Unemployment Insurance Expansions and Family Structure: Evidence from AFDC-UP in the 1970s and 1980s

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Abstract

This paper evaluates the short-run family structure-stabilizing effects of public assistance for poor two-parent families. 1961 amendments to the Social Security Act gave states the option to extend the Aid to Families with Dependent Children program to provide a segment covering the families of an unemployed parent (AFDC-UP), and access to benefits remained contingent on state of residence for the next three decades. AFDC-UP functioned similarly to the U.S. unemployment insurance (UI) system but also largely targeted the families of individuals who were UI-ineligible. Leveraging Current Population Survey Annual Social and Economic Supplement data as early as 1968 and through 1988, the difference in differences research design compares measures of program participation and family structure between recently unemployed men and those not experiencing unemployment, in AFDC-UP versus non-AFDC-UP states, and by tercile of previous earnings. I find that AFDC-UP provided a large degree of protection, preventing upwards of 50 percent of increases in divorce, separation, or spousal absence and declines in cohabitation that would have otherwise occurred. These effects are concentrated primarily among poor families of long-term unemployed men. The findings in this paper indirectly suggest that the welfare gains of extending UI benefits to ineligible workers and their families would be large, highlight the historical importance of complementary safety net programs in protecting against unemployment, and contribute to our understanding of the increasing association between unemployment and two-parent family dissolution since the 1960s.

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U.S. unemployment insurance (UI) is the primary system that provides income support to eligible unemployed individuals and their families. Yet, both historically and today, many unemployed workers do not qualify for benefits, with ineligibility disproportionately common among the more economically disadvantaged.[1] Intended to help address such deficiencies — which came to the fore during the Covid-19 Pandemic — the 2020 temporary Pandemic Unemployment Assistance (PUA) program constituted the largest federal expansion to eligibility in the history of the program, with PUA cases quickly accounting for over 40 percent of total claims (Ganong et al., 2022). The successes of the PUA and other pandemic-related temporary UI expansions have renewed calls for converting UI to a fully federal system with less stringent eligibility criteria (e.g., Dube (2021)). While there is much evidence on how benefit *levels* impact myriad outcomes for those who receive them, little is known about the ramifications of extending UI to a group of ineligible (and typically poorer) households.

This paper investigates the benefits of expanding UI eligibility by turning to an overlooked program that functioned similarly for three and a half decades, but which also largely targeted families of the UI-ineligible unemployed. In 1961 Congress authorized states to pay Aid to Families with Dependent Children (AFDC) benefits — the primary 20th century cash public assistance program for poor single parent families — to two-parent families if the "main breadwinner" was unemployed. AFDC-Unemployed Parent (AFDC-UP) eligibility required the principle household earner to be currently unemployed, have a minimum work and earnings history, and be actively seeking employment, which are all central tenets of UI eligibility. At the same time, surveys of early recipients (Bureau of Family Services, 1962) and the results in this paper indicate that AFDC-UP largely provided benefits to the

[1] The Fraction of Insured Unemployment (FUI) – or average fraction of unemployed workers who are weekly insured for UI – declined from roughly 49 percent in 1960 (Carter et al., 2006, series Ba485; Bf485) to 28 percent in 2019 (U.S. Department of Labor, 2019; Edwards and Smith, 2020). Even as statutory coverage expanded markedly over this period to nearly 100 percent of the workforce, stricter requirements on earnings and work history has meant that a lower fraction of those working in covered employment are eligible for benefits (Blaustein et al., 1993; Baicker et al., 1998). In addition to the positive correlation between UI eligibility, recipiency, and income, recent evidence also points to lower UI eligibility and recipiency among black individuals (Kuka and Stuart, 2021; Skandalis et al., 2022).

families of workers who were UI-ineligible. Envisioned as a program to encourage stability among poor two-parent families (Lansdale, 1967), the key eligibility differences between the two programs are that AFDC-UP was also contingent on having minor children and was means tested.

The outcomes I focus on relate to two-parent family structure. Economic models predict that AFDC-UP should reduce incentives for couples with children to separate during unemployment spells (e.g., Becker, 1981).[2] In light of the well-established empirical association between unemployment and divorce or separation (Charles and Stephens, 2004; Doiron and Mendolia, 2012; Eliason, 2012; Lindo et al., 2022), this paper conceptualizes the benefits of AFDC-UP as helping to mitigate separation that would have occurred absent the program. Because evidence of such responses would implicitly represent an increase in household welfare during unemployment under the choice to remain in a union, the results in this paper also serve as indirect tests of the degree to which expanding UI eligibility would increase household welfare.

