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Wage Inequality and the Rise in Labor Force Exit: The Case of US Prime-Age Men

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Abstract:
This article offers the first empirical evidence that labor force exit rates rise when workers' relative earnings fall. The model takes into account that a job not only provides economic security but also affirms a worker's social status, which is tied to their relative position in the labor market. Based on the results, the decline in relative earnings for non-college prime-age men over the last four decades is estimated to have raised their labor force exit propensity by 0.49 percentage point, accounting for 44 percent of the total growth in their labor force exit rate during this period.

JEL Classifications: J21, J24, J31
Keywords: labor force exit rates, prime-age male labor force participation, relative earnings, wage inequality

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Reflecting a decline that began more than a half-century ago, the United States had the fourth-lowest prime-age (25 to 54) male labor force participation rate among the 38 Organisation for Economic Co-operation and Development (OECD) countries in 2019.[1] Prime-age men traditionally constitute the backbone of the US workforce. The decline thus carries direct implications for both the economy and the well-being of individual workers. However, despite the significant scholarly interest in the causes of this decline, the literature offers limited explanations for it. Neither changes in real wages nor growth in alternative income sources explain the persistence and pervasiveness of the decline. As pointed out by Elsby et al. (2019), a closer look at the data reveals that more frequent labor force exit among men without a four-year college degree (hereafter non-college men) has been driving the observed participation decline.[2] At the same time, the educational and occupational gradients in earnings have gradually widened in the US labor market, and as a result, non-college men's earnings measured as a share of the average earnings of all prime-age workers (hereafter relative earnings) had fallen steadily by more than 30 percent by 2019, with considerable variation across occupation groups.

In light of the parallel trends, this paper investigates whether prime-age non-college men are more inclined to leave the labor force when their expected earnings fall relative to the earnings of other workers in their labor market. The empirical model takes into account that a job not only provides economic security but also affirms a worker's social status, which is tied to their position relative to their age-range peers. The model uses a difference-in-differences framework and for identification relies on the variation in non-college men's relative earnings across states and occupations over time. The sample consists of state-occupation-level panel data on labor force exit rates, occupation earnings, and job loss risk matched with information on the state-level earnings distribution and a set of state socioeconomic controls over the period 1980 through 2019.

The results from the ordinary least squares (OLS) regression analysis show that

[1] Based on OECD official statistics.

[2] In this study, labor force exit is defined as the transition from being employed or searching for a job to neither working nor job searching.

labor force exit rates decline with a worker group's expected earnings but increase with their reference earnings, defined as the average earnings in a state across all prime-age workers. According to the estimate, a 10 percent growth in expected earnings has an associated 0.12 percentage point decrease in the exit rate. Contrarily, a 10 percent growth in reference earnings has an associated 0.13 percentage point increase in the exit rate, fully discounting the earnings effect. These coefficients offer suggestive evidence that non-college men's labor market exit behavior is tied to the relative values of their earnings. Over the course of the study period, non-college men's relative earnings declined 30 percent on average. Based on the estimates, this decline in relative earnings had an associated 49 percentage point increase in the exit rate, accounting for 44 percent of the total growth in the exit rate among non-college men over this period. In contrast, changes in real earnings alone account for only 18 percent of the total growth in exit rate.

The relationship between relative earnings and labor force exit rates varies by worker demographics. Across races and ethnicities, a correlation between labor force exit rate and relative earnings is identified among non-Hispanic White men but absent among non-Hispanic Black and Hispanic men. Across age groups, the magnitude of the correlation between exit rate and relative earnings declines with age, with the strongest correlation observed among men aged 25 to 34. The correlation between relative earnings and exit rate also weakens when the sample is stratified by a worker's marital status. These findings suggest that a decline in social status among White men and marriage market sorting are potential channels through which relative earnings affect men's labor force exit decisions. A supplementary analysis further reveals that the labor force exit rate of non-college men increases with both the level and the skewness of the reference group's earnings distribution. Conditional on the state median earnings not increasing or decreasing, non-college men are more likely to leave the labor force when the top earners in a state make disproportionately more than the other workers, showing further evidence that earnings inequity may undermine non-college men's work incentives.

Because the OLS method is subject to the influence of selection and unobserved labor demand and supply shocks, it likely creates biases for both the earnings and the reference earnings coefficients. To assess if the observed correlations are driven by such biases, I leverage two sets of instrumental variables (IVs) that exploit changes in national pay levels to predict the state-level changes in occupation earnings and reference earnings. Specifically, occupation earnings in a state are instrumented by the median occupation earnings in other US states, and reference earnings are instrumented by a cubic function of the inner product between a state's initial 1980–1984 occupation employment shares and the average earnings of each occupation in other US states. By using changes in the national pay level alone to predict state pay levels, the instruments diminish the influence of unobserved state-specific labor market shocks on the observed earnings. The result under the 2SLS method remains consistent with the OLS result.

By leveraging the panel nature of the data and controlling for possible confounding factors, the analysis eliminates unobserved time trend, selection, local labor market shocks, cost of living, and changes in job loss risks in driving the systematic correlation between relative earnings and labor force exit rates. The finding offers evidence that the decline in relative earnings is a plausible key factor behind the trend decline in US non-college men's labor force participation. The result is also consistent with prior studies that find direct linkages between relative earnings and workers' job satisfaction, marriage market return, productivity, and health, all of which are possible mechanisms through which relative earnings influence men's labor force exit behavior.

This article contributes to two areas of the literature. First, this is the first study to examine the relationship between relative earnings and workers' labor force participation behavior. The finding offers a possible explanation for the persistent decline in US male labor force participation, which is not fully explained by the existing literature. Earlier studies attribute the long-term decline in the US prime-age male labor force participation rate to the falling real wages of lower-skilled men (Juhn et al., 1991; Juhn, 1992). However, the parallel trends between real wages and participation paused in the 1990s, when real wages rebounded but labor force participation continued to fall, creating a puzzle in the literature (Juhn et al., 2002; Binder and Bound, 2019). Similar patterns were observed in the 2010s, when labor force participation endured a long-term decline despite a strong recovery in real wages. Contrary to real wages, non-college men's relative wages have followed a consistent downward trend over the last 40 years as a result of the rise in the college wage premium and the overall wage dispersion. While a couple of studies acknowledge the overlapping trends in wage dispersion and the fall in prime-age male labor force participation, surprisingly, there has been little investigation into the theoretical or empirical relationship between the two. Council of Economic Advisers (2016), the only other paper to examine this connection, shows that state-level labor participation is negatively correlated with measures of wage inequality. Supplementing the suggestive evidence in Council of Economic Advisers (2016), this paper connects the growth in wage dispersion and the decline in male labor force participation by introducing relative earnings into non-college men's labor force exit decisions. The finding shows a negative correlation between relative earnings and non-college men's labor force exit rate, resolving the earlier contradictory evidence around wages.

