate at entry, it would yields a sample selection bias. Thus, I also examine the effects on employment and the likelihood of having a valid wages in Section 5. In brief, the unemployment rate at entry does not have a statistically significant negative effect on the likelihood of having a valid wage. On the contrary, the unemployment rate at entry has a weak, *positive* effect on employment, although the effect on the likelihood of having wages seems negligible.

The state unemployment rates based on the CPS are available only for 1976-2000. However, dropping individuals who entered the labor market before 1975 may unbalance the composition of birth-year cohorts by making older cohorts more educated on average. Thus, I used the state unemployment rates based on the Unemployment Insurance records for 1973-1975. Since the UI series tend to be lower than the CPS series, I rescaled it, multiplying by the coefficient of the CPS series in a regression of the UI series between 1976 and 1982 on the CPS series of the same year and constant.

4.2 Initial Evidence

Figure 1 plots the average log real hourly wages over potential experience (years since entry) for the four demographic groups: white men, black men, white women and black women. To see how the observed average wage profiles differ across those who entered during booms and recessions, the sample are split into two groups, one with higher and one with lower unemployment rates at entry. First, the upper left panel shows a gradually fading but significantly persistent wage gap for white men, as expected; this is also consistent with existing results by Oreopoulos *et al.* (2006) and by Kahn (2006). For black men, the lower left panel shows a significant wage gap in the first few years, but it disappears sooner than

the gap among white men. This difference in persistence across race seems to fit in with the prediction by theory reviewed in Section 2.

Turning to women, the upper right panel shows a very small gap among white women. The smaller gaps for women than men are consistent with the theories which predict a weak effect on those with weak labor force attachment. Surprisingly, black women who left school during recessions earn more in their fifth year or later. This counter-intuitive positive effect may be a spurious one due to accidental positive selection on educational background, which is shown in Table 3, or a state specific component that happened to correlate with the unemployment rates. Except for this, the observed patterns in Figure 1 agree with the predictions of theory.

Table 3 shows summary statistics of educational background and the AFQT scores by race, gender and whether unemployment rate at entry was higher than the median of the sample. The highest grades completed are almost equal among those who entered the labor market during recessions and booms, except that black women who entered during recessions have slightly more education than black women who entered during booms. The AFQT score is the percentile score of Armed Forces Qualification Test, which was attached to the NLSY in 1981. The AFQT score is often used as a measure of ability in studies using NLSY. Again, black women in the group with high unemployment rates at entry are slightly positively selected. Except for this, there is little difference within each race and gender group. At least, people who left during recessions are not negatively selected on education or pre-market human capital measured by the AFQT.

Table 3 also shows that black people are more likely to have graduated when the unemployment rate was low. This difference mainly comes from the difference in race compositions across states, suggesting that it is essential to control for the entry-state fixed effects. Also, timing and place of entry may be endogenous because people may try to avoid entering the labor market during recessions. The next section describes strategies to deal with these issues.

| 5 | Empirical Strategy

My goal is to identify the effect of the unemployment rate at labor market entry on subsequent wages for each gender and race group. To see the differences in persistence, I also allow the effect of the unemployment rate at entry to vary with potential experience (years since graduation from school).

Since the state unemployment rate may correlate with unobserved characteristics of each state, controlling for the entry-state fixed effects is essential. This is especially crucial for the comparison of blacks and whites, given the considerably different racial composition across states. Also, since less educated people graduated earlier than more educated people in the same birth-year cohort, temporary macro shocks in the year of entry may also bias the estimates. It is important to keep in mind when using the NLSY that the majority of the sample who graduated from school during the recession in 1982-1983 are college graduates who have graduated from high school in the late 1970s. To control for these fixed effects, I include dummy variables for the state of entry and for the year of entry. The identification

of the effect of high unemployment at entry relies on the variations in unemployment rates at the year- and state- level net of those fixed components. Given the strong auto-correlation of unemployment rates, I also need to control for current business cycles so that the estimated coefficients do not pick up cyclical ups and downs.

Specifically, consider the following wage function:

$$\log w_{it\tilde{s}ys} = \beta_{(t-y)} u_{ys} + \gamma_{(t-y)} u_{t\tilde{s}} + \delta' X_{it} + \phi_t + \eta_s + \mu_y + \varepsilon_{itys} \qquad (1)$$

where $w_{it\tilde{s}ys}$ is the real hourly wage of individual i who left school in year y and in state s observed in calendar year t and state $\tilde{s}$, u_{ys} is the unemployment rate of the year and the state of entry into the labor market, $u_{t\tilde{s}}$ is the current unemployment rate in the state of current residence, and X_{it} is other control variables including potential experience, the highest grade completed and the age adjusted AFQT score. Unobserved error components are decomposed into a calendar year fixed effect ϕ_t, an entry-state fixed effect η_s, an entry-year fixed effect μ_y and the remaining error ε_{itys}.

If the sample size were large enough, (1) could be estimated separately for each gender and race. However, as shown in Table A2, some states have too few black people to include entry-state dummies in the blacks-only regressions. Therefore, I employ a parsimonious specification that pools black and whites together. Yet, I run separate regressions for men and women. It is necessary to take into account differences in the wage-experience profiles and sensitivity to aggregate labor market conditions across race groups, since misspecifications of them will contaminate the estimate of β. For example, steeper experience-wage profiles for men leads to an upward bias on the β of older men. Also, the effect of the contemporaneous unemployment rate may well differ across demographic groups and potential experience. Therefore, I allow potential experience to enter non-parametrically as $\alpha_{(t-y),g}$ and let the coefficients of unemployment rate at entry and at present vary across race and with potential experience. Furthermore, to incorporate the difference in racial wage gaps across regions, I include dummy variables for the ten BEA regions interacted with a dummy variable for blacks, θ_{rg}. On the other hand, δ, ϕ_t, η_s and μ_y are left common to both white and black men:

$$\log w_{it\tilde{s}ysg} = \alpha_{(t-y),g} + \beta_{(t-y),g} u_{ys} + \gamma_{(t-y),g} u_{t\tilde{s}} + \delta' X_{it} + \phi_t + \eta_s + \theta_{rg} + \mu_y + \varepsilon_{itysg} \qquad (2)$$

X_{it} here includes a dummy variable for blacks, years of schooling and the age adjusted AFQT score. (2) is estimated for women in the same manner.