The empirical analysis is based primarily on pooled Current Population Survey Annual Social and Economic Supplement (CPS-ASEC) repeated cross-sections between 1977-1988. The CPS-ASEC contains measures of current and prior-year unemployment and unemployment duration, prior year earnings and program participation, and current measures of family structure. Given that recent or current unemployment spells and recent earnings of household heads are both necessary for assigning AFDC-UP eligibility, the cross-sectional nature of the CPS requires focusing on family structure responses among men.[3] The primary

[2]Specifically, AFDC-UP removed the incentive for children and mothers to respond to unemployment by pursuing benefits under the traditional AFDC program, eligibility for which was contingent on being both poor and residing in a single parent household. Through helping to mitigate earnings losses during unemployment (Jacobson et al., 1993; Ganong and Noel, 2019), AFDC-UP may also have reduced stressors shown to correlate both with job loss and with separation (Kuhn et al., 2009; Deb et al., 2011; Black et al., 2015). The existence of such effects would be consistent with a growing literature documenting the positive effects of UI for consumption smoothing (Gruber and Madrian, 1997; East and Kuka, 2015), well-being (Eliason and Storrie, 2009; Kuhn et al., 2009; Sullivan and Von Wachter, 2009; Deb et al., 2011; Classen and Dunn, 2012; Black et al., 2015), and, in a particularly closely related paper, for preventing separation (Lindo et al., 2022).

[3]Specifically, I cannot observe former spouses of currently or recently unemployed men who have separated

sample begins in 1977 because most states are not identifiable in the CPS in earlier years and ends in 1988 because the Family Support Act (FSA) of 1988 mandated all states adopt the program. Additional analyses leverage a restricted sample of states as early as 1968.

AFDC-UP's optionality before 1988 motivates the difference in differences research design that compares measures of program participation and family structure between recently or currently unemployed versus employed men (first difference) and in AFDC-UP versus non-AFDC-UP states (second difference). The analysis effectively estimates short-run responsiveness to unemployment (within roughly 15 months). Among ever-married men, I test for differences in probabilities of being currently married; divorced, separated, or reporting spousal absence (henceforth together referred to as "separated"); living with one's own minor child(ren); and living with one's own minor child(ren) and their mother (henceforth referred to as "cohabitation").[4] Due to difficulty in ascertaining rules over income and asset eligibility tests (which vary both across states and over time), I take a straightforward approach and conduct analyses separately by tercile of the previous year's income, expecting both welfare receipt and family structure to respond primarily in the lowest income tercile.[5]

The primary empirical specification produces two interpretable coefficients per outcome: the association of unemployment with family structure in control states and the protectiveness of AFDC-UP against unemployment-associated family dissolution. I derive identification conditions of Average Treatment Effects on the Treated (ATTs) that require equality between average untreated potential outcomes across AFDC-UP and non-AFDC-UP states, and then use comparisons in the middle and upper terciles (where AFDC-UP is largely un-

(and are therefore recorded as living in a separate household), although short short-term effects of AFDC-UP on separation among men must translate into similar effects for women (Moffitt, 1992). While the share of two parent families headed by women increased over the period, roughly 90 percent of AFDC-UP households between 1977-1988 were headed by men (as determined in the CPS-ASEC).

[4]Cohabitation is of central interest because, as emphasized by Winkler (1995) and Moffitt et al. (1998), the eligibility conditions for AFDC-UP relating to family structure considered whether both natural or adoptive parents of the child were in the home, not whether the parents were married.

[5]Because the interest lies in comparing men who have similar full time earnings, I use the Least Absolute Shrinkage and Selection Operator (LASSO) to predict "potential" earnings from men who worked 52 weeks in the previous year (described further in Section II).

available) as tests of these assumptions. I also investigate a different empirical strategy that leverages a Doubly Robust estimator (Mercatanti and Li, 2014) to re-weight men in control groups to resemble their treated counterparts.

The results confirm that unemployment is associated with a large, heightened risk of separation in the short-term, with the preferred estimate in the lowest income tercile of 12 percentage points constituting a 57 percent increase relative to rates among lower income employed men in non-AFDC-UP states. The results also provide novel evidence that such associations are present across the income distribution, are present when examining whether men cohabit or live with children, and are increasing in unemployment duration.