On the labor demand side, in addition to the studies focused on wages, there is a sizable literature that connects the decline in manufacturing employment to the rising import competition with China (Bernard et al., 2006; Autor et al., 2013; Pierce and Schott, 2016; Acemoglu et al., 2016). Autor et al. (2013) and Acemoglu et al. (2016) estimate that the US employment-to-population rate dropped 1.1 to 1.2 percent as a result of Chinese import competition over the 1990–2007/2011 period, accounting for 18 percent of the male labor force participation decline during this period.[3] Yet, the employment effect from Chinese import competition appears to have halted after 2007, while the downward trend in men's labor force participation continued (Bloom et al., 2019). Another demand-side factor that has garnered much public interest is whether jobs have been replaced by new technologies, such as computers and robots. The few academic studies directly testing this hypothesis generally conclude that technology has played a minor role in the aggregate employment decline (Autor et al., 2015; Acemoglu and Restrepo, 2020; Abraham

[3]Following the estimates by Autor et al. (2013), here I assume the employment effect was the same for men and women and that 70 percent of the employment loss led to labor force exits.

and Kearney, 2020). These findings suggest that a sizable participation gap remains unexplained by changes in labor demand. Beyond the demand-side explanations, researchers have also examined whether the participation decline has been due to changes in men's labor supply. Factors such as increased spousal income (Juhn and Potter, 2006; Tüzemen, 2018) or public transfers (Bound, 1989; Autor and Duggan, 2003; Von Wachter et al., 2011; Maestas et al., 2013; Council of Economic Advisers, 2016; Abraham and Kearney, 2020; Charles et al., 2019), changes in drug use (Krueger, 2017; Aliprantis et al., 2019; Harris et al., 2020; Currie et al., 2019; Charles et al., 2019) or time use preferences (Aguiar et al., 2021), and the growing incidence of criminal records (Pager, 2003; Holzer et al., 2006; Pager et al., 2009; Raphael, Raphael; Mueller-Smith, 2015; Sakala, 2014; Carson, 2020) could all potentially reduce work incentives or raise work barriers, leading to lower participation rates. Compared with studies of demand-side factors, investigations of labor supply-side explanations generally find more modest or mixed evidence.[4]

This study's second contribution lies in its extension of the existing literature on relative earnings. Using randomized controlled experiments, recent studies find a worker's perceived relative earnings increase their job satisfaction (Card et al., 2012), productivity (Breza et al., 2018), and, in a laboratory setting, labor supply (Bracha et al., 2015). The causal evidence corroborates earlier findings on the systematic correlations between relative earnings and a worker's job satisfaction (Clark and Oswald, 1996), subjective well-being (Luttmer, 2005; Solnick and Hemenway, 1998), marriage rate (Watson and McLanahan, 2011), women's employment (Neumark and Postlewaite, 1998), health status (Marmot, 2006), and mortality (Eibner and Evans, 2005; Daly et al., 2013). These results offer consistent evidence that a worker's subjective well-being is partially tied to their relative earnings compared with their peers'. This relationship stands in contrast to a traditional neoclassical labor supply model in which the utility from work depends only on the real returns. Yet, to date the literature offers little insight into the effect of relative earnings on the labor supply decisions of workers in a real-world setting. Extending this literature,

[4]Interested readers can also refer to Abraham and Kearney (2020) for a thorough review of the current evidence on the aggregate employment decline in the United States.

this paper presents evidence that the labor force exit rate declines with non-college men's relative earnings, suggesting that relative earnings carry macroeconomic implications beyond individual workers' job satisfaction and well-being.

The paper proceeds as follows. Section 1 describes the trend decline in non-college men's relative labor market status over the last 40 years and the inverse relationship between relative earnings and labor force exit rates. Section 2 outlines the theoretical framework and empirical estimation strategy. Section 3 summarizes the regression variables and their data sources. Section 4 reports the estimation results and discusses the finding. Section 5 concludes.

1 Background

1.1 The Decline in Less Educated Men's Labor Market Status

From 1980 to 2019, the US prime-age non-college men's labor market return steadily declined relative to that of college-educated workers. During this period, non-college men's median weekly earnings fell 17 percent, while college-educated men experienced a 20 percent earnings gain, increasing the weekly earnings gap nearly 150 percent from $280 to $687. In addition, non-college men's earnings premium over women's quickly dissipated as gender equity improved. Over this period, the median earnings of women grew 32 percent, significantly outpacing the growth rate of men's earnings. Most of the growth was from college-educated women, whose earnings increased a stellar 39 percent. An extensive literature attributes this change to the skill-biased technological change, which disproportionately favors high-skilled workers over low-skilled workers (Katz and Murphy, 1992; Bound and Johnson, 1992; Goldin and Katz, 2009). Other factors behind the unequal earnings growth include increased impact of import competition from low-wage countries (Autor et al., 2014, 2013; Ebenstein et al., 2014; Feenstra, 2008) and, relatedly, the waning influence of labor unions (Card, 1998; Hirsch, 2008; Baldwin, 2003). The earnings gain for workers with at least a four-year college degree thus contributed to much of the earnings growth during this period.

In this paper, I measure a worker group's relative labor market gains by its relative earnings. The term is defined as a worker group's median weekly earnings divided by the average weekly earnings across all prime-age workers. Because earnings have grown faster at the upper tail of the earnings distribution, using the average earnings of all prime-age workers as a benchmark implicitly puts higher weight on the high earners. This approach highlights the wage dispersion between a median earner of any worker group and the high earners in the economy. Under this definition, Figure 1 plots how relative earnings have changed over time for US men and women by college education status.

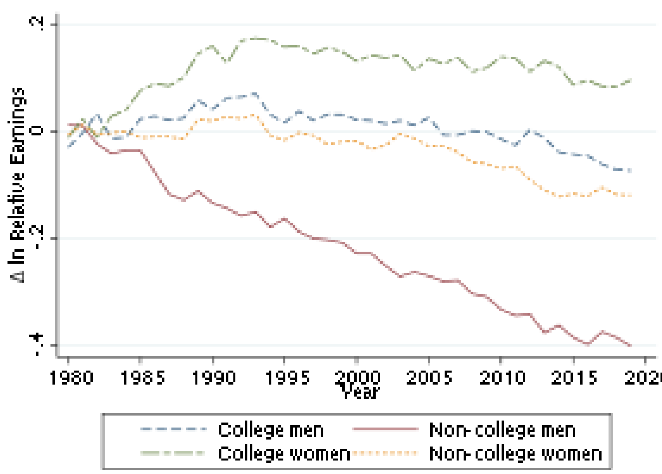

Figure 1: Changes in Relative Earnings

Note.– Using 1980 through 1982 as the base period, the figure plots the change in US prime-age workers' log relative earnings over time from their base-period value. Relative earnings are measured by a worker group's median weekly earnings divided by average weekly earnings across all US prime-age workers and are log-transformed. All earnings records are retrieved from the CPS Earner Study, 1980 through 2019.

Compared with their 1980–1982 level, non-college men's relative earnings have followed a consistent downward trend over the last 40 years, despite periodic real earnings gains. Specifically, from 1980 to 2019, the value of non-college prime-age men's median earnings fell from 104 percent to only 69 percent of the average earnings across their prime-age peers. In contrast, relative earnings for college-educated men, college-educated women, and non-college women grew initially, from 1980 to 1993, and then saw a modest downward trend reflecting the accelerated earnings growth at the upper tail of the earnings distribution and the overall stagnation of earnings at the median during the later years of the sample period. Nevertheless, the decline rate was substantially smaller compared with the rate at which relative earnings for non-college men fell. Together, these patterns highlight the earnings dispersion across education levels as well as the increased skewness in the earnings distribution over the last four decades.

1.2 Relative Earnings and Labor Force Exit Rate: An Inverse Relationship

For many workers, a job not only offers financial security, it also affirms their status, which is tied to their position relative to their age peers and many social outcomes. The social psychology and sociology literatures document extensively the relevance of relative income, or wage equality, on worker well-being (Adams, 1963; Festinger, 1954; Davis, 1959; Pollis, 1968; Runciman and Runciman, 1966). Using random controlled experiments, recent economics studies also show that a perceived wage inequality decreases job satisfaction (Card et al., 2012), depresses productivity (Breza et al., 2018), and, in a laboratory setting, reduces labor supply (Bracha et al., 2015). Similarly, studies report the correlation between relative income and job satisfaction (Clark and Oswald, 1996), subjective well-being (Luttmer, 2005; Solnick and Hemenway, 1998), women's employment (Neumark and Postlewaite, 1998), and health (Marmot, 2006). In these studies, workers evaluate their income against a reference group in their immediate surroundings, such as family members, colleagues, workers of similar demographic characteristics, or fellow experiment participants. However, with advancing information technology and mobility, the status of a worker may also depend on their relative position among a broader pool of workers in their city, state, or country. If this is true, the diminishing relative earnings of non-college men could reduce their utility gain from working, undermining their labor force participation incentives.

