Minimum Wages, Retirement Timing, and Labor Supply*†

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Abstract

We use linked survey-administrative data to study the impact of minimum wage increases on employment, permanent labor force exit, and retirement benefit claiming for low-wage, older workers. The share of the labor force working for a rate of pay near the minimum wage increases for older ages near retirement, yet this population is typically ignored in the minimum wage literature. We find that minimum wage increases lead to increased employment, delayed permanent labor force exit, and delayed retirement benefit claiming. Increased employment and delayed permanent labor force exit are consistent with labor supply substitution effects, while the delayed benefit claiming appears to be driven by an interaction between minimum wages and the Social Security earnings test. Own-wage elasticity estimates of the employment effect are larger than those found in the existing minimum wage literature but are consistent with a growing literature on labor supply effects of minimum wages and evidence from life-cycle models of labor supply. The delay in permanent labor force exit and benefit claiming are similar to estimates from prior work on business cycle effects. These results suggest that the labor supply and benefit claiming of retirement-age individuals are very responsive to changes in labor market opportunities.

JEL codes: H55, J22, J26, J38
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1 Introduction

We test whether minimum wage policy influences employment, permanent labor force exit, and retirement benefit claiming for low-wage, retirement-age workers. Most minimum wage research focuses on teenagers and restaurant workers. These two groups are disproportionately made up of minimum wage workers, but the proportion of individuals who work for a rate of pay near the minimum wage also increases for ages near retirement (see Figure 1). Theoretical predictions for the effect of minimum wages on employment, permanent exit, and benefit claiming are ambiguous, as both labor demand and labor supply play a role and can have opposite effects. Work incentives associated with the Social Security earnings test may also be a factor, as working full-time for the minimum wage corresponds to earnings near the earnings test threshold. In addition to analyzing employment, permanent exit, and benefit claiming outcomes, we also attempt to differentiate between mechanisms related to labor demand, labor supply, and the Social Security earnings test.

Existing evidence related to minimum wages and older workers is limited and somewhat mixed in its conclusions. Fang and Gunderson (2009) find large increases in employment for older workers after minimum wage increases in Canada.[1] Borgschulte and Cho (2020) find mixed evidence for either no employment effect or small positive effects.[2] Lordan and Neumark (2018) find that minimum wage increases reduce employment in automatable jobs and that older workers in industries such as manufacturing may be particularly vulnerable to automation. Borgschulte and Cho (2020) also analyze Social Security retirement benefits and find evidence of reduced Social Security county-year retirement beneficiary and payout totals. However, the authors include a robustness check in the online appendix based on Current Population Survey data and find mixed evidence: they find no relationship between minimum wages and Social Security retirement beneficiary or payout totals, but there is a reduction in other retirement

[1] The marginal effect in Table 3 and mean employment rate for the treatment group in Table 2 imply a minimum wage-employment elasticity of 0.18.

[2] Their preferred minimum wage-employment elasticity estimates range from 0.059-0.126 depending on the sample (see columns (2) and (5) in Table 3). The estimates are marginally significant in some specifications and not significant in others. Interactive fixed effects estimates in the online appendix show no evidence of increased employment, but do show evidence of increased full-time work.

income. In summary, the existing research on employment effects includes evidence for negative effects, little or no effect, and large positive effects, while the only study on retirement benefit claiming finds results that are not entirely robust to the data source.

The contribution of our study stems from the use of linked survey-administrative data. We use data on individuals in the Survey of Income and Program Participation (SIPP) linked to individual data from the Internal Revenue Service (IRS) and Social Security Administration (SSA). Whereas prior work has relied on self-reported survey data or aggregate county-year data, the linked data provides several decades of objective measures of individual-level labor supply and benefit receipt. These data allow us to better account for individual-level confounders that may plague other studies using aggregate data, study some new outcomes such as permanent exit from employment, and study dynamics such as how long individuals delay permanent exit from employment or claiming of retirement benefits.

We use earnings information from the IRS to measure employment. The long history of IRS earnings data allows us to study the transitions of individuals into and out of employment and include individual fixed effects in doing so. We also use changes in annual earnings from an individual's lifetime maximum amount to define full-time vs part-time employment and full vs partial permanent exit from employment, thus providing the first evidence on minimum wages and permanent exit from employment. Finally, the benefit receipt data from the SSA allows us to identify the timing of when individuals claim retirement benefits down to the exact month, thus allowing for precision in accounting for the effect of minimum wages on claiming while adjusting for things such as the month when individuals reach their full retirement age (FRA).

We find that a 10% increase in the minimum wage is associated with a 1.51 percentage point increase in the probability of employment in a given year after the minimum wage change (2% increase from the sample mean), a 0.512 percentage point decrease in the probability of full permanent employment exit in a given year after the minimum wage change (6.4% reduction from the sample mean), and a 0.380 percentage point decline in the probability of claiming retirement benefits in a given month after the minimum wage change (9% reduction from the sample mean). Distributed lag models show that the employment effect persists for multiple years after the minimum wage increase, whereas the delay in full permanent exit and benefit claiming last for one year and six months, respectively.

Increased employment and delayed permanent employment exit are consistent with labor supply substitution effects. This could also explain the reduction in retirement benefit claiming if individuals tend to tie the timing of when they claim retirement benefits to when they exit the labor force. However, the Social Security earnings test may also play a role as working full-time at the minimum wage results in annual earnings near the earnings test threshold. We test this by exploiting the Senior Citizens Freedom to Work Act of 2000, which removed the Social Security earnings test for individuals past their FRA, and find that individuals not subject to an earnings test are still more likely to be employed after a minimum wage increase, but no longer show delayed claiming. Instead, minimum wage increases are associated with increased claiming. This suggests that the delay in claiming associated with minimum wage increases is due to a binding earnings test threshold, which is consistent with prior work showing that individuals respond to the earnings test threshold (Gelber et al., 2020b,a) and showing that the removal of the earnings test for post-FRA individuals caused increases in claiming (Gruber and Orszag, 2003; Loughran and Haider, 2005; Song and Manchester, 2007).

Our results described so far are based on individuals with survey-reported wages near the minimum wage. We also construct an alternative sample based on individuals with somewhat higher wages. This sample could be seen as a falsification sample in the sense that they should not be directly impacted by minimum wage increases, but factors such as general equilibrium effects, fairness related to relative pay structures, and firm adjustments could cause ripple effects up the wage distribution (Clemens, 2021). The sample can therefore also be used to check for these ripple effects and firm adjustments. We find no relationship between minimum wage increases and wages in the SIPP or employment and permanent exit in the IRS data for this higher wage sample, but we do find some evidence of small reductions in part-time employment and small increases in retirement benefit claiming. These results are consistent with a substitution by firms away from part-time, higher-wage individuals in response to higher minimum wages and increased labor supply of minimum wage workers and a subsequent increase in benefit claiming by affected workers to offset lost earnings income.

A key takeaway from our results is that low-wage individuals near retirement are very responsive to changes in labor market opportunities. In order to compare our results to other studies, we translate our preferred minimum wage-employment elasticity into an own-wage elasticity that is equal to 1.042. A recent survey of the literature found that the median own-wage elasticity of employment due to minimum wages is -0.17 (Dube, 2019). We replicate our analysis on a younger sample and estimate an own-wage elasticity of -0.144. Thus, our own-wage employment elasticity for retirement-age individuals is large compared to the rest of the literature and this appears to be due to unique labor supply behavior for older workers. This is consistent with growing evidence that minimum wage increases can induce positive labor supply effects for certain populations (Agan and Makowsky, 2018; Godoy, Reich and Allegretto, 2019) and evidence from life-cycle models of labor supply which generate elasticity of labor supply with respect to wage estimates in the range of 1-2 around the time of retirement (French, 2005; French and Jones, 2011).

In summary, our own-wage employment elasticity estimate is among the largest we are aware of associated with minimum wages.[3] Our finding of delayed full permanent exit from employment provides new evidence on labor market effects of minimum wage increases and is similar to findings on permanent exit and the unemployment rate (Gorodnichenko, Song and Stolyarov, 2013). Our finding of delayed claiming is similar in direction and magnitude to other research on minimum wages (Borgschulte and Cho, 2020) and the unemployment rate (Haaga and Johnson, 2012). Finally, it is worth noting that Social Security rules and thresholds have been adjusted many times over the years to encourage more work and later retirement. Our study provides more evidence that minimum wage increases can help achieve the same outcomes.

The remainder of the paper is organized as follows: Section 2 provides background information on minimum wages and the labor market, the Social Security system, and potential

[3]Godoy, Reich and Allegretto (2019) estimate an own-wage elasticity of employment of 0.471 for single mothers and 0.346 for married fathers. The largest imputed own-wage elasticity for retirement-age individuals in Borgschulte and Cho (2020) is 0.811, while other imputed estimates from the paper are much smaller. See Section 6 for more details.

mechanisms. Section 3 describes the data and sample. Section 4 presents the econometric methods. Section 5 details the results. Section 6 discusses the relation between our results and similar studies. Section 7 concludes.

2 Background

2.1 Minimum wage variation and the labor market

Figure 2 shows the number of states with a minimum wage increase in each year from 1978-2014. There are four different periods of federal minimum wage increases during this time, each of which was phased in over multiple years (1978-1981, 1990-1991, 1996-1997, 2007-2009). Federal minimum wage increases affected all or most states, although they have affected fewer states in more recent years due to states increasingly having their own minimum wages that are much higher than the federal. There are also many instances of states raising their own minimum wage in years between federal minimum wage increases. These tend to be more common in years leading up to a federal minimum wage increase, but there is at least one minimum wage change in every year except 1982-1984. In 2014, 18 states increased their minimum wage and 24 had a minimum wage higher than the federal level.

A large literature has studied the effects of minimum wages on labor market outcomes. There is robust evidence that minimum wage increases raise average earnings for the lower part of the wage distribution, but there has been much debate in the literature about whether that comes at the expense of employment for some groups (Allegretto, Dube and Reich, 2011; Allegretto et al., 2017; Dube, Lester and Reich, 2010; Neumark, Salas and Wascher, 2014b,a). Panel regressions with two-way fixed effects usually show reductions in employment, with minimum wage-employment elasticities in the range of -0.1 to -0.3. Adding controls for regional heterogeneity in employment trends produces smaller elasticities not distinguishable from zero. Recent work using more flexible methods to account for unobserved heterogeneity or directly comparing the count of missing jobs below versus excess jobs above the minimum wage after an increase are consistent with small employment effects not distinguishable from zero (Cengiz et al., 2019; Totty, 2017).[4]

Minimum wage-employment elasticities have the limitation that the bite of the wage increase can vary across settings, making it difficult to compare elasticities across studies. Own-wage elasticity estimates address this limitation by comparing the employment change for a given group to the average wage change that occurred for that group from a minimum wage increase. An own-wage elasticity of -1, for example, indicates that the wage gains and employment losses fully cancel out, and the given group experiences no net increase in total earnings. Dube (2019) summarized 55 studies where an own-wage employment elasticity from the minimum wage could be calculated from the reported results. The medium own-wage elasticity was -0.17, indicating that employment losses are only about 1/6 as large the wage gains. As the evidence against large reductions in employment from minimum wage increases has grown, so has the fraction of researchers who conclude that employment effects of minimum wages are close to zero (Belman and Wolfson, 2019).

Much of this literature has focused on teenage employment. Figure 1 shows why: teenagers and individuals in their early twenties are very likely to work for a rate of pay near the minimum wage. However, Figure 1 also shows that individuals around retirement age are more likely to work for a rate of pay near the minimum wage than individuals of prime working age.[5] There are many factors that may contribute to this rise in the share of individuals working near the minimum wage at older ages. Declining productivity due to aging could lead to lower wages (Ben-Porath, 1967; Lazear, 1979) and changing work preferences as workers age may lead them to trade-off wages for other amenities such as schedule flexibility, non-wage compensation, and location preferences (Ameriks et al., 2020; Giandrea, Cahill and Quinn, 2009; Maestas et al., 2017; Ruhm, 1990; Rutledge and Wettstein, 2020). Alternatively, workers with higher rates of pay may retire earlier on average.

[4]There is also a large literature on the reasons why there me may little or no effect on employment. Possible explanations include offsets along other margins, such as a reduction in fringe benefits, the provision of health-care, and/or schedule flexibility; monopsony or imperfect competition; labor-labor substitution; labor-capital substitution; reduced turnover costs; increased productivity; cost pass-through in the form of higher prices; reduced profits; increased labor supply; and skill-upgrading. See Clemens (2021) and Manning (2021) for a summary of this research.

[5]This fact was first shown by Borgschulte and Cho (2020) using the Current Population Survey. We replicated this pattern using monthly wage data from the SIPP, which we describe in more detail in Section 3.

Figure 3 shows average annual wage growth by age. Wage growth is large for young workers who are just beginning their careers and transitioning out of part-time or seasonal employment while in school. Wage growth then steadily declines with age until approximately age 62, at which point average wage growth rises again. Rising average annual wage growth during retirement age in Figure 3 and rising rates of work near the minimum wage during retirement age in Figure 1 suggests that there are selection effects occurring with respect to retirement timing. It also suggests a possible mechanism through which minimum wages may influence retirement timing. Individuals who experience greater wage growth may be more likely to continue working past the age at which they qualify for retirement benefits, while individuals who experience lower (or negative) wage growth may choose retirement. French (2005) and French and Jones (2011) find evidence consistent with this behavior. In other words, the labor supply of individuals near retirement age may be particularly sensitive to wage changes, and thus potentially impacted by minimum wage increases.

2.2 The Social Security program

Social Security helps older Americans, workers who become disabled, and families in which a spouse or parent dies. As of June 2019, about 177 million people worked and paid Social Security taxes and about 64 million people received monthly Social Security benefits, with about 48 million of these being retirees. Social Security benefits are intended to supplement one's own retirement income from personal savings once an individual's labor supply decreases later in life.

The Social Security Administration uses tax dollars of workers to fund benefits received by retirees and other beneficiaries. Regular employees pay a Social Security tax of 6.2 percent of their earnings, which employers also have to match, while those that are self-employed pay a 12.4 percent rate. A majority of these tax dollars go into a trust fund that pays benefits to retirees. To become eligible for Social Security benefits, individuals work and earn Social Security "credits." A maximum of four credits can be earned in a given year, and most people need 40 credits (10 years of work) to qualify for benefits.

Social Security replaces a percentage of a person's pre-retirement income. The amount of average wages that Social Security replaces depends on two factors: 1) a person's earnings prior to retirement; and 2) the timing of when a person chooses to start benefits. The benefit amount is based on an individual's 35 highest earning years, but is adjusted depending on when the individual claims their benefits. If an individual starts their benefits at the FRA ($\geq$65 for those born before 1943, $\geq$66 for those born between 1943-1959, and 67 for those born after 1959), this percentage can range from 75 percent for low earners, to about 40 percent for medium earners and 27 percent for high earners.[6] If individuals elect to start benefits prior to reaching the FRA, then the percentages are lower.