Among lower income unemployed men and their families, state AFDC-UP variation strongly predicts reported welfare receipt. AFDC-UP's potential to encourage two-parent families partly stems from the fact that recipients gained categorical eligibility for Medicaid and food stamps. This not only expanded the set of families statutorily eligibility for these programs, but also likely reduced the type of non-pecuniary costs shown to influence participation in transfer programs across different contexts (Rossin-Slater, 2013; Armour, 2018; Finkelstein and Notowidigdo, 2019; Deshpande and Li, 2019; Herd and Moynihan, 2019). Further results show that Medicaid undoes roughly 80 percent of the health insurance coverage loss associated with long-term unemployment, that the average cash value of food stamps is in excess of 30 percent of the average AFDC-UP benefit, and that the total value of benefits inclusive of program interactions replaced upwards of 60 percent of previous wages. The results also show that most AFDC-UP recipients did not also receive UI benefits and that, while there were some families who received both, there is little evidence of disproportionate UI access among the non-welfare receiving poor unemployed.

Turning to family structure, the results routinely indicate that AFDC-UP meaningfully contributed to two-parent family stability, with preferred estimates of protectiveness of around 50 percent of unemployment-associated separation and all of unemployment-associated changes in cohabitation. AFDC-UP disproportionately provided benefits to the families of the long-term unemployed (greater than 6 months) — families in a particularly high marginal utility state of the world (Ganong and Noel, 2019) — and additional results show that both recipiency and the protectiveness of AFDC-UP are increasing in unemployment duration, reflecting the link between need, eligibility, and the family structure-stabilizing effects of the program. The effects of AFDC-UP are present only in the lowest tercile, are driven entirely by couples with children, and are generally robust to a variety of additional sample restrictions, variable definitions, and specification choices. Using the alternative Doubly Robust empirical strategy leads to similar estimates for AFDC-UPs protectiveness of cohabitation and *higher* estimates for separation.

Relative to the literature on UI's income effects, this paper's primary contribution is the focus on extensive-margin eligibility under a UI-type program. Research in this area typically leverages variation in intensive margin generosity by exploiting state-level measures of UI average or maximum benefits (e.g., see Gruber and Madrian (1997); Hsu et al. (2018); Lindo et al. (2022)). As noted by Moffitt (2003) in the context of AFDC, however, variation in benefit levels across states that all operate a program makes extrapolations to zero inherently speculative. Furthermore, higher UI benefits lead to higher recipiency (Blank and Card, 1991; Anderson and Meyer, 1997) — conflating potentially distinct responses along each margin — and UI maximums may also be less reflective of benefits received for poorer workers. My geographic-based research design identifies responses only at the margin of eligibility, while the targeting of the program also allows for new evidence on the ramifications of UI-type benefits for a group less able to privately insure against adverse economic shocks.[6] The effect sizes are close to what would be implied by Lindo et al. (2022) were average UI benefits equal to average UI maximums, which is consistent both with high replacement

[6]The targeting of disadvantaged families under AFDC-UP also resembles the PUA, which was designed to assist those ineligible for regular UI — primarily including the self-employed, those seeking part-time employment, and individuals lacking sufficient work history — and disproportionately aided lower income and marginally attached workers and their families (Ganong et al., 2022). Other differences between AFDC-UP and UI concern benefit determination, which in general was not a function of previous earnings for AFDC-UP (conditional on satisfying the means test) and contained functionally no limitations on length of recipiency in most states (U.S. ASPE, 1996).

rates when accounting for categorical eligibility and with the targeting of high marginal utility households by AFDC-UP.

This paper also contributes to the large and often inconclusive literature on AFDC and family structure (for reviews see Groeneveld et al., 1983; Moffitt, 1992, 1998, 2003). It is particularly related to the much smaller strand studying AFDC-UP (Schram and Wiseman, 1988; Schultz, 1994; Winkler, 1995; Hoynes, 1997), which typically leverages cross-sectional state-level designs and finds zero or *positive* effects of AFDC-UP on single motherhood. This paper helps to clarify these findings by emphasizing that the incentives provided by AFDC-UP affected family structure for a particular minority of families — those who are poor and headed by men currently experiencing long-term unemployment — and by documenting quite large effects only among this group. Apart from the greater need of the longer-term unemployed and their families, the magnitudes are also consistent with theoretical predictions that family structure should respond more to more permanent changes to welfare availability (Rosenzweig, 1999; Moehling, 2007; Kearney, 2004) and with AFDC-UP providing "double protection" by counteracting incentives under the traditional AFDC program for couples to separate.