Consistent with this hypothesis, the historical data exhibit an inverse relationship between changes in a worker group's relative earnings and changes in their labor force exit rate. Figure 2 plots the changes in college-educated and non-college men's relative earnings (panel A) and labor force exit rate (panel B) over time from their 1980–1982 values. Compared with college-educated men, non-college men have experienced a much steeper fall in relative earnings as well as a more sizable rise in their labor force exit rate, showing the negative correlation between the two.

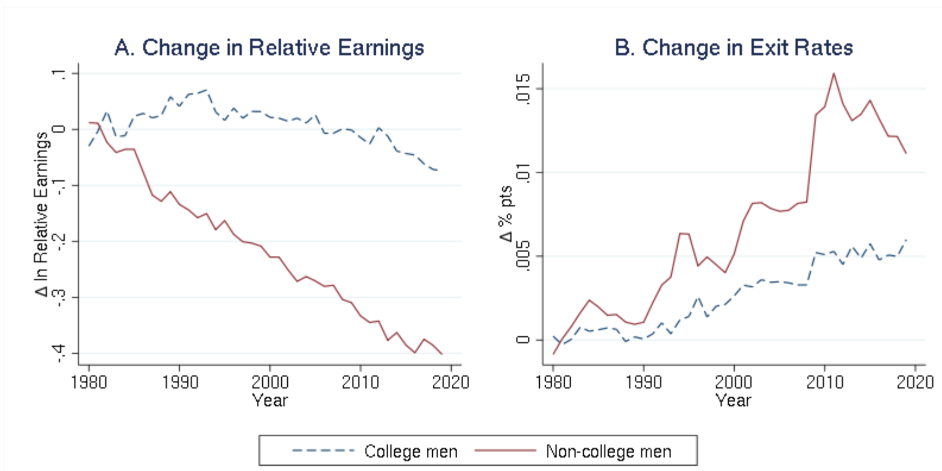

Figure 2: Changes in Relative Earnings and Labor Force Exit Rate by Education Attainment

Note.– Using 1980 through 1982 as the base period, panel (A) plots the change in US prime-age men's log relative earnings over time from their base-period value by college education status. Relative earnings are measured by a worker group's median weekly earnings divided by average weekly earnings across all US prime-age workers and are log-transformed. All earnings records are retrieved from the CPS Earner Study, 1980 through 2019. Similarly, panel (B) plots the change in US prime-age men's labor force exit rate over time from the base-period value by college education status. The labor force exit rate is measured by the average share of workers who transitioned from being in the labor force to neither working nor searching in each month estimated using the linked monthly basic CPS data from 1980 through 2019.

This inverse relationship also appears across subgroups of non-college men. To illustrate this, below I compare the changes in relative earnings and labor force exit rates for non-college men by occupation and geographic location.

Aside from noting the widening educational gradient in earnings, recent studies document a polarized growth trend across the occupation spectrum. Job demand has expanded for high- and low-skill jobs but shrunk for middle-skill jobs, a phenomenon for which the term ”job polarization” has been coined (Autor and Dorn, 2013; Acemoglu and Autor, 2011; Goos and Manning, 2007; Goos et al., 2014). Specifically, Acemoglu and Autor (2011) report that wages remained mostly stagnant from 1979 to 2009 for male workers in middle-skill blue-collar and clerical/sales occupations, while they increased more than 10 percent for high-skill professional, managerial, and technical occupations and low-skill service occupations. Among non-college men, the extent of the decline in relative earnings, therefore, varied across occupations.

Panel A of Figure 3 plots the changes in relative earnings for non-college men by their occupations' skill-level requirements. Following the 2010 Standard Occupational Classification System (SOC) and the convention in this literature, high-skill occupations here include management, professional, and public safety occupations (SOC: 11–31, 331–333). Middle-skill occupations include sales and administrative support occupations (SOC: 41, 43); production, installation, and repair occupations (SOC: 49, 51); and transportation and construction occupations (SOC: 45, 47, 53). Low-skill occupations comprise various service occupations (SOC: 33–39).[5] While non-college men in all three occupation groups experienced sizable declines in relative earnings, the decline was steeper for workers in the middle-skill occupations (–33 percent) than for those in the low-skill occupations (–26 percent) or the high-skill occupations (–20 percent). Panel B of Figure 3 shows that non-college men in the middle-skill occupations also experienced the most rapid increase in labor force exit rate, followed by workers in the low-skill occupations and in the high-skill occupations.

[5] Following Autor and Dorn (2013), I categorize public safety occupations (SOC: 331–333) as high-skill professional jobs. Additionally, because of the nature of the job, workers in the aviation transportation field (SOC: 532) are also categorized as high-skill professionals.

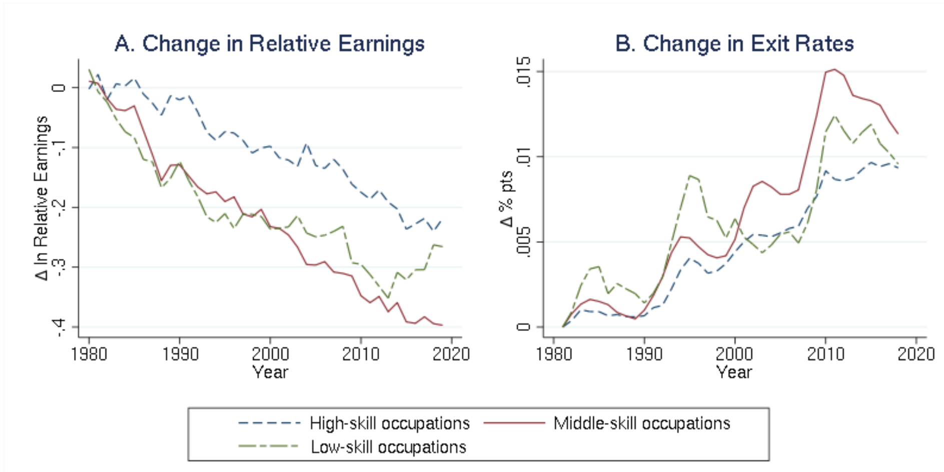

Figure 3: Changes in Relative Earnings and Labor Force Exit Rate by Occupation

Note.– Using 1980 through 1982 as the base period, panel A plots the change in US prime-age non-college men's log relative earnings over time from their base-period values by occupation. Relative earnings are measured by a worker group's median weekly earnings divided by average weekly earnings across all US prime-age workers and are log-transformed. All earnings records are retrieved from the CPS Earner Study, 1980 through 2019. Similarly, panel B plots the change in US prime-age non-college men's labor force exit rates over time from their base-period values by occupation. The changes are measured in three-year moving averages to reduce data volatility. The labor force exit rate is measured by the average share of workers who transitioned from being in the labor force to neither working nor searching in each month and estimated using the linked monthly basic CPS data from 1980 through 2019. High-skill occupations here consist of management and professional occupations (SOC: 11–31, 331–333, 532). Middle-skill occupations include sales and administrative support occupations (SOC: 41, 43); production, installation, and repair occupations (SOC: 49, 51); and transportation and construction occupations (SOC: 45, 47, 53). Low-skill occupations comprise various service occupations (SOC: 33–39).

The decline in non-college men's relative earnings during the 1980–2019 period also varied spatially. Because of state-level differences in industrial composition, the college wage premium rose at different speeds across states, and the extent of the decline in non-college men's relative earnings ranged from 10 percent to 64 percent depending on where a worker lived. To illustrate the spatial variation, I divide the 50 states and Washington, D.C., into two groups by the average decline rate in non-college men's relative earnings in each state. As in the two previous examples, panels A and B of Figure 4 show an inverse relationship between relative earnings and the labor force exit rate, such that non-college men in states with a more modest decline in relative earnings also exhibit less of an increase in their labor force exit rate.