A potential source of bias that is still remaining is endogeneity of labor market entry. That is, some people may determine their timing and place of entry so that they can avoid entry during recessions. In theory, a high unemployment rate may increase enrolment to schools for two reasons: the relatively low opportunity cost of not working and avoidance of entering the thin labor market. Concerned mainly with the first aspect, Card and Lemieux (2000) have shown that a temporary rise of the local unemployment rate increases high

school enrollment for age 15-17, although the effect on college enrollment is weak, indecisive for men and slightly negative for women, using the Current Population Survey. A similar concern applies to the geographical mobility. Using the Census data, Wozniak (2006) reports that college graduates tend to move to states with high labor demand when they enter the labor market, although less educated people do not move as often.

To see if the unemployment rate at entry or age 18 affects the year and the state of entry, I estimate several regressions over persons (not multiple observations per person) in the following form:

$$Dept.\text{var}_i = \beta_g u_i + \delta' X_i + FEs + \varepsilon_i \qquad (3)$$

The first column of Table 4 indicates the dependent variables and the second column indicates what u_i stands for: the unemployment rate at actual entry, at age 18 or at predicted entry. The predicted year of graduation is calculated as follows: year of birth + 18 for those without any degree/diploma and those with high school diploma or equivalent, year of birth +20 for those with AA and year of birth + 22 for those with BA, BS, master and doctor. I use the state of residence at age 14 for the state of residence at age 18 and at predicted entry. X_i includes the race dummy and the age adjusted AFQT score, and FEs are the set of fixed effects appropriate for the corresponding u_i (see the notes below Table 4). The dependent variables in rows (2)-(4) are discrete choice indices, and the estimates reported in Table 4 are from the linear probability model for the sake of intuitive grasp. The results from probit model are similar in terms of both statistical significance and marginal effects measured at the mean of the explanatory variables.

Row (1) suggests weak positive effects of the unemployment rates on years of schooling, which is consistent with the findings by Card and Lemieux who use a larger dataset, although the effect of the unemployment rates at age 18 is not statistically significant. The effect on the likelihood of having a college degree (row (2)) is not statistically significant, either, and seems to be economically small as well. Furthermore, rows (3) and (4) show no evidence that people avoid graduating during a recession.

Overall, the correlation between labor market conditions and options upon graduation is very weak for this sample from the NLSY. Yet, this may be due to the small sample size. If the tendency to adjust the timing and the place of entry were correlated with earning ability, it could still cause bias. Therefore, it is worth trying to correct for this endogeneity of entry by instrumenting for u_{ys} with the unemployment rate in the predicted year of graduation based on the highest degree attained (in the same manner as in Table 4) and the state of residence at age 14.

Since the state of residence at age 14 is obviously exogenous in this context, the question is whether the predicted year of entry is exogenous. Admittedly, the decision to attain a degree is a choice and those who proceed to college during a recession might be different from those who proceed to college during a boom. However, from a practical view point, using the unemployment rate at particular age would be difficult. First, the unemployment rates at age 18 and younger do not have strong enough predictive power strong enough to work as an instrument for college graduates. Second, unemployment rates at older ages are

unlikely to be independent of the error term, since the business cycle conditions in the first few years after entry may affect subsequent wages for high school graduates. Further, as shown in Table 4, the probability of having a college degree does not seem to depend on the unemployment rate at age 18. Card and Lemieux (2000) also report that effects of the unemployment rate on the likelihood of proceeding to college for students in the twelfth grade are statistically insignificant and variable in sign. Also, recall that information on college degree is not based on the years of schooling but is directly asked. Thus, there is evidence that the decision to obtain a college degree is not correlated with the labor market conditions.

On the other hand, I assign the year of birth + 18 to high school drop-outs because Card and Lemieux (2000) find positive effects of the unemployment rate on high school completion; likewise I ignore graduate degrees because Kahn (2006) reports that a recession at graduation from college slightly increases the likelihood of attaining a graduate degree. Since the number of high school drop outs and the number of those with graduate degrees are small, the instrument has sufficient explanatory power in the first stage regressions. The details of the first stage regressions are presented in the appendix; in brief, the coefficients of the instrument are around 0.5 and always statistically significant at the 1% level.

| 6 | RESULTS

The first two columns of Table 5 reports $\beta_{(t-y),g}$ in equation (2), the coefficients of the unemployment rate at entry interacted with race and potential experience, estimated by separate OLS regressions for men and women. Not surprisingly, there are negative effects right after entry for all demographic groups. However, the negative effect disappears in 6 years for black men, while the effect remains negative (though statistically insignificant) for white men. Also, for whites, the size of the effect is smaller for women; in fact, the effect is not statistically significant for white women even right after entry. Except that the effect for black women is stronger than that for black men, these observations are basically the same as what Figure 1 suggested.

