

Relative Earnings and Labor Force Exit Rate: Evidence from US Prime-Age Men

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Abstract

This study presents the first empirical evidence that labor force exit rate rises when workers' relative earnings fall. The analysis posits that workers form earnings expectations through peer comparison and evaluate their earnings against those of their age-range peers. Based on the results, the decline in the earnings of less-educated US prime-age men relative to the overall prime-age earnings from 1980 to 2019 leads to a 0.34 to 0.48 percentage point increase in their likelihood of leaving the labor force, accounting for 24 to 33 percent of the total growth in US prime-age male labor force exit rate over this period.

JEL Codes: J21, J24, J31

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1 Introduction

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 Organization for Economic Co-operation and Development (OECD) countries in 2019.¹ Prime-age men traditionally constitute the backbone of the US workforce. The decline hence carries direct implications for both the economy and the well-being of individual workers. However, despite significant scholarly interest in this phenomenon, the literature has yet to fully explain the causes of this observed decline. Neither changes in real earnings nor growth in alternative income sources account for its persistence and pervasiveness.

A closer look at the data reveals that the observed decline in labor force participation was driven by more frequent exits among men without a four-year college degree (hereafter "non-college men").² Meanwhile, from 1980 to 2019, the median earnings of non-college men measured as a share of the median earnings of all prime-age workers fell steadily by 25 percent, exhibiting a trend that paralleled this group's declining labor force participation. These parallel trends are consistent with recent studies that, using randomized controlled experiments, find that workers form their earnings expectations through peer comparison, such that their relative earnings have been shown to influence job satisfaction (Card et al., 2012), productivity (Breza et al., 2018), and labor supply in a lab setting (Bracha et al., 2015). These findings suggest that a job not only provides economic security, but also affirms a worker's status, which is tied to their position relative to their peers.

Considering these parallel trends and drawing on insights from the literature, this study investigates whether prime-age men are more inclined to leave the labor force when their expected earnings fall relative to the earnings of other prime-age workers in their state. The sample consists of worker-level labor force transition records retrieved from the Current Population Survey (CPS) matched with information on state-level earnings distribution from 1980 through 2019. A worker's expected earnings are measured by the mean log earnings of their skill group in their state and the

¹Based on [OECD official statistics](#).

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

reference group earnings are defined as the mean log earnings across all prime-age workers in their state.³ Consistent with the hypothesis that a worker's return on labor force participation is connected to their relative earnings and not real earnings alone, the analysis finds that, holding a prime-age man's expected earnings constant, his likelihood of exiting the labor force increases with the reference group earnings in his state. Based on the more conservative ordinary least squares (OLS) estimates, a 10 log point increase in a worker's expected earnings is associated with a 0.123 percentage point decline in the likelihood of labor force exit, while a 10 log point increase in the reference group earnings is associated with a 0.118 percentage point increase in the likelihood of labor force exit. The symmetry of the point estimates indicates that prime-age men nearly fully discount their expected earnings by the reference group earnings when evaluating the return on participation. This result is robust to a rich set of controls of potential correlates to state average earnings, including state-level variation in housing cost, unemployment rate, poverty rate, public benefit generosity, and other family income sources. It is also robust to a two-stage least squares (2SLS) specification that leverages a state's initial industrial structure and regional pay level as an instrument for the reference group earnings, in order to rule out the possibility that state-specific selection or latent demand shocks are driving the observed correlation.

After establishing the robustness of the baseline result, the study investigates the relationship between relative earnings and a set of labor market outcomes related to a worker's exit decision. The supplementary analysis is motivated by the following argument: If the rise in prime-age men's labor force exit rate reflects their labor supply response to deteriorating labor market status, their reduced labor supply should be observed in both more labor force exits as well as more voluntary quits or reduction of work hours for workers who remain in the labor force. Conversely, if the relationship is driven by latent demand shocks, the increase in labor force exit should be observed along with more involuntary job loss or hour reduction. Assessing these related labor market outcomes therefore lends further credence to the underlying mechanisms driving the baseline result. The finding from this supplementary analysis shows a coherent relationship between relative earn-

³In the empirical analysis, all earnings measures are based on log-transformed individual weekly earnings from the CPS Earner Study.

ings and the labor supply intensity of prime-age male labor force participants, as evidenced by a positive correlation between the reference group earnings and voluntary job departures or reduction of work hours. However, the analysis does not find that changes in the reference group earnings are statistically significantly correlated with demand-driven unemployment or part-time work.

To gain deeper insight into the behavioral mechanisms underlying the baseline outcome, the study then examines worker sensitivity to alternative measurements of the reference group earnings. Conditional on their own expected earnings, non-college men and men with a four-year college degree (hereafter “college men”) show different sensitivities to the various quantiles and moments of the state earnings distribution. Non-college men’s likelihood of labor force exit correlates most strongly with the median state earnings. By contrast, college men have a more muted response to changes in the median state earnings, but their labor force exit rate increases with the higher state earnings quantiles. This difference indicates that prime-age men not only evaluate their earnings against their peers’, but also place greater weight on the peers who have earnings profiles similar to or above their own. The analysis further reveals that both non-college and college men’s labor force exit rate increases with the skewness of the state earnings distribution, suggesting that prime-age men’s non-pecuniary return on employment is related to both how much their peers earn and how equitably the earnings are distributed.

This study contributes to two areas of the literature. First, this is the first study to examine the causal relationship between relative earnings and labor force participation. The topic relates in part to an analysis in Council of Economic Advisers (2016) that documents a negative correlation between wage inequality and men’s labor force participation. Building on the CEA’s finding, this study introduces relative earnings into workers’ functions for labor market returns and offers a mechanism that directly connects the rise in wage dispersion with the decline in men’s labor force participation. Earlier studies on wages and labor force participation 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 and again in the 2010s, 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). This study resolves the contradictory evidence around real wages and labor force participation by showing that prime-age men’s labor force exit rate declines with their relative earnings, which fell steadily for less-educated men despite periodic gains in real earnings over the last 40 years. The finding therefore offers a plausible explanation for the long-term trend in US prime-age male labor force participation.

In addition to the research on wages, prior studies on labor demand connect rising import competition with China with the decline in US employment and labor force participation in the 1990s and 2000s (Bernard et al., 2006; Autor et al., 2013; Pierce and Schott, 2016; Acemoglu et al., 2016). 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 the replacement of jobs by new technologies, such as computers and robots. Studies 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 and Kearney, 2020). Researchers have also examined whether changes in men’s labor supply cause declines in labor force participation. 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 reduce work incentives or raise barriers to employment, leading to lower participation rates. Compared with studies of demand-side factors, investigations of labor supply generally find more modest or mixed evidence, and a sizable participation gap remains unexplained.⁴

⁴For more information, Abraham and Kearney (2020) and Aaronson et al. (2014) offer thorough reviews of recent findings concerning aggregate employment and labor force participation decline in the United States.

As its second contribution, the study extends the literature on relative earnings. Recent studies using randomized controlled trials show causal evidence that relative earnings increase a worker's job satisfaction (Card et al., 2012) and productivity (Breza et al., 2018), as well as labor supply in a lab setting (Bracha et al., 2015). These findings corroborate earlier evidence of associations between relative earnings and a worker's job satisfaction (Clark and Oswald, 1996), quit rate (Brown et al., 2008), subjective well-being (Luttmer, 2005; Ferrer-i Carbonell, 2005), marriage rate (Watson and McLanahan, 2011), women's employment decision (Neumark and Postlewaite, 1998), and health outcomes (Marmot, 2006; Eibner and Evans, 2005; Daly et al., 2013). These results indicate consistently that a worker's non-pecuniary return on work is partially tied to their relative earnings. This relationship supplements the traditional neoclassical labor supply model, in which a worker's utility function depends only on the real return, and has gained increasing acknowledgement in mainstream economic theory. Yet, the literature offers limited insight into the relationship between relative earnings and aggregate labor market outcomes. Extending this literature, this study presents the first evidence that relative earnings affect men's labor force participation choices and carry macroeconomic implications beyond individual workers' well-being.

The paper proceeds as follows. Section 2 explains the rationale for using labor force exit as the outcome variable in this study. Section 3 describes the trend decline in less-educated men's relative labor market status over the past four decades. Section 4 outlines the study's theoretical framework and empirical estimation strategy. Section 5 summarizes the regression variables and their data sources. Section 6 reports the estimation results and discusses the key findings. Section 7 concludes the paper.

2 Why Study The Labor Force Exit Rate?

This study uses labor force exit rate as the primary outcome variable. While the exit rate is not equivalent to the labor force participation rate, a closer look at the data reveals that it is a practical proxy for understanding the dynamic change in labor force participation but offers some advan-

tages over investigating labor force participation rate directly.

Intuitively, the change in labor force participation rate depends on two factors: the outflow from the labor force and the inflow to the labor force. The outflow measures the share of workers leaving the labor force in each period, while the inflow measures the share of workers returning to the labor force after a previous exit event. From 1980 to 2019, prime-age men’s labor force exit rate doubled; meanwhile, the probability of a worker returning to the labor force remained a stable survival function that declines non-linearly with the time elapsed since the exit event.⁵ As a result, changes in prime-age men’s labor force exit rate drove the changes in their labor force participation rate over this period, as evidenced by the overlapping trends between the percentage change in prime-age men’s labor force non-participation rate and labor force exit rate (Figure 2.1).