While most individuals are eligible to receive retirement benefits at age 62, there are incentives put into place to encourage a delay in retirement until reaching the FRA. Workers that have reached age 62 but have not yet reached the FRA for Social Security are subject to a retirement earnings test. That is, workers earning more than a determined threshold will have a portion of their benefits deferred until reaching the FRA at a rate of one dollar per every two dollars of earnings that lie above the annual limit (in 2021, the limit is $18,960).[7] Once a person reaches their FRA, continuing to work will not reduce their benefits no matter their earnings. It is important to note that any benefits withheld due to the earnings test are not lost. Upon reaching the FRA, benefits are increased permanently in an actuarially fair way to account for the months in which benefits were withheld.

Traditionally, the earnings test applied to individuals above the FRA as well. These individuals were subject to an earnings test through age 71 from 1975-1982, then age 69 from 1983-1999. The post-FRA earnings test thresholds were higher than the pre-FRA thresholds beginning in 1978. In 2000, President Clinton signed a bill that eliminated the limits on what post-FRA individuals on Social Security could earn, known as the Senior Citizens Freedom to Work Act of 2000. This bill effectively eliminated the earnings test among those that had reached the FRA, allowing these individuals to continue working without being penalized by

[6] Traditionally, the FRA was 65, however over time it has increased to 66, and will gradually rise to 67 for those born 1960 and later.

[7] The threshold is higher for the year during which a person reaches their FRA and applies only to the months before such attainment. That amount was $50,520 in 2021.

decreased Social Security benefits.

Social Security retirement claiming rules have been adjusted many times over the years for the purpose of incentivizing more work later in life. Raising the FRA, increasing the earnings test threshold for individuals past the FRA at a faster rate than the standard wage index adjustment beginning in 1996, removing the earnings test for individuals past the FRA in 2000, and increasing the delayed retirement credits associated with delaying past the FRA were all done in order to incentivize more work and reduce strain on the Social Security system (Blundell, French and Tetlow, 2016).[8]

2.3 Minimum wages, employment, and retirement benefits: potential mechanisms

2.3.1 Employment mechanisms

We first discuss potential mechanisms for a relationship between minimum wages and employment before turning to benefit claiming. The relationship between minimum wages and employment for older workers could be impacted by both the demand for and supply of labor. Higher minimum wages may decrease the demand for low-wage labor, which in turn may lead to decreased employment for older workers. Alternatively, labor supply effects could lead to decreased or increased employment for older workers. The labor supply mechanism can be characterized as an income versus substitution effect. On the one hand, workers could respond to a wage increase through an income effect, in which case workers reduce their amount of work because they can achieve a target level of income faster. This could subsequently lead to earlier labor force exit and reduced employment. On the other hand, workers could respond through a substitution effect, in which case they increase their amount of work due to an increase in the cost of leisure. This could subsequently lead to delayed labor force exit and increased em-

[8]Prior work has found that retirement and labor supply decisions are influenced by a variety of factors including unemployment and the business cycle (Coile and Levine, 2010; Haaga and Johnson, 2012), increasing the FRA (Behaghel and Blau, 2012; Deshpande, Fadlon and Gray, 2020), changes in Medicare eligibility and rules (French and Jones, 2011; Wettstein, 2020), employer-sponsored health insurance (Nyce et al., 2013; Shoven and Slavov, 2014), the Social Security earnings test (Engelhardt and Kumar, 2014; Gelber et al., 2020b), and increasing delayed retirement credits (Duggan et al., 2021), to name a few.

ployment. Hence, theoretical predictions for the effect of minimum wages on employment are multi-faceted and ambiguous.

Empirically, prior work suggests labor supply substitution effects may be the most relevant for our study. As discussed in Section 2.1, most recent research and a growing fraction of economists conclude that minimum wages have little or no effect on employment, which runs counter to the labor demand channel. Similarity in employment trends of older workers across countries, despite differences in labor market institutions and norms, also suggests that labor demand may be less important in determining employment for older workers than labor supply (Blundell, French and Tetlow, 2016). The evidence in Section 2.1 and from French (2005) and French and Jones (2011) that individuals who experience larger wage growth are more likely to continue working past age 62 suggests that labor supply substitution effects may be greater than labor supply income effects. Collectively, the literature on minimum wage effects and labor supply around retirement jointly suggests that decreased employment around retirement age due to minimum wage increases is unlikely and that increased employment is realistic.

The model of labor supply and retirement developed in French (2005) and extended in French and Jones (2011) is useful for thinking about the relationship between minimum wages and labor supply in greater detail. French (2005) models labor supply, retirement, and savings and attempts to match life-cycle moments of labor force participation, hours worked, and assets. Individuals can choose consumption, labor force participation and work hours, and whether to apply for pension benefits. There is a fixed cost of work and individuals cannot borrow against future labor, pension, or Social Security income.

The fixed cost of work along with the inability to borrow against future Social Security or pension income, the tax structure of Social Security and pensions, and uncertainty related to health and wages results in many older individuals being near the participating margin, which helps match both the dramatic drop in labor force participation around retirement age and the clustering of hours worked at 0 and 2,000. The model also shows that groups closer to the participation margin have higher labor supply elasticities with respect to wage because of the fixed cost of work. For example, the labor supply elasticity from a permanent wage change rises from 0.17 at age 40 to 1.17 at age 60 in the year of the change; it rises from 0.26 at age 40 to 2.24 at age 60 in years after the change (French and Jones, 2011). These are elasticities of hours worked with respect to wage, but the large elasticities are related to employment changes as can be seen by the tendency of hours to cluster at 0 and 2,000. Thus, these large elasticities of labor supply with respect to wages suggest we might expect increased employment for retirement-age workers following minimum wage increases, and that the increase might be large relative to employment effects for younger workers.

2.3.2 Claiming mechanisms

The theoretical relationship between minimum wages and retirement benefit claiming is also multi-faceted. Minimum wages may indirectly impact claiming through the employment effect described above. If reduced labor demand leads to reduced employment, then workers may turn to retirement benefits to supplement their income. Similarly, the net labor supply effect could impact the timing of retirement benefit claiming if individuals tend to tie the timing of when they claim retirement benefits to the timing of when they exit the labor force. Both the labor demand and labor supply channels therefore imply that a reduction (increase) in employment could lead to an increase (reduction) in retirement benefit claiming. Building on the above discussion about employment mechanisms, we might expect a reduction in retirement benefit claiming after a minimum wage increase due to increased labor supply if individuals prefer to begin receiving retirement benefits when they exit the labor force.

That being said, incentives for claiming retirement benefits are not directly related to employment on the extensive margin: individuals can claim retirement benefits while continuing to work without penalty.[9] Incentives for claiming retirement benefits are, however, directly related to an individual's amount of earnings. Individuals who earn above the Social Security earnings test threshold while receiving retirement benefits have their benefits temporarily with-

[9]While studies have shown that benefit claiming and intensive margin labor supply in the form of earnings or hours responds to the Social Security earnings test, evidence on extensive margin labor supply adjustments is more mixed. Friedberg (2000), Gruber and Orszag (2003), and Song and Manchester (2007) find no effect of the Social Security earnings test on employment, but recent work by Gelber, Jones and Sacks (2020) and Gelber et al. (2020a) does find evidence of an employment effect.

held. If minimum wages increases cause an individual's earnings to surpass or go further beyond the earnings test threshold then their benefits will be reduced, which reduces the immediate financial return to claiming

Figure 4 shows the level of the Social Security earnings test threshold over time, as well as the amount of earnings an individual would make if they worked full-time (40 hours per week, 50 weeks per year) at the federal minimum wage. This figure illustrates the relevance of the earnings test for near-minimum-wage workers. The post-FRA earnings threshold tracked the full-time federal minimum wage earnings level closely through the mid-1980s, with the pre-FRA threshold a little below. From the mid-1980s onward, the pre-FRA threshold tracks the full-time federal minimum wage level closely, while the post-FRA threshold diverges somewhat in the late 1990s before its removal in 2000. Thus, regardless of the effect of minimum wages on employment, the interaction of minimum wages with the Social Security earnings test could lead to reductions in retirement benefit claiming if minimum wages push individuals across or further beyond the earnings threshold, thus reducing their benefit payments.

In the analysis of retirement benefit claiming to follow, we attempt to distinguish between the employment effect mechanism and the interaction between minimum wage increases with the Social Security earnings test. We do this by exploiting the removal of the Social Security earnings test for individuals past their FRA in 2000, which affected the earnings test mechanism but not the employment effect mechanism.

3 Data and Sample Selection

3.1 Linked survey-administrative data

Our data source is the SIPP Gold Standard File (GSF), a Census Bureau product that integrates person-level microdata from the SIPP with person-level administrative tax and benefit data from the IRS and SSA.[10] The GSF includes all individuals and a subset of longitudinally

[10] We use version 7.0 of the GSF. Outside researchers can access a synthetic version of the GSF, known as SIPP Synthetic Beta (SSB). Researchers can then have their results validated on non-synthetic data. More information is available in Benedetto, Stanley and Totty (2018).

harmonized variables from nine different SIPP panels (1984, 1990, 1991, 1992, 1993, 1996, 2001, 2004, and 2008). The SIPP includes monthly survey response information for a short period of time, usually a span of two-to-four years.

Individuals from the SIPP are linked to administrative records from the IRS and SSA in order to merge tax and benefit information. Linkage is based on a protected identification key (PIK), which is a confidentiality-protected version of a social security number. The various SIPP panels all have a 70-90% successful link to the administrative records data, except for the 2001 panel, which has a 47% successful link rate.

For individuals who were successfully linked, the GSF includes tax information from W-2 and Schedule C records along with Social Security benefit application date, receipt date, and amount information. Unlike the SIPP information, the linked administrative data cover a long time frame. The W-2 and Schedule C earnings information includes annual earnings reported to the IRS from 1951-2014.[11] Similarly, the benefit application, receipt, and amount information covers 1951-2014. Thus, while we only know information collected in the SIPP while the person was in the survey, the tax and benefit information cover a much longer time frame. A visual representation of the GSF including its data sources and variable availability over time is shown in Figure 5.

The GSF is well-suited to study the impact of minimum wages on employment and benefit claiming for retirement-age individuals. The SIPP includes information that is useful for identifying potential minimum wage workers, including monthly earnings, monthly hours worked, birth date, and state of residence. The earnings information from the IRS and SSA provide annual employment information which allows us to study labor supply responses. The benefit information from the SSA includes an indicator for receipt of OASDI retirement benefits, the date when receipt of the benefits began, and the amount received in the first month of payment.

[11]There are two sets of annual earnings records in the GSF. The Summary Earnings Record (SER) is available from 1951-2014 and is capped at the taxable maximum each year. The Detailed Earnings Record (DER) is not capped but is only available from 1978-2014. Both the SER and DER are provided by the SSA. The SSA receives the tax information in the SER and DER form the IRS. The SSA performs some cleaning of the data before the Census Bureau receives it. We use the DER because the amounts are not capped.

3.2 Sample construction

Individuals cannot receive Social Security benefits before reaching age 62 and can increase their monthly amount received by delaying only until age 70. Our main analysis is therefore restricted to ages 62-70. Because the SIPP is our source of state of residence as well as the earnings and hours information that we use to identify minimum wage changes and low-wage workers, we limit the sample to individuals who are observed in the SIPP at some point during ages 62-69. We then link monthly and annual historical minimum wage information from Vaghul and Zipperer (2016) and the website of David Neumark.[12] We also link covariate information on state-month and state-year unemployment rates from the Bureau of Labor Statistics and state-year population from the Bureau of Economic Analysis. We adjust all wage, earnings, and minimum wage variables for inflation to 1999 dollars using the R-CPI-U-RS.[13]

In order to identify low-wage workers, we construct an hourly wage variable equal to monthly earnings divided by monthly hours for each month an individual is in the SIPP. We then take an average of these monthly wage observations in order to identify individuals who consistently work for a rate of pay near the minimum wage, similar to the approach in Clemens and Wither (2019). For individuals who experience a change in their state minimum wage while in the SIPP, we take the average of all monthly wage observations prior to the first minimum wage change.[14] We only use months prior to the first minimum wage change in order to avoid conditioning the sample on a variable that is directly affected by the minimum wage change. For individuals who do not experience a minimum wage change while in the SIPP, we take the average of all monthly wage observations while in the SIPP. We then select individuals whose average monthly wage is less than or equal to the prevailing minimum wage before any change plus two dollars.[15] We

[12]http://www.economics.uci.edu/ dneumark/datasets.html.

[13]https://www.bls.gov/cpi/research-series/r-cpi-u-rs-home.htm.

[14]A limitation of the GSF data is that we only observe state of residence at the beginning of the SIPP panel; we do not have state of residence from the annual IRS/SSA data. We therefore have to hold state of residence constant based on what was first reported in the SIPP. If an individual moves to a different state before or after their SIPP interview but during the time frame of our analysis, then state will be measured with error.

[15]Clemens and Wither (2019) study the effect of the series of federal minimum wage increases that occurred during the Great Recession. To create their low-wage sample, they take the average of SIPP monthly wage observations prior to the first federal minimum wage change. They define the "target group" as individuals whose average wage was less than or equal to $7.50, which is equivalent to the federal minimum wage before the changes plus $2.35.

extend the cutoff above the minimum wage because of measurement error in wages and because it is well-documented that minimum wage effects can spill over to individuals who earn a little above the minimum wage (e.g., Gopalan et al, 2021; Lopresti and Mumford, 2016). We chose the cutoff of two dollars because several studies suggest that these spillovers exist up to about two dollars above the minimum wage value prior to the minimum wage change (Brochu et al., 2018; Cengiz et al., 2019; Fang and Gunderson, 2009; Gopalan et al., 2021).

After identifying low-wage, retirement-age workers, we construct our sample for the employment analysis based the IRS data. We use the annual W-2 and Schedule C earnings from administrative tax records to measure employment. We start with the retirement-age, near-minimum-wage sample described above and make two additional sample restrictions. First, the individual must be successfully linked to the administrative records. Second, the individual must have no record of dying prior to the year during which they turned 71. We then use the linked annual minimum wage information and birth date to build a balanced person-year panel that spans from the year during which the individual turned 62 through the year during which the individual turned 70.[16]

We then turn to the OASDI retirement benefit claiming analysis. We make three additional sample restrictions to the retirement-age, near-minimum-wage sample. First, the individual from the SIPP must be successfully linked to the administrative records data as described above. Second, the individual must have no record of dying prior to the month they turned 71 without having already claimed retirement benefits. Third, the individual must have at least 40 quarters of covered work prior to reaching age 62 so that they qualify for retirement benefits.[17] We identify retirement benefit claiming in the GSF by using the variable that indicates when OASDI retirement benefit payments began.[18] We use this variable along with

[16]Information on date of death is obtained using a hierarchy of administrative sources: (i) SSA's Master Benefits Record file, (ii) SSA's Supplemental Security Record file, and (iii) the Census Person Characteristics File with death information coming from the SSA Numident and Master Death Files.