The last two columns of Table 4 report the estimates by IV in the same way as the OLS results. Here, the unemployment rate in the state and the year of actual entry is instrumented by the unemployment rate that the person would have experienced if she had stayed in the same state since she was fourteen years old and gone straight to her final degree attained. This instrumental variable should correct biases from endogenous choice of timing and place of graduation conditional on the highest degree attained. Note that instrumental variable estimators in general have larger standard errors than the corresponding OLS estimators; in fact the standard errors in the last two columns in Table 5 are much larger than those in the first two columns. The endogeneity of entry seems to bias the OLS estimates towards zero for men, although the standard errors are also boosted.

Table 4 also reports the coefficients of years of schooling and the age adjusted AFQT score for comparison reason. The coefficients of the contemporaneous unemployment rate are not statistically significant, except that those for black men are significantly negative. This confirms that black men are the most affected group by the contemporaneous labor market

conditions. Each set of fixed effects is jointly statistically significant.

To grasp the result more intuitively, Figure 2a and 2b plot the coefficients of the unemployment rate at entry over years since leaving school for each race gender group. Figure 2a compares the OLS estimates of $\beta_{g(t-y)}$ and Figure 2b compares the IV estimates of $\beta_{g(t-y)}$. The gaps between men and women are larger in Figure 2b as the graphs for men shift downward while those for women shift slightly upward from Figure 2a. Yet, the same pattern across race hold for both OLS and IV estimates. These figures confirm the basic observations: the negative effect of high unemployment at entry is stronger for men than for women, and the slope is steeper (i.e. the effect fades faster) for blacks than for whites.

If the unemployment rate at entry has a negative effect on wages, it may well have a negative effect on employment. Therefore, I estimate the effect on employment by replacing the dependent variable in equation (2). Table 6 reports the results from the linear probability model.[6] Surprisingly, the estimated coefficients are mostly *positive*, although statistically insignificant. If those who were not employed were negatively selected on their potential wages, this positive effect would cause upward bias on the estimated effect. However, it is not very plausible that those with low potential wages are more likely to be employed when there is a negative shock to wages at the cohort level. Furthermore, the estimates tend to be larger for women, suggesting that decisions on marriages and fertility may matter. Thus, the direction of potential biases is ambiguous.

In any case, recall that non-employment does not necessarily imply missing wages because respondents who have worked since last interview report wages even if they are not employed in the week of survey. Thus, to examine this issue more directly, table 7 reports the effect on the likelihood of having a valid wage. The estimated coefficients are still positive but on average smaller and noisier. Also, except for the IV estimates of black women, the estimated effects are economically small compared to the effects of years of schooling and the AFQT score. Further, I tried replacing the missing wage data with imputed wages based on several slightly different versions of wage equations with individual fixed effects, and found little differences in the estimated coefficients in the wage regressions. Therefore, I believe that, even if there are biases from non-random selection of those lacking valid wages, they will not alter the qualitative conclusions.

| 7 | CONCLUDING REMARKS

I have found the following facts. First, the initial impact of a recession at entry is stronger for African Americans than for whites, but the effect fades faster. Second, the negative effect is weaker for women than for men. Also, endogeneity of entry seems to cause underestimation of the effect for men. These findings are consistent with the economic theory reviewed in Section 2: the effect of a recession at entry will be stronger for people with stronger labor force attachment and more persistent for more skilled workers who are

[6]The probit estimates are similar to the corresponding OLS estimates in Tables 6 and 7, in both statistical significance and marginal effects evaluated at the mean of explanatory variables.

more likely to be on stable employment contracts.

To put it the other way around, the benefit from a tight labor market at entry lasts only for white men, who are more likely to be employed in a stable job with good training opportunities. Although the lower persistence of the effect of a recession for blacks might sound like an advantage, actually it implies that they are more vulnerable to fluctuations in labor demand later in their career. Likewise, the weaker effect for women suggests that women cannot exploit the benefit from obtaining a good job at an early stage as much as men, presumably because of intermittence for marriage and childbirths.

The observed difference between across races raises a further question. If the difference in the effect of a recession at entry is attributed to differences in skill, a comparison between less and more educated whites will be similar to a comparison between blacks and whites. Although it is hard to implement with the NLSY given its sample size, in another project I have confirmed the same pattern between less educated and more educated whites using the Current Population Survey (Kondo, Genda and Ohta 2007). Therefore, differences in skill are important factor of differential effects of graduating during a recession across races, although the possibility remains that racial discrimination plays an important role as well.

It is difficult to distinguish one theory from the others. Note that skill level matters for both access to stable employment in high-paying jobs and returns to training. In theory, once a worker finds a good job after entry, a recession prior to the job change no longer matters in absence of human capital accumulation through the past on-the-job training. Therefore, controlling for labor demand at the beginning of the current job will cancel out the effect of a recession at entry if the loss of training opportunity does not matter. However, it is very difficult to simultaneously include unemployment rates in the year of leaving school and in the year of beginning of the current job, because they are highly collinear. Since this colinearity amplifies bias caused by misspecification of the effect of tenure and experience, it is necessary to include many interaction terms in the regression. The required sample size is far larger than the NLSY; this remains for future work.

APPENDICES

A: Sample Restrictions and the Definition of Entry

On the course of defining year of entry, I drop people who dropped out from the survey before turning to 30 years old. Then, I drop people with less than 6 years of schooling. The year of entry cannot be defined for people who are still in school at age 30, and those who never go to school since 1979 and miss the year of last enrollment before 1979. To merge with state unemployment rate, people have to live in the contiguous United States in the year of entry. I used state of residence at age 14 as the state of residence at entry for those entered before 1979. I also delete people who left school before getting 16 years old and those who have ever served in the army sometime within 12 years since leaving school. Lastly, I delete people whose state of residence at entry is too small to control state fixed effects. The number of people deleted and remaining is summarized in Table A1, and the number of individuals and the average unemployment rate by the state of entry is shown in Table A2.