This paper is also related to a growing literature on the patchwork safety net. In one regard, I emphasize the high degree of *complementary* between AFDC-UP and UI in providing protection against unemployment — particularly for the population of poor, two-parent families — which closely mirrors the low degree of substitutability between UI and public assistance today (Leung and O'leary, 2020).[7] On the other hand, I also show the importance of accounting for program interactions when inferring the overall transfer value of AFDC-UP, and for public assistance programs more generally. These results indicate that ignoring these interactions may produce misleading estimates of how family structure responds to a

[7]There is evidence of substitutability across other types of transfer programs, including between AFDC and Supplemental Security (SSI) (Schmidt and Sevak, 2004; Goodman-Bacon and Schmidt, 2020), between Social Security and Social Security Disability Insurance (SSDI) (Duggan et al., 2007), and between Social Security and public assistance (Coe and Wu, 2014; Fetter and Pesner, 2021).

particular transfer program.

Finally, this paper helps to inform long-run trends in separation and cohabitation and how such trends have differed by the employment status of men. Figure 1, Panel A shows marked increases in the share of ever-married men currently separated beginning in the 1970s, a long-run increasing association between unemployment and separation since the early 20th century, and a relatively steeper (age-adjusted) increase among those currently unemployed since the 1960s.[8] Panel B shows reverse trends in rates of cohabitation and a negative association between unemployment and cohabitation, but also shows no recent increase in the magnitude of the association before the 1990s.[9] I show in Section VI that, while trends were very similar among employed men in AFDC-UP and non-AFDC-UP states, differences by unemployment status begin in the 1960s (after the implementation of AFDC-UP), are reduced in the 1990s (after AFDC-UP ceased to exist), and are larger for cohabitation than separation. Taken together, the econometric results and patterns in separation and cohabitation by employment and by AFDC-UP status strongly suggest that AFDC-UP helped mitigate even larger increases in aggregate associations between unemployment and two-parent family dissolution than observed since the 1960s.

[8] Data for Figure 1 come from 1 percent decennial census samples (Ruggles et al., 2023) and are restricted to ever-married men ages 24-58 who are in the labor force, to match the primary analysis sample described in Section II.

[9] Corresponding increases in nonmarital births and rates of single motherhood (Ventura and Bachrach, 2000; Curtin and Martinez, 2014), and the general decoupling of marriage and motherhood, are well documented (Bailey et al., 2013). Various explanations include the passage of unilateral divorce laws (Wolfers, 2006), decreased labor market opportunities for men and increased labor market opportunities for women (Wilson, 1996; Weiss and Willis, 1997; Blau et al., 2000; Black et al., 2005; Wilson, 2012; Autor et al., 2019; Shenhav, 2021), the availability of contraceptive technology in the 1960s and the legalization of abortion in the early 1970s (Akerlof et al., 1996; Goldin and Katz, 2002; Guldi, 2008; Bailey, 2010; Bailey et al., 2013), increases in male incarceration (Charles and Luoh, 2010), and other shifts in the "price" of marriage (Buckles et al., 2011) or marital dissolution (Goodman-Bacon and Cunningham, 2019). Kearney and Wilson (2018) provide evidence that increased labor market opportunities for men affect fertility, but not marriage. The increasingly stark relationship between unemployment and dissolution is a feature of U.S. families documented by economists at least as early as Ross et al. (1975) and Becker et al. (1977).

I Background and expected effects

I.A Brief history of AFDC-UP

The Social Security Act (SSA) of 1935 established the Aid to Dependent Children (ADC) program, defining an eligible dependent child as one "deprived of parental support or care by reason of the death, continued absence from the home, or physical or mental incapacity of a parent." Formed in response to the 1958 recession, the Advisory Council on Public Assistance noted that such a definition "penalized" poor children living with two able-bodied parents (U.S. Advisory Council on Public Assistance, 1960, p. 15) and recommended the adoption of a new program for children of unemployed fathers, which Congress passed in 1961 (ADC-Unemployed Father). A second parent was allowed in benefit determination in 1962 and the name of the program was changed to Aid to Families with Dependent Children (U.S. DHHS, 1998) to mark a commitment to the family unit (Wexler and Engel, 1999). AFDC-UP remained optional until 1988, when the FSA mandated all remaining states adopt the program by 1990, and all AFDC programs were subsequently replaced with Temporary Assistance for Needy Families (TANF) in 1996. Henceforth I refer to the primary component of AFDC for single parent families as "AFDC-basic" to distinguish it from AFDC-UP.