[17]Our data from the SSA reports the total quarters of qualified work per year beginning in 1951. The minimum amount of earnings required in order to qualify for a quarter of covered work changes over time based on average wages. The amount in 2021 is $1,470.

[18]We refer to our outcome as the *claiming* of retirement benefits throughout the paper, although our outcome is actually the date of first *receipt*. This is common in the literature (e.g., Behaghel and Blau, 2012; Deshpande et al., 2020). The date of first receipt, not application, is what determines payout amounts and thus is the variable of interest for financial well-being. Changes in claiming behavior are therefore inferred from changes

the linked monthly minimum wage and birth date information to build a person-month panel that spans from the month an individual turns 62 through the month they claim retirement benefits or the month prior to turning 70.

Finally, we construct an alternative version of the two samples described above based on individuals with higher wages. The higher-wage sample construction is based on the same sample selection criteria described above, except that we select individuals whose average wage is \$5-\$10 above the minimum wage instead of less than or equal to the minimum wage plus \$2. These individuals have low enough wages that they still belong to the lower part of the wage distribution, but high enough wages that their outcomes should not be directly effected by changes in the minimum wage. However, other research has shown that minimum wage increases can have ripple effects up the wage distribution and that firms may adjust to higher minimum wages along several different margins including labor-labor substitution (e.g., Clemens (2021) and Gopalan et al. (2021)). We use this sample to test for such adjustments. Prior work suggests that ripple effects up the wage distribution in terms of workers who see wage increases only exist up to about \$2 above the minimum wage, so we do not expect to see wage effects for this higher-wage sample. But we could see effects on employment or claiming if there are labor demand effects in the form of labor-labor substitution. For example, firms could adjust to higher minimum wages by employing fewer minimum wage workers and replacing them with workers higher up the wage distribution. Or, they could employ the same or a larger amount of minimum wage workers and offset costs by employing fewer higher-wage workers.

Appendix A shows a validation or first-stage test of our sample design in which we use only the monthly SIPP data to estimate short-run effects of minimum wage increases on wages, employment, hours, and earnings. For our main low-wage sample, we find increased wages and earnings with no effect on employment or hours. For the higher-wage sample, we find no effect on wages, earnings, employment, or hours. These results validate that our main sample is in fact impacted by minimum wage increases in months immediately following the increase, while our higher-wage sample is not directly impacted, at least in terms of these outcomes.

in receipt timing. The SSA recommends that individuals apply for retirement benefits four months before they would like to begin receiving them.

3.3 Summary statistics

Table 1 shows summary statistics for the employment and OASDI retirement benefit claiming samples in columns (1)-(2) and (3)-(4), respectively. The outcomes of interest for the employment analysis are three different annual employment indicators. One is an indicator for any employment, based on the presence of any amount of earnings in the IRS data. The others are indicators for full-time and part-time employment, based on the amount of earnings reported in the IRS data. We follow an approach similar to Gorodnichenko, Song and Stolyarov (2013) in how we define full-time and part-time work using earnings reported to the IRS. Full-time work is defined as an individual earning at least 50% of their lifetime highest annual earning amount observed in the data, in inflation-adjusted dollars. Part-time work is defined as earning less than 50% of their highest observed earnings year but still at least $5,000 in inflation-adjusted dollars, which equates to working approximately 20 hours per week at the minimum wage for a full year or working 40 hours per week at the minimum wage for six months.

We also use the employment sample to study the timing of permanent labor force exit. Here we use the exact approach in Gorodnichenko, Song and Stolyarov (2013) to define what they refer to as partial retirement and full retirement. If an individual's annual earnings permanently fall to less than 50% of their lifetime inflation-adjusted maximum but still at least $5,000 in inflation-adjusted dollars then they classify the individual as having partially retired. If an individual's annual earnings permanently fall to less than 50% of their lifetime inflation-adjusted maximum and less than $5,000 in inflation-adjusted dollars then they classify the individual as having fully retired.[19] Because the definition of "retired" can vary across studies, we avoid this term and instead use a more explicit phrase: partial and full "employment exit." Summary statistics for these two variables are shown under the employment sample.

The outcome variable for the retirement claiming sample is an indicator variable for whether

[19] We compute an individual's lifetime highest inflation-adjusted annual earnings amount using the SER amounts from 1951-1977 and the DER amounts from 1978-2014. As described in Section 3.1, the SER is capped at the taxable maximum amount each year. This causes measurement error in our estimate of an individual's lifetime highest earning amount if their highest amount occurred during 1951-1977 and was above the taxable maximum. This is less of a concern for our sample because we are analyzing low-wage individuals in the SIPP who are therefore less likely to have high earnings. Less than 6% of our sample had their highest annual earnings occur before 1978 and equal to the taxable maximum.

the individual began receiving retirement benefits in the current month. In addition to showing the mean and standard deviation for this variable in the *Time-Varying Variables* section of the table, the *Time-Invariant Variables* section also shows the share of the sample that receives retirement benefits by age 70 and the average age at which individuals first receive benefits.

The covariates used in the employment and claiming analyses include indicators for sex, race, Hispanic status, highest education level, marital status, birth year, health insurance, health insurance via employer, defined-benefit pension, and defined-contribution pension.[20] The continuous covariates include total net worth, total non-housing wealth, state-year unemployment rates, and total annual earnings reported to the IRS.

4 Econometric Methods

4.1 Employment

The employment analysis is based on ordinary least squares (OLS) regressions where the dependent variable is an indicator variable for employment (either any employment, full-time employment, or part-time employment depending on the analysis) in a given year. Individuals can move into and out of employment over time. Because non-employment is not an absorbing state, we can include individual fixed effects. Our analysis is therefore based on a balanced panel of observations from the year during which an individuals turned 62 through the year during which an individual turned 70.

The independent variable of interest is the log of the minimum wage. An alternative approach could be to include a binary indicator variable for minimum wage change events that permanently changes from 0 to 1 after a minimum wage change. The event-based approach is challenging in this setting because minimum wage changes are recurring events and often recur in windows as short as 6 or 12 months. The event-based approach is therefore not easily applicable to minimum wage settings; at least not to the analysis of outcomes over the time span of

[20]Health insurance coverage and health insurance via employer are reported longitudinally in the SIPP. Our covariate is equal to one if the respondent ever indicated coverage and zero otherwise. We also explored using covariates for always having coverage or the share of the individual's responses over time that indicate coverage. The results were very similar.

more than 1-2 years. The employment and benefit claiming analyses, on the other hand, cover a time span of up to nine years and most individuals will be exposed to multiple minimum wage changes over that time span. The log minimum wage approach is common in the literature (see, e.g., the discussion in Cengiz et al. (2019) and Neumark (2018)) and allows us to not only analyze a longer time span, but also account for the size of changes in the minimum wage.

The primary regression model is:

$$Employed_{iast} = \beta log(MW_{st}) + X_{ist}\psi + \tau_a + \alpha_i + \delta_t + u_{iast}. \tag{1}$$

$log(MW_{st})$ is the log of the minimum wage in state s at time t. X_{ist} is a vector of covariates for macroeconomic conditions and individual characteristics.[21] τ_a, α_i and δ_t are age (in years), individual, and time period (in years) fixed effects, respectively. We also report robustness checks that add state-specific linear time trends, age-by-time fixed effects, and Census region-by-time fixed effects.

Age fixed effects adjust for differences in employment rates across ages and other unobserved age effects. Year fixed effects adjust for unobserved common time effects that influence employment. Finally, individual fixed effects adjust for time-invariant individual characteristics that influence employment. The ability to include individual fixed effects is especially valuable as it allows us to compare individual outcomes before and after exposure to minimum wage changes. Under the conditional parallel trends assumption that minimum wage increases are uncorrelated with unobserved individual or state-year characteristics that influence employment and are not already accounted for in the model, β represents the effect of a given percent change in the minimum wage on the probability of a given type of employment in each time period after a minimum wage increase.

We also estimate a distributed-lag model in order to observe the evolution in employment around the time of a minimum wage change. This model shows how long differences in employ-

[21]These covariates include continuous variables for the log of the state-year unemployment rate, total net worth, and non-housing wealth; and indicator variables for the year during which an individual reaches FRA, sex, race, Hispanic status, highest education level (less than high school, high school, some college, bachelor's, graduate), marital status (never married, married, divorced or widowed), has a defined-benefit pension, has a defined-contribution pension, removal of the social security earnings test after the FRA in 2000, and birth year.

ment (if any) last after the minimum wage change, in addition to providing a visual check of the conditional parallel trends assumption. An event-study model would be a natural approach for this type of question, but is complicated by the continuous and recurring nature of minimum wage change events over a nine year time span as discussed above. We therefore adopt the approach from Dube, Lester and Reich (2010), who applied a distributed lag model to regressions of employment and earnings on the log of the minimum wage. We include leads up to 3 years prior to the minimum wage increase and lags up to 3 years after the minimum wage increase in one-year intervals. Including the leads/lags as one-year differences except for the last lag produces coefficients that represent the cumulative response of employment to minimum wage changes over time. The model is:

$$Employed_{iast} = \sum_{j=-3}^{2} \beta_{-j} \Delta log(MW_{s,t-j}) + \beta_{-3} log(MW_{s,t-3}) + X_{ist}\psi + \tau_a + \theta_s + \delta_t + u_{iast}, \quad (2)$$

where Δ represents a one-year difference operator. We expand the balanced panel from ages 62-70 to ages 59-73 in order to include three years of leading and lagged values of the log minimum wage for the full sample.

4.2 Permanent exit from employment

For the analysis of partial and full permanent employment exit we estimate hazard models of exit timing using OLS. The dependent variable is equal to zero in years prior to the year of permanent exit and one in the year of permanent exit. We include age fixed effects and drop all observations after the year of exit, thus effectively converting the specification to a hazard model. If an individual does not permanently exit employment by the year they turned 70 then their time-series stops at that year.

The primary regression model is:

$$Exit_{iast} = \beta log(MW_{st}) + X_{ist}\psi + \tau_a + \theta_s + \delta_t + u_{iast}. \quad (3)$$

$Exit_{iast}$ is either full exit or partial exit depending on the analysis. Because observations are dropped after individuals permanently exit, exit is an absorbing state, which means that we cannot include individual fixed effects like we did for the employment analysis. We replace the individual fixed effects with state fixed effects, θ_s.[22] Under the conditional parallel trends assumption that minimum wage increases are uncorrelated with unobserved individual or state-year characteristics that influence employment exit and are not already accounted for in the model, β represents the effect of a given percent change in the minimum wage on the probability of permanent employment exit in each time period after a minimum wage increase.

We also estimate a distributed-lag hazard model in order to observe the evolution in the probability of employment exit around the time of a minimum wage change. This model shows how long differences in exit rates (if any) last after the minimum wage change, in addition to providing a visual check of the conditional parallel trends assumption. Following the approach from Dube, Lester and Reich (2010), we include leads up to 3 years prior to the minimum wage increase and lags up to 3 years after the minimum wage increase in one-year intervals. Including the leads/lags as one-year differences except for the last lag produces coefficients that represent the cumulative response of exit behavior to minimum wage changes over time. The model is:

$$Exit_{iast} = \sum_{j=-3}^{2} \beta_{-j} \Delta log(MW_{s,t-j}) + \beta_{-3} log(MW_{s,t-3}) + X_{ist}\psi + \tau_a + \theta_s + \delta_t + u_{iast}, \quad (4)$$

where Δ represents a one-year difference operator. We expand the balanced panel from ages 62-70 to ages 59-73 in order to include three years of leading and lagged values of the log minimum wage for the full sample.

4.3 Hazard models of OASDI retirement benefit claiming

Our retirement benefit claiming analysis is also based on estimating hazard models using OLS. The dependent variable is equal to zero in months prior to receiving retirement benefits and one

[22]State fixed effects were excluded from equation (1) because state of residence is only measured once in the SIPP GSF and thus drops out with individual fixed effects.

in the month of first receipt, following recent papers on retirement claiming such as Behaghel and Blau (2012) and Wettstein (2020). We include age (in months) fixed effects and drop all observations after the month of retirement, thus effectively converting the specification to a hazard model.[23] If an individual does not receive retirement benefits by the month before turning 70 then their time-series stops at that age.

The primary regression model is:

$$Receipt_{iast} = \beta log(MW_{st}) + X_{ist}\psi + \tau_a + \theta_s + \delta_t + u_{iast}. \tag{5}$$

$log(MW_{st})$ is the log of the minimum wage in state s at time t. X_{ist} is a vector of covariates for macroeconomic and individual characteristics.[24] τ_a, θ_s, and δ_t are age (in months), state, and time period (in months) fixed effects, respectively.[25] Under the conditional parallel trends assumption that minimum wage increases are uncorrelated with unobserved individual or state-month characteristics that influence retirement benefit claiming and are not already accounted for in the model, β represents the effect of a given percent change in the minimum wage on the probability of claiming retirement benefits in each time period after a minimum wage change.

We also estimate a distributed-lag hazard model in order to observe the evolution in the probability of claiming retirement benefits around the time of a minimum wage change. If an individual chooses not to claim retirement benefits due to a change in the minimum wage, then this is a temporary delay in claiming and we would like to be able to estimate the length of such a delay. Following the approach from Dube, Lester and Reich (2010), we include leads up to 24 months prior to the minimum wage increase and lags up to 24 months after the minimum wage increase in six-month intervals. Including the leads/lags as six-month differences except for the

[23]Appendix B shows robustness checks where we estimate the hazard models using probit, logit, and contemporaneous log-log estimation techniques instead of OLS. We prefer the OLS method because it is more common in the literature on retirement benefit claiming (e.g., Behaghel and Blau (2012) and Wettstein (2020))

[24]These covariates include continuous variables for the log of the state-year unemployment rate, total net worth, non-housing wealth, total earnings reported to the IRS; and indicator variables for the month in which an individual reaches FRA, sex, race, Hispanic status, highest education level (less than high school, high school, some college, bachelor's, graduate), marital status (never married, married, divorced or widowed), has a defined-benefit pension, has a defined-contribution pension, removal of the social security earnings test after the FRA in 2000, and birth year.