B. First Stages

Table B1 is cross tabulation of age of actual and predicted entry. Many people actually enter the labor market several years after their predicted entry based on the highest degree ever received. This is partly because some people leave colleges without receiving degrees: about 24% of those with only high school diploma report one or more years in college in answer to the question on the highest grade completed. Also, about 20% of those assigned 22 as the age of predicted entry have graduate degrees. Other people simply did not go straight to their final degree. Even among those who completed exact the twelfth grade, about 10% have not finished schooling by the age of 20.

For the state, the proportion of people who had moved across states since the age of 14 when they entered the labor market in each race-gender group is 13.6 % for white men, 9.6% of black men, 14.5% of white women and 11.6% of black women.

To see the correlation between the unemployment rates at actual entry and at predicted entry defined in Section 4, I estimate the following collapsed regression:

$$u_{ys} = \beta_g u_{\tilde{y}\tilde{s}} + \delta' X_i + \eta_{\tilde{s}} + \theta_{rg} + \mu_{\tilde{y}} + \varepsilon_i \qquad (5)$$

where $\tilde{y}$ is the year of predicted entry and $\tilde{s}$ is the state of residence at age 14. Equation (5) is estimated separately for men and women. β_g varies across race. X_i includes the race dummy, the age adjusted AFQT score and years of schooling. $\eta_{\tilde{s}}$, θ_{rg}, $\mu_{\tilde{y}}$ are the fixed effects corresponding to the subscripts. The dataset consists of persons, not multiple observations per person. The upper panel of Table B2 reports estimated β_g: they are in the range of 0.37-0.51, and t-statistics are about 10. The correlation between the endogenous variable and the instrument is strong enough for all four race-gender groups.

The lower panel of Table B2 reports the coefficient of the instrument in the actual first stage regressions common to the IV estimates reported in Tables 5-7. Since the unemployment rate at entry is interacted with the four race-gender groups and the four 3-year potential experience brackets, there are sixteen pairs of endogenous variables and corresponding instruments. Therefore, there are sixteen separate first stage regressions. Table B2 reports only the coefficient of the instrument corresponding to the endogenous variable in each regression. For example, the upper right cell means that a unit increase of the instrument is associated to increase of the unemployment rate at actual entry by 0.454 for white men with 1-3 years of potential experience. The differences over potential experience are probably due to change in the composition of the sample caused by entry before 1979, which affect the group with 1-3 years of experience the most.

REFERENCES

Altonji, J.G. and R. M. Blank. 1999. Race and Gender in the Labor Market. Handbook of Labor Economics vol.3, Ch48.

Baker, G., M. Gibbs and B. Holmstrom. 1994a. The Internal Economics of the Firm: Evidence from Personnel Data. Quarterly Journal of Economics, 109, 881-919.

Baker, G., M. Gibbs and B. Holmstrom. 1994b. The Wage Policy of a Firm. Quarterly Journal of Economics, 109, 921-955.

Beaudry, P. and J. DiNardo. 1991. The Effect of Implicit Contracts on the Movement of Wages over the Business Cycle: Evidence form Micro Data. Journal of Political Economy, 99 (4) 665-688.

Bertrand, M. 2004. From the Invisible Handshake to the Invisible Hand? How Import Competition Changes the Employment Relationship. Journal of Labor Economics, 22 (4), 723-765.

Bowlus, A. J. 1995. Matching Workers and Jobs: Cyclical Fluctuations in Match Quality. Journal of Labor Economics, 13, 335-350.

Card, D. and T. Lemieux. 2000. Dropout and Enrollment Trends in the Post-War Period: What Went Wrong in the 1970s?. Working Paper no. 7658, National Bureau of Economic Research, Cambridge, MA.

Clark, K. and L. H. Summers. 1981. Demographic Differences in Cyclical Employment Variation. The Journal of Human Resources, 16 (1), 61-79.

Cutler, D. M. and L. Katz. 1993. Macroeconomic Performance of the Disadvantaged. Brookings Papers on Economic Activity part 2, 285-239.

Devereux, P. J. 2004. Cyclical Quality Adjustment in the Labor Market. Southern Economic Journal, 70 (3), 600-615.

Doeringer, P. and M. Piore. 1971. Internal Labor Markets and Manpower Analysis. Lexington, MA: D. C. Heath and Company.

Grant, D. 2003. The Effect of Implicit Contracts on the Movement of Wages over the Business Cycle: Evidence from the National Longitudinal Surveys. Industrial and Labor Relations Review, 56 (3).

Hynes Jr., J. R, H. Hoines and A. B. Krueger. 2001. Another look at whether a rising tide lifts all boats. Working Paper, no. 8412, National Bureau of Economic Research, Cambridge, MA.

Kahn, L. B. 2006. The Long-term Labor Market Consequences of Graduating College in a Bad Economy. Unpublished manuscript, Harvard University, Cambridge, MA.

Kondo, A., Y. Genda and S. Ohta. 2007. Long term effects of a recession at entry in Japan and the United States. Unpublished manuscript, Columbia University, New York.

Neumark, D. 2002. Youth Labor Markets in the United States: Shopping around vs. Staying put. The Review of Economics and Statistics, 84 (3) 462-482.

Okun, A. 1973. Upward Mobility in a High-pressure Economy. Brookings Papers on Economic Activity, 1973(1), 207-261.

Oreopoulos, P., T. von Wachter and A. Heisz. 2006. The Permanent and Transitory Effects of Graduating in a Recession: An Analysis of Earnings and Job Mobility using Matched Employer-Employee Data. Working Paper no. 12159, National Bureau of Economic Research, Cambridge, MA.