Figure 2, Panel A maps the states that ever adopted the program before the FSA, by era of first-adoption. Of the 31 contiguous states that would ever adopt the program through 1988, 13 states adopted in the first year AFDC-UP was available (1961), another 13 states adopted between 1962-1969, 4 states adopted between 1970-1976, and just 1 state adopted for the first time after 1975 (South Carolina, in 1985).[10]

Figure 3 shows that the 1960s were characterized by a gradual and largely monotonic decline in unemployment, even as AFDC-UP caseloads normalized per 1,000 cohabiting couples rose quickly in the early 1960s due to adoption, and then remained relatively flat

[10]I partition first-time adoption years into these groups for the sake of interpretability. I use the term first-time adoption because many states would drop and restart the program, a feature I discuss below in Section II and further in Appendix B.

until 1969.[11] In contrast, the 1970s and 1980s were characterized by multiple recessions and lingering elevated unemployment, and caseloads track the unemployment rate quite closely. Alternatively, Figure A.1 shows that caseloads spike during each recession (classified according to the National Bureau of Economic Research) beginning with the recession of 1969, achieving a pre-FSA maximum of over 300,000 cases and 1.3 million recipients in March of 1984. In 1980, the last year for which I have found spending information, total expenditure was roughly $5 billion in 2019 dollars (Bureau of Public Assistance, 1980).

I.B AFDC-UP as a complement to UI

The name of AFDC-UP suggests its use as a form of unemployment insurance, while the time series' in Figure 3 and the similarity of requirements to those of UI support this view. Under the original 1961 act, states were required to deny assistance if an unemployed parent refused to accept work without "good cause" (U.S. DHHS, 1998). 1971 federal guidelines specified that a parent must work fewer than 100 hours in a month to be considered eligible.[12] Furthermore, the "principle earner" must have worked for 6 of any 13-quarter period ending within a year before application or have received or been eligible for UI within the year before, with quarters defined as those with at least $50 of earnings (U.S. ASPE, 1996 p. 396).[13] Workers similarly qualify for UI when they have lost a job through no fault of their own, are available to work and are actively seeking work, and obtained above a specified level of earnings in covered employment during the "reference period" (typically 4 of the 5 quarters preceding the claim). UI benefits statutorily replace 50 percent of covered earnings in most

[11] The series' are restricted to households in which there is only one such couple. I measure cohabiting couples in the CPS-ASEC beginning in 1968 — the first year that the age of the youngest child in the household was collected — and also present a series using decennial censuses beginning in 1960 and linear interpolations in intercensal years.

[12] This constraint was relaxed if the work was considered temporary or intermittent, in which case they must have worked less than 100 hours in the two preceding months and "expect" to also work fewer than 100 hours in the next month (U.S. ASPE, 1996 p. 395).

[13] At State option, elementary or secondary school attendance, vocational or technical training, or participation in a job training program may have substituted for up to 4 of the 6 required quarters of work.

states subject to minimum and maximum monthly amounts, which leads to significantly lower average replacement rates (East and Kuka, 2015). The majority of states have a maximum duration of 26 weeks.[14]

UI has remained unavailable to a large fraction of unemployed worker since its creation as part of the SSA. This was by design; President Roosevelt's Committee on Economic Security did not assume UI could "adequately handle the entire unemployment problem", instead arguing it was "but a complementary part of an adequate program for protection against the hazards of unemployment and that public-work programs and the 'modernization' of public assistance programs were also essential" (Cohen, 1960, pp. 311-312). Figure 4, Panel A shows that the unemployment rate sat well above the fraction of the labor force weekly insured under UI (defined as the average number weekly insured for UI benefits as a share of the labor force) between 1948 and 2000. Panel B shows the Fraction of Insured Unemployment (FIU; Blank and Card (1991)), which is the average weekly number of UI-insured workers as a share of the total number of unemployed workers, or ratio of lines in Panel A. Even as statutory coverage of workers increased markedly in the 1970s — from less than 60 percent to close to 90 percent (Price, 1985) — the FIU continued a broad trend of decline.[15]