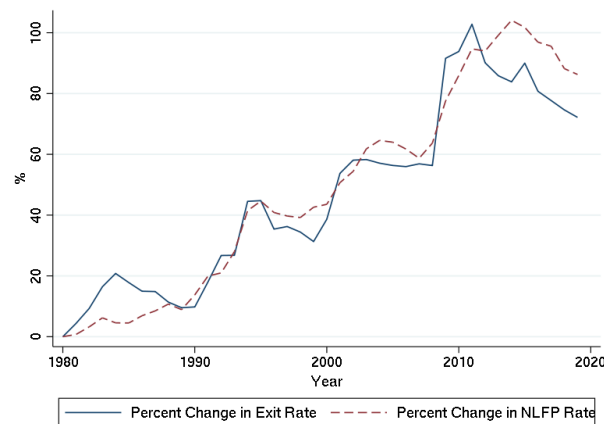


Figure 2.1: Percentage Changes in Prime-Age Men’s Labor Force Non-Participation Rate and Labor Force Exit Rate: 1980–2019

Source(s): IPUMS-CPS 1980–2019.

Note(s): The figure plots the percentage change in the labor force non-participation rate and labor force exit rate for prime-age US men from 1980 to 2019.

The key to understanding changes in the prime-age male labor force participation rate therefore lies in understanding the factors affecting men’s labor force exit decision. However, this task is not as straightforward as it may seem, because the observed labor force participation rate is a function of workers’ labor force exit decisions in both the current and preceding periods. Failing to address these lagged effects could inadvertently introduce biases into the regression estimates.

⁵See Appendix A.1 for a technical discussion of prime-age men’s labor force participation rate and labor force exit rate from 1980 to 2019.

To illustrate this, I start with a simple accounting equation between labor force participation rate and labor force exit rate. Assume that all prime-age men begin in the labor force and can transition in and out of the labor force afterward.⁶ The share of the prime-age men in the labor force, $LFPR_t$, is a function of the labor force exit rate, ε_t , measuring the share of the population exiting the labor force at t and $s(\tau)$, a time-invariant survival function of the probability that a worker remains outside of the labor force for longer than τ periods, where $s(0) = 1$ and $s(\tau) \rightarrow 0$ as $\tau \rightarrow T$, and their lagged terms:⁷

$$LFPR_t = 1 - \sum_{k=0}^t \varepsilon_{t-k} \cdot s(k). \quad (2.1)$$

For the purpose of this example, assume that the labor force exit rate is a function of the expected earnings of workers, such that

$$\varepsilon_t = \alpha + \beta\omega_t. \quad (2.2)$$

Equation 2.1 and 2.2 imply that the labor force participation rate can be re-written as a function of the expected earnings and its lagged terms,

$$LFPR_t = 1 - \sum_{k=0}^t (\alpha + \beta\omega_{t-k}) \cdot s(k). \quad (2.3)$$

If we regress the labor force participation rate on the expected earnings but omit the lagged terms in the estimating equation, the size of the estimated coefficient will depend on the serial correlation of the expected earnings. In the extreme case where the expected earnings are perfectly correlated across periods, the labor force participation rate converges to a function of the current expected

⁶Some prime-age men may have never entered the labor force, but the share never in the labor force was negligible during the study period.

⁷Over time, exiting workers either return to the labor force, age out of the prime-age range, or die. Therefore, the survival function reaches zero within a finite time horizon T .

earnings,

$$LFPR_t = 1 - \alpha \sum_{k=0}^t s(k) - \beta \sum_{k=0}^t s(k) \cdot \omega_t. \quad (2.4)$$

In this case, the coefficient on the expected earnings reflects the sum of the contemporary and lagged effects of the expected earnings on the exit rate, weighted by the survival rate, and is several magnitudes larger than β . Using the labor force participation rate as the outcome variable therefore inflates the earnings coefficient and it can be challenging to interpret the coefficient in terms of its implications on the exit and entry margin of workers.

In light of this issue, the empirical analysis in this study uses the labor force exit rate, rather than the labor force participation rate, as the primary outcome variable. Nevertheless, the qualitative findings remain highly consistent regardless of the choice of outcome variables. Quantitatively, however, as anticipated, using the labor force participation rate as the outcome variable yields substantially larger coefficients in absolute terms. For readers interested in comparing the results, Section 6.4 provides the corresponding estimation results when the labor force participation rate is used as the outcome variable.

3 The Decline in Less-Educated Men’s Labor Market Status

From 1980 to 2019, the mean log weekly earnings of US prime-age non-college men declined by 14 log points. In comparison, the mean log earnings for their college-educated counterparts rose by 14 log points, while college-educated and non-college women’s earnings grew by 37 and 13 log points, respectively. These divergent trends reflect both an increase in the college wage premium and an improvement in gender pay equity. Over the same period, the share of prime-age workers with a four-year college degree grew from 24 percent to 43 percent, and women’s employment share grew from 42 percent to 47 percent.⁸ As a result of these earnings and compositional changes, the earnings of non-college men steadily declined compared to their peers.

Figure 3.1 plots relative earnings over time for US men and women by college education status.

⁸Author calculation based on the Current Population Survey.

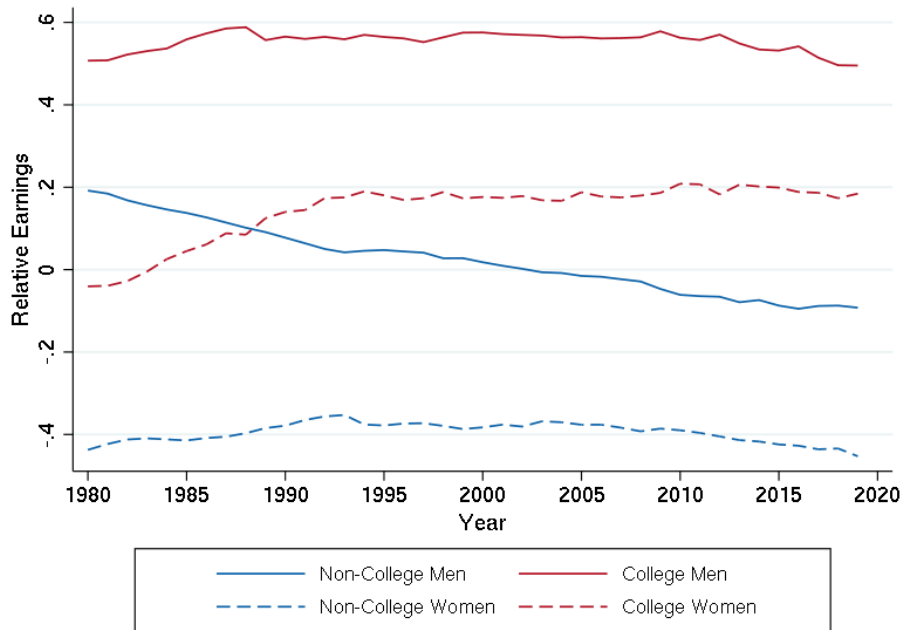


Figure 3.1: Changes in Relative Earnings

Source(s): CPS Earner Study 1980-2019.

Note(s): The figure plots US prime-age workers' relative earnings over time. Relative earnings are measured by a worker group's mean log weekly earnings minus the mean log weekly earnings across all US prime-age workers.

Relative earnings are defined as a worker group's mean log weekly earnings minus the mean log weekly earnings across all prime-age workers. Compared to their 1980 levels, the relative earnings of non-college men have followed a consistent downward trend over the past four decades, despite occasional real earnings gains. By contrast, the relative earnings of college-educated men and non-college women have remained largely stagnant, while the relative earnings of college-educated women have grown and surpassed those of non-college men during this period.

A growing economic literature finds that workers form earnings expectation by peer comparison, such that their productivity or job satisfaction increases with their relative earnings instead of real earnings alone (Card et al., 2012; Breza et al., 2018; Bracha et al., 2015; Clark and Oswald, 1996; Brown et al., 2008). In these studies, workers evaluate their income against a reference group in their immediate surroundings, such as family members, colleagues, or fellow experiment participants. However, with advances in information technology and mobility, a worker's perceived status may also depend on their relative position within a broader pool of workers in their city, state, or country. Therefore, the diminishing relative earnings of non-college men could potentially undermine their utility gain from working.

Consistent with this hypothesis, the historical data exhibit an inverse relationship between changes in prime-age men's relative earnings and changes in their labor force exit rate. Panels A and B of figure 3.2 plot the change in college-educated and non-college men's relative earnings and labor force exit rate over time from their base-period values. Compared with college men, non-college men experienced a much steeper fall in relative earnings as well as a more sizable rise in their labor force exit rate, showing a negative correlation between the two.

These patterns, however, are subject to the influence of other confounding factors and may be coincidental. The next section outlines this study's empirical strategy to causally identify whether prime-age men are more inclined to leave the labor force when their expected earnings fall relative to the earnings of other prime-age workers. The strategy leverages panel data estimation techniques, an extensive set of individual- and state-level controls, and earnings information from nearby states to rule out demand-side factors or selection driving the association.