Rosen, S. 1972. Learning and Experience in the Labor Market. Journal of Human Resources 7(3).

Welch, F. 1979. Effects of Cohort Size on Earnings: The Baby Boom Babies' Financial Bust. The Journal of Political Economy, 87(5), S65-S97.

Wozniak, A. 2006. Educational Differences in the Migration Responses of Young Workers to Local Labor Market Conditions. Discussion paper No. 1954, Institute for the Study of Labor, Bonn, Germany.

Table 1: Number of Individuals and Mean Unemployment Rate by Year of Entry

Year of entry	U. rates	White men	White women	Black men	Black women	Total
1973	4.88	3	5	3	1	12
1974	5.97	12	16	9	10	47
1975	10.90	57	79	35	41	212
1976	7.39	86	116	46	48	296
1977	6.84	103	138	62	62	365
1978	6.03	116	133	76	101	426
1979	5.81	155	202	94	117	568
1980	7.41	151	183	103	121	558
1981	7.89	174	185	87	125	571
1982	10.26	155	182	94	132	563
1983	10.12	124	126	61	83	394
1984	7.61	97	110	56	66	329
1985	7.34	92	95	34	49	270
1986	7.15	81	98	28	34	241
1987	6.43	55	65	16	34	170
1988	5.67	34	41	9	17	101
1989	5.27	31	25	9	15	80
1990	5.96	16	23	8	14	61
1991	6.83	15	17	3	9	44
1992	7.61	9	8	4	6	27
1993	6.66	3	7	3	0	13
Total	7.64	1,569	1,854	840	1,085	5,348

Table 2: Number of Observations with and without Valid Wages (% of Total Number of Observations in the Corresponding Race-Sex-Potential Experience Category)

	Exp.	White men	Black men	White women	Black women
With valid wages	1-3	3,715	1,666	3,962	1,853
		(88.7%)	(75.4%)	(81.6%)	(63.3%)
	4-6	4,248	2,101	4,418	2,220
		(93.2%)	(85.4%)	(81.8%)	(70.0%)
	7-9	4,110	2,066	4,147	2,288
		(94.7%)	(86.2%)	(80.6%)	(74.8%)
	10-12	3,603	1,824	3,588	2,054
		(94.5%)	(84.8%)	(78.0%)	(74.7%)
Missing wages due to non-employment	1-3	330	466	744	1,002
		(7.9%)	(21.1%)	(15.3%)	(34.2%)
	4-6	171	293	820	885
		(3.8%)	(11.9%)	(15.2%)	(27.9%)
	7-9	97	258	854	716
		(2.2%)	(10.8%)	(16.6%)	(23.4%)
	10-12	81	255	872	629
		(2.1%)	(11.9%)	(19.0%)	(22.9%)
Missing wages though being employed	1-3	145	78	151	71
		(3.5%)	(3.5%)	(3.1%)	(2.4%)
	4-6	140	65	161	65
		(3.1%)	(2.6%)	(3.0%)	(2.1%)
	7-9	131	74	145	55
		(3.0%)	(3.1%)	(2.8%)	(1.8%)
	10-12	130	72	140	68
		(3.4%)	(3.3%)	(3.0%)	(2.5%)

Figure 1: Average wage profiles (no control for state/year)

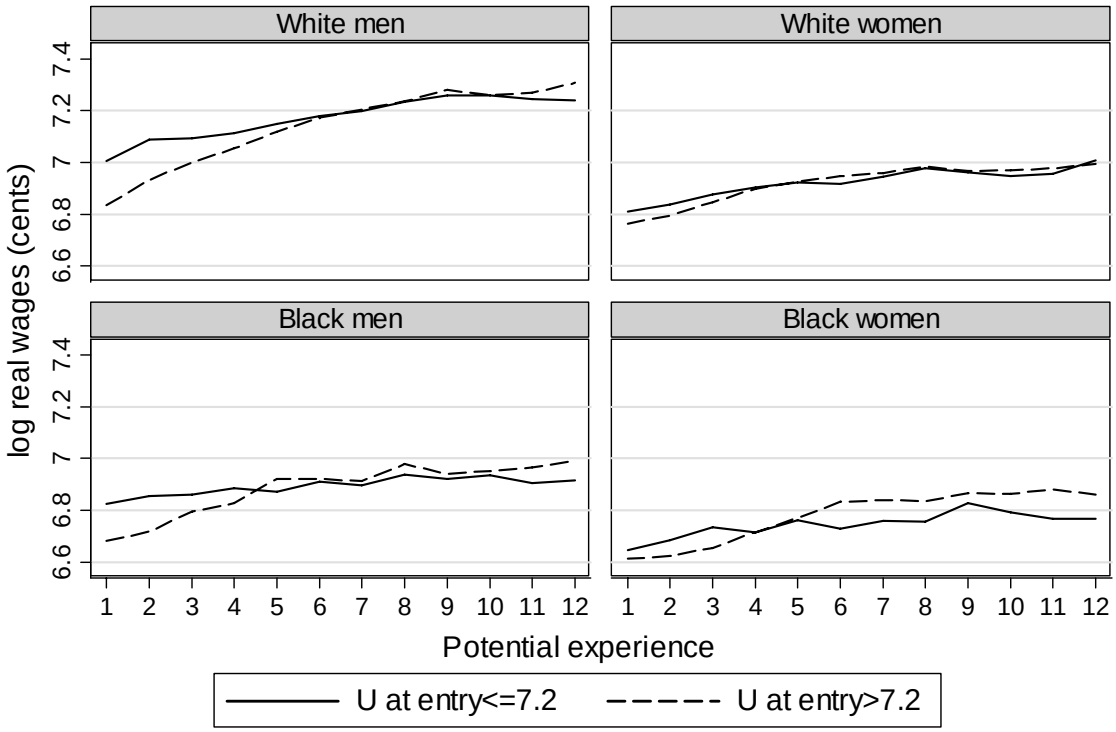

Table 3: Mean of Selected Predetermined Variables by Unemployment Rate at Entry

		White men	Black men	White women	Black women
Highest grades completed	U<=7.2	13.4	12.4	13.4	12.7
	U>7.2	13.4	12.3	13.4	13.0
Age adjusted AFQT score	U<=7.2	0.25	-0.93	0.16	-0.93
	U>7.2	0.27	-0.95	0.19	-0.82
# individuals	U<=7.2	778	442	938	593
	U>7.2	791	398	916	492

Note: 7.2% is the median of the unemployment rates at entry into the labor market among the entire sample (based on number of individuals).