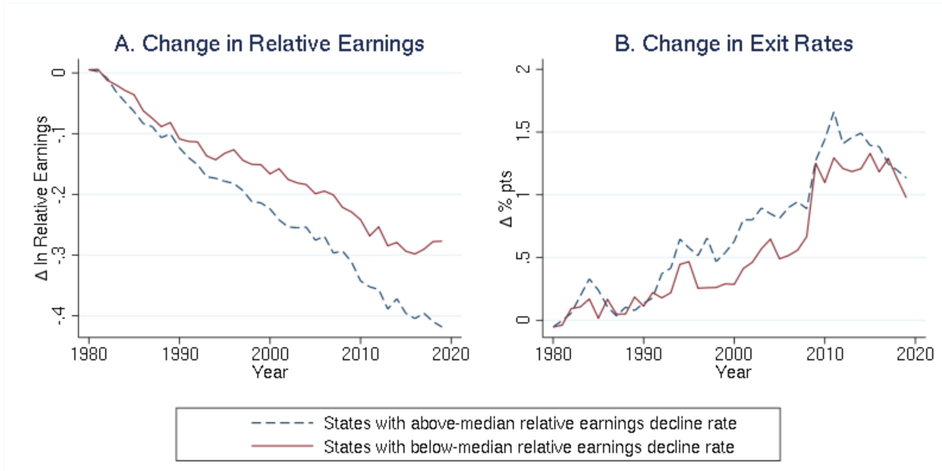

Figure 4: Changes in Relative Earnings and Labor Force Exit Rate by State
Note.– Using 1980 through 1982 as the base period, panel A plots the change in US prime-age non-college men's log relative earnings over time from their base-period value by workers' residence state (50 states and Washington, D.C.). States are divided into two groups according to whether their average decline rate in relative earnings is above or below the national median. Relative earnings are measured by a worker group's median weekly earnings divided by average weekly earnings across all prime-age workers in their state and are log-transformed. All earnings records are retrieved from the CPS Earner Study, 1980 through 2019. Panel B plots the change in US prime-age non-college men's labor force exit rate over time from the base-period value by the two state groups. The labor force exit rate is measured by the average share of workers who transitioned from being in the labor force to neither working nor searching in each month estimated using the linked monthly basic CPS data from 1980 through 2019.

Together, these descriptive statistics offer suggestive evidence that the labor force exit rate increases when relative earnings fall. These patterns, however, are subject to the influence of other confounding factors and may be only coincidental. To verify whether prime-age non-college men are more inclined to exit the labor force when their expected earnings fall relative to the earnings of other prime-age workers, Section 2 outlines the empirical strategy, leveraging the panel data, instrumental variables, and an extensive set of controls in the identification.

2 Empirical Method

2.1 Conceptual Framework

Consider an economy with N locations and Z occupations. Workers start in a fixed location-occupation pair and are immobile between locations and occupations but can choose whether to stay in the labor force for the next period. Their decision depends on the expected earnings from labor force participation, which is a product of the job-finding probability and the expected earnings from employment. Workers, nonetheless, evaluate the value of their earnings against the earnings of a reference group in the same location. Hence, they discount the utility gain from their expected earnings by the earnings of the reference group. The discount factor, γ, reflects the extent to which relative status and wage equity matters in workers' labor force participation decision. When the discount factor approaches zero, workers consider only the real value of their earnings when deciding whether to stay in the labor force. When the discount factor increases, workers place a higher weight on the relative value of their earnings in making the decision. To examine how these factors affect the exit propensity, I start with a baseline ordinary least squares difference-in-differences equation:

$$\begin{aligned}\text{Exit Rate}_{n,z,t} &= \alpha - \beta \ln \frac{\omega_{n,z,t}}{\widetilde{\omega}^{\gamma}_{n,t}} + \delta \ln Emp_{n,z,t} + \lambda_{n,z} + \mu_t + \epsilon_{n,z,t} \\ &= \alpha - \beta \ln \omega_{n,z,t} + \beta\gamma \ln \widetilde{\omega}_{n,t} + \delta \ln Emp_{n,z,t} + \lambda_{n,z} + \mu_t + \epsilon_{n,z,t}\end{aligned} \tag{2.1}$$

where the outcome variable is the labor force exit rate for prime-age non-college men in occupation z and US state n at time t; $\ln \omega_{n,z,t}$ is the expected earnings from employment, measured by the log median occupation earnings for prime-age non-college men in occupation z, state n at time t; $\ln \widetilde{\omega}_{n,t}$ is the expected reference group earnings, measured by the log mean earnings for all prime-age workers in state n at t; $\ln Emp_{n,z,t}$ is the probability of being employed, measured by the log employment rate in occupation z in state n at time t for prime-age non-college men;

$\lambda_{n,z}$ is the occupation-state fixed effect; μ_t is the time fixed effect; and $\epsilon_{n,z,t}$ is an idiosyncratic error term.

2.2 Endogeneity Concerns

The parameter of interest in equation 2.1 is β, measuring how responsive workers are to expected earnings from employment, and γ, the wage discount factor. Nevertheless, a few endogeneity concerns arise from the equation.

First, the coefficients of occupation earnings, $\ln \omega_{n,z,t}$, and reference earnings, $\ln \widetilde{\omega}_{n,t}$, may be biased away from zero as a result of unobserved demand shocks. For example, if an adverse demand shock reduces both wages and job vacancies for workers in occupation z and state n, without a proper control for changes in job loss risk, this correlation creates bias away from zero for the occupation earnings coefficient, β. In another instance, if computerization simultaneously raises the employment share in higher-earning computer engineer occupations and depresses the employment share in lower-earning blue-collar occupations, workers in the blue-collar occupations will see an increase in reference earnings and a decrease in their job vacancies at the same time, which introduces bias away from zero for the reference earnings coefficient, $\beta\gamma$.

Second, unobserved supply shocks or selection effects may bias the coefficients of occupation earnings and the coefficient of reference earnings. On one hand, the data show that workers with lower earnings prospects have a higher exit propensity. Because of the selection, an increase in the exit rate could increase the median occupation earnings, creating a bias toward zero for the occupation earnings coefficient. On the other hand, an unobserved labor supply shock could drive up the exit rate and, subsequently, the equilibrium occupation wage rate, depressing the estimated effect of occupation earnings. By underestimating the effect of occupation earnings, the model in turn creates a bias toward zero for $\beta\gamma$, which measures the counteracting effect of reference earnings.

Besides unobserved supply and demand shocks, other factors may influence the observed exit behavior of workers. At the state level, the condition of the state labor market, population growth, availability of public benefits, and housing costs could affect work incentives. At the occupation level, the utility gain from non-participation, or leisure, may vary across the worker groups depending on workers' alternative income resources. For example, if workers in higher-earning occupations are more likely to have higher investment income or spousal income, they may be more likely to exit the labor force, which creates a bias toward zero for the earnings coefficient.

2.3 Empirical Strategy

To address the endogeneity concerns, I take three steps in the empirical estimation. First, to account for changes in job vacancies and job loss risk, the model includes, besides the log employment rate, two additional controls on employment risk for each state-occupation pair: the job displacement rate and the employment-to-unemployment transition rate. Still, these controls may not fully account for all the demand-side changes in job loss risk. To test for such factors, Appendix A.2 reports the test statistics of selection on the observables and unobservables following the methodology developed by Oster (2019) based on the earlier work of Altonji et al. (2005). The result shows that the degree of bias from unobserved changes in job-finding rates or job loss risk is low and unlikely to explain the size of the estimated effects.

Second, to control for other state- and occupation-level factors that could potentially influence workers' exit behavior, I added the following covariates to the regression model: the average household income from sources other than own earnings for workers in occupation z in state n at time t and a vector of state-level economic and social characteristics, $K_{n,t}$. The vector $K_{n,t}$ includes the state unemployment rate, state housing cost index, the share of prime-age men receiving Social Security benefits (SS) or Supplemental Security Income (SSI), state poverty rate, and the log of state population.