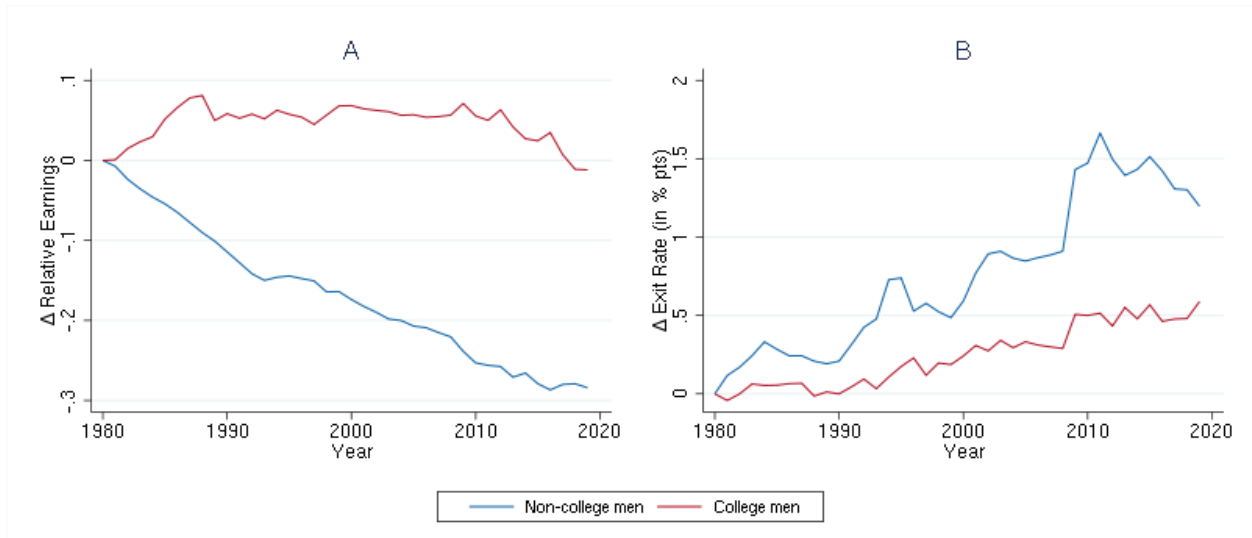


Figure 3.2: Changes in Relative Earnings and Labor Force Exit Rate

Source(s): IPUMS-CPS 1980–2019.

Note(s): Relative earnings are measured by a worker group’s mean log weekly earnings minus the mean log weekly earnings across all US prime-age workers. 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 job searching in each month. Using 1980 as the base period, panel A and B plot the change in US prime-age men’s relative earnings and labor force exit rate over time from their base-period value by college education status.

4 Method

4.1 Conceptual Framework

This section develops a simple conceptual framework to illustrate the relationship between relative earnings and a worker’s labor force exit decision. According to this framework, workers start in the labor force in a fixed location n . They cannot relocate to another location but can choose whether to stay in the labor force for the next period. Their decision depends on the probability of being employed in the next period, governed by $e_{n,i,t}$, and on their expected earnings from employment, $W_{n,i,t}$. Workers evaluate their expected earnings against other workers’ earnings in their location, $\widetilde{W}_{n,t}$, or the reference group earnings.⁹ The utility gain from a worker’s expected earnings is thus discounted by the reference group earnings, and the extent of the discount is governed by a discount factor, γ . When the discount factor approaches zero, workers consider only the real value of their

⁹A labor force participant’s expected earnings are not necessarily their current earnings as some labor force participants are unemployed with zero earnings.

earnings when deciding whether to stay in the labor force. When the discount factor approaches one, workers place greater weight on the relative value of their earnings in the decision. If a worker chooses to leave the labor force, they derive a fixed utility gain, $l_{n,i,t}$, from leisure, which represents the opportunity cost of labor force participation.

Using lowercase letters to denote logs, worker i 's expected return on labor force participation at t is defined as the following:

$$\begin{aligned} R_{n,i,t} &= \ln \frac{W_{n,i,t}}{\widetilde{W}_{n,t}^\gamma} + e_{n,i,t} - l_{n,i,t} \\ &= \omega_{n,i,t} - \gamma \widetilde{\omega}_{n,t} + e_{n,i,t} - l_{n,i,t} \end{aligned}$$

The worker chooses to leave the labor force at t if participation produces a negative return. Denote $\text{Exit}_{n,i,t}$ as an indicator variable equal to one if the worker decides to exit the labor force at t , then

$$\text{Exit}_{n,i,t} = 1 \text{ iff } R_{n,i,t} < 0.$$

Conditional on having the same expected earnings from employment, the model predicts that a worker is more likely to leave the labor force when the reference group earnings increase, the probability of finding a job decreases, or the opportunity cost of participation increases.

4.2 OLS Model

Based on the intuition derived from the conceptual framework, I estimate a reduced-form ordinary least squares (OLS) model of relative earnings and a worker's likelihood of leaving the labor force:

$$\text{Exit}_{n,cz,i,t} = \alpha - \beta \omega_{n,cz,t} + \beta \gamma \widetilde{\omega}_{n,t} + Z_{n,cz,t} + X_{n,t} + K_{i,t} + \delta_n + \lambda_t + \epsilon_{n,cz,i,t}, \quad (4.1)$$

where $Z_{n,cz,t}$, $X_{n,t}$ and $K_{i,t}$ are state-by-skill-, state- and individual-level controls of the worker's job-finding likelihood and the opportunity cost of his labor force participation, and δ_n and λ_t are state and year fixed effects.

In the model, a prime-age man's probability of leaving the labor force falls with his expected earnings $\omega_{n,cz,t}$, measured by the mean log weekly earnings across men with his college attainment status c (college or non-college) in his occupation z in state n , but rises with the reference group earnings $\tilde{\omega}_{n,t}$, defined as the mean log weekly earnings across all prime-age workers in state n . The key parameters of interest are $\beta\gamma$, the coefficient on the reference group earnings, and γ , the ratio between the coefficient on the reference group earnings ($\beta\gamma$) and the coefficient on a worker's expected earnings (β). The ratio γ measures the extent to which workers discount their expected earnings by the reference group earnings.

An endogeneity concern with the OLS model is that the earnings variables may reflect other unobserved changes in labor supply or demand that affect both earnings levels and the likelihood of prime-age men's labor force exit. The direction of the potential bias can go both ways, depending on the specific scenario being considered. For this study, however, the primary threat to identification is the presence of latent skill-biased demand shocks that could simultaneously increase the reference group's earnings and reduce job opportunities for men marginal to the labor force exit decision. Specifically, a skill-biased demand shock can introduce bias into the estimation when it disproportionately expands the employment share of high-skill industries or, within an industry, high-skill workers, while suppressing the job demand for low-skill industries or workers. This selection effect, either across or within industries, results in a bias away from zero for the reference group earnings coefficient.

4.3 2SLS Model

To validate that the observed correlation between prime-age men's labor force exit rate and the reference group earnings is not driven by such selection, I supplement the baseline OLS model with a two-stage least-squares (2SLS) analysis. In the 2SLS analysis, the reference group earnings are instrumented by the inner product between a state's initial industry shares from 1976 to 1979 and the industry-level mean log weekly earnings of workers residing in the other states of the same

census division (d),

$$\tilde{\omega}_{n,t}^{IV} = \sum_{\iota=1}^I \pi_{n,\iota,76-79} \cdot \omega_{\iota,d,-n,t}. \quad (4.2)$$

Similarly, a worker's expected earnings are instrumented by the mean log weekly earnings of prime-age men in his occupation z and college attainment c in the other states of the same census division (d),

$$\omega_{n,cz,t}^{IV} = \omega_{d,-n,cz,t}. \quad (4.3)$$

The underlying rationale behind the instruments is that wages in a state are benchmarked against the wage rates for the same industry/occupation in the broader regional labor market. However, by fixing a state's industry structure and leveraging earnings records from nearby states to predict earnings, this specification removes the influence of selection across industries on the reference group earnings variable, as well as any selection within the industries/occupations that is specific to the state. Because the instruments preserve regional pay differences, the predicted reference group earnings remain subject to biases from selection within industries that are common across states in a region.¹⁰ Therefore, additional tests are needed to provide support for the use of the instrument in identifying meaningful causal effects.

4.3.1 Testing the Identifying Assumptions

The instrument defined in Equation 4.2 departs from standard shift-share instruments by allowing industry-level wage shocks to vary across regions. The validity of the instrument therefore rests on two assumptions. First, a state's initial industry shares and the region it is located in are unrelated to factors predicting growth in its labor force exit rate (Goldsmith-Pinkham et al., 2020). Second, conditional on the model controls, differences in regional industry earnings levels are unrelated to changes in job demand for prime-age men at the margin of labor force exit (Borusyak et al., 2022, 2025).

¹⁰Hypothetically, the risk of endogenous shocks could be reduced by having common shocks and constructing the instrumental variable with nationwide earnings records. However, conditional on the model controls, the instrument based on all other US states has a weak correlation with the observed reference group earnings, rendering it an unsuitable choice.

Building on the suggestions and methods developed by Goldsmith-Pinkham et al. (2020), I test the plausibility of the first assumption by identifying the top industries and regions that are the primary drivers of the 2SLS estimator. I then examine the relationship between the initial shares of the industries or region indicators and selected 1976–1979 state-level characteristics. These characteristics include factors that predict more substantial increases in prime-age male labor force exit rate, such as the share of prime-age men without a four-year college degree, the share who are unmarried, the share who are non-Hispanic Black, and the share who are employed in sales and administrative support occupations. A detailed description of the test and the results are presented in Appendix A.2. Based on the results in Appendix A.2 Table A2, the region indicators and industry shares show weak correlations with these state characteristics.

Next, to evaluate the validity of the second assumption, I investigate the correlation between the instrument and a set of related outcomes that would signal the presence of potential confounders. If the 2SLS estimator is driven by unobserved changes in job demand for prime-age men, an increase in the predicted reference group earnings should correlate not only with more workers leaving the labor force, but also with higher probabilities of job loss or part-time work due to slack labor demand. Conversely, if the relationship reflects prime-age men’s labor supply response to deteriorating labor market status, their reduced labor supply should be observed in both more labor force exits as well as more voluntary quits or reduction of work hours. Regardless of the exact force driving the observed correlation, it is unlikely that it is manifested solely in the labor force exit margins, without also impacting other labor market outcomes for workers who remain in the labor force. Assessing the relationship between the predicted reference group earnings and these outcomes therefore lends further credence to the underlying mechanisms driving the 2SLS estimator.