[25]Individual fixed effects from equation (1) are excluded here because OASDI receipt is an absorbing state.

last lag produces coefficients that represent the cumulative response of retirement claiming to minimum wage changes over time. The model is:

$$Receipt_{iast} = \sum_{j=-4}^{3} \beta_{-6j} \Delta_6 log(MW_{s,t-6j}) + \beta_{-24} log(MW_{s,t-24}) + X_{ist}\psi + \tau_a + \theta_s + \delta_t + u_{iast}, \quad (6)$$

where Δ_6 represents a six-month difference operator. We expand the balanced panel from ages 62-69 to ages 59-72 in order to include two years of leading and lagged values of the log minimum wage (in six-month differences) for the full sample.

5 Results

5.1 Employment

Results for the effect of minimum wage increases on the employment of low-wage, retirement-aged individuals are shown in Table 2. Panel A shows results for the main sample, whereas Panel B shows results for the higher-wage sample. The results in Panel A show a positive and statistically significant effect of minimum wage increases on employment. The coefficient for the log minimum wage variable ranges from 0.214 when only state, year, and age fixed effects are included in the specification to 0.151 when covariates and individual fixed effects are added. Our preferred model is the one with full covariates and individual fixed effects in column (3). The coefficient of 0.151 represents a 0.0151, or 1.5 percentage point, increase in the probability of employment (i.e., having positive IRS earnings) in a given year after a 10% increase in the minimum wage, statistically significant at the 1% level. This corresponds to a 2% increase from the sample mean probability of employment in a given year. The higher-wage sample in Panel B shows no relationship between minimum wage increases and employment for retirement-aged individuals with slightly higher wages.

The coefficient on the indicator for the year in which a person reaches FRA is small and not statistically significant. The FRA year is only relevant for determining the monthly OASDI payment amount if a person claims retirement benefits. Furthermore, if an individual stopped working as soon as they reached full retirement age, the individual would still be coded as employed in the year of their FRA in our analysis because we are using an annual employment indicator due to the nature of the IRS earnings data. Deshpande, Fadlon and Gray (2020) also find no relationship between the FRA and the presence of positive IRS earnings. So, it is perhaps not surprising to see that reaching FRA is not associated with reductions in annual employment in our setting, but is important to note that this does not rule out that there are labor supply effects of reaching the FRA that our data and method do not capture.

Table 3 shows results for employment broken out by full-time and part-time employment rather than any employment. The results suggest that the increase in employment is made up of increases in both full-time and part-time employment. The preferred specification with individual fixed effects shows a 1.1 percentage point increase in the probability of both types of employment after a 10% increase in the minimum wage, statistically significant at the 10% level. This corresponds to a 6% increase from the sample mean for full-time employment and a 4% increase from the sample mean for part-time employment.

The higher-wage sample in Panel B of Table 3 shows no significant relationship between minimum wages and full-time employment for retirement-aged individuals with slightly higher wages. There are, however, negative and statistically significant reductions in part-time employment for these relatively higher, but still low-wage, workers after an increase in the minimum wage. The specification with individual fixed effects shows a 0.651 percentage point reduction in the probability of part-time employment after a 10% increase in the minimum wage, statistically significant at the 10% level. This result is the opposite sign of the results for the main sample, so this does not suggest that unobserved confounders are driving the increase in employment for near-minimum-wage workers. It may suggest a labor-labor substitution effect by firms in order to offset some of the additional labor cost related to the minimum wage increase. If businesses have to pay higher wages to the lowest part of the wage distribution, and perhaps also experience an increase in the supply of these workers, then they may offset the additional cost associated with these workers by reducing employment of slightly-higher wage, part-time workers.

Figure 6 shows results for the distributed lag model for employment. The results are based on the specification for any employment with individual fixed effects in column (3) of Table 2. The leading coefficients are generally flat and are not statistically significant. The contemporaneous and lag coefficients are all positive and statistically significant. The contemporaneous coefficient is approximately 0.19, indicating a 1.9 percentage point increase in the probability of employment in the year of a 10% increase in the minimum wage. The lag coefficients are even larger, suggesting that the increase in the probability of employment after an increase in the minimum wage lasts for at least three years after the minimum wage increase.

5.2 Permanent exit from employment

Results for the effect of minimum wage increases on permanent employment exit for low-wage, retirement-age individuals are shown in Table 4. Panel A shows results for the main sample, whereas Panel B shows results for the higher-wage sample. Columns (1)-(2) of Panel A show a negative relationship but no statistically significant relationship between minimum wages and partial exit. Columns (3)-(4) show a negative and statistically significant effect of minimum wage increases on full exit from employment. The coefficient for the log minimum wage ranges from -0.0641 when only state, year, and age fixed effects are included in the specification to -0.0512 when covariates are added. Our preferred model is the one with covariates in columns (2) and (4). The coefficient of -0.0512 in column (4) represents a 0.512 percentage point reduction in the probability of fully exiting employment in a given month after a 10% increase in the minimum wage, statistically significant at the 10% level. This corresponds to a 6.4% reduction from the sample mean rate of full exit in a given month. There is no relationship between the year an individual reaches FRA and partial exit or full exit from employment.[26] Finally, the higher-wage sample in Panel B shows no relationship between minimum wage increases and employment exit.

The distributed-lag results are shown in Figure 7. There is no visual evidence of leading

[26]Our permanent exit hazards are most similar to the self-reported retirement hazards in Behaghel and Blau (2012), which find no statistically significant relationship between the FRA and retirement.

effects prior to the minimum wage change. The probability of full exit then reduces by approximately 1.2 percentage points for a 10% increase in the minimum wage in the month of the minimum wage change, which is a 15% reduction from the sample mean monthly full exit rate. The reduction in exit associated with minimum wage increases then steadily goes away in years following the minimum wage increase. Only the year of the minimum wage increase is associated with a statistically significant reduction in full exit.

5.3 Retirement benefit claiming

Results for the effect of minimum wage increases on retirement benefit claiming for low-wage, retirement-age individuals are shown in Table 5. Panel A shows results for the main sample, whereas Panel B shows results for the higher-wage sample. The results in Panel A show a negative and statistically significant effect of minimum wage increases on benefit claiming. The coefficient for the log minimum wage variable ranges from -0.0351 when only state, time, and age fixed effects are included in the specification to -0.0380 when covariates are added. Our preferred model is the one with covariates in column (2). The coefficient of -0.0380 represents a 0.380 percentage point reduction in the probability of claiming retirement benefits in a given month after a 10% increase in the minimum wage, statistically significant at the 1% level. This corresponds to a 9% reduction from the sample mean rate of claiming in a given month. The coefficient for the indicator for the month when an individual reaches their FRA is positive and statistically significant: the month of reaching FRA is associated with a 5.15 percentage point increase in the probability of claiming retirement benefits.

Unlike the negative and statistically significant relationship between minimum wages and claiming retirement benefits in Panel A, the higher-wage sample in Panel B shows either no relationship or a small positive relationship depending on the specification. The coefficient in column (1) is 0.0047 while the coefficient in column (2) is 0.0142 and statistically significant at the 10% level. These results are both small relative to the results in Panel A and of the opposite sign. One interpretation for a possible small positive effect is related to the evidence for a small labor-labor substitution effect from Table 3 discussed in Section 5.1: if some firms respond to higher minimum wages and increased labor supply of retirement-age, near-minimum-wage workers by reducing employment of part-time, higher-wage individuals, then those higher-wage individuals who lose their employment may turn to retirement benefits to offset their lost employment income. Finally, the coefficient for the month during which an individual reaches their FRA is very similar to the result in Panel A: reaching FRA is associated with a 5.15 percentage point increase in the probability of claiming retirement benefits for near-minimum-wage workers and a 5.42 percentage point increase for higher-wage workers.

The distributed-lag results are shown in Figure 8. There is no visual evidence of leading effects prior to the minimum wage change. The probability of claiming then reduces by approximately 0.6 percentage points for a 10% increase in the minimum wage in the month of the minimum wage change, which is a 15% reduction from the sample mean monthly claiming rate. The reduction remains at 0.6 percentage points and statistically significant for the six-month lag. Beyond six months, the difference in the probability of claiming associated with a minimum wage increase returns to near zero and is not statistically significant. The lack of evidence for spurious leading effects and the evidence for temporary rather than permanent reductions in claiming suggest that we are capturing real delays in claiming that result from changes in the minimum wage.

5.4 Robustness to geographic unobservables

Prior work has shown that estimates of the effect of minimum wages on employment can be sensitive to the inclusion of state-specific time trends and geographic-specific time fixed effects (Allegretto, Dube and Reich, 2011; Allegretto et al., 2017; Dube, Lester and Reich, 2010; Neumark, Salas and Wascher, 2014b,a; Totty, 2017). Table 6, Table 7, and Table 8 show robustness of our employment, employment exit, and benefit claiming results, respectively, to the inclusion of such controls. We separately add state-specific linear time trends, age-by-time fixed effects, and Census Region-by-time fixed effects to our preferred specification from the main results to ensure that the coefficient on the log minimum wage is not sensitive to particular sets of controls. We then also show results with all three sets of robustness controls included together. State-specific linear time trends account for possible heterogeneous trends in employment, exit, or retirement benefit claiming across states that may be correlated with minimum wage policy. Census Region-specific time effects account for regional macroeconomic shocks that may influence employment, exit, or claiming and be correlated with minimum wage policy. We also add age-by-time fixed effects which can account for changes in trends towards later or earlier retirement over time.

Unlike prior research on minimum wages and employment, our results on minimum wages and retirement-related outcomes are not sensitive to the inclusion of these controls. All of the employment results in Table 6 are larger than our preferred specification in Table 2. The coefficient increases from 0.151 in column (3) of Table 2 to 0.189 for the most saturated specification in column (4), representing a 1.89 percentage point increase in employment after a 10% increase in the minimum wage. The results for partial and full exit in Table 7 are also similar to the main results in Table 4. There is some evidence of a reduction in partial exit from employment after minimum wage increases in columns (1)-(4), but the result is only statistically significant in column (1) when linear state-specific trends are included. The results for full employment exit in columns (5)-(8) show more consistent evidence of a reduction in full exit. Similar to the employment results, the coefficient increases in magnitude from -0.0512 in column (4) of Table 4 to -0.0974 for the most saturated specification in column (8), representing a 0.974 percentage point reduction in full employment exit after a 10% increase in the minimum wage. The statistical significance level also increases from 10% to 5%. Finally, the results for benefit claiming in Table 8 also show a similar pattern. The coefficient increases in magnitude from -0.0380 in column (3) of Table 5 to -0.0475 for the most saturated specification in column (4), representing a 0.475 percentage point reduction in retirement benefit claiming after a 10% increase in the minimum wage.

5.5 Potential mechanisms

In this section we explore possible mechanisms for the results described so far. As discussed in Section 2.3, minimum wage increases could decrease employment via reduced labor demand due to higher labor costs or reduced labor supply via income effects. Our findings of increased employment and delayed full employment exit are consistent with neither of these hypotheses. Instead, our findings suggest that labor supply substitution effects via higher wages are the dominant force. These results are consistent with prior empirical and theoretical evidence discussed in Section 2.3 suggesting that labor supply substitution effects would be an important mechanism for the relationship between minimum wages and labor market outcomes for retirement-age workers.

The mechanism for reduced retirement benefit claiming after minimum wage increases is not as straightforward. As described in Section 2.3, delayed claiming is consistent with both an employment channel mechanism, in which individuals delay claiming retirement benefits because they continue working longer, and a mechanism involving the interaction between minimum wages and the Social Security earnings test, in which individuals delay claiming because higher minimum wages push their earnings across or further beyond the earnings test threshold thereby reducing the immediate return to claiming. We distinguish between these two mechanisms by exploiting the removal of the post-FRA Social Security earnings test in the year 2000. Splitting our sample of analysis into pre- vs post-FRA ages and pre- vs post-2000, three of the four groups are subject to a Social Security earnings test. We therefore add a post-FRA variable, post-2000 variable, and interaction between the two variables into our model and then also interact them with the log minimum wage.

The results are shown in column (1) of Table 9. The full marginal change in the probability of claiming after a change in the minimum wage for each of the four age-by-time groups is shown at the bottom of the table based on the linear combination of the appropriate coefficients in the top of the table. The marginal change is negative and statistically significant for all three of the groups that are subject to an earnings test. For the post-FRA, post-2000 group, the result is the opposite: an increase in the minimum wage is associated with a statistically significant increase in claiming retirement benefits. This finding is similar to research showing that the removal of the post-FRA earnings test in 2000 accelerated retirement benefit claiming for individuals who had reached their FRA (Gruber and Orszag, 2003; Loughran and Haider, 2005; Song and

Manchester, 2007).

Column (2) of Table 9 implements the same mechanism test, but for employment rather than the claiming model. As described earlier, the earnings test threshold is not explicitly associated with the extensive margin of labor supply. We therefore do not necessarily expect the post-FRA, post-2000 group to show a different result for the relationship between minimum wages and employment. The results are consistent with this hypothesis. The marginal change in the probability of employment after an increase in the minimum wage is positive for all four groups. The pre-2000 groups show smaller changes that are not statistically significant, but this is potentially explained by prior work suggesting that effects of minimum wages differ between the 1980s through early-1990s time period versus later time periods (Cengiz et al., 2019; Godoy, Reich and Allegretto, 2019).[27]

In summary, these results suggest that the mechanism driving delays in retirement benefit claiming is the interaction between minimum wages and a binding Social Security earnings test rather than a behavioral tendency to tie the timing of benefit claiming with labor force exit.

5.6 Own-wage elasticity

Our results so far show that minimum wage increases lead to increased employment for low-wage, retirement-age workers. Prior evidence on minimum wages and labor market outcomes for older workers is limited and mixed but also finds some evidence of increased employment for retirement-age workers (Borgschulte and Cho, 2020; Fang and Gunderson, 2009). However, as described in Section 2.1, an issue in comparing minimum wage-employment elasticities across different studies is that the bite of the minimum wage may differ across the populations studied. In order to make our results more comparable to past work, we calculate own-wage elasticity estimates of employment for our low-wage, retirement-age population. For comparison, we also compute an own-wage elasticity estimate for younger (ages 40-54) low-wage workers in our data using an otherwise identical sample construction and methodology.

[27] The nature of minimum wage policy was different in the 1980s and 1990s than today in terms of the value of the minimum wage relative to median wages and the relative frequency of state versus federal minimum wage changes. The poolability of 1980s-1990s and 2000s-onward data for minimum wage analysis is beyond the scope of this paper, but merits future work.