Lastly, to gauge the extent of the bias created by unobserved local labor market shocks or selection effects, I estimate, as a robustness check, a 2SLS model where occupation earnings and reference earnings are instrumented using changes in the median occupation earnings in the country and a state's initial occupation employment share. By using national earnings alone to predict state earnings trajectories, the exercise reduces spurious correlations between the labor force exit rate and the state-level earnings measures.

Instrumental Variables

Specifically, in the 2SLS model, state occupation earnings are instrumented using the median occupation earnings across all other US states in period t:

$$\ln \omega_{n,z,t}^{IV} = \ln \omega_{-n,z,t}$$

Following the Bartik tradition, in constructing the variable, earnings records from a worker group's own state n are excluded to reduce the endogeneity of the instrumental variable (IV). The idea of the instrument is that the national median occupation earnings level serves as a pay standard for an occupation, directly influencing the pay level in each state. Yet, it is not subject to state-specific labor market shocks or selection effects that may create biases for the estimates.

Under the same logic, in the 2SLS estimation, reference earnings are instrumented using a cubic function of the inner product between a state's initial 1980–1984 occupation employment shares and the mean occupation earnings across all other states in period t:

$$
\begin{aligned}
\ln \widetilde{\omega}_{n,t}^{IV} = \ln \sum_{z=1}^{Z} &\text{Employment Share}_{n,z,1980-84} \times \widetilde{\omega}_{-n,z,t} \\
&+ (\ln \sum_{z=1}^{Z} \text{Employment Share}_{n,z,1980-84} \times \widetilde{\omega}_{-n,z,t})^2 \\
&+ (\ln \sum_{z=1}^{Z} \text{Employment Share}_{n,z,1980-84} \times \widetilde{\omega}_{-n,z,t})^3
\end{aligned}
$$

By fixing the occupation employment share to its initial 1980–1984 level, the instrument projects a state's average earnings using only changes in the national occupation earnings level. The cubic specification allows the initial employment share to have a non-linear relationship with a state's earnings trajectory. The projected average state earnings are thus less subject to state-specific labor market shocks yet reflect the increase in spatial earnings dispersion over time.

The instruments, however, come with a few limitations. First, by using national earnings to predict state-level earnings, it remains possible that the predicted earnings are correlated with unobserved labor supply or demand shocks common to the entire US labor market. Second, by design, the instruments eliminate the inter-state variances of earnings within each occupation. The 2SLS estimation hence relies on the remaining variance across occupations and over time. If the treatment effect varies across different dimensions, the 2SLS result may reflect the heterogeneous effects, on top of any bias correction. For these reasons, Section 4 focuses primarily on the OLS result and uses the 2SLS estimate as supplementary evidence to highlight the consistent wage discounting effect under both specifications.

Estimating Equations

Adopting the steps outlined in Section 2.3, the final OLS estimating equation for the empirical analysis is:

$$\begin{aligned} \text{Exit Rate}_{n,z,t} = \alpha - \beta \ln \omega_{n,z,t} + \beta\gamma \ln \widetilde{\omega}_{n,t} \\ + X_{n,z,t} + K_{n,t} + \lambda_{n,z} + \mu_t + \epsilon_{n,z,t} \end{aligned} \tag{2.2}$$

The equation uses the variables specified in the baseline equation. In addition, $X_{n,z,t}$ is a vector of time-variant controls for each state-occupation pair (n, z), including the log employment rate, the displacement rate, the employment-to-unemployment transition rate, and the log median household income from sources other than own earnings; and the vector $K_{n,t}$ is a vector of state-level time-variant covariates that include the state unemployment rate, state housing cost index, state poverty rate, SS/SSI receipt rate, and log state population in period t.

3 Data

The regression sample consists of a panel of occupation-state-level observations over the period 1980 through 2019. Following the convention in the literature, the occupation classification is derived from the 2010 Standard Occupational Classification (SOC) system and includes five occupation groups: management, professional, and public safety (SOC: 11–31, 331–333); sales and administrative support (SOC: 41, 43); production, installation, and repair (SOC: 49, 51); transportation and construction (SOC: 45, 47, 53); and service (SOC: 33–39).[6] To reduce the incidence of zero exit rate as a result of a small sample size, each period t in the analysis refers to a five-year interval within the 1980–2019 period. The final sample thus includes 255 unique state-occupation pairs and 2,040 observations.

[6] Because of the nature of the jobs, workers in the aviation transportation industry (SOC: 532) are categorized in the management, professional, public safety group.

3.1 Measures

Labor Force Exit Rate

The Integrated Public Use Microdata Series-Current Population Survey (IPUMS-CPS) 1980–2019 constitutes the main data source for this study (Flood et al., 2020). The CPS gathers employment information on members of about 70,000 housing units in the sample each month. Individual respondents who remain in the same housing unit are interviewed monthly for four months, followed by an eight-month break, and interviewed again for four consecutive months. This design allows researchers to observe the employment transitions of individuals over time. Using the matched CPS data, I construct the monthly labor force exit rate by dividing the number of workers in occupation n who have left the labor force in month m by the total number of workers in occupation n in the labor force in month $m-1$. Moreover, I restrict the exit rate sample to prime-age non-college men who were not in the first and fifth CPS interview months, reported an occupation at the time of the interview, and were in the labor force in month $m-1$.[7] Because information on the most recent occupation affiliation is available for 99.9 percent of prime-age male labor force participants but only 6.5 percent of non-participants in IPUMS-CPS, the choice of using labor force exit rate as the outcome variable allows me to stratify the sample by a worker's previous occupation affiliation, a key determinant of his expected return from labor force participation.

Occupation Earnings and Reference Earnings

Each month, besides collecting employment information, the CPS conducts an Earner Study by asking a sub-sample of its respondents additional questions regarding their weekly earnings from their current job. Because of the short recall period, the Earner Study provides earnings information that is less subject to recall

[7] Workers in the labor force who were chronically unemployed for more than five years have no occupation information in the CPS data and therefore are dropped from the sample. These workers represent less than 0.01 percent of the sample and less than 2 percent of the exiting workers.

bias than surveys with longer recall periods. The IPUMS-CPS offers Earner Study data for the years 1982 through 2019. For years 1980 and 1981, earnings records are retrieved from the original CPS data provided by the National Bureau of Economic Research (NBER). Using the Earner Study, I measure occupation earnings by the median weekly occupation earnings in state z at t for prime-age non-college men. Similarly, I measure reference earnings by the mean weekly earnings in state z at t across all prime-age workers.[8] All earnings are inflation-adjusted using the CPI-U and measured at the 2018 dollar value.

State-Occupation-Level Controls

All state-occupation level controls in this study are estimated using the IPUMS-CPS. Using the IPUMS-CPS basic monthly data, I construct the occupation employment rate as the employed share of non-college men who were in the labor force in occupation n in state z in period t. To gauge the dynamic change in employment in each period, the employment-to-unemployment transition rate is estimated using the IPUMS-CPS linked basic monthly data and defined as the average monthly share of non-college men in occupation n in state z who transitioned from employment to job-searching/unemployment in time period t. As a measure of job loss shocks, the job displacement rate is estimated using the IPUMS-CPS Displaced Worker Supplement and defined as the number of workers who lost their job due to plant or company closure, abolished position, or insufficient workload, measured as a share of the total number of employed workers in occupation z in state n and averaged over time period t.[9] Lastly, the median value of annual household income from sources other than own earnings is obtained from the IPUMS-CPS Annual Social and Economic Supplement, in which workers are asked about their total family income and own earnings in the previous calendar year.[10] The household income

[8] All of the weekly earnings variables in the Earner Study are adjusted for top-coding.

[9] The original CPS survey records job displacement over the three- to five-year period preceding the survey period. To reduce recall bias, in constructing the variable, I counted only the number of workers who lost their jobs in the calendar year preceding the survey period.