Table 4: Correlation with Unemployment Rates at Entry or Age 18 and Options Upon Graduation

Dependent var.	U at	White men	Black men	White women	Black women
(1) Years of schooling	Actual entry	0.024	-0.009	0.059	0.062
		[0.027]	[0.033]	[0.026]**	[0.030]**
	Age 18	0.047	0.027	-0.006	0.008
		[0.045]	[0.047]	[0.039]	[0.040]
(2) Having a college degree (AA, BA, BS)	Age 18	0.013	0.007	-0.003	-0.006
		[0.009]	[0.009]	[0.009]	[0.009]
(3) Actual entry is later than predicted entry	Predicted entry	0.011	0.009	0.013	0.021
		[0.014]	[0.014]	[0.014]	[0.015]
(4) Having changed state of residence	Actual entry	-0.003	0.004	-0.004	-0.002
		[0.006]	[0.007]	[0.005]	[0.006]

Note: OLS regressions (the linear probability model for rows (2)-(4)). The regressions on the unemployment rate at actual entry include dummies for year of entry and state of residence at entry. The regressions on the unemployment rate at age 18 include dummies for year of birth and state of residence at age 14. The regression on the unemployment rate at predicted entry includes dummies for year of predicted entry and state of residence at age 14. Standard errors are in brackets, clustered by the relevant pair of the year and the state. ***, **, * indicate statistically significant in 1%, 5% and 10%.

Table 5: The Effect of the Unemployment Rate at Entry on Wages

	OLS		IV	
	Men	Women	Men	Women
U at entry				
White exp = 1-3	-0.016 [0.006]**	-0.008 [0.006]	-0.030 [0.016]*	0.006 [0.015]
White exp = 4-6	-0.011 [0.006]*	-0.006 [0.006]	-0.027 [0.011]**	-0.002 [0.011]
White exp = 7-9	-0.008 [0.006]	-0.003 [0.006]	-0.019 [0.012]	-0.002 [0.011]
White exp = 10-12	-0.007 [0.006]	-0.005 [0.007]	-0.029 [0.012]**	-0.004 [0.014]
Black exp = 1-3	-0.019 [0.008]**	-0.025 [0.007]***	-0.051 [0.023]**	-0.034 [0.028]
Black exp = 4-6	-0.013 [0.009]	-0.014 [0.007]**	-0.049 [0.019]**	-0.015 [0.019]
Black exp = 7-9	0.000 [0.009]	-0.007 [0.007]	-0.034 [0.020]*	0.003 [0.018]
Black exp = 10-12	0.008 [0.009]	-0.001 [0.007]	-0.028 [0.023]	0.024 [0.018]
Highest grade completed	0.062 [0.005]***	0.080 [0.004]***	0.057 [0.005]***	0.077 [0.005]***
Age adjusted AFQT score	0.104 [0.010]***	0.105 [0.010]***	0.108 [0.010]***	0.108 [0.011]***
Observations	22947	24233	22333	23404
R-squared	0.30	0.30	0.30	0.29

Note: The dependent variable is log real hourly wages (cents). Constant, current unemployment rates interacted with race and the three-year experience brackets, a race dummy, non-linear controls for potential experience differentiated by race are also included in the regressions though omitted from the table. OLS regressions include dummies for year of entry, state of residence at entry and year of the survey; IV regressions include dummies for year of predicted entry, state of residence at age 14 and year of the survey. Standard errors are in brackets, clustered by the year and the state of entry (predicted entry for IV). ***, **, * indicate statistically significant in 1%, 5% and 10%.

Figure 2a: Coeffcients of U at entry; OLS
(Taken from Table 5)