In this test, I adopt the same 2SLS model as the main analysis but focus on four alternative labor market outcomes for workers who remain in the labor force at t :¹¹

¹¹For the unemployment analysis, I restrict the sample to workers who are in the labor force at time t . For the part-time work analysis, I restrict the sample to workers who are employed at time t . All the model controls are otherwise the same as the main model as specified in Equation 4.1.

1. $\text{Unemployed}^{\text{demand}} = 1$ if a worker is unemployed due to demand-side factors, including job loss, layoff, or the termination of a temporary position.
2. $\text{Unemployed}^{\text{supply}} = 1$ if a worker is unemployed due to voluntary departure from their previous position.
3. $\text{Part-time}^{\text{demand}} = 1$ if a worker works part-time due to demand-side factors, including slack business conditions or limited availability of full-time work.
4. $\text{Part-time}^{\text{supply}} = 1$ if a worker works part-time due to supply-side reasons, such as personal obligations or health limitations.

Section 6.2 Table 6.2 presents the estimation results pertaining to these alternative labor market outcomes. The estimates show that the instrumented reference group earnings have small (or negative) and statistically insignificant correlations with demand-driven unemployment or part-time employment at the 10% significance level. By contrast, they have larger and positive correlations with supply-driven unemployment and part-time employment, significant at 1% and 5% level, respectively. The results do not establish a relationship between the instrument and unobserved demand shocks. Instead, they are consistent with the hypothesis that higher reference group earnings affect prime-age men's labor force exit decision through supply-side adjustment.

Together, these findings provide weak evidence that the instrument captures pre-existing determinants of labor force exit rate growth or confounding labor demand shocks, lending support to the two identifying assumptions for the instrument. Section 6 presents both the OLS and 2SLS estimation results. The 2SLS estimates tend to be larger in absolute magnitude compared to the OLS estimates. This suggests that any potential bias in the OLS model acts to attenuate the coefficients of interest, and that the OLS results likely represent more conservative estimates compared to the 2SLS estimates.

5 Data

5.1 Data Sources and Sample

The primary data source for this study is the US Census Bureau’s Current Population Survey (CPS) (Flood et al., 2020). The CPS gathers employment information on members of approximately 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 match individuals over time and observe their monthly employment transitions during a four-month window. Based on the matched CPS, the analysis sample consists of prime-age men who reported being in the labor force in the previous calendar month over the survey years 1980 to 2019.¹² The individual CPS records are augmented with information on state-level earnings from the CPS Earner Study and state-level controls retrieved from the CPS (Basic Monthly files and Annual Social and Economic Supplements (ASEC)) and administrative sources, including the Federal Housing Finance Agency (FHFA) and the US Department of Labor (USDOL).

5.2 Analysis Variables

Labor force exit is an indicator variable that is equal to 100 (%) if a worker reported having left the labor force since the previous calendar month. In the supplementary analysis of alternative labor market outcomes, the outcome variables are also indicator variables jointly determined by a worker’s employment status in the current calendar month and self-reported reasons of unemployment/part-time work.

A worker’s expected earnings are measured by the mean log weekly earnings in his occupa-

¹²The sample excludes CPS respondents in the first and fifth month-in-sample (MIS), whose labor market activity in the previous calendar month is not observed. A CPS respondent can show up multiple times in the sample in MIS 2 through 4 and MIS 6 through 8. To assess whether this data feature introduces bias into the estimated coefficients, I estimate the baseline model separately for each of the six CPS MIS groups. The estimated coefficients are qualitatively comparable to results based on the pooled sample, but some subtle differences exist across the MIS groups due to panel conditioning effects. See Appendix A.3 for more details.

tion across prime-age men of his college attainment status (college or non-college) in his state of residence. 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).¹³ The reference group earnings are defined as the mean log weekly earnings across all prime-age workers in a worker’s state of residence. In the 2SLS specification pertaining to Equation 4.2 and 4.3, a worker’s expected earnings are instrumented by the mean log weekly earnings across prime-age men of his occupation and college attainment status in other states of the same census division, and the reference earnings are instrumented by the inner products between a state’s initial industry share from 1976 to 1979 and the mean prime-age log weekly earnings of each industry in other states of the same census division based on a harmonized scheme of the IPUMS-CPS *ind1990* variable of 134 industries. All earnings variables are estimated using the CPS Earner Study, weighted by the CPS Earner weight, and adjusted for top-coding and for inflation using the Consumer Price Index for All Urban Consumers (CPI-U).

In addition to the labor market outcome and earnings variables, the model includes individual-level controls retrieved from the CPS on a worker’s age, age squared, educational attainment, marital status, race/ethnicity, and whether the worker was unemployed in the previous calendar month. The model also includes state-by-skill- and state-level covariates to control for spatial variations in a worker’s expected risk of unemployment and the opportunity cost of labor force participation. The state-by-skill-level control includes the unemployment rate of a worker’s occupation across prime-age men of his college attainment status in his state (source: basic monthly CPS). The state-level controls include prime-age male unemployment rate (source: basic monthly CPS), prime-age poverty rate (source: CPS-ASEC), prime-age men’s mean log family income from sources other than their own earnings (source: CPS-ASEC), the share of prime-age men receiving any Social

¹³Due to the nature of the jobs, workers in the aviation transportation industry (SOC: 532) are categorized in the management, professional, public safety group.

Security payments, which primarily reflects access to Disability Insurance (source: CPS-ASEC), the maximum weekly state Unemployment Insurance (UI) benefit (source: USDOL), and the state all-transactions house price index (source: FHFA).

5.3 Summary Statistics

Table 5.1 summarizes the sample of the empirical analysis. Of the 7,903,696 labor force participants in the baseline sample, 119,724 (1.5%) left the labor force in the next calendar month. Workers who left the labor force had lower expected earnings and slightly higher reference group earnings compared with workers who stayed in the labor force. Aside from earnings, the exiters faced higher unemployment rate in their state and skill group and higher poverty rate than the stayers. The two groups worked under otherwise similar state economic and policy environments, with comparable family income from sources other than their own earnings, maximum UI benefit levels and access to Social Security benefits, though the exiters faced modestly higher housing costs. Most of the differences between the stayers and exiters are manifested in their individual characteristics. In terms of worker demographics, the exiters, relative to the stayers, were 0.97 years younger, 14 percent less likely to have a four-year college degree, and 17 percent more likely to be non-white. In terms of previous labor market activities, 39 percent of the exiters were unemployed before they left the labor force compared with 4 percent of the stayers. The pattern is consistent with evidence from longitudinal survey showing that labor force exiters on average have weak labor market attachment and repeated unemployment stretches before leaving the labor force.¹⁴ These demographic differences highlight potential confounders in the analysis and the importance of accounting for individual-level characteristics when estimating the relationship between earnings and labor force participation. All of the mean differences reported in this table between exiters and stayers are statistically significant at the 0.05 level under simple t tests.

¹⁴See Appendix A.4 for a summary of US prime-age men's employment and income pathways to labor force exit based on the Survey of Income and Program Participation (SIPP).

Table 5.1: Sample Summary Statistics by Labor Force Exit Decision

	Stayer (N=7,783,972)		Exiter (N=119,724)	
	Mean	S.D.	Mean	S.D.
A. Relative Earnings				
Expected earnings	6.88	0.33	6.73	0.32
Reference group earnings	6.68	0.10	6.69	0.10
B. State-by-Skill and State Controls				
Prime-age male unemployment rate	5.07	2.10	5.27	2.20
Prime-age male unemployment rate by college and occupation	4.94	3.80	6.48	4.25
Prime-age poverty rate	10.37	2.71	10.84	2.72
Other family income	9.57	0.21	9.57	0.21
Share of prime-age men receiving Social Security	0.03	0.01	0.03	0.01
Maximum UI benefit (in \$1,000)	0.47	0.13	0.47	0.13
Log FHFA House Price Index	5.48	0.50	5.57	0.49
C. Individual Controls				
Age	38.78	8.44	37.81	8.99
Education				
Less than high school	0.12	0.32	0.22	0.42
High school diploma/GED	0.32	0.47	0.37	0.48
Associates degree/some college	0.25	0.43	0.24	0.42
Bachelor's degree or higher	0.31	0.46	0.17	0.38
Race				
Non-Hispanic whites	0.72	0.45	0.55	0.50
Non-Hispanic blacks	0.10	0.30	0.20	0.40
Hispanics	0.13	0.34	0.18	0.38
Others	0.05	0.23	0.08	0.26
Unemployed in the previous month	0.04	0.20	0.39	0.49

Source(s): IPUMS-CPS 1980–2019.

Note(s): The sample consists of prime-age men aged 25 to 54 over the study period from 1980 to 2019. The summary statistics are weighted by CPS individual longitudinal weight for linking between two adjacent calendar months.

6 Estimation Results

6.1 Relative Earnings and Labor Force Exit Rate

Table 6.1 presents the baseline estimation results. Columns 1 to 3 report the OLS estimates, while column 4 provides the 2SLS results. To assess the effects of the control variables, the state-level and individual-level controls are added successively across the model specifications.

Column 1 reports the estimated coefficients and standard errors of the expected and reference group earnings variables, with only state and year fixed effects included as controls. The coefficient on expected earnings of -2.17 indicates that a 10 log point increase in a worker's expected earnings is associated with a 0.217 percentage point decrease in the likelihood of their labor force exit. By contrast, the coefficient on reference group earnings of 1.37 suggests that a 10 log point increase in reference group earnings is associated with a 0.137 percentage point increase in the likelihood of labor force exit.