Our own-wage elasticity estimates are the ratio of the percent change in employment from a minimum wage increase to the percent change in wage from a minimum wage increase. The percent change in employment is available from our main results presented earlier: the change in employment represented from the coefficient in column (3) of Table 2 divided by the mean employment rate in Table 1. The percent change in wage is calculated following equation (2) in Cengiz et al. (2019). Specifically, we first calculate the elasticity of the wage bill, or total earnings, with respect to the minimum wage. We do this by estimating the same regression as we did for employment, except we replace the binary outcome variable for the presence of positive IRS earnings with the log of the amount of IRS earnings. We then use the wage bill elasticity along with the employment elasticity to compute a wage elasticity following Cengiz et al. (2019).

The own-wage elasticity of employment from the minimum wage for low-wage, retirement-age workers is shown in column (1) of Table 10. Panel A reproduces the preferred employment regression from Table 2 along with the sample mean employment rate from Table 1. Panel B reports the wage bill regression, which shows a statistically significant increase in log earnings associated with minimum wage increases. Panel C reports the resulting elasticities. The own-wage employment elasticity for low-wage, retirement-age workers is 1.042. Column (2) shows results when we repeat the analysis, but instead use low-wage individuals who are ages 40-54. This prime-age group shows a small negative effect on employment and a positive effect on log earnings for an own-wage elasticity of -0.144. These results suggest that retirement-age low-wage workers have a significant labor supply response to minimum wage increases, while younger low-wage workers experience a reduction in employment that is small relative to the wage gains.

6 Discussion

Our results show that minimum wage increases lead to increased employment, delayed full permanent exit from employment, and delayed claiming of retirement benefits for low-wage, retirement-age individuals. We now provide a discussion of the results presented so far in relation to the literature. We attempt to relate our results to prior evidence on minimum wages and employment, labor supply around retirement, and retirement benefit claiming.

Our evidence of increased employment stands in mild contrast to research on minimum wages and non-retirement-age populations which tends to find either no effect on employment or small negative effects. However, our own-wage elasticity estimate of -0.144 for the younger low-wage group is in line with prior work on minimum wages and employment. For example, Dube (2019) found that the median own-wage elasticity across 55 minimum wage studies was -0.17. These studies typically use younger populations relative to our retirement-age population. This suggests that our large own-wage elasticity of 1.042 for low-wage, retirement-age workers is a result of something unique about retirement-age individuals rather than an issue with our data, sample construction, or methodology.

Borgschulte and Cho (2020) provide the only other evidence on own-wage elasticity for retirement-age workers. The employment and weekly earnings elasticities in column (5) of Table 3 in their paper impute an own-wage elasticity of 0.811 after converting the weekly earnings elasticity into a wage elasticity using equation (2) in Cengiz et al. (2019). Like ours, this own-wage elasticity is large relative to the non-retirement-age studies, although column (5) of Table 3 is the largest employment elasticity in their paper: it is 0.126, statistically significant at the 10% level, while the estimate from their preferred specification in column (2) is 0.059 and not statistically significant.[28,29]

Thus, while there is prior evidence that minimum wage increases may lead to increased employment for low-wage, retirement-age workers, our results are the largest and most consistent in this direction. Our results also contribute to an emerging literature showing that there can be positive labor supply effects of minimum wage increases. Godoy, Reich and Allegretto (2019) find that minimum wage increases lead to increased parental labor supply, with own-wage elasticity estimates of 0.471 for single mothers and 0.346 for married fathers. Agan and

[28]We use the results in column (5) of their paper to compute an own-wage elasticity, instead of their preferred specification in column (2), because column (5) has a corresponding earnings elasticity needed to compute an own-wage elasticity.

[29]The interactive fixed effects specifications in their online appendix all generate much smaller own-wage elasticities. For example, when 6 factors are included, which is when the estimates appear to stabilize, the imputed own-wage elasticity would be -0.057.

Makowsky (2018) find that minimum wage increases lead to reduced criminal recidivism due to labor-crime substitution effects that draw individuals back into the legal labor market.

We hypothesized in Section 2.3 that minimum wages may lead to increased employment for retirement-age workers due to labor supply substitution effects. As it turns out, our own-wage elasticity estimates are similar in magnitude to life-cycle model estimates of the intertemporal elasticity of labor supply to wages around retirement age. The life-cycle model in French (2005) estimates the labor supply elasticity to be 1.04-1.33 at age 60. French and Jones (2011) extend the model and estimate the labor supply elasticity from a permanent wage change to be 1.17 in the year of the wage change and 2.24 in years after the wage change. Imai and Keane (2004) estimate an intertemporal labor supply model with human capital accumulation and find that the intensive labor supply elasticity with respect to wage rises to 2 by age 60. Ameriks et al. (2020) use strategic survey questions to estimate the impact of job opportunities for flexible schedules on how much older Americans work and in doing so estimate a median intertemporal elasticity of substitution of labor supply with respect to wage to be 0.85 around the time of retirement. It is reassuring that our result for the own-wage elasticity of employment from the minimum wage is similar to results on the elasticity of labor supply with respect to wage.

Our results on employment and labor force exit are also related to Gorodnichenko, Song and Stolyarov (2013), from whom we obtained our definitions of partial and full exit from employment based on annual earnings from Social Security Administration data.[30] They find that increases in the unemployment rate and inflation both lead to increases in partial and full permanent exit from employment. They report that a one percentage point increase in the unemployment rate leads to a 1-2 percentage point decrease in employment for older workers, depending on the exact age range, which is made up of increased flows into both partial and full permanent exit. We find that a 10% increase in the minimum wage leads to a 1.5 percentage point increase in employment for older low-wage workers, which is made up mostly of reductions in full permanent exit from employment. Thus, our results suggest that, for low-wage workers, a 10% increase in the minimum wage has a similar effect on employment and labor force exit

[30] What we refer to as partial exit and full exit they refer to as partial retirement and full retirement.

as a one percentage point reduction in the unemployment rate.

Finally, our results for retirement benefit claiming are comparable to the results in Borgschulte and Cho (2020) and Haaga and Johnson (2012) on the effects on minimum wages and the unemployment rate on claiming. Borgschulte and Cho (2020) study the relationship between minimum wages and total annual OASDI retirement beneficiaries over 1983-2016 and estimate an elasticity of -0.033 in their preferred specification, while their full set of specifications that account for unobserved geographical heterogeneity range from -0.011 to -0.033.[31] This implies that a 10% increase in the minimum wage lowers total annual OASDI retirement beneficiaries by 0.11-0.33%. We estimate that a 10% increase in the minimum wage lowers the individual-level likelihood of claiming by 9% using data from 1978-2014. In order to compare our elasticity to Borgschulte and Cho (2020), we need to translate our monthly flow elasticity into an annual stock elasticity. We do this via some back-of-the-envelope calculations. The average number of *new* retirement beneficiary awards per year from 1978-2014 was 1.9 million.[32] Assuming, based on Figure 1, that approximately 25% of retirement-age workers are within $2 of the minimum wage and thus would fall into our sample of analysis, a 9% reduction in annual flows into retirement benefits for 25% of the 1.9 million average annual new recipients would be 42,750 fewer new recipients in a given year. The average annual *total* beneficiaries from 1983-2016 was 32.5 million.[33] A reduction of 42,750 from 32.5 million would be a reduction of 0.13%. Thus, our monthly claiming hazard elasticity of -0.9 translates to an annual total beneficiaries elasticity of -0.013, which is similar to the -0.011 to -0.033 range in Borgschulte and Cho (2020).

Haaga and Johnson (2012) study the relationship between the unemployment rate and OASDI retirement benefit claiming. They find that a one percentage point increase in the unemployment rate increases OASDI retirement benefit claiming by 0.4 percentage points. We find that a 10% increase in the minimum wage reduces retirement benefit claiming by 0.38 percentage points. Thus, our results suggest that, for low-wage workers, a 10% increase in the minimum wage has a similar effect on retirement benefit claiming as a one percentage point

[31] See columns (2)-(4) of Table 5 in Borgschulte and Cho (2020).

[32] See the "Retired workers" column on page 6.1 at https://www.ssa.gov/policy/docs/statcomps/supplement/2020/supplement20.pdf.

[33] See the "Retired workers and dependents" column at https://www.ssa.gov/oact/STATS/OASDIbenies.html.

reduction in the unemployment rate.

To summarize, we find positive employment effects that are larger than the existing literature on minimum wages and employment, but are consistent with a recent emerging literature on labor supply effects of minimum wages and life-cycle models of labor supply around retirement. We also find delays in permanent exit from employment and retirement benefit claiming that are consistent with and similar in magnitude to prior work on minimum wages and the unemployment rate. Together, these results suggest that the labor supply and benefit claiming of retirement-age workers is very responsive to changes in labor market opportunities.

7 Conclusion

We used linked data between the SIPP, IRS, and SSA to study the effect of minimum wage increases on employment, permanent exit from the labor force, and retirement benefit claiming for low-wage, retirement-age individuals. We find that minimum wage increases lead to increased employment, delayed full permanent exit from employment, and delayed claiming of retirement benefits. A primary takeaway of these results is that low-wage, retirement-age individuals are very responsive to changes in labor market opportunities.

Increased employment and delayed benefit claiming associated with minimum wage increases are beneficial to workers for numerous reasons. Continuing to work while of retirement age, even after claiming retirement benefits, can increase an individual's monthly benefit payment if their annual earnings are among their 35 highest earning years. Furthermore, choosing to delay the claiming of retirement benefits increases the monthly payment once benefits are claimed. The amount of the increase for each month of delay depends on year of birth and whether the delay occurs before or after reaching the FRA, but is generally around one-half of one percent. Delaying the claiming of retirement benefits by six months would therefore increase the monthly payment by approximately 3%. Thus, minimum wage increases appear to improve the financial well-being of low-wage workers not only by increasing their earnings via higher wages while they are working, but potentially by also increasing their OASDI retirement benefit payments via additional years of earnings and delayed claiming.

The results also suggest that minimum wage increases may be helpful as another policy lever for encouraging more work and later retirement. The Social Security eligibility rules and thresholds have been adjusted many times in recent decades in order to encourage more work and later retirement. Our results suggest that minimum wage increases can achieve the same goals. Labor supply effects for older workers are not typically listed among the primary reasons for raising the minimum wage, but they could be valuable given other research showing, for example, the "stickiness" of retirement at age 65 despite adjustments in the FRA to encourage later retirement (Deshpande, Fadlon and Gray, 2020) and potential adverse selection in response to adjustments in delayed retirement credits that may actually increase Social Security payouts and worsen OASDI funding issues (Duggan et al., 2021). Future work should more carefully consider minimum wages as another policy lever for influencing employment and retirement benefit claiming, including how minimum wages interact with other factors that influence retirement and their net effect on the Social Security system.

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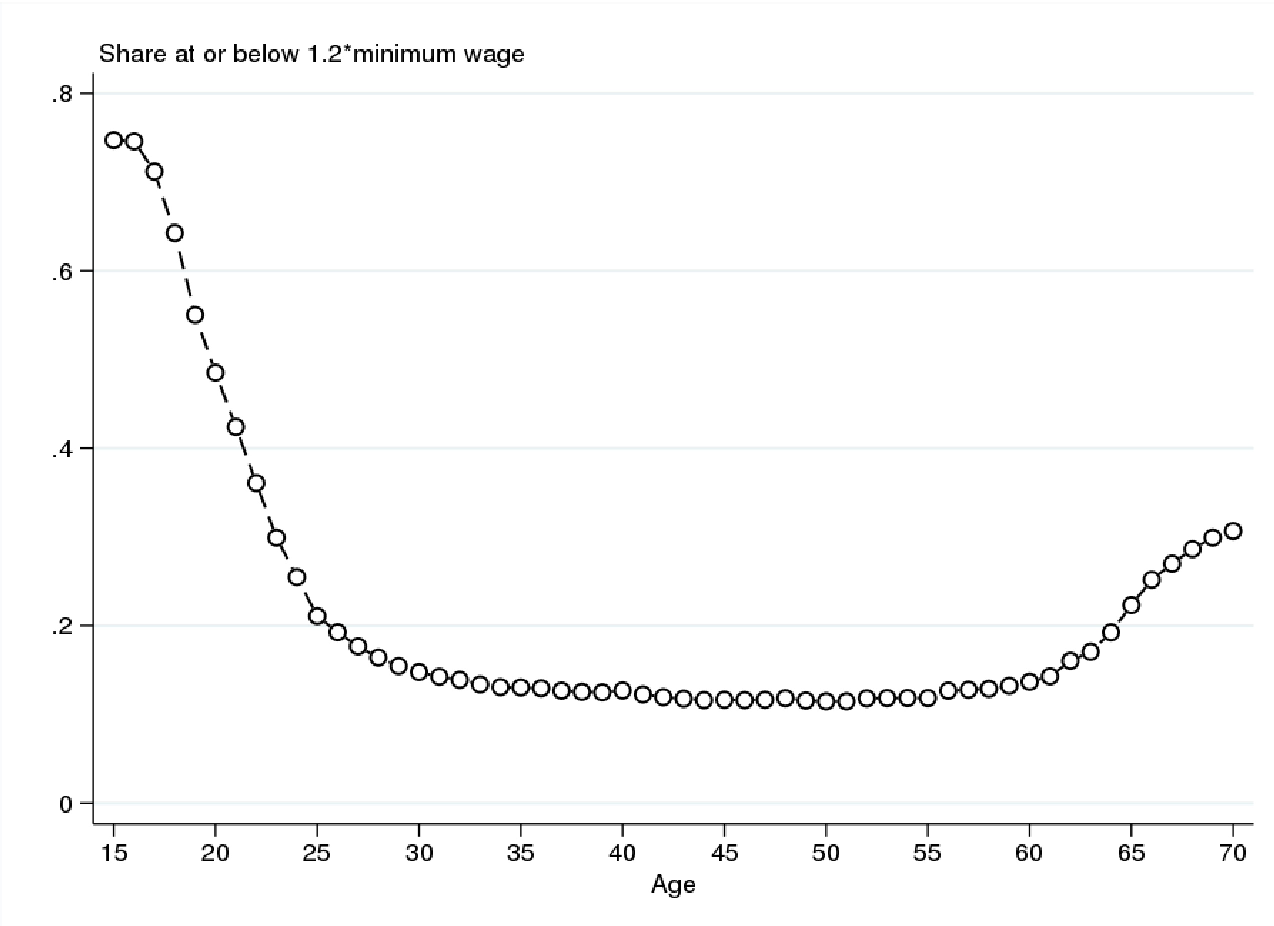
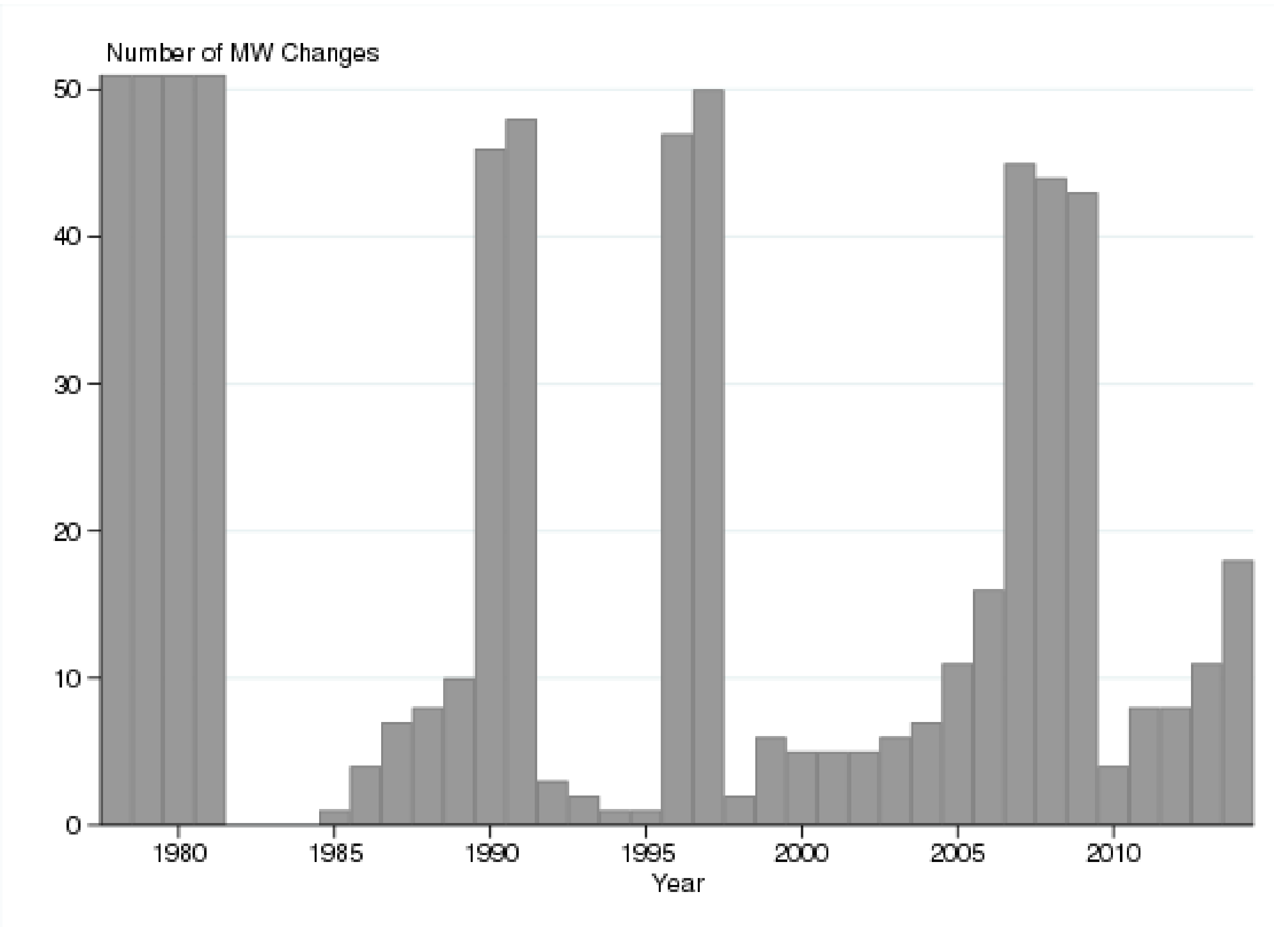
8 Figures and Tables