AFDC-UP largely targeted a population ineligible for UI. Of the original recipient cohorts in the early 1960s, roughly 70 percent had not received UI nor had pending UI applications (Bureau of Family Services, 1962, p. 8). Of the remaining, 19 percent had already claimed

[14] See Blaustein et al. (1993) pp. 278-282 for a useful discussion of changes to state qualifications regarding prior work. Of particular relevance to this paper, fourteen states provided additional allowances for dependent children, although the amount was generally $30 or less per week (U.S. Department of Health and Human Services , DHHS). Since 1970, there have also been automatic federal extensions of UI for states when the covered unemployed rate exceed statutory thresholds.

[15] There are three distinct eras. A precipitous decline occurred in the early 1960s — widely believed to be the result of changing demographics of the labor force (Blaustein et al., 1993) — and then again in the early 1980s. This latter decline resulted in large part from statutory changes leading to increased qualifying requirements, stricter disqualification rules, and more restrictive extended UI benefits (Blaustein et al., 1993), while others have argued the 1980s FIU decline is entirely due to declining take up (Blank and Card, 1991).

and 12 percent were currently claiming.[16] Consistent with the idea that both programs were complementary, a report on the 1958 recession found that unskilled laborers were much more likely to be unemployed relative to skilled and semi-skilled workers, had longer unemployment spells, were less likely to receive UI, and were less likely to receive UI for the entire duration of their unemployment conditional on receiving it (Cohen et al., 1960, p. 31). In more modern settings, research also suggests that UI and public assistance are largely complementary programs that target distinct populations (Leung and O'leary, 2020).

One primary difference between UI and AFDC-UP is the lack of a maximum duration under the latter. In 1973, for example, in 40 percent of recipient families the father had been currently unemployed for longer than 6 months, and of these another 50 percent had been unemployed for over one year (U.S. NCSS, 1975b p. 34).[17] Indeed, as argued in the introduction, the apparent disproportionate targeting of the long-term unemployed, and the high insurance value that access to AFDC-UP benefits in such a state of the world provides, is a central explanation for why AFDC-UP is expected to provide meaningful protection against unemployment-associated family dissolution.

In sum, AFDC-UP worked similarly to UI and appears to have largely affected a UI-ineligible population. Nevertheless, it is important to note that there was some degree of substitutability between the programs. As already noted, roughly 30 percent of the original cohorts had received or were awaiting UI benefits. Furthermore, the rise in AFDC-UP caseloads during the early 1980s recessions (Figure 3) appears to be beyond what would have occurred absent the tighter UI eligibility conditions imposed around that time.[18] As

[16]It is impossible to know whether the original recipients were ineligible or were eligible but did not apply, as well as whether ineligibility was due to working in uncovered jobs or due to insufficient history. Given the near ubiquity of UI coverage in later years and contemporaneous declining rates of recipiency, it is probably the case that most AFDC-UP recipients worked in covered employment but did not satisfy the earnings or work requirements for eligibility.

[17]States experimented in the 1990s with time limits on eligibility and, as of February 1996, 12 States had such restrictions (U.S. ASPE, 1996 p. 395). None of these states are "treated" under the primary definition. While the average AFDC-UP duration is longer than the average UI duration, it is also much shorter than the basic program. In 1983, for example, the majority of families enrolled in AFDC-basic were in the midst of spells that lasted over 8 years (U.S. ASPE, 1985).

[18]The 1979-1985 percent change in the overall unemployment rate was 23 percent as compared to the percent

such, the rise in caseloads may reflect program substitution to some extent, and suggests the degree to which the programs acted as substitutes or complements is best determined by empirical evidence.

I.C Expected effects of AFDC-UP on family structure

The idea that rising welfare participation drove historical increases in rates of single motherhood has a long history in public discourse, exemplified through a 1960 report on ADC and "illegitimacy" prepared by the Bureau of Public Assistance (Bureau of Public Assistance, 1960, p. 1).[19] Upon signing AFDC-UP into law, President Kennedy reaffirmed this position in a 1961 address to Congress, stating "too many fathers, unable to support their families, have resorted to real or pretended desertion to qualify their children for help." (Kennedy, 1977).