[10] The income of the top 2 percent highest-earning families in each state is replaced by the 98th percentile income to avoid distortion generated by outliers.

Table 1: Descriptive Statistics

	Mean	S.D.
Labor force exit rate (%)	1.94	0.93
ln Occupation earnings	6.72	0.23
ln Reference earnings	6.89	0.11
State-occupation-level controls		
ln Employment rate	-0.06	0.03
Displacement rate (%)	4.22	2.50
Employment-to-unemployment transition rate (%)	1.74	1.06
ln Other household income	10.21	0.21
State-level controls		
Poverty rate (%)	13.48	3.03
Unemployment rate (%)	6.23	1.79
SS/SSI receipt rate (%)	3.70	1.37
ln State population	15.95	0.91
House price index	1.23	0.35

Note.– The summary statistics are weighted by the sum of the CPS weights for the individuals in each cell.

variable is log-transformed when included in the model.

State-Level Controls

The state-level controls on the unemployment rate, poverty rate, and population were retrieved from the UKCPR National Welfare Data Series (University of Kentucky Center for Poverty Research, 2019). Shares of prime-age men receiving Social Security benefits or SSI in each state are computed using the IPUMS-CPS Annual Social and Economic Supplement. In addition to the above controls, the model controls for state-level changes in house price level using the inflation-adjusted annual all-transactions state-level house price index acquired from the Federal Housing Finance Agency.

4 Estimation Results

4.1 Labor Force Exit Rate and Relative Earnings

Table 2: Relative Earnings and Labor Force Exit Rate

	(1)	(2)	(3)
ln Occupation earnings	-1.29***	-1.16***	-2.69***
	(0.26)	(0.23)	(0.20)
ln Reference earnings	1.40**	1.26**	2.78**
	(0.46)	(0.40)	(1.11)
Occupation-state fixed effects	✓	✓	✓
Time fixed effects	✓	✓	✓
State-occupation-year controls		✓	✓
State-year controls		✓	✓
Specification type	OLS	OLS	2SLS
Within R^2	0.64	0.68	0.67
Kleibergen-Paap Wald 1st stage F statistic	-	-	24.65
Observations	2040	2040	2040
$\hat{\gamma}$	1.09	1.09	1.03

Note.– The dependent variable of the regression is labor force exit rate measured in % pts. The sample was retrieved from the IPUMS-CPS and consists of US prime-age non-college men for the period 1980 through 2019. The results are weighted by the sum of the CPS weights for the individuals in each cell. Standard errors are clustered at the state level. See Appendix A.1 for the full estimation result.

* for P $<$ 0.05, ** for P $<$ 0.01, *** for P $<$ 0.001.

Table 2 reports the estimation results. To assess how the controls and the estimation method affect the estimates, column 1 reports the result of the baseline model without any additional controls, column 2 the full OLS result of equation 2.2, and column 3 the 2SLS result of equation 2.2 with both occupation earnings and reference earnings instrumented using the IVs.

According to the baseline estimate (column 1), a 10 percent increase in both occupation earnings and reference earnings is associated with a 0.13 percentage point *decrease* and a 0.14 percentage point *increase* in the exit rate, respectively. Column 2 shows that the coefficient estimates decline modestly after including the controls. This confirms that the coefficients are subject to biases away from zero as changes in earnings and job vacancies are correlated. The extent of the bias, however, is limited. The degree of bias from omitted variables also appears to be low and unlikely to explain the observed correlation.[11] The estimated earnings discount factor, $\hat{\gamma}$, is 1.09 under both specifications. Earnings growth in an economy therefore has two opposing relationships with the labor force exit rate. By increasing the expected earnings for each worker group, earnings growth is associated with lower exit rates, and, by increasing the reference earnings level, it is associated with higher exit rates. This double-edged relationship cannot be reflected if the exit rate is modeled as a function of occupation earnings alone.

As discussed in Section 2.2, the occupation earnings and reference earnings are endogenous and the OLS estimates are subject to estimation biases. To gauge the severity of the biases, column 3 reports the 2SLS estimates where occupation earnings and reference earnings are instrumented using earnings records from other US states, as specified in Section 2.3. The coefficients of the occupation earnings and reference earnings under the 2SLS method are approximately two times the size of the OLS estimates, with an implied $\hat{\gamma}$ of 1.03. Consistent with the OLS results, the value of $\hat{\gamma}$ suggests that workers' own expected earnings and the reference earnings bear equal weights in their labor force exit decision. The exercise therefore offers limited evidence that uncontrolled local labor market shocks have driven the observed discounting effect of the reference earnings. The large numerical increase

[11] Appendix A.2 reports the test statistics from the test of selection on the observables and unobservables developed by Oster (2019) based on the work of Altonji et al. (2005) The results suggest that accounting for unobserved job loss risks modestly attenuates the occupation earnings coefficient to –1.16 and the reference earnings coefficient to 1.09. The omitted factors need to be 6 to 112 times more strongly correlated with the exit rate in order to explain the size of the coefficients.

in the point estimates, however, raises the concern that earnings may have heterogeneous treatment effects across different estimation margins (see Section 2.3).[12] To err on the side of caution, this section focuses on the more conservative OLS estimate, which likely represents a lower bound of the relationship between relative earnings and the labor force exit rate.

In sum, the estimation result shows that non-college men are more likely to leave the labor force when their earnings fall relative to other prime-age workers. An increase in real earnings hence does not necessarily lead to a lower rate of labor force exit if the growth lags behind the average earnings growth rate in the economy. The finding is consistent with the continual decline in non-college men's labor force participation over the last 40 years, during which time there was no monotone decline in their real earnings while their relative earnings fell persistently.

4.2 Robustness Tests

Alternative Reference Earnings

The estimation results from equation 2.2 show that non-college men are more likely to leave the labor force when their own earnings fall relative to the reference earnings, measured by the state average earnings. The choice of state average earnings as the reference group is quite arbitrary. In reality, workers could be comparing themselves with a variety of reference groups when evaluating their relative labor market status, such as other workers in the same neighborhood, the median worker in their state, or with only men or women. While these measures are strongly correlated, which may explain the observed relevance of the choice reference group, here

[12] The presence of heterogeneous treatment effects in a standard two-way fixed effects environment can potentially distort the OLS estimates when combined with negative weights. To assess the influence of such distortion, Appendix A.3 reports relative earnings' average treatment effect across all state-occupation groups and time periods using the estimator developed by De Chaisemartin and d'Haultfoeuille (2020) which is robust to heterogeneous effects. The result shows that a 10 percent increase in relative earnings is associated with a 0.10–0.23 percentage point average decrease in the exit rate, falling within the range of estimated effects under the OLS and 2SLS method.

I test if any of the alternative groups stands out as a more or less probable reference group than the others.

Table 3: Relative Earnings and Labor Force Exit Rate with Alternative Reference Earnings

	(1)	(2)	(3)	(4)
ln Occupation earnings	-1.19***	-1.50***	-1.00***	-1.15***
	(0.23)	(0.19)	(0.21)	(0.22)
ln Reference earnings	1.30***	1.83***	0.67†	
	(0.37)	(0.27)	(0.40)	
$\mu_{n,t}$				1.22**
				(0.39)
$\sigma_{n,t}$				1.97***
				(0.56)
Occupation-state fixed effects	✓	✓	✓	✓
Time fixed effects	✓	✓	✓	✓
State-occupation-year controls	✓	✓	✓	✓
State-year controls	✓	✓	✓	✓
Reference group	Men	Non-college men	Women	All
Specification type	OLS	OLS	OLS	OLS
Within R^2	0.68	0.69	0.68	0.69
Observations	2040	2040	2040	2040
$\hat{\gamma}$	1.09	1.22	0.67	-

Note.– The dependent variable of the regression is labor force exit rate measured in % pts. The sample was retrieved from the IPUMS-CPS and consists of US prime-age non-college men for the period 1980 through 2019. The results are weighted by the sum of the CPS weights for the individuals in each cell. Standard errors are clustered at the state level.
* for P < 0.05, ** for P < 0.01, *** for P < 0.001.