Figure 1: Share of Workers Earning Near Minimum Wage By Age



Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.
Note: Share of person-month wage observations less than or equal to 120% of the minimum wage in the SIPP Gold Standard File, by age.

Figure 2: Distribution of Minimum Wage Change Events

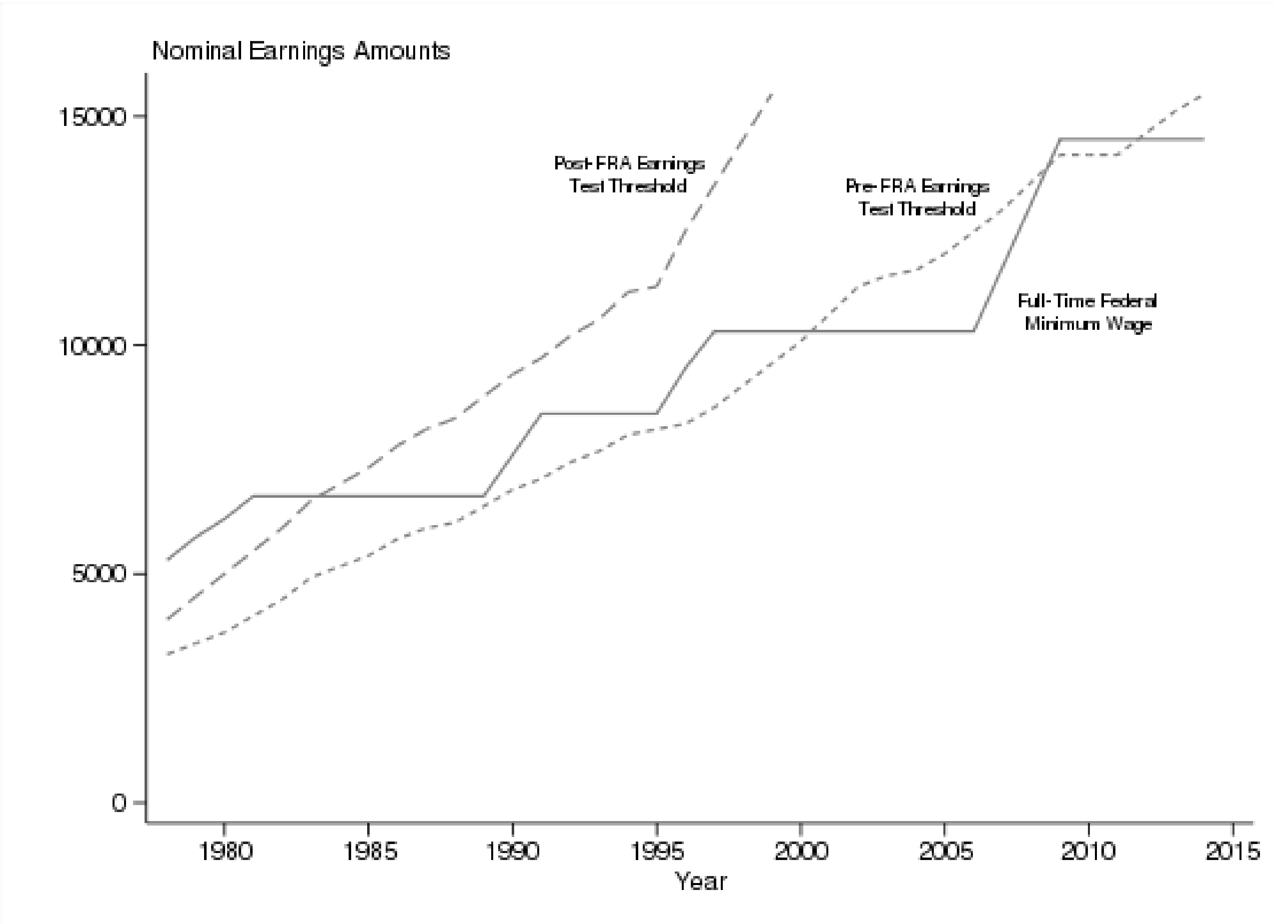


Source: Minimum wage data from Vaghul and Zipperer (2016).
Note: Number of states (including District of Columbia) with a change in the binding minimum wage from 1978-2014. The binding minimum wage is the higher of the state and federal minimum wage.

Figure 3: Average Annual Log Wage Growth By Age

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY21-CED001-B0004.
Note: Annual wage is computed as annual earnings divided by annual hours. Annual wages for years that are not fully in scope for the survey are excluded. Annual wages less than \$1/hour or more than \$100/hour are excluded. The bottom and top 5% of annual log wage changes are also excluded.

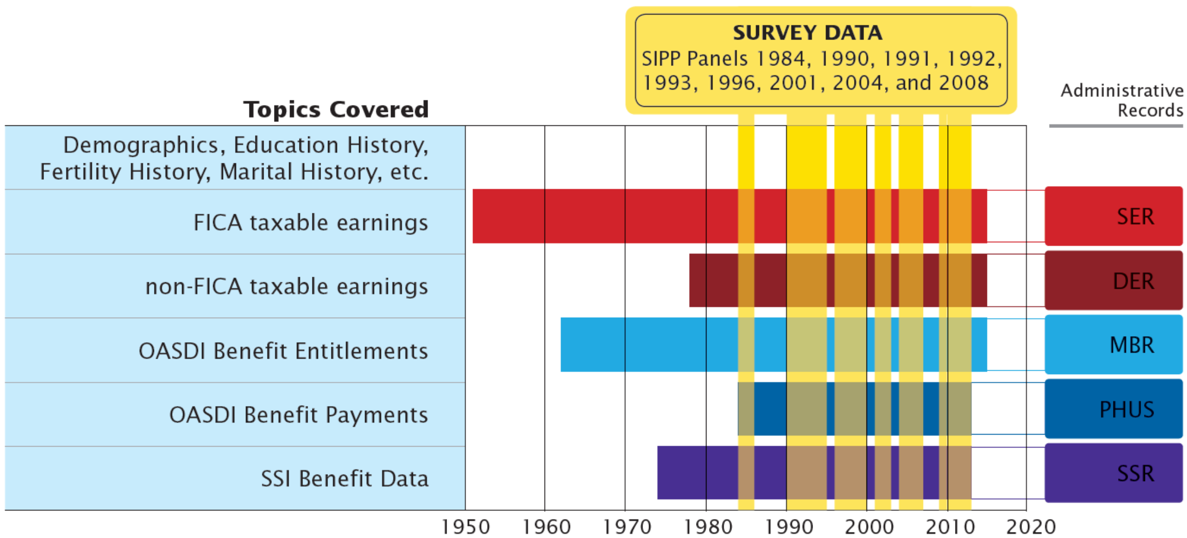
Figure 4: Comparison of Full-Time Minimum Wage and Social Security Earnings Test Amounts Over Time



Source: Minimum wage data from Vaghul and Zipperer (2016). Earnings test threshold data from ssa.gov/OACT/COLA/rteahistory.html and ssa.gov/OACT/COLA/rtea.html.
Note: Full-time minimum wage earnings are based on working for the federal minimum wage, 40 hours per week, 50 weeks out of the year. Pre-FRA earnings test threshold amounts for 2000 onward are based on the threshold amounts for years before the year in which the individual reaches FRA.

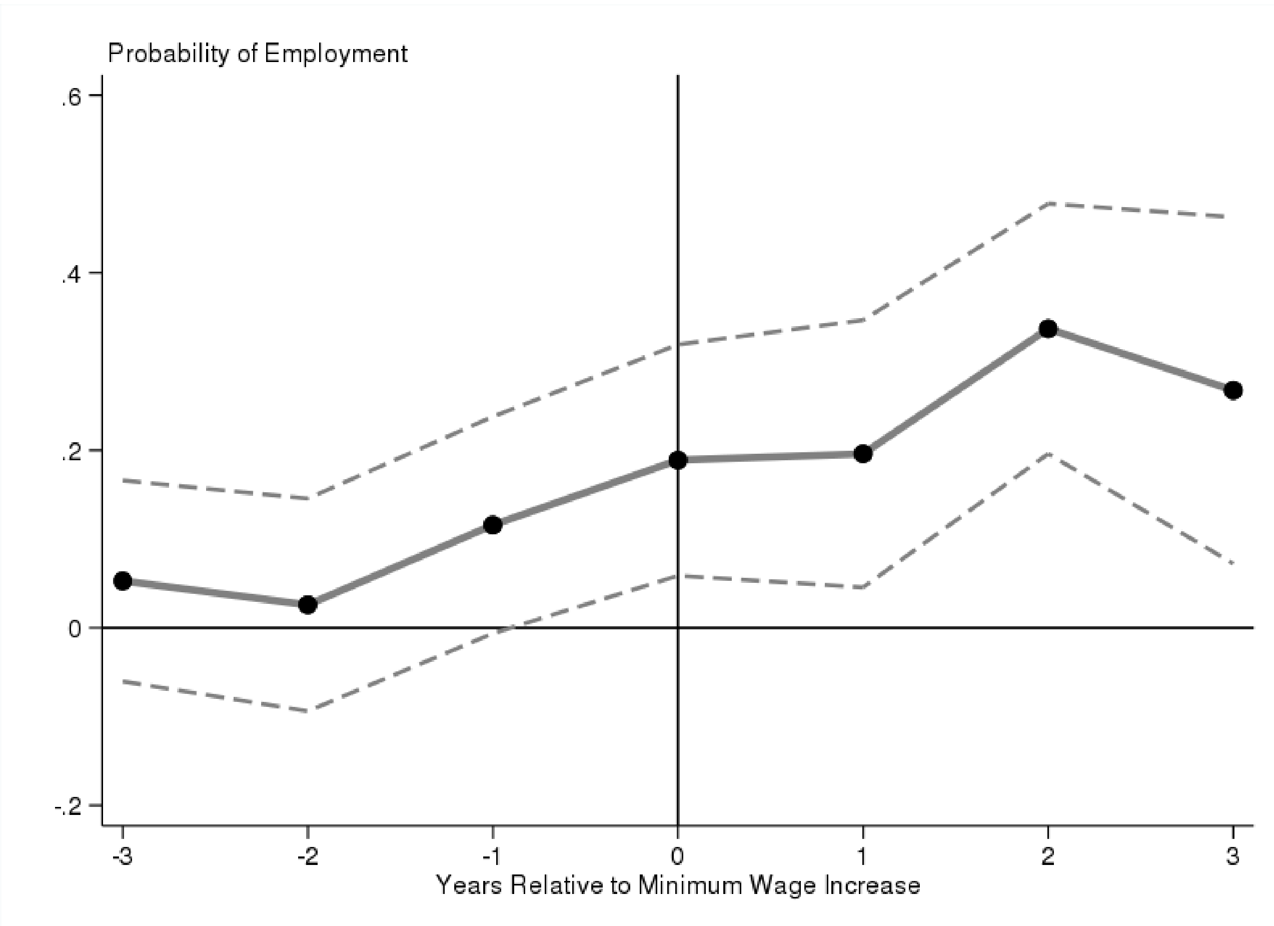
Figure 5: Gold Standard File (GSF) Data Sources and Variable Availability Over Time