Column 2 adds state-by-skill- and state-level controls to the model. Incorporating these controls attenuates the absolute value of the coefficient on expected earnings from -2.17 to -1.37, suggesting that a worker's expected earnings are partially correlated with the risk of unemployment and the opportunity cost of labor force participation, which independently affect his labor force exit probability. Specifically, the coefficients on the covariates indicate that prime-age men are more likely to leave the labor force in states with higher unemployment rates for their skill group, lower overall unemployment rates, or higher poverty rates. Other state-level characteristics do not appear to have strong correlations with prime-age men's labor force exit decision.

Column 3 further adds individual-level controls into the model. The results indicate that a man's likelihood of labor force exit increases when he is younger, identifies as non-Hispanic Black, lacks a high school diploma, or was unemployed in the previous month. However, the inclusion of these individual-level controls has only modest effects on the earnings coefficients. After accounting for both state-level and individual-level covariates, a 10 log point increase in a worker's expected earnings is associated with a 0.118 percentage point decline in the likelihood of labor

Table 6.1: Relative Earnings and Prime-age Men's Labor Force Exit Rate

	(1)	(2)	(3)	(4)
A. Relative Earnings				
Expected earnings	-2.17*** (0.08)	-1.37*** (0.06)	-1.23*** (0.06)	-1.38*** (0.08)
Reference group earnings	1.37*** (0.25)	1.03*** (0.26)	1.18*** (0.26)	1.99* (0.88)
B. State-by-Skill and State Controls				
Prime-age male unemployment rate by college and occupation		0.12*** (0.00)	-0.01+ (0.00)	-0.01** (0.00)
Prime-age male unemployment rate		-0.06*** (0.01)	-0.04*** (0.01)	-0.04*** (0.01)
Prime-age poverty rate		0.03*** (0.01)	0.02*** (0.01)	0.03*** (0.01)
Share of prime-age men receiving Social Security		0.98 (0.84)	1.12 (0.85)	0.96 (0.84)
Maximum UI benefit		-0.20 (0.17)	-0.32+ (0.17)	-0.36+ (0.19)
Log FHFA House Price Index		-0.03 (0.09)	-0.09 (0.09)	-0.14 (0.11)
Other family income		0.15+ (0.08)	0.15+ (0.08)	0.16+ (0.08)
C. Individual Controls				
Age			-0.25*** (0.01)	-0.25*** (0.01)
Age-squared			0.00*** (0.00)	0.00*** (0.00)
High school diploma/GED			-0.77*** (0.09)	-0.77*** (0.09)
Associates degree/some college			-0.91*** (0.09)	-0.90*** (0.09)
Bachelor's degree or higher			-0.71*** (0.07)	-0.65*** (0.06)
Non-Hispanic blacks			1.26*** (0.13)	1.25*** (0.13)
Hispanic			-0.03 (0.09)	-0.04 (0.09)
Others			0.73*** (0.07)	0.73*** (0.07)
Unemployed in the previous month			11.51*** (0.21)	11.51*** (0.21)
Method	OLS	OLS	OLS	2SLS
Observations	7,903,696	7,903,696	7,903,696	7,903,696
R^2	0.00	0.00	0.04	0.04
$\hat{\gamma}$	0.63	0.75	0.96	1.45
First-stage Kleibergen-Paap rk Wald F statistic				11.69

Source(s): IPUMS-CPS 1980–2019.

Note(s): The dependent variable of the regression is the probability of labor force exit measured in % pts. The sample consists of US prime-age men for the 1980–2019 period. All models include state and year fixed effects. The results are weighted by the CPS longitudinal weight for linking individuals between adjacent months. Standard errors are clustered at the state level.

† for $P < 0.10$, * for $P < 0.05$, ** for $P < 0.01$, *** for $P < 0.001$.

force exit, while a 10 log point rise in the reference group earnings is linked to a 0.123 percentage point increase. The symmetry of these point estimates suggests that prime-age men nearly fully discount their expected earnings by the reference group earnings, with an indicated discount factor of 0.96. Column 4 presents the corresponding 2SLS estimation results with the full set of model controls. The findings are not statistically different from the OLS results in column 3 but the point estimates and standard errors for the earnings variables are moderately larger under this specification. The consistency between the OLS and 2SLS results provides weak evidence that the correlation between prime-age men's labor force exit rate and the earnings variables is driven by state-specific selection, changes to a state's industrial structure, or latent demand shocks that adversely affect prime-age men's employment opportunities.

Together, results from Table 6.1 indicate that prime-age men discount their expected earnings by the value of their peers' earnings when evaluating the return on labor force participation, with an estimated discount factor of 0.63 to 1.45. The range of the estimated discount factor tracks closely with the literature's findings on relative earnings and the subjective well-being of workers. Using a similar research design, prior studies find that the earnings discount factor, defined as the absolute value of the ratio between the coefficient on the reference group earnings and the coefficient on a worker's own expected earnings, lands between 0.40 and 2.17 (Clark and Oswald, 1996), 0.73 and 1.23 (Ferrer-i Carbonell, 2005), 0.82 and 2.23 (Luttmer, 2005), or 0.13 and 0.91 (Brown et al., 2008). Notably, the lower (< 0.50) estimates of the discount factor in the literature correspond to outcome variables that directly measure a worker's satisfaction with their pay level, while the larger estimates correspond to outcome variables measuring a worker's satisfaction with the non-pecuniary return on employment, such as respect, happiness, achievement, or overall satisfaction.¹⁵ The comparability between the estimated discount factors in this paper and past findings on the non-pecuniary return on employment suggests that a potential behavioral mechanism linking, at least in part, relative earnings to prime-age men's labor force exit decision occurs through the

¹⁵Luttmer (2005) notes that the coefficient on a person's own income can be biased downward when the regression model includes other individual-level controls that are correlated with personal income, leading to discount factors that significantly exceed one in some cases.

Table 6.2: Relative Earnings and Related Labor Market Outcomes

	Supply-related			
	Unemployment		Part-time work	
	(1)	(2)	(3)	(4)
Expected earnings	-0.34*** (0.03)	-0.41*** (0.04)	-5.01*** (0.25)	-5.57*** (0.24)
Reference group earnings	0.55*** (0.11)	0.89** (0.30)	2.24*** (0.40)	2.55* (1.19)
Method	OLS	2SLS	OLS	2SLS
Observations	7,783,972	7,783,972	7,456,790	7,456,790
R^2	0.04	0.04	0.01	0.01
$\hat{\gamma}$	1.62	2.14	0.45	0.46
First-stage Kleibergen-Paap rk Wald F statistic	11.73		11.67	
	Demand-related			
	Unemployment		Part-time work	
	(5)	(6)	(7)	(8)
Expected earnings	1.18*** (0.06)	1.36*** (0.07)	-4.00*** (0.17)	-4.02*** (0.20)
Reference group earnings	-0.95** (0.27)	-0.96 (0.63)	0.66 (0.54)	1.83 (1.23)
Method	OLS	2SLS	OLS	2SLS
Observations	7,783,972	7,783,972	7,456,790	7,456,790
R^2	0.35	0.35	0.04	0.04
$\hat{\gamma}$	0.80	0.71	0.16	0.45
First-stage Kleibergen-Paap rk Wald F statistic	11.73		11.67	

Source(s): IPUMS-CPS 1980–2019.

Note(s): The dependent variables are as follows: (1)/(2) is the probability of supply-related unemployment due to voluntary job separation. (3)/(4) is the probability of part-time work for non-economic personal reasons. (5)/(6) is the probability of demand-related unemployment due to involuntary job loss. (7)/(8) is the probability of part-time work for demand-related economic reasons. All dependent variables measured in % pts. The sample consists of US prime-age men for the 1980–2019 period. All models include state- and individual-level controls as well as state and year fixed effects. The results are weighted by the CPS longitudinal weight for linking individuals between adjacent months. Standard errors are clustered at the state level.

† for $P < 0.10$, * for $P < 0.05$, ** for $P < 0.01$, *** for $P < 0.001$.

non-pecuniary return on employment. Lower relative earnings could reduce a worker’s sense of respect, achievement, and equity, in turn undermining their labor force participation incentives.

6.2 Relative Earnings and Related Labor Market Outcomes

To supplement the baseline result and as a test of the identifying assumption for the instrument, I investigate the relationship between relative earnings and a set of related labor market outcomes

for workers who stay in the labor force.¹⁶ Evidence from longitudinal data indicates that labor force exit frequently follows a period of weakened attachment to the labor force, characterized by reduced work hours and repeated unemployment stretches.¹⁷ It is therefore unlikely to observe an increase in labor force exit rate without also seeing increases in part-time work and/or unemployment.

Using the "reason for unemployment/part-time work" questions in the CPS, I distinguish between supply-driven and demand-driven unemployment and part-time work. If the observed relationship between reference group earnings and labor force exit rate reflects prime-age men's labor supply response to their declining non-pecuniary return to employment, the contracted supply should be evident in more voluntary hour reductions and job departures as well. Conversely, if the relationship is driven by latent demand shocks, higher reference group earnings should positively correlate with more demand-driven unemployment and involuntary hour cuts. Differentiating between these supply-side and demand-side mechanisms provides additional empirical tests on the factors underlying the relationship between reference group earnings and prime-age men's labor force exit decision.

Table 6.2 presents the estimation results. The top panel shows the OLS and 2SLS results pertaining to the supply-related outcomes, while the bottom panel shows the demand-related outcomes. According to columns 1 to 4, conditional on a worker's own expected earnings, his probability of voluntarily leaving his job or working a reduced-hour schedule increases when the reference group earnings rise. By contrast, the results in columns 5 to 8 indicate that the reference group earnings have smaller (or negative) and statistically insignificant association with involuntary job losses or hour reductions at the 10 percent level in both the OLS and 2SLS settings. The findings provide coherent evidence supporting the hypothesis that prime-age men consider their earnings in relation to their peers' when making labor supply decisions, along both the participation margin and the intensive margin.