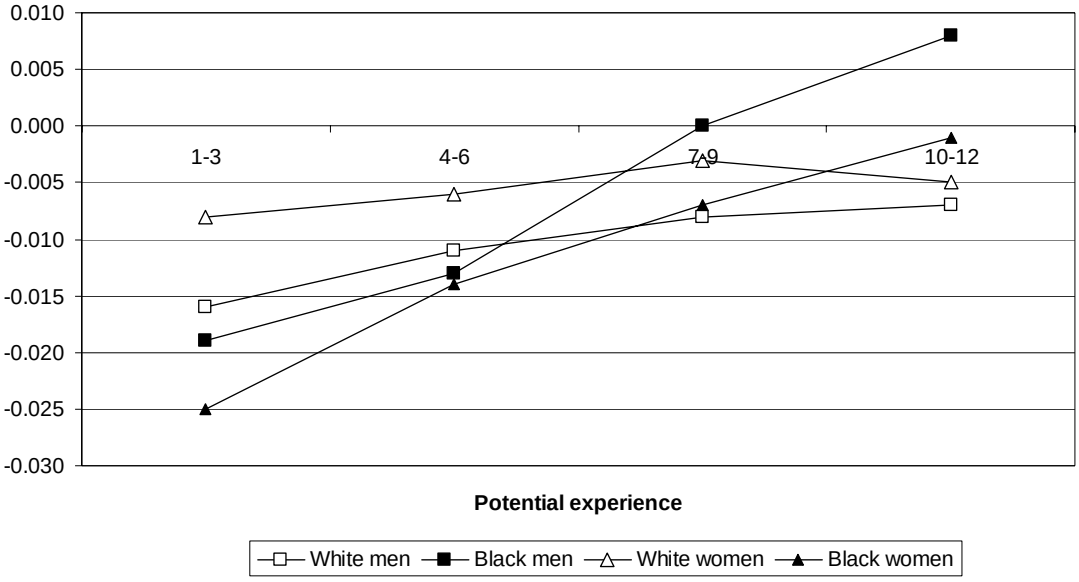

Figure 2b: Coeffcients of U at entry; IV
(Taken from Table 5)

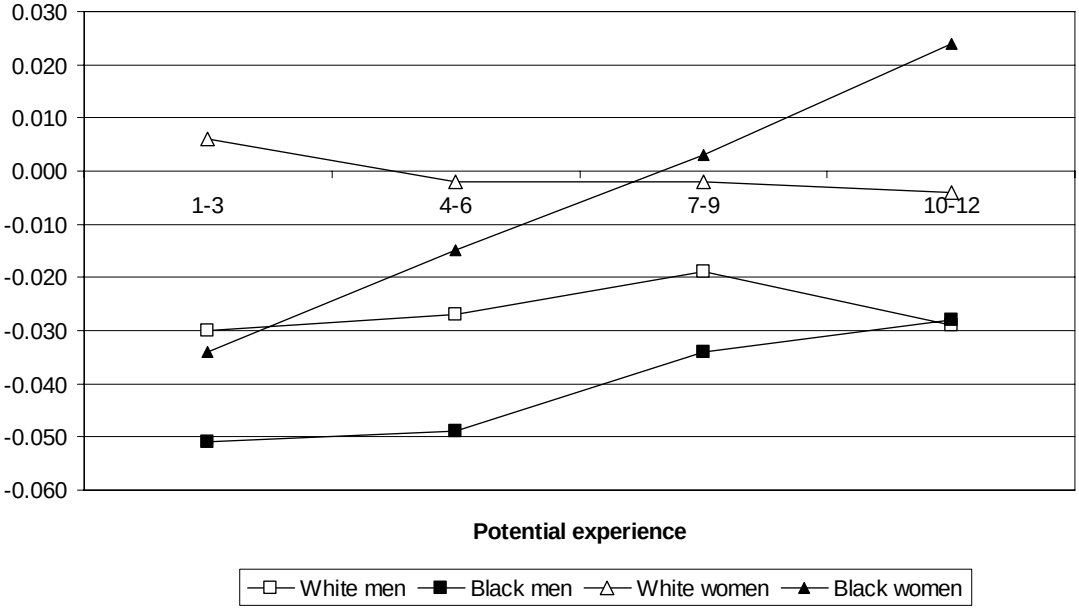

Table 6: The Effect of the Unemployment Rate at Entry on Employment

	OLS		IV	
	Men	Women	Men	Women
U at entry				
White exp = 1-3	0.000 [0.005]	0.009 [0.004]*	-0.002 [0.011]	0.003 [0.012]
White exp = 4-6	0.000 [0.004]	0.002 [0.004]	0.002 [0.007]	0.004 [0.009]
White exp = 7-9	0.003 [0.004]	0.005 [0.005]	0.006 [0.007]	0.007 [0.009]
White exp = 10-12	0.004 [0.004]	0.005 [0.005]	0.001 [0.008]	0.008 [0.010]
Black exp = 1-3	0.002 [0.007]	0.001 [0.007]	-0.003 [0.017]	0.033 [0.022]
Black exp = 4-6	0.005 [0.006]	-0.001 [0.007]	0.005 [0.013]	0.010 [0.015]
Black exp = 7-9	0.010 [0.006]	0.005 [0.006]	0.003 [0.014]	-0.003 [0.016]
Black exp = 10-12	0.012 [0.006]*	0.011 [0.007]	-0.008 [0.015]	-0.011 [0.016]
Highest grade completed	0.023 [0.002]***	0.045 [0.003]***	0.023 [0.003]***	0.047 [0.003]***
Age adjusted AFQT score	0.035 [0.006]***	0.043 [0.008]***	0.037 [0.006]***	0.051 [0.008]***
Observations	25688	31460	24992	30430
R-squared	0.13	0.13	0.13	0.13

Notes: The dependent variable takes 1 if the respondent is employed. Linear probability model with the same right hand side variables as the OLS regressions in Table 5; see notes for Table 5.

Table 7: The Effect of the Unemployment Rate at Entry on Likelihood of Having a Valid Wage

	OLS		IV	
	Men	Women	Men	Women
U at entry				
White exp = 1-3	0.002 [0.004]	0.004 [0.004]	0.008 [0.009]	-0.002 [0.011]
White exp = 4-6	0.000 [0.003]	0.005 [0.004]	0.004 [0.006]	0.007 [0.008]
White exp = 7-9	0.002 [0.003]	0.007 [0.004]*	0.005 [0.006]	0.009 [0.008]
White exp = 10-12	-0.001 [0.003]	0.006 [0.005]	0.004 [0.006]	0.007 [0.009]
Black exp = 1-3	0.000 [0.006]	0.009 [0.007]	-0.01 [0.016]	0.047 [0.021]**
Black exp = 4-6	0.002 [0.005]	0.009 [0.007]	0.002 [0.011]	0.016 [0.014]
Black exp = 7-9	0.008 [0.005]	0.008 [0.006]	0.000 [0.011]	0.008 [0.014]
Black exp = 10-12	0.004 [0.005]	0.008 [0.006]	-0.015 [0.012]	-0.015 [0.015]
Highest grade completed	0.007 [0.002]***	0.034 [0.003]***	0.007 [0.003]***	0.035 [0.003]***
Age adjusted AFQT score	0.026 [0.005]***	0.035 [0.007]***	0.026 [0.005]***	0.041 [0.008]***
Observations	25,688	31,460	24,992	30,430
R-squared	0.09	0.13	0.09	0.13

Notes: The dependent variable takes 1 if the observation has a valid (positive) wage. Linear probability model with the same right hand side variables as the OLS regressions in Table 5; see notes for Table 5.