Data Sources New and Improved.PNG



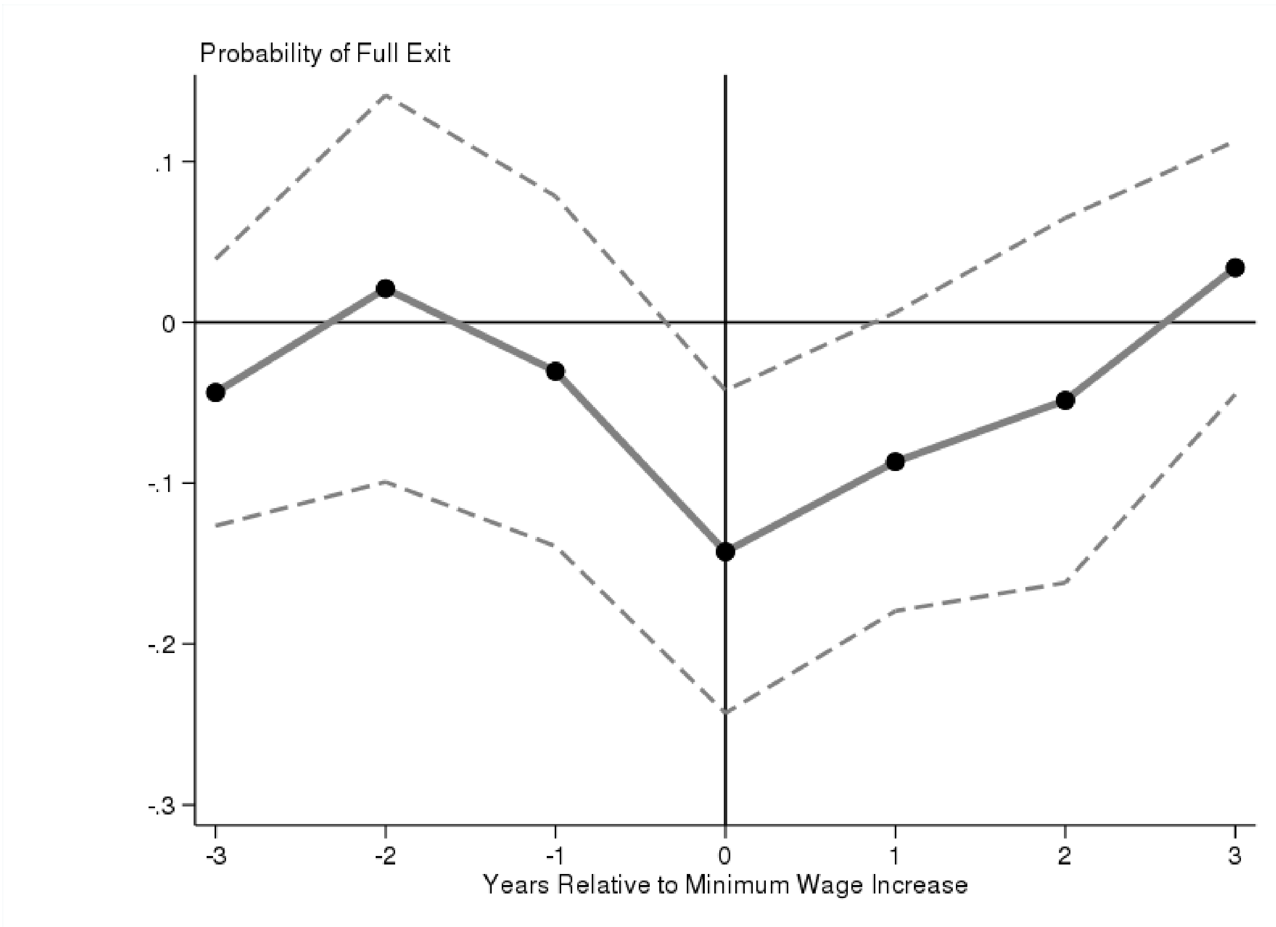
Source: U.S. Census Bureau.
Note: .

Figure 6: Distributed Lag Model: Employment



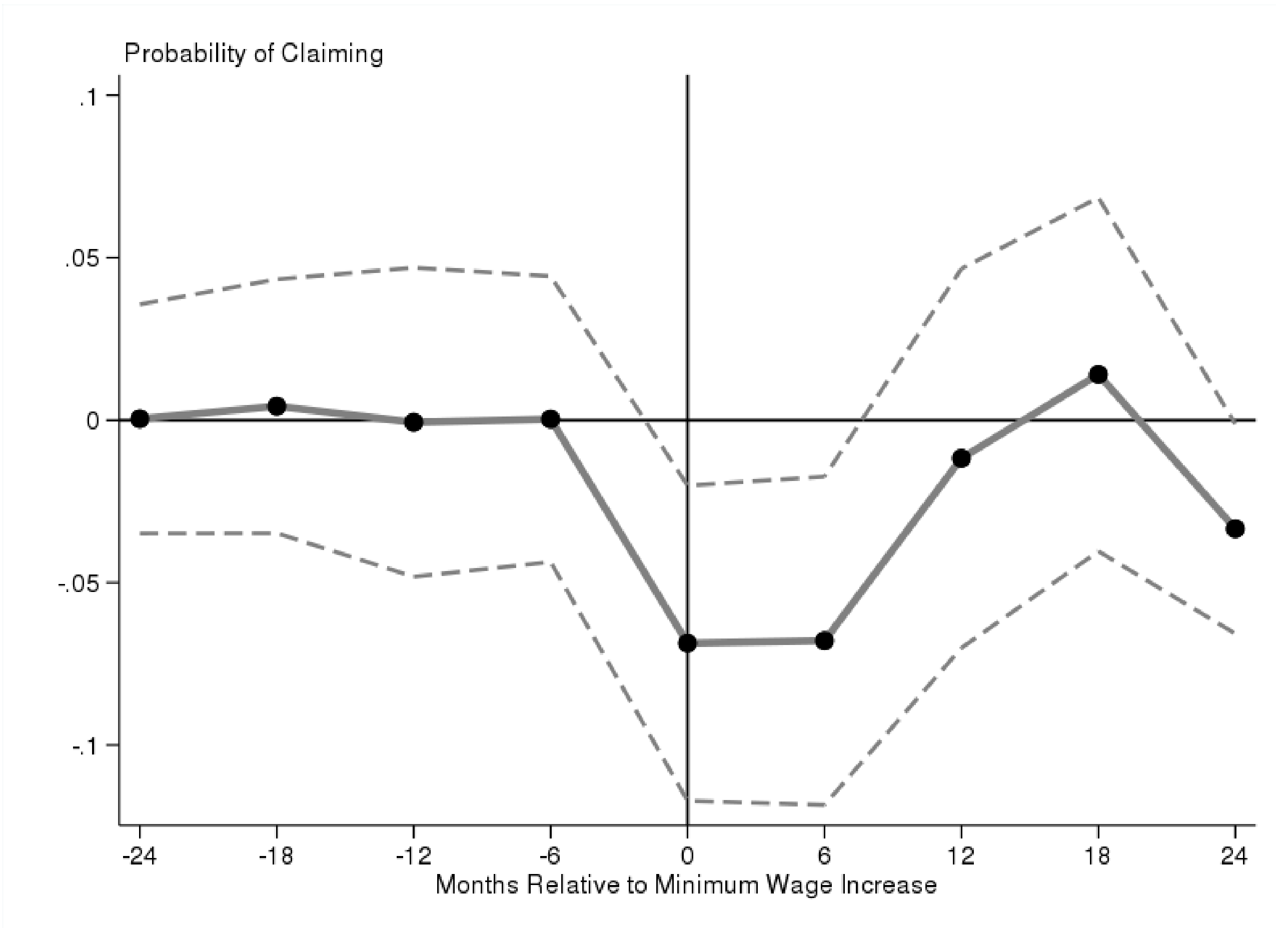
Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.
Note: The dots are OLS regression coefficients from a balanced panel model for employment (presence of positive earnings reported to the IRS) over ages 62-70 with a distributed-lag in the log of the minimum wage. The dashed lines are 90% confidence intervals based on standard errors clustered at the state level.

Figure 7: Distributed Lag Model: Full Labor Force Exit



Source: U.S. Census Bureau Gold Standard File, CBDRB-FY21-CED001-B0002.
Note: The dots are OLS regression coefficients from a hazard model for full labor force exit over ages 62-70 with a distributed-lag in the log of the minimum wage. The dashed lines are 90% confidence intervals based on standard errors clustered at the state level.

Figure 8: Distributed Lag Model: Hazard Model for OASDI Retirement Benefit Claiming



Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.
Note: The dots are OLS regression coefficients from a hazard model for retirement benefit receipt with a distributed-lag in the log of the minimum wage. The dashed lines are 90% confidence intervals based on standard errors clustered at the state level.

Table 1: Summary Statistics

	(1)	(2)	(3)	(4)
	Employment Sample		Benefit Claiming Sample	
	Mean	SD	Mean	SD
Time-Invariant Variables				
Male	0.378	0.485	0.408	0.492
White	0.837	0.369	0.839	0.368
Black	0.127	0.333	0.135	0.341
Other Race	0.0357	0.185	0.0264	0.16
Hispanic	0.0716	0.258	0.0717	0.258
Less than High School Degree	0.310	0.463	0.290	0.454
High School Degree	0.371	0.483	0.365	0.482
Some College	0.207	0.405	0.231	0.422
Bachelor's Degree	0.0673	0.251	0.0704	0.256
More than Bachelor's	0.0447	0.207	0.0437	0.204
Never Married	0.0523	0.223	0.0556	0.229
Currently Married	0.609	0.488	0.605	0.489
Widowed or Divorced	0.339	0.473	0.339	0.474
Birth Year	1930	7.46	1930	8.76
Health Insurance Coverage	0.876	0.329	0.876	0.33
Employer-Provided Insurance	0.297	0.457	0.351	0.477
Defined Benefit Pension	0.0513	0.221	0.0556	0.229
Defined Coverage Pension	0.0327	0.178	0.0402	0.196
Total Net Worth	166,000	578,000	161,000	569,000
Non-Housing Wealth	94,200	55,4000	91,700	544,000
Received Retirement Benefits by Age 70	-	-	0.912	0.283
Age at First Receipt	-	-	65.1	1.59
Individuals	3,000		3,100	
Time-Varying Variables				
Employed	0.679	0.467	-	-
Full-Time Employed	0.199	0.399	-	-
Part-Time Employed	0.276	0.447	-	-
Full Labor Force Exit	0.080	0.271	-	-
Partial Labor Force Exit	0.043	0.204	-	-
Claimed Retirement Benefits (First Receipt)	-	-	0.0414	0.1992
Log Minimum Wage	4.92	1.24	4.94	1.37
Unemployment Rate	5.90	1.84	6.04	1.95
Inflation-Adjusted Annual IRS Earnings	7,640	47,300	16,000	99,100
Observations	27,000		68,500	

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003, CBDRB-FY20-CED001-B0005.

Note: Summary statistics are shown separately for the two main samples used in the analysis. The top part of the table shows cross-section mean and standard deviation for the time-invariant variables used in the model, as well as two descriptive retirement variables in the first two rows. The bottom part shows the statistics for the time-varying variables over the full panel sample.

Table 2: The Effect of Minimum Wages on Employment of Low-Wage, Retirement-Age Individuals

	(1)	(2)	(3)
	Employed (Has IRS Earnings)		
Panel A: Main Sample			
Log Minimum Wage	0.214***	0.175***	0.151***
	(0.0654)	(0.0646)	(0.0429)
Year of FRA		0.00652	0.00658
		(0.00891)	(0.00890)
Observations	27,000	27,000	27,000
Panel B: Higher-Wage Sample			
Log Minimum Wage	-0.0691	-0.0592	0.00184
	(0.0573)	(0.0612)	(0.0587)
Year of FRA		0.0129*	0.0127*
		(0.00734)	(0.00731)
Observations	23,000	23,000	23,000
State, Year, Age Fixed Effects	Yes	Yes	Yes
Covariates	No	Yes	Yes
Individual Fixed Effects	No	No	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.

Note: Each column reports OLS regression results based on a balanced person-year panel over ages 62-70. The dependent variable in columns (1)-(2) is equal to one if the individual had any earnings reported to the IRS and zero otherwise. Age fixed effects are in years. Covariates include an indicator for the year in which the individual reached FRA, the log of the state-year unemployment rate, sex, race, Hispanic status, highest education level, health insurance coverage, health insurance coverage through employer, defined-benefit pension, defined-contribution pension, marital status, removal of the post-FRA Social Security earnings test in 2000, total net worth, non-housing wealth, and birth year fixed effects. The individual fixed effects specification retains the indicator for the year in which the individual reached FRA, the log of the state-year unemployment rate, and removal of the post-FRA Social Security earnings test in 2000; the other covariates drop out of the model. State fixed effects also drop out of the model. Standard errors are clustered at the state level.

Table 3: The Effect of Minimum Wages on Employment Broken Out By Full-Time and Part-Time Employment Based on IRS Earnings Amounts

	(1)	(2)	(3)	(4)	(5)	(6)
	Full-Time Employed			Part-Time Employed		
Panel A: Main Sample						
Log Minimum Wage	0.207**	0.160**	0.112*	0.106	0.113*	0.111*
	(0.0849)	(0.0755)	(0.0660)	(0.0897)	(0.0678)	(0.0640)
Year of FRA		-0.00111	-0.00115		0.00534	0.00536
		(0.00635)	(0.00632)		(0.0123)	(0.0123)
Observations	27,000	27,000	27,000	27,000	27,000	27,000
Panel B: Higher-Wage Sample						
Log Minimum Wage	0.0482	0.0937	0.0635	-0.129**	-0.144***	-0.0651*
	(0.0966)	(0.0982)	(0.0681)	(0.0504)	(0.0474)	(0.0356)
Year of FRA		0.000807	0.000886		0.00939	0.00924
		(0.00837)	(0.00837)		(0.00914)	(0.00913)
Observations	23,000	23,000	23,000	23,000	23,000	23,000
State, Year, Age Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Covariates	No	Yes	Yes	No	Yes	Yes
Individual Fixed Effects	No	No	Yes	No	No	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.
Note: Each column reports OLS regression results based on a balanced person-year panel over ages 62-70. The dependent variable in columns (1)-3) is equal to one if the individual had inflation-adjusted earnings equal to at least 50% of their lifetime max dating back to 1951 and zero otherwise. The dependent variable in columns (4)-(6) is equal to one if the individual had inflation-adjusted earnings of less than 50% of their lifetime max but more than $5,000 and zero otherwise. See section 3.3 for more information on the measurement of full-time and part-time employment. See the notes for Table 2 for more information about standard errors and covariates.