¹⁶See Section 4.3 for the rationale behind using this analysis as a test of the identifying assumptions for the instrumental variable.

¹⁷See Appendix A.4.

6.3 Alternative Reference Group Earnings

The results presented in Tables 6.1 and 6.2 establish the relationship between prime-age men's labor force exit decision and the mean log earnings of their peers. To gain deeper understanding of how workers compare themselves to other workers, the study then tests worker sensitivity to alternative measurements of reference group earnings, separately for men with and without college degrees. Table 6.3 reports the estimation results.

In columns 1 through 4 and 6 through 9, the reference group earnings are represented by the 90th, 75th, 50th, and 25th percentiles of the log prime-age earnings in a worker's state of residence. Since state-level earnings closely follow a log-normal distribution, the distribution can be sufficiently captured by the mean of the log earnings and the standard deviation of the log earnings. In columns 5 and 10, the estimation results pertain to a modified model in which the reference group earnings are measured by these two moments instead of a single earnings quantile.

For both non-college and college men, the estimated coefficients on a worker's expected earnings are qualitatively equivalent across the five models. However, holding own expected earnings constant, workers show different sensitivities to the different quantiles of the state earnings distribution. Compared with the median earnings, non-college men have more muted response to changes in the higher or lower end of the state earnings distribution, as evidenced by the smaller point estimates of the 90th, 75th, and 25th earnings percentiles. By contrast, college men's labor force exit rate has a positive correlation with the 90th earnings percentile, but the correlation declines as the reference group earnings move down to lower quantiles of the state earnings distribution. These differences imply that the earnings of a median worker are more salient or relevant to non-college men, whereas college men put more weight on other high earners when evaluating their expected earnings.

Furthermore, conditional on the same mean log earnings, non-college men are more likely to leave the labor force when the state earnings distribution becomes more skewed, as indicated by the positive coefficient on the standard deviation of the log state earnings in column 5. In comparison, college men have low sensitivity to the mean log earnings but their labor force exit rate also

Table 6.3: Alternative Reference Group Earnings

	Non-college men				
	(1)	(2)	(3)	(4)	(5)
Expected earnings	-1.12*** (0.08)	-1.13*** (0.08)	-1.15*** (0.08)	-1.14*** (0.08)	-1.15*** (0.08)
Reference Group Earnings					
P90 log state earnings	0.69* (0.32)				
P75 log state earnings		1.16** (0.34)			
P50 log state earnings			1.42*** (0.31)		
P25 log state earnings				0.94*** (0.20)	
Mean log state earnings					1.59*** (0.32)
S.D. log state earnings					0.75** (0.27)
Method	OLS	OLS	OLS	OLS	OLS
Observations	5459380	5459380	5459380	5459380	5459380
R^2	0.05	0.05	0.05	0.05	0.05
gamma	0.62	1.03	1.23	0.82	1.39
	College men				
	(6)	(7)	(8)	(9)	(10)
Expected earnings	-1.01*** (0.05)	-1.00*** (0.05)	-1.00*** (0.05)	-1.00*** (0.05)	-1.00*** (0.05)
Reference Group Earnings					
P90 log state earnings	0.51* (0.24)				
P75 log state earnings		0.45+ (0.23)			
P50 log state earningsj			0.28 (0.26)		
P25 log state earnings				-0.17 (0.21)	
Mean log state earnings					0.37 (0.29)
S.D. log state earnings (sigma)					0.77*** (0.20)
Method	OLS	OLS	OLS	OLS	OLS
Observations	2444316	2444316	2444316	2444316	2444316
R^2	0.03	0.03	0.03	0.03	0.03
gamma	0.51	0.45	0.28	0.17	0.37

Source(s): IPUMS-CPS 1980–2019.

Note(s): The dependent variable of the regression is the probability of labor force exit measured in % pts. The sample consists of US prime-age men for the 1980–2019 period. All models include state- and individual-level controls as well as state and year fixed effects. The results are weighted by the CPS longitudinal weight for linking individuals between adjacent months. Standard errors are clustered at the state level.

† for $P < 0.10$, * for $P < 0.05$, ** for $P < 0.01$, *** for $P < 0.001$.

Table 6.4: Relative Earnings and Prime-age Men’s Labor Force Participation Decision

	Exit Decision		Participation Decision	
	(1)	(2)	(3)	(4)
Expected earnings	-1.62*** (0.15)	-3.14*** (0.37)	8.27*** (1.33)	21.84*** (2.80)
Reference group earnings	1.32*** (0.32)	2.97** (1.11)	-7.90*** (1.47)	-20.94*** (5.34)
Method	OLS	2SLS	OLS	2SLS
Observations	7,910,756	7,910,756	8,599,023	8,599,023
R^2	0.01	0.01	0.04	0.04
$\hat{\gamma}$	0.81	0.95	0.96	0.96
First-stage Kleibergen-Paap rk Wald F statistic		12.76		12.59

Source(s): IPUMS-CPS 1980–2019.

Note(s): The dependent variables of the regressions are the probability of labor force exit (1) (2) and the probability of labor force participation (3) (4), both measured in % pts. The sample consists of US prime-age men for the 1980–2019 period. All models include state- and individual-level controls as well as state and year fixed effects. The results are weighted by the CPS longitudinal weight for linking individuals between adjacent months. Standard errors are clustered at the state level.

† for $P < 0.10$, * for $P < 0.05$, ** for $P < 0.01$, *** for $P < 0.001$.

increases with the skewness of the state earnings distribution (column 10). The skewness of the earnings distribution reflects the dispersion of the upper earnings in a state. This finding is consistent with the argument that college men’s labor force exit rate is disproportionately influenced by the pay distribution of other high earners in the economy. It also suggests that non-college men’s labor force participation incentives are related to both how much a median peer earns and how equitably the earnings are distributed between the high and median earners.

6.4 Relative Earnings and Labor Force Participation Rate

Section 2 discusses the merits of using labor force exit rate as the outcome variable for this study. Nevertheless, the key question this study aims to answer is whether the declining relative earnings of prime-age men contributed to their falling labor force participation over the last four decades. To this end, it is helpful to show the corresponding results on the relationship between relative earnings and labor force participation.

In this supplementary analysis, I use a modified version of Equation 4.1 as the estimating equation. Because information on occupation affiliation is not available for individuals who do not have an observed labor force participation spell, in this modified model, a worker’s expected

earnings are measured by the mean log weekly earnings across prime-age men of his education attainment in his state of residence. For the same reason, the skill-specific state unemployment rate in the model is measured by education-specific state unemployment rate, and the individual control of unemployment status in the previous month is dropped from the model. Aside from these changes, the model covariates are the same as Equation 4.1. To facilitate a fair comparison, I present the estimation results under this modified model using both indicators for labor force exit and labor force participation as outcome variables.

Table 6.4 columns 1 and 2 report the estimation results pertaining to the outcome of labor force exit under the modified model. Compared to the baseline results in Table 6.1, the modified model yields larger absolute values of the coefficients on expected and reference group earnings. Additionally, under the modified model, the coefficients on a worker's own expected earnings are relatively larger in absolute terms, as reflected by the smaller wage discount factors. These differences likely reflect the model's less precise control on unemployment risks, which are correlated with a worker's own expected earnings. The findings are otherwise comparable between the two models.

Moving to columns 3 and 4, these present the estimation results where the outcome is labor force participation. According to the OLS and 2SLS estimates, a 10 log point increase in a worker's relative earnings is predicted to lead to a 0.8 to 2.2 percentage point decline in their labor force participation rate. The magnitudes of the coefficients are 6 to 7 times larger compared to the results in columns 1 and 2. This difference confirms that using labor force participation as the outcome variable can significantly inflate the earnings coefficients, as they reflect the sum of the contemporary and lagged effects of earnings on workers' labor force exit decision (see Section 2). Notably, despite differences in the absolute magnitudes of the earnings coefficients, the symmetry between the coefficients on expected earnings and reference group earnings remains coherent across the model specifications, with an estimated discount factor ranging between 0.81 and 0.96.

These empirical analysis results provide coherent evidence that the labor force participation decision of prime-age men is affected by their earnings relative to others in the economy, rather

than just their own expected earnings level. It is hence important to place prime-age men's earnings in the context of the economy's entire earnings distribution to understand how their labor force participation incentives may change in the absence of significant changes in their real earnings.

7 Conclusion

The steady decline in labor force participation among prime-age US men over the last 40 years has spurred much interest and discussion among scholars. Earlier research finds that changes in real wage rate or supply-side factors cannot sufficiently account for the magnitude and persistence of the decline. This paper posits that the growing earnings dispersion and the resulting drop in the relative earnings of less-educated men explain a significant portion of the observed decline. Specifically, from 1980 to 2019, the mean log earnings of prime-age non-college men fell by 0.14 log points, while the mean log earnings across all prime-age workers increased by 0.15 log points. According to the baseline OLS and 2SLS estimates (Table 6.1, column 3 and 4), these changes are associated with a 0.34 to 0.48 percentage point increase in non-college men's labor force exit rate, accounting for 24 to 33 percent of the total rise in the exit rate of prime-age men over this period. These results suggest that the deteriorating labor market status of less-educated men is a plausible key factor driving prime-age men's declining labor force participation. The results of this study point to an underexamined area in the existing literature. Despite the widespread interest in earnings dispersion, its impact on the aggregate labor market behavior of workers remains unclear. If the widening wage gap between high and low earners influences men's labor supply incentives, then earnings dispersion may carry broader implications for the economy than previously acknowledged.