Table A1: Sample Restrictions and Sample Size (Number of Individuals)

	Deleted	Remaining
Original NLSY 79 sample		12,686
Keep cross sectional white and black and supplemental black samples only; delete Hispanics, poor white and military sample	4,847	7,839
Dropped out from the survey before age 30	1,837	6,002
Highest grade completed at age 30 < 6	6	5,996
Enrolled at age 30 (except for two year college)	125	5,871
Not enrolled in 79 and missing last enrollment	46	5,825
State of residence at entry is missing	90	5,735
Age at entry <16	50	5,685
Served in the army sometime within 12 years since leaving school	293	5,392
State of residence at entry was too small (w/ <20 people in sample)	44	5,348

Table A2: Sample Size by State of Residence at Entry (Number of Individuals)

State	Mean U	Total	Black men	Black women	White men	White women
Alabama	8.97	234	52	70	60	52
Alaska	9.53	22	0	0	10	12
Arizona	6.75	46	4	2	20	20
Arkansas	7.91	80	20	37	9	14
California	7.60	359	48	56	113	142
Colorado	6.15	87	0	0	41	46
Connecticut	5.80	106	8	8	36	54
DC	8.08	48	23	22	2	1
Florida	7.16	201	21	51	53	76
Georgia	6.78	255	106	104	21	24
Illinois	8.11	190	30	36	67	57
Indiana	8.05	96	11	5	42	38
Iowa	6.39	55	0	0	27	28
Kansas	4.55	40	4	5	17	14
Louisiana	8.85	57	14	24	7	12
Maryland	6.10	61	6	14	17	24
Massachusetts	6.31	78	10	7	24	37
Michigan	11.04	357	32	42	134	149
Minnesota	5.94	129	0	1	72	56
Mississippi	9.36	54	18	32	2	2
Missouri	6.79	145	34	35	30	46
Montana	6.97	49	0	2	19	28
New jersey	7.56	200	24	34	65	77
New York	7.71	296	63	64	91	78
North Carolina	6.57	241	51	61	57	72
Ohio	8.53	379	42	68	126	143
Oklahoma	5.82	56	16	16	13	11
Oregon	8.36	28	0	1	10	17
Pennsylvania	8.48	258	27	36	82	113
South Carolina	7.23	193	59	83	14	37
Tennessee	8.19	94	7	15	33	39
Texas	6.17	287	59	84	56	88
Virginia	5.54	162	36	41	42	43
Washington	8.55	56	0	4	23	29
West Virginia	9.98	122	4	11	42	65
Wisconsin	6.75	227	11	14	92	110

Note: States with less than 20 people (deleted): Delaware, Hawaii, Idaho, Kentucky, Maine, Nebraska, Nevada, New Hampshire, New Mexico, North Dakota, Rhode Island, South Dakota, Utah, Vermont and Wyoming.

Table B1: Age of Predicted Entry and Actual Entry

	White men			White women		
	Age of predicted entry			Age of predicted entry		
Age of actual entry	18	20	22	18	20	22
16	54	0	0	68	0	0
17	61	1	0	99	0	0
18	368	7	1	474	9	3
19	244	6	2	211	16	4
20	82	17	3	97	26	11
21	45	15	13	36	24	19
22	31	9	125	33	14	189
23	29	11	105	27	21	83
24	18	9	57	25	12	46
25-30	63	20	129	88	35	114
Total	995	95	435	1,158	157	469
	Black men			Black women		
	Age of predicted entry			Age of predicted entry		
Age of actual entry	18	20	22	18	20	22
16	28	0	0	31	0	0
17	67	0	0	58	1	0
18	172	4	1	241	2	0
19	188	3	0	176	6	0
20	96	5	2	88	16	4
21	39	7	2	80	16	13
22	28	7	22	46	12	36
23	13	5	24	29	9	30
24	15	2	17	21	5	12
25-30	33	7	28	64	21	39
Total	679	40	96	834	88	134

Table B2: Coefficients of Instruments in the First Stage Regressions

	White men	Black men	White women	Black women
Collapsed regressions	0.503	0.455	0.508	0.378
	[0.042]***	[0.054]***	[0.043]***	[0.047]***
Actual first stage (16 separate regressions)				
Potential Exp = 1-3	0.454	0.441	0.455	0.316
	[0.006]***	[0.006]***	[0.006]***	[0.006]***
Potential Exp = 4-6	0.574	0.523	0.589	0.404
	[0.007]***	[0.007]***	[0.006]***	[0.006]***
Potential Exp = 7-9	0.562	0.508	0.588	0.428
	[0.007]***	[0.007]***	[0.006]***	[0.006]***
Potential Exp = 10-12	0.557	0.534	0.589	0.454
	[0.006]***	[0.006]***	[0.006]***	[0.006]***

Note: See Appendix B for the detail. Standard errors are in brackets, clustered by the year of predicted entry and the state of residence at age 14. ***, **, * indicate statistically significant in 1%, 5% and 10%.

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