Table 4: The Effect of Minimum Wages on Labor Force Exit Based on IRS Earnings Amounts

	(1)	(2)	(3)	(4)
	Partial Exit Hazard		Full Exit Hazard	
Panel A: Main Sample				
Log Minimum Wage	-0.0233	-0.0247	-0.0641**	-0.0512*
	(0.0247)	(0.0225)	(0.0274)	(0.0303)
Year of FRA		-0.00866		0.00865
		(0.0121)		(0.0165)
Observations	14,500	14,500	17,000	17,000
Panel B: Higher-Wage Sample				
Log Minimum Wage	0.0102	0.00840	-0.0142	-0.0105
	(0.0296)	(0.0283)	(0.0352)	(0.0386)
Year of FRA		0.00487		-0.0177*
		(0.0105)		(0.0104)
Observations	16,000	16,000	15,500	15,500
State, Year, and Age Fixed Effects	Yes	Yes	Yes	Yes
Covariates	No	Yes	No	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY21-CED001-B0002.
Note: Each column reports regression results for labor force exit hazards based on a person-year panel. The dependent variable is equal to zero in years prior to exiting the labor force and one in the year of exit. Individuals drop out of the sample after the year in which they first exit, or once they reach the year in which they turn 70 without having already exited. Partial exit is defined as when an individual's inflation-adjusted earnings permanently fall to less than 50% of their lifetime inflation-adjusted maximum but still at least $5,000 inflation-adjusted dollars. Full exit is defined as when an individual's inflation-adjusted earnings permanently fall to less than 50% of their lifetime inflation-adjusted maximum and less than $5,000 inflation-adjusted dollars (including zero). See section 3.3 for more information on the measurement of full-time and part-time employment. See the notes for Table 2 for more information about standard errors and covariates. covariates.

Table 5: The Effect of Minimum Wages on OASDI Retirement Benefit Claiming

	(1)	(2)
	Claimed Hazard	
Panel A: Main Sample		
Log Minimum Wage	-0.0351**	-0.0380***
	(0.0151)	(0.0139)
Month of FRA		0.515***
		(0.0240)
Observations	68,500	68,500
Panel B: Higher-Wage Sample		
Log Minimum Wage	0.0047	0.0142*
	(0.0083)	(0.0081)
Month of FRA		0.542***
		(0.0146)
Observations	95,500	95,500
State, Time, Age Fixed Effects	Yes	Yes
Covariates	No	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.

Note: Each column reports regression results for benefit claiming hazards based on a person-month panel. The dependent variable is equal to zero in months prior to first receipt of OASDI retirement benefits and one in the month of receipt. Individuals drop out of the sample after the month in which they first receive benefits, or once they reach the month in which they turn 70. Standard errors are clustered at the state level and shown in parentheses. Time fixed effects are at the year-month level. Age fixed effects are in months. Covariates include an indicator for the month in which the individual reached FRA, the log of the state-year unemployment rate, sex, race, Hispanic status, highest education level, health insurance coverage, health insurance coverage through employer, defined-benefit pension, defined-contribution pension, marital status, removal of the post-FRA Social Security earnings test in 2000, total net worth, non-housing wealth, annual IRS earnings, and birth year fixed effects.

Table 6: Employment Robustness Checks for Time and Geographic Heterogeneity

	(1)	(2)	(3)	(4)
	Employed			
Log Minimum Wage	0.154***	0.163***	0.197***	0.189***
	(0.0488)	(0.0452)	(0.0582)	(0.0599)
Observations	27,000	27,000	27,000	27,000
State, Time, Age Fixed Effects	Yes	Yes	Yes	Yes
Covariates	Yes	Yes	Yes	Yes
Individual Fixed Effects	Yes	Yes	Yes	Yes
State Linear Time Trends	Yes	No	No	Yes
Age-by-Time Fixed Effects	No	Yes	No	Yes
Region-by-Time Fixed Effects	No	No	Yes	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0005.

Note: The regressions are the same as column (3) of Table 2, except with additional fixed effects as indicated in the table. Region corresponds to the four Census Regions.

Table 7: Labor Force Exit Robustness Checks for Time and Geographic Heterogeneity

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Partial Exit Hazard				Full Exit Hazard			
Log Minimum Wage	-0.0595**	-0.0282	0.0320	-0.0463	-0.0961**	-0.0486	-0.0510	-0.0974**
	(0.0263)	(0.0223)	(0.0380)	(0.0306)	(0.0407)	(0.0298)	(0.0471)	(0.0485)
Observations	14,500	14,500	14,500	14,500	17,000	17,000	17,000	17,000
State, Year, and Age Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Covariates	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
State Linear Time Trends	Yes	No	No	Yes	Yes	No	No	Yes
Age-by-Time Fixed Effects	No	Yes	No	Yes	No	Yes	No	Yes
Region-by-Time Fixed Effects	No	No	Yes	Yes	No	No	Yes	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY21-CED001-B0002.
Note: The regressions are the same as columns (2) and (4) of Table 4, except with additional fixed effects as indicated in the table. Region corresponds to the four Census Regions.

Table 8: Claiming Hazard Robustness Checks for Time and Geographic Heterogeneity

	(1)	(2)	(3)	(4)
	Claiming Hazard			
Log Minimum Wage	-0.0396***	-0.0441**	-0.0462**	-0.0475**
	(0.0142)	(0.0176)	(0.0194)	(0.024)
Observations	68,500	68,500	68,500	68,500
State, Time, Age Fixed Effects	Yes	Yes	Yes	Yes
Covariates	Yes	Yes	Yes	Yes
State Linear Time Trends	Yes	No	No	Yes
Age-by-Time Fixed Effects	No	Yes	No	Yes
Region-by-Time Fixed Effects	No	No	Yes	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0005.
Note: The regressions are the same as column (2) of Table 5, except with additional fixed effects as indicated in the table. Region corresponds to the four Census Regions.

Table 9: Claiming Result Mechanism: Interaction Between Minimum Wages and the Social Security Earnings Test

	(1) Claiming Hazard	(2) Employment
Log Minimum Wage	-0.0633***	0.0870
	(0.0215)	(0.0710)
Log Minimum Wage * post-FRA	-0.0277***	0.0107
	(0.0037)	(0.0126)
Log Minimum Wage * post-2000	0.0105	0.161
	(0.0225)	(0.1080)
Log Minimum Wage * post-FRA * post-2000	0.129***	-0.129
	(0.0277)	(0.0909)
Observations	68,500	27,000
State, Time, and Age Fixed Effects	Yes	Yes
Covariates	Yes	Yes
Individual Fixed Effects	No	Yes
Log Minimum Wage Marginal Effect:		
pre-FRA, pre-2000	-0.0633***	0.0870
post-FRA, pre-2000	-0.0910***	0.0977
pre-FRA, post-2000	-0.0528***	0.2480***
post-FRA, post-2000	0.0485**	0.1300**

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0005.
Note: The regressions are the same as column (2) of Table 5 and column (3) of Table 2, except with controls for post-FRA, post-2000, their interaction, and the interaction of all three with the log minimum wage variable. "post-FRA" includes the month/year in which the FRA was reached. "post-2000" includes the year 2000. Marginal effects at the bottom of the table are based on linear combinations of the coefficients shown above.

Table 10: Younger Sample and Own-Wage Elasticity

	(1) Ages 62-70	(2) Ages 40-54
Panel A: Employment		
Log minimum wage	0.151***	-0.0468*
	(0.0429)	(0.0254)
Dependent variable mean	0.679	0.851
Observations	27,000	80,500
State, Year, and Age Fixed Effects	Yes	Yes
Covariates	Yes	Yes
Individual Fixed Effects	Yes	Yes
Panel B: Log Earnings		
Log minimum wage	0.482***	0.305**
	(0.167)	(0.121)
Observations	18,500	68,500
State, Year, and Age Fixed Effects	Yes	Yes
Covariates	Yes	Yes
Individual Fixed Effects	Yes	Yes
Panel C: Elasticities		
% change in employment:	0.222	-0.055
% change in wage:	0.213	0.381
Own wage elasticity:	**1.042**	**-0.144**

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003, CBDRB-FY21-CED001-B0004.

Note: Panel A shows regression results where the dependent variable is an indicator for positive IRS earnings in the DER. Column (1) of Panel A is a reproduction of the results in Table 2 for the regression results and Table 1 for the dependent variable mean. Column (2) reports results from the same specifications and sample, except for individuals who were observed during ages 40-54 during the SIPP instead of 62-70. Panel B reports regression results for the same samples and specifications, except with log earnings as the dependent variable. Panel C report elasticities based on Panel A and Panel B. The employment elasticity is the Panel A coefficient divided by the sample mean employment rate reported in Panel A. The wage elasticity is computed using equation (2) in Cengiz et al. (2019). Own-wage elasticity is equal to the % change in employment divided by the % change in wage.

Appendix

A First-stage test: short-run labor market outcomes

As a validation of our sample design and construction, we use the monthly SIPP information to estimate the short-term effect of minimum wages on hourly wages, earnings, employment, and hours worked. This is primarily intended to be a first-stage test, ensuring that we find the expected short-run effects of minimum wage increases on hourly wages in particular, to validate the sample selection. We restrict our low-wage main sample and higher-wage alternative sample described in Section 3.2 of the main text the subset of individuals who have a balanced sample of non-missing SIPP observations for each given outcome of interest over a 19 month time-span around their first minimum wage increase while in the SIPP. The time span ranges from six months before the minimum wage increase to 12 months after. If the individual did not experience a minimum wage increase while in the SIPP, then we use their first 19 months with non-missing information. Using a balanced sample prevents the results from being complicated by changing composition of the sample before and after the minimum wage increase, which is always a concern in difference-in-differences style analyses, but especially so in this setting given that minimum wage increases may influence retirement behavior and thus exit from the labor market altogether.

For these short-run outcomes, we follow recent minimum wage work by Clemens and Wither (2019) and Cengiz et al. (2019) and use an event-based difference-in-differences analysis. We analyze outcomes in a 19 month window ranging from six months before the minimum wage change event to 12 months after the event and regress the outcome of interest on an indicator that corresponds to being in a treated state and the post-treatment months.

The primary regression model is:

$$y_{iast} = \beta\Gamma_{st} + X_{st}\psi + \tau_a + \alpha_i + \delta_t + u_{iast}, \tag{A.1}$$

where y_{iast} is equal to one of log(wage), an employment indicator, log(earnings), or log(hours).

Γ_{st} is an indicator variable equal to one if the minimum wage was increased in state s during or prior to time period t. X_{st} includes covariates for the state-year unemployment rate and state-year population. τ_a, α_i, and δ_t are age, individual, and time-period fixed effects, respectively.[1] We also estimate specifications with state-specific linear time trends and Census Division-by-period fixed effects.

We perform this analysis for the main low-wage sample described in Section 3.2 of the main text as well as the alternative higher-wage sample. Panel A of Table A1 shows results for the main sample. Wages increase by approximately 9-11% after a minimum wage increase, depending on the specification. This suggests that our sample of low-wage workers are in fact working near the minimum wage and are affected by increases in the minimum wage. Employment and hours do not respond to increases in the minimum wage. Consistent with a wage increase and no evidence of a reduction in employment or hours, we find an increase in earnings almost as large as the increase in wages.

While these results are encouraging for the construction of our sample, they could be driven by unobserved confounders that influence the lower part of the wage distribution in general. Panel B of Table A1 repeats the analysis, except with the alternative sample of individuals whose average wage is \$5-\$10 above the minimum wage. All the coefficients are close to zero relative to Panel A and lack statistical significance. This suggests that results for the main sample in Panel A are not driven by unobserved confounders that affect the broader low-wage part of the wage distribution and that there are no spillover effects that exist as far up the wage distribution as \$5-\$10 above the minimum wage.

[1] The other covariates shown in Table 1 of the main text drop out because of the individual fixed effects.

Table A1: Short-Run Labor Market Outcomes

	(1) Employed	(2) Employed	(3) Log Wage	(4) Log Wage	(5) Log Earnings	(6) Log Earnings	(7) Log Hours	(8) Log Hours
Panel A: Main Sample								
Treat*Post	0.00479	0.0162	0.112***	0.0875*	0.0934***	0.0717**	0.000961	-0.00203
	(0.0254)	(0.0244)	(0.0329)	(0.0459)	(0.0218)	(0.0342)	(0.0247)	(0.0261)
Observations	96,000	96,000	31,000	31,000	34,000	34,000	34,000	34,000
Panel B: Higher-Wage Sample								
Treat*Post	-0.00518	-0.00903	0.00381	-0.0170	0.00647	-0.0208	0.00562	-0.00285
	(0.0190)	(0.0172)	(0.0156)	(0.0138)	(0.0165)	(0.0175)	(0.00852)	(0.0104)
Observations	92,000	92,000	46,000	46,000	49,500	49,500	48,500	48,500
Individual, Time, Age Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Covariates	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Additional Controls	No	Yes	No	Yes	No	Yes	No	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.

Note: Each column reports the OLS coefficient from a balanced person-month panel model for the given outcome over a 19-month window from six months before to 18 months after a minimum wage change. Standard errors are clustered at the state level and shown in parentheses. The Treat variable is equal to one for individuals who were exposed to a minimum wage change while in the SIPP. The Post variable is equal to one in the period of the minimum wage change and periods after the change. Time fixed effects are at the year-month level. Age fixed effects are in years. Covariates include the log of the state-year unemployment rate and the log of the state-year population. Additional controls include state-specific time trends and Census Division-by-period fixed effects.

B Robustness of claiming results to alternative estimation methods

Table B1: Robustness of Hazard Models for OASDI Retirement Benefit Claiming to Alternative Estimation Methods

	(1)	(2)	(3)	(4)	(5)	(6)
	Probit Hazard		Logit Hazard		CLogLog Hazard	
Panel A: Main Sample						
Log Minimum Wage	-0.0232**	-0.0176**	-0.0198**	-0.0148**	-0.0168**	-0.0125**
	(0.0095)	(0.0070)	(0.0080)	(0.0060)	(0.0072)	(0.0058)
Month of FRA		0.0633***		0.0459***		0.0428***
		(0.0048)		(0.00394)		(0.0037)
Observations	68,500	68,500	68,500	68,500	68,500	68,500
Panel B: Higher-Wage Sample						
Log Minimum Wage	0.0026	0.0056	0.0021	0.0052	0.0014	0.0042
	(0.0056)	(0.0047)	(0.0043)	(0.0038)	(0.0042)	(0.0035)
Month of FRA		0.0512***		0.0360***		0.0312***
		(0.0021)		(0.0018)		(0.0016)
Observations	95,500	95,500	95,500	95,500	95,500	95,500
State, Time, Age Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Covariates	No	Yes	No	Yes	No	Yes

Source: U.S. Census Bureau Gold Standard File, CBDRB-FY20-CED001-B0003.
Note: This table shows robustness of the results in Table 5 of the main text to the use of different estimation methods instead of OLS. See the notes in Table 5 for more details about the sample and specifications.