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Appendix

A.1 A Technical Note on Labor Force Participation and Exit Rate

Equation 2.1 indicates that labor force non-participation rate, defined as the share of population not in the labor force, can be written as a function of the labor force exit rate, ε_t , measuring the share of the population exiting the labor force at t and $s(\tau)$, the survival function of the probability that a worker remains outside of the labor force for longer than τ periods, where $s(0) = 1$ and $s(\tau) \rightarrow 0$ as $\tau \rightarrow T$, and their lagged terms:

$$NLFPR_t = \sum_{k=0}^t \varepsilon_{t-k} \cdot s(k) \quad (\text{A1})$$

Based on the equation, changes in either the survival function $s(\tau)$ or the labor force exit rate ε_t can lead to changes in the non-participation rate. However, according to longitudinal data from the CPS, $s(\tau)$ has been largely stable over the last 40 years. Figure A1 plots the labor force exit survival function according to the linked IPUMS-CPS basic monthly data series from 1980 through 2019. The function shifts slightly between 2010 and 2019, likely reflecting worker movement after the Great Recession, but otherwise exhibits minimal changes over time. Consequently, empirically, changes in the labor force exit rate drive the changes in prime-age men's non-participation rate. A similar observation was recently reported in Elsby et al. (2019).

Assume hereafter that $s(\tau)$ is time invariant. The changes in the non-participation rate from period t' to t , where $T < t' < t$, can be written as:¹⁸

$$\begin{aligned} \Delta NLFPR_{t',t} &= \sum_{k=0}^{t'} (\varepsilon_{t-k} - \varepsilon_{t'-k}) \cdot s(k) + \sum_{k'=t'+1}^t \varepsilon_{t-k'} \cdot s(k') \\ &= \sum_{k=0}^{t'} (\varepsilon_{t-k} - \varepsilon_{t'-k}) \cdot s(k) \end{aligned} \quad (\text{A2})$$

¹⁸The second part of equation A2 can be dropped because of the assumption that $T < t' < t$ so that $s(\tau) \rightarrow 0 \forall \tau > t'$.

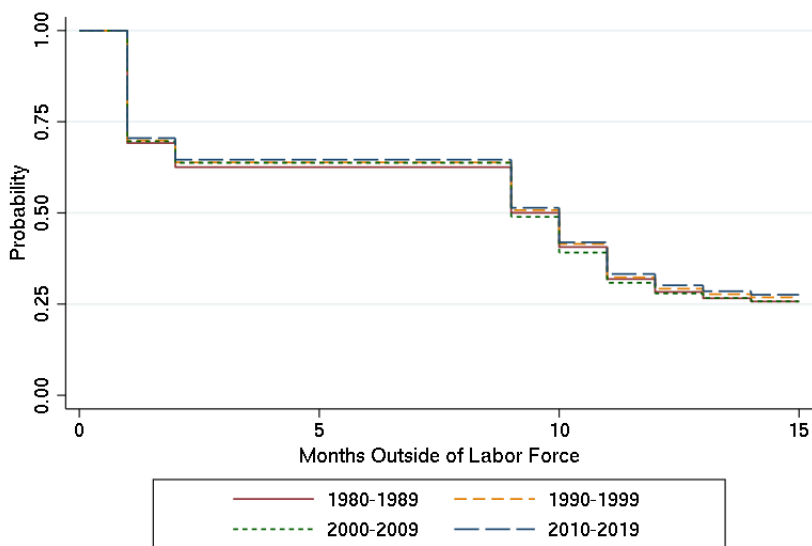


Figure A1: Labor Force Exit Survival Function, 1980–2019

Source(s): IPUMS-CPS 1980-2019.

Note(s): The figure plots the labor force exit survival function using the linked IPUMS-CPS basic monthly data series. The survey design of the Current Population Survey does not allow the observation of the employment transitions between the 5th and the 12th month during the 16-month survey period, which is the reason the segment between month 3 and month 9 appears flat.

In other words, the change in $NLFPR$ from t' to t depends on $\Delta\varepsilon_{t',t}$, the changes in exit rate from t' to t , and the changes in the exit rate in previous periods, $\Delta\varepsilon_{t'-k,t-k}$. An increase in the exit rate hence has both a contemporary effect on the current non-participation rate and lagged effects on the future non-participation rates from workers staying outside of the labor force past the current period.

Equation A1 also indicates that the percentage change in the non-participation rate directly relates to the percentage change in the exit rate. The percentage change in the non-participation rate from period t' to t , where $T < t' < t$, can be derived as:

$$\% \Delta NLFPR_{t',t} = \frac{\sum_{k=0}^{t'} (\varepsilon_{t-k} - \varepsilon_{t'-k}) \cdot s(k)}{\sum_{k=0}^{t'} \varepsilon_{t'-k} \cdot s(k)} \quad (\text{A3})$$

Under the assumption that the exit rate grows at a fixed rate r , such that $\varepsilon_{t+k} = \varepsilon_t(1+r)^k$, the percentage change in the non-participation rate is equal to the percentage change in the exit rate:

$$\begin{aligned} \% \Delta NLFPR_{t',t} &= \frac{\sum_{k=0}^{t'} (\varepsilon_{t-k} - \varepsilon_{t'-k}) \cdot s(k)}{\sum_{k=0}^{t'} \varepsilon_{t'-k} \cdot s(k)} \\ &= \frac{\sum_{k=0}^{t'} \varepsilon_{t'-k} [(1+r)^{t-t'} - 1] \cdot s(k)}{\sum_{k=0}^{t'} \varepsilon_{t'-k} \cdot s(k)} \\ &= (1+r)^{t-t'} - 1 \\ &= \% \Delta \varepsilon_{t',t} \end{aligned} \quad (\text{A4})$$

A.2 Relationship Between State Characteristics and IV Exposure Shares

In the 2SLS analysis, the reference group earnings are instrumented by the inner product between a state's initial industry shares in the period between 1976 and 1979 and the industry-level wage rates in their region, defined by census divisions. This instrument can be viewed as a special case of standard shift-share instruments. In this case, the exposure shares of the instrument are a state's initial region-industry shares, defined as the interaction between a state's initial industry shares and region dummies and the earnings shocks are common within each region-industry in each period. The instrument is therefore analogous to instrumenting using a state's initial region-

industry shares and their interactions with time period dummies. Accordingly, this reconfiguration allows the adoption of test techniques derived by Goldsmith-Pinkham et al. (2020) to identify the key regions or industries associated with the largest Rotemberg weights, which measure their relative importance in determining the 2SLS estimator.

If states in these key regions or with high initial shares of the key industries exhibit other baseline characteristics that predict more substantial growth in labor force exit rate, it would suggest that the exposure shares are not exogenous to changes in the exit rate. To test this, I follow the approach suggested by (Goldsmith-Pinkham et al., 2020) and regress the initial industry shares (or region indicators) on selected 1976–1979 state-level characteristics that could predispose the states to experience more substantial increases in prime-age male labor force exit rate. These characteristics include the share of prime-age men without a four-year college degree, not married, non-Hispanic Black, or employed in sales and administrative support occupations.¹⁹

For computational feasibility, I adopt a modified sample and model specification for this exercise. The sample is collapsed from individual-level to state-by-skill level observations, spanning the years 1980, 1985, 1990, 1995, 2000, 2005, 2010, 2015, and 2019. In this simplified model, the only endogenous variable in the model is the reference group earnings, with the expected earnings variable included as an exogenous covariate. The instrument for the reference group earnings is constructed with a more parsimonious industry classification scheme of 35 industries. Despite the modification, the OLS and 2SLS results of the original and modified model remain substantially comparable (Table A1) and this consistency allows me to draw insights from the simplified model to inform the primary analysis.

Table A2 presents the test results. In panel A, the dependent variables are a state’s initial shares of the top four industries with the largest Rotemberg weights. In panel B, the dependent variables are indicator variables of the top four census divisions. Across the set of key industries and regions, there appears to be limited correlations between the exposure shares and these state characteris-

¹⁹These four factors are selected from a broader set of 11 demographic characteristics spanning race/ethnicity, marital status, educational attainment, and occupations. A OLS regression analysis indicates that this subset of characteristics is associated with statistically significant higher growth trends in the prime-age male labor force exit rate.

tics. In a few cases, negative correlations are observed, suggesting that the direction of bias, if any, would likely attenuate the 2SLS estimator towards zero. The only exposure share positively associated with factors predicting larger exit growth is the indicator for the Middle Atlantic census division. Excluding states in the Middle Atlantic census division from the sample, however, has limited impact on the OLS and 2SLS estimators. These findings provide supportive evidence for the exogeneity of the exposure shares of the instrument.

Table A1: Results from Original and Modified Models

	Original		Modified	
	(1)	(2)	(3)	(4)
Expected earnings	-1.23*** (0.06)	-1.38*** (0.08)	-1.26*** (0.14)	-1.29*** (0.14)
Reference group earnings	1.18*** (0.26)	1.99* (0.88)	1.53** (0.35)	2.12** (0.66)
Method	OLS	2SLS	OLS	2SLS
Observations	7,903,696	7,903,696	4,583	4,583
R^2	0.04	0.04	0.57	0.53
$\hat{\gamma}$	0.96	1.45	1.22	1.64
First-stage Kleibergen-Paap rk Wald F statistic		11.69		93.41

Source(s): IPUMS-CPS 1980–2019.