Neoclassical economic models generally support these assertions for the AFDC-basic program (Becker, 1981; Moffitt, 1998) and are also largely unambiguous in their prediction that AFDC-UP should provide protection against unemployment-associated separation. In the context of household bargaining models, AFDC-UP provides protectiveness by effectively increasing men's relative wages during unemployment (Lundberg and Pollak, 1996; Willis, 1999), which is consistent with evidence that declines in relative wages increase separation in a variety of settings (Wilson, 2012; Bertrand et al., 2015; Autor et al., 2019; Shenhav, 2021).[20] Central to the value of AFDC-UP is the insurance it provides against unemployment consumption risk, particularly for households experiencing long-term unemployment. These effects may be quite large; for example, decreased expenditure risk after the introduction

change in caseloads per thousand cohabiting couple, which was 129 percent (Figure 3).

[19] Murray (1984) is a prominent advocate of this idea in the academic sphere and wrote particularly about black families, although his research has been subject to stark and enduring critique (e.g., see McLanahan et al. (1985) and the research summarized therein). In part because AFDC-UP did not disproportionately aid black families as in the AFDC-basic program (Hoynes, 1996 Table 1) and in part due to sample size limitations, this paper emphasizes trends by income, rather than race.

[20] Similar to Autor et al. (2019), I cannot distinguish between the potentially distinct effects of AFDC-UP protecting against a fall in absolute family income from a fall in the relative wages of husbands to wives.

of elderly public health insurance (Medicare) four years after AFDC-UP may alone have covered 40 percent of its cost (Finkelstein and McKnight, 2008).

The targeting of poor two-parent families under AFDC-UP also provided "double" protection against separation in the sense that, absent the program, one outside option would be for children and mothers to go on the AFDC-basic program if they formed a single parent household. At the same time, we would still expect the benefits of AFDC-UP to prevent separation in a state of the world without the basic program. This is because, apart from income loss, job loss is associated with various indicators of stress that are also associated with separation, including increased risky behavior (Black et al., 2015; Deb et al., 2011) and increased prevalence of mental health conditions (Kuhn et al., 2009).

Through changing the definition of a dependent child "deserving" of benefits, AFDC-UP constituted one of the largest changes to extensive margin transfer income eligibility since the creation of ADC.[21] Relative to the introduction of ADC, the implementation of AFDC-UP is useful in part because it occurred in a more data-rich period, which allows likely-impacted families to be classified and observed.[22] The relative permanence in the geographic variation of AFDC-UP is another useful feature of the policy environment, given that behavioral responses to public assistance are likely driven by large perceived changes in benefits (Rosenzweig, 1999; Moehling, 2007; Kearney, 2004) and family structure takes time to respond to statutory changes in welfare incentives (Moffitt, 1998).[23] This stands in contrast to research on the AFDC-basic program, which has largely focused on cross-sectional

[21] Goodman-Bacon and Cunningham (2019) show that access to legal services in the early 1960s effectively increased AFDC recipiency (absent statutory changes in eligibility) and led to increased rates of divorce and nonmarital births.

[22] By 1948 all states had an AFDC program save for NV, well before intercensal microdata containing information on family structure is available.

[23] Prior to AFDC-UP, two parent impoverished families could only receive assistance from state and locally financed General Assistance (GA) programs in some states and specific cases. GA excluded any families that contained an "employable person" in 15 states in 1959 and 18 states in 1969 (Bureau of Public Assistance, 1959, 1969). In the remaining states, eligibility was often left up to local discretion and was "not available in most communities to those unemployed who exhaust their unemployment insurance benefits" (Cohen, 1960 p. 311). Benefits were also sparse in relative terms (Bureau of Public Assistance, 1960) and "grossly inadequate" for those who received them.

or year to year changes in benefit levels.[24]

The few papers studying the impact of AFDC-UP on family structure largely use cross-sectional methodologies that parallel that of the early AFDC-basic literature (Schram and Wiseman, 1988; Schultz, 1994; Winkler, 1995).[25] Yet, because AFDC-UP caseloads correlate strongly with unemployment (Figure 3), state-level associations between single motherhood and AFDC-UP may be wrong-signed because the broader labor market environment masks mitigating effects in state level aggregates. Thus, relative to prior research on AFDC-UP, a central contribution of this paper is the use individual comparisons across employment status in addition to comparisons between AFDC-UP adopting and non-adopting states.

Each of the above theoretical underpinning are amplified when considering that AFDC-UP conferred categorical eligibility for Medicaid and food stamps. Apart from wage losses during unemployment, many