Table 3 presents the regression results where reference earnings are measured by the state-level average earnings across all prime-age men (column 1), non-college men (column 2), all prime-age women (column 3), and the level and skewness of the state prime-age earnings distribution to test the correlation between the exit rate and these alternative reference earnings measures.

Columns 1 and 2 show a strong correlation between the labor force exit rate and the average earnings of prime-age men and non-college men, respectively. The correlation, however, becomes weaker when the reference group is comprised of women only, suggesting that men's relative status in the labor market primarily depends on the earnings distribution of other men. While columns 1 and 2 report a similar strength of correlation, there remain some subtle differences between the measures. During this period, men's average earnings steadily grew from increases in college wage premium. In contrast, non-college men's average earnings were mostly stagnant. When restricting the reference group to non-college men only, the model leaves a more sizable unexplained time trend residual after accounting for the changes in relative earnings. Empirically, the increasing gap between college and non-college earnings is crucial in accounting for the trend increase in labor force exit among non-college men.

Column 4 tests how the exit rate relates to other moments of the state prime-age earnings distribution. Because state earnings closely follow a log-normal distribution, earning moments on the distribution can be sufficiently captured by the mean of the log earnings, μ, and the standard deviation of the log earnings, σ. In this exercise, instead of picking a specific earnings moment, I substitute reference earnings with the two parameters, μ and σ. The result shows that, all else being equal, the labor force exit rate increases with both the level of state earnings and the skewness of the state earnings distribution, indicating that workers' labor participation incentives are related to both how much their peers earn and how equitable the earnings are distributed. This finding adds further evidence that earnings equity may affect the utility gain from work for non-college men.

Results by Demographic Characteristics

To examine if the relationship between relative earnings and the labor force exit rate varies across worker demographics, Table 4 stratifies the estimation results by race (columns 1 through 3), age (columns 4 through 6), and marital status (columns 7 and 8).

Across race/ethnicity groups, the positive correlation between reference earnings and the labor force exit rate is identified among non-Hispanic White men but absent among non-Hispanic Black or Hispanic men. The deteriorating labor market status of non-college men appears to have chiefly impacted non-Hispanic Whites. This racial specificity is consistent with prior evidence on economic despair and worker health (Case and Deaton, 2015; Stein et al., 2017; Case and Deaton, 2021), and a possible explanation is that non-Hispanic Black and Hispanic men's labor supply was disproportionately influenced by other non-labor-market factors over the study period, such as mass incarceration (Pager, 2003; Holzer et al., 2006; Pager et al., 2009; Raphael, Raphael; Sakala, 2014; Mueller-Smith, 2015; Carson, 2020) or immigration (Borjas, 2017; Bureau of Labor Statistics, 2022). Across age groups, the correlation between relative earnings and the labor force exit rate declines with age. The correlation also weakens significantly for both single and married men when the sample is stratified by a worker's marital status. Together, these results by demographics indicate that some of the possible channels through which relative earnings affect men's labor supply behavior include changes in non-Hispanic White men's social status and marriage market sorting, which is more relevant to younger men than older men.

Table 4: Relative Earnings and Labor Force Exit Rate by Demographic Characteristics

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
ln Occupation earnings	-1.12***	-0.20	-0.75**	-1.28***	-0.98***	-0.73**	-1.21***	-0.66***
	(0.25)	(0.41)	(0.25)	(0.27)	(0.20)	(0.24)	(0.35)	(0.18)
ln Reference earnings	1.14***	-1.10	-0.76	1.42*	1.10*	0.92	0.44	1.07**
	(0.32)	(1.28)	(0.87)	(0.56)	(0.49)	(0.58)	(0.67)	(0.36)
Occupation-state fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
Time fixed effects	✓	✓	✓	✓	✓	✓	✓	✓
State-occupation-year controls	✓	✓	✓	✓	✓	✓	✓	✓
State-year controls	✓	✓	✓	✓	✓	✓	✓	✓
Demographics	White	Black	Hispanic	25-34	35-44	45-54	Single	Married
Specification type	OLS	OLS	OLS	OLS	OLS	OLS	OLS	OLS
Within R^2	0.61	0.33	0.22	0.59	0.45	0.40	0.39	0.50
Observations	2008	1871	1845	2040	2040	2040	2040	2040
$\hat{\gamma}$	1.02	-5.50	-1.01	1.11	1.12	1.26	0.36	1.62

Note.– The dependent variable of the regression is labor force exit rate measured in % pts. The sample was retrieved from the the IPUMS-CPS and consists of US prime-age non-college men for the period 1980 through 2019. The results are weighted by the sum of the CPS weights for the individuals in each cell. Standard errors are clustered at the state level.
† for $P < 0.10$, * for $P < 0.05$, ** for $P < 0.01$, *** for $P < 0.001$.

4.3 Back-of-the-Envelope Calculation

To assess the extent to which relative earnings account for the observed variation in labor force exit rates over time and across groups, I compare the estimation results with the descriptive patterns reported in Section 1.2.[13]

Over the course of the study period, non-college men's relative earnings declined 0.30 log point (see Figure 2), which was the combined result of a 0.17-log-point decline in their real median earnings and a 0.23-log-point growth in the reference earnings. A back-of-the-envelope calculation based on the OLS estimates suggests that this decline is associated with a 0.49 percentage point increase in the exit rate, accounting for 44 percent of the net growth in non-college men's labor force exit rate during this period. In contrast, changes in real occupation earnings alone account for only 18 percent of the total increase in the labor exit rate for non-college men, leaving a larger unexplained growth trend.

Non-college men's labor force exit rate trajectory also varied across occupations and states over the study period (see Figure 3 and Figure 4 in Section 1.2). Across occupations, the OLS model prediction accounts for 88 percent of the observed difference between the high- and middle-skill occupations and 59 percent of the difference between the low- and middle-skill occupations. Because the value of reference earnings is constant across occupational lines, the predicted difference here is driven entirely by variations in real earnings. Across states, the states with an above-median relative earnings decline rate experienced a steeper increase in the labor force exit rate than the states with an below-median relative earnings decline rate. Based on the model prediction, changes in relative earnings account for 107 percent of the observed between-state gap, while real earnings and reference earnings each contribute to approximately half of the projected difference. These calculations show that the empirical model adequately captures the dynamic change in non-college men's labor force exit rate over time as well as the heterogeneity across worker groups and states.

[13] As in Section 1.2, the changes in labor force exit rates and earnings are calculated using year 1980–1982 as the base period.

5 Conclusion

The decline in the US prime-age men's labor force participation rate over the last 40 years has spurred much interest and discussion among scholars. Prior research finds that changes in real wage rates cannot sufficiently account for the magnitude and persistence of the decline. Changes in supply-side factors explain an even smaller share of the trend decline. This paper posits that the increased educational and occupational gradients in earnings and the resulting fall in non-college men's relative earnings explains a sizable share of the trend. Specifically, the results indicate that changes in relative earnings account for 44 percent of the total growth in labor force exit among non-college men during the 1980–2019 period. This finding suggests that deteriorating social status is a plausible key factor driving prime-age men's declining labor force participation.