Note(s): The dependent variables of the regressions are the probability of labor force exit measured in % pts. The sample consists of US prime-age men for the 1980–2019 period. All models include state- and individual-level controls as well as state and year fixed effects. The results are weighted by the CPS longitudinal weight for linking individuals between adjacent months. Standard errors are clustered at the state level.

† for $P < 0.10$, * for $P < 0.05$, ** for $P < 0.01$, *** for $P < 0.001$.

Table A2: Relationship Between Instrument Shifters and State Characteristics

	A. Top Four Industries			
	(1) Professional Services	(2) Retail Trade	(3) Construction	(4) Agriculture
Non-college	-0.31*** (0.06)	0.01 (0.06)	-0.11* (0.06)	-0.14 (0.12)
Non-married	0.02 (0.06)	0.04 (0.06)	-0.07 (0.06)	-0.15 (0.12)
Non-Hispanic Black	0.03 (0.03)	-0.05+ (0.03)	0.02 (0.03)	-0.09+ (0.05)
Sales and administration	-0.08 (0.16)	-0.07 (0.17)	-0.78*** (0.15)	-0.97** (0.32)
Observations	51	51	51	51
R^2	0.52	0.09	0.41	0.31
	B. Top Four Census Divisions			
	(5) New England	(6) East South Central	(7) Middle Atlantic	(8) West South Central
Non-college	-1.49 (1.19)	1.31 (1.07)	3.06* (1.50)	-1.15 (1.37)
Non-married	0.04 (1.25)	-0.79 (1.13)	0.44 (1.58)	-2.36 (1.45)
Non-Hispanic Black	-0.65 (0.55)	0.82 (0.49)	-0.68 (0.69)	0.83 (0.63)
Sales and administration	-3.27 (3.25)	-2.15 (2.94)	18.20*** (4.12)	-1.91 (3.78)
Observations	51	51	51	51
R^2	0.10	0.20	0.30	0.08

Source(s): IPUMS-CPS 1980–2019.

Note(s): The dependent variable of the regression is the Rotemberg weights for the top four industries and top four Census regions. The sample consists of US states for the 1980–2019 period. The results are weighted by the male population in each state.

† for $P < 0.10$, * for $P < 0.05$, ** for $P < 0.01$, *** for $P < 0.001$.

A.3 Results by CPS Month-in-sample

Because of the rotating panel design of the CPS, a worker can appear in this study's primary analysis sample multiple times in their second to fourth and sixth to eighth month-in-sample (MIS 2 through 4 and MIS 6 through 8). This feature raises concerns about estimation bias when unobserved individual characteristics are correlated with the outcome and independent variables. Conceptually, the problem can be addressed by restricting the sample to a single MIS group so that each worker is included in the sample no more than once during their CPS survey window. In practice, survey methodologists have found that survey responses to CPS labor force participation questions vary non-randomly across CPS MIS groups due to panel conditioning effects (Warren and Halpern-Manners, 2012). It is unclear whether using a single MIS introduces additional sources of bias.

As a robustness test to assess whether using a single MIS group as well as choice of the MIS affect the estimation results, I estimate the baseline model separately for each of the six MIS groups. Table A3 reports the results. The point estimates based on the sub-samples of MIS 2, 3, 6, and 7 are similar and qualitatively comparable to the baseline results in Table 6.1. By contrast, the point estimates are smaller based on the sub-samples of MIS 4 and 8. The lower marginal effects likely reflect the general lower exit rate reported by these two MIS groups.

The difference shown in MIS 4 and 8 samples is consistent with the observation in Warren and Halpern-Manners (2012) that CPS survey respondents are disinclined to repeat socially stigmatized responses in consecutive surveys, which results in lower unemployment estimates based on MIS 4 or 8 compared with other months in survey. Aside from the anomaly in results based on MIS 4 and 8, there is no strong evidence that using a pooled sample qualitatively affects the estimation results.

Table A3: Estimation Results by CPS Month in Survey (MIS)

	(1)	(2)	(3)	(4)	(5)	(6)
Expected earnings	-1.56*** (0.10)	-1.36*** (0.11)	-0.96*** (0.11)	-1.33*** (0.08)	-1.20*** (0.09)	-0.99*** (0.10)
Reference group earnings	1.64** (0.47)	1.39*** (0.35)	0.96* (0.46)	1.46*** (0.39)	1.19** (0.40)	0.42 (0.44)
MIS	2nd	3rd	4th	6th	7th	8th
Observations	1,296,172	1,328,726	1,331,211	1,302,716	1,321,230	1,323,641
R^2	0.05	0.05	0.04	0.05	0.04	0.04
$\hat{\gamma}$	1.05	1.03	1.00	1.10	0.99	0.43
Average exit rate (%)	1.88	1.65	1.51	1.76	1.54	1.44

Source(s): IPUMS-CPS 1980–2019.

Note(s): The dependent variable of the regression is the probability of labor force exit measured in % pts. The sample consists of US prime-age men for the 1980–2019 period. All models include state- and individual-level controls as well as state and year fixed effects. The results are weighted by the CPS longitudinal weight for linking individuals between adjacent months. 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.4 Pathways to Labor Force Exit

This section describes the employment and income pathways through which prime-age men transition from active labor force participation to both long-term and short-term withdrawal from the workforce. The analysis leverages longitudinal data from the Survey of Income and Program Participation (SIPP) 2008 panel to gain insights into this dynamic process. The dataset comprises US prime-age men who were in the labor force during one or more months of the SIPP 2008 sample period.

Conditional on being in the labor force in month t , the observations are grouped based on the individuals' labor force exit decision in the subsequent month $t + 1$ and their labor force participation status in months $t + 2$ through $t + 13$. This categorization results in three mutually exclusive groups: long-term dropouts, short-term dropouts, and stayers. The long-term dropouts are workers who left the labor force in $t + 1$ and remained outside the workforce for more than 12 months. The short-term dropouts are workers who exited the labor force in $t + 1$ but returned within 12 months. Finally, the stayers are individuals who remained in the labor force in $t + 1$.

For each of the three groups, I summarize their employment patterns, personal income, transfer benefits, and health status in the 12 months preceding the labor force exit event ($t - 11$ through t), as well as the 12 months following the exit event ($t + 2$ through $t + 13$). This provides an illustration of the trajectories leading up to and resulting from a prime-age man's decision to withdraw from the labor force.

Table A4 reports the results. Columns 1 and 2 in Panel A summarize the employment activities of long-term and short-term dropouts, respectively. On average, long-term dropouts spent 27.5 percent of their time employed full-time in the year preceding the exit event, while short-term dropouts spent 38.8 percent of their time to full-time work during this period. The remainder of their time was divided into part-time employment (19.5 percent and 19.6 percent), job search (31.2 percent and 30.3 percent), and short-term exit from the labor force (24.6 percent and 13.0 percent). By contrast, workers who stayed in the labor force dedicated approximately 80 percent of their time to full-time employment, with significantly less time spent job searching or temporarily non-

participating. These differences highlight that labor force exit, particularly long-term exit, often follows an extended period of weak attachment to the workforce.

Reflecting their weak attachment to the workforce, both long-term and short-term dropouts had low personal incomes in the 12 months preceding the exit event (Panel B). The opportunity cost of non-participation therefore tends to be smaller for the dropouts compared to the stayers. For some long-term dropouts, their lost earnings were partially replaced by Social Security benefits following the exit event, as indicated by an 8.3 percent increase in the share receiving such benefits after t . There is no strong evidence that the short-term dropouts' decision to exit the labor force was correlated with changes in the availability of transfer benefits. Additionally, the long-term dropouts reported a higher rate of work-limiting disability compared to the other two groups prior to the exit event, and this gap significantly widened after they left the labor force. This increase suggests that the decision to sever from the labor force nearly permanently often coincides with self-reported disability for many workers.

These descriptive patterns suggest that workers tend to leave the labor force when the expected returns from participation are small. Lengthy unemployment, low income, and poor physical condition are strong predictors of both a worker's decision to exit the labor force and the duration of their non-participation spell.

Table A4: Prime-Age Men’s Employment and Income Pathways to Labor Force Exit

	Long-term dropouts	Short-term dropouts	Stayers
	(1)	(2)	(3)
A. Employment			
Percent of time full-time employed			
Months 1-12 before event	27.5	38.8	80.0
Months 1-12 after event	0.0	30.4	80.3
Percent of time part-time employed			
Months 1-12 before event	19.5	19.6	11.6
Months 1-12 after event	0.0	13.8	11.4
Percent of time searching for work			
Months 1-12 before event	31.2	30.3	6.0
Months 1-12 after event	0.0	24.0	5.9
Percent of time out of labor force			
Months 2-12 before event*	24.6	13.0	2.6
Months 1-12 after event	100.0	31.8	2.5
B. Income and Benefits			
Median monthly personal income			
Months 1-12 before event	\$980	\$1,404	\$3,533
Months 1-12 after event	\$109	\$987	\$3,516
Percent receiving Social Security benefits			
Months 1-12 before event	8.7	1.5	0.7
Months 1-12 after event	17.0	2.1	0.8
Percent receiving unemployment benefits			
Months 1-12 before event	6.6	10.5	3.7
Months 1-12 after event	2.9	10.3	3.8
C. Health and Education			
Percent with work-limiting disability			
Months 1-12 before event	38.9	11.6	4.6
Months 1-12 after event	57.8	14.3	5.1

* For “Share of time out of labor force”, the summary statistics reflect labor market activities from month $t - 11$ through $t - 1$ because, by sample definition, a worker has to be in the labor force at t .

Source(s): SIPP 2008 panel.

Note(s): The sample consists of prime-age men who reported participating in the labor force in one or more months from September 2008 through August 2012. The results are weighted by the SIPP sample weight.