Why would men leave the labor force when their relative status declines? The stratified result by demographics suggests that marriage market sorting among younger workers is a possible link between the two. Research and data point to a few other contributing factors besides marriage market sorting: low job satisfaction, disability, and schooling. A recent paper by Krause and Sawhill (2017) finds that prime-age men who were not in the labor force spent twice as much time on leisure activities and sleeping compared with labor force participants. This suggests that most non-participants did not leave the labor force to engage in other non-paid work. Instead, their choice likely reflects a dissatisfaction about work. Labor force exit also appears to have an intertwined relationship with disability. Nearly 30 percent of the prime-age men who left the labor force reported a work-limiting disability condition at the time of the exit event.[14] Previous studies find that the stress and low self-esteem associated with lower social status contribute to poorer health and higher mortality rates (Marmot, 2006; Eibner and Evans, 2005; Daly et al., 2013). Lastly, the percentage of exiting workers enrolled in school at the time of the exit event increased modestly from 11 percent to 14 percent over the last two decades.[15]

[14] Author's calculation using the Survey of Income and Program Participation's 2008 panel.

[15] Author's calculation using the Survey of Income and Program Participation's 1996 and 2008

As earnings diverge across skill sets, workers could be more inclined to temporarily leave the labor force to seek retooling or upskilling opportunities.

The findings of this paper point to an area less attended to in the literature. If the increasing wage gap between high and low earners directly or indirectly affects men's aggregate labor supply, wage inequality might have carried wider implications to the economy than previously believed. More research is needed to fully understand how earnings distribution affects a worker's perception about work, the social return from work, and their aggregate labor market behavior.

panels.

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Appendix A

A.1 Full Estimation Result

Table A1: Relative Earnings and Labor Force Exit Rate: Full Estimation Result

	(1)	(2)	(3)
Treatment variables			
ln Occupation earnings	-1.29***	-1.16***	-2.69***
	(0.26)	(0.23)	(0.20)
ln Reference earnings	1.40**	1.26**	2.78*
	(0.46)	(0.40)	(1.11)
State-occupation-year controls			
ln Employment rate		-8.16***	-7.98***
		(0.84)	(0.86)
Job displacement rate		0.31	0.13
		(0.61)	(0.62)
Employment-to-unemployment transition rate		-5.94†	-4.88
		(3.14)	(3.09)
Other family income		0.22**	0.29***
		(0.08)	(0.08)
State-year controls			
Poverty rate		0.03**	0.02†
		(0.01)	(0.01)
ln Population		-0.15	-0.02
		(0.19)	(0.21)
Unemployment rate		-0.03	-0.03
		(0.02)	(0.02)
House price index		-0.09	-0.18
		(0.08)	(0.12)
SS/SSI receipt rate		3.32†	4.68**
		(1.71)	(1.66)
Occupation-state fixed effects	✓	✓	✓
Time fixed effects	✓	✓	✓
Specification type	OLS	OLS	2SLS
Within R^2	0.64	0.68	0.67
Observations	2040	2040	2040

Note.– The dependent variable of the regression is labor force exit rate measured in % pts. The results are weighted by the sum of the CPS weights for the individuals in each cell. Standard errors are clustered at the state level.
† for P $<$ 0.10, * for P $<$ 0.05, ** for P $<$ 0.01, *** for P $<$ 0.001.

A.2 Bias from Selection on Unobservables

Based on the work of Altonji et al. (2005), Oster (2019) develops a test to estimate the bias from unobserved selection under an OLS model environment. The idea of the test is to project the bias from unobserved selection leveraging two pieces of information: (1) the degree of bias correction from the selection observed and (2) the share of the variation in the outcome variable that is possibly attributable to the unobserved selection. For the present study, the goal of this exercise is to estimate the extent to which unobserved job loss risk not captured by the occupation-level controls in the model may further attenuate the coefficient estimate of occupation earnings and reference earnings.

To gauge the maximum share of variation in the outcome variable attributable to the unobserved job loss risk, I augment equation 2.2 with the occupation employment share of all workers, male workers, and non-college male workers in state z at t. These variables reflect the equilibrium employment share and are functions of both the exit rate and the job loss risk. Because they are directly affected by the exit rate, the employment shares are not suitable controls for job loss risk in the model. However, by including these variables in the model, I can obtain an upper bound for the R^2 that is attributable to the unobserved changes in job loss risk. Using this upper bound of 0.90, I then conduct the test as detailed by Oster (2019) to estimate a bias-adjusted coefficient of reference earnings.

According to the test statistics reported in Table A2, the occupation earnings coefficient is projected to decrease by another 0.006 to 1.16 and the reference earnings coefficient is projected to decrease by another 0.17 to 1.09 after the selection on the unobservables is taken into consideration. The test statistics also suggest that the omitted factors need to be 6 to 112 times more strongly correlated with the exit rate to explain the total size of the coefficients. The results offer evidence that the degree of the bias from the unobservables is low and unlikely to be driving the observed correlation between the labor force exit rate and the two explanatory variables.

Table A2: Selection on Unobservables

	(1)	(2)
Treatment variables		
ln Occupation earnings	-1.18***	-1.16***
	(0.26)	(0.25)
ln Reference earnings	1.74***	1.26**
	(0.44)	(0.43)
State-occupation-year controls		
Other family income	✓	✓
Job loss risks		✓
State-year controls	✓	✓
Occupation-state fixed effects	✓	✓
Time fixed effects	✓	✓
Specification type	OLS	OLS
R^2	0.89	0.90
Observations	2040	2040
Oster (2019) Test Statistics		
ln Occupation earnings		
$\beta*$	-1.16	
δ	112.28	
ln Reference earnings		
$\beta*$	1.09	
δ	5.56	

Note.– The dependent variable of the regression is labor force exit rate measured in % pts. The results are weighted by the sum of the CPS weights for the individuals in each cell. Standard errors are clustered at the state level.

* for P $<$ 0.05, ** for P $<$ 0.01, *** for P $<$ 0.001.

A.3 De Chaisemartin and d'Haultfoeuille (2020) Estimator

Table A3 reports the estimated treatment effect of relative earnings using the standard OLS and 2SLS method (row 1 and 2) and the DID$_M$ estimator (row 3-6) developed by De Chaisemartin and d'Haultfoeuille (2020) which is robust to heterogeneous treatment effects. To accommodate the constraints associated with the De Chaisemartin and d'Haultfoeuille (2020) estimator, the estimation is based on a simplified version of equation 2.2 where the treatment variable is relative earnings, defined as $\ln \omega_{n,z,t} - \ln \widetilde{\omega}_{n,t}$, and there is no controls beyond the group and time fixed effects. The purpose of the exercise is to verify if relative earnings' estimated effect is significantly distorted by the potential presence of heterogeneous treatment effects and negative weights.

Because the treatment variable is continuous and not stable over time, the DID$_M$ estimator requires specifying a stable treatment threshold value, below which any change in the treatment variable is assumed to have no effect on the outcome. Table A3 reports the DID$_M$ estimates based on alternate thresholds ranging from 0.03 to 0.09 log point to demonstrate the influence of threshold choices on the estimated effect.

Based on the estimates, a 10 percent increase in relative earnings is associated with a 0.13 percentage point decrease in the labor force exit rate under the OLS method, a 0.25 percentage point decrease under the 2SLS method, and a 0.10–0.23 percentage point decrease using the DID$_M$ estimator. The estimated average effect increases modestly with the value of the threshold, suggesting that smaller changes in relative earnings have more limited impact on the labor force exit rate. Still, the DID$_M$ results fall closely within the range of effect suggested by the OLS and 2SLS estimate. The finding therefore upholds the validity of the main estimation results reported in Section 4.

Table A3: Result using the De Chaisemartin and d'Haultfoeuille (2020) Estimator

Method	Estimate	Standard error	Observations	Stable treatment threshold
OLS	-1.30	0.28	2040	-
2SLS	-2.50	0.30	2040	-
DID$_M$	-1.37	0.85	1297	0.03 log point
DID$_M$	-1.02	0.52	1314	0.05 log point
DID$_M$	-2.16	0.62	1150	0.07 log point
DID$_M$	-2.26	0.80	888	0.09 log point

Note.– The dependent variable of the regression is labor force exit rate measured in % pts. The treatment variable is log relative earnings. The results are weighted by the sum of the CPS weights for the individuals in each cell. Standard errors are clustered at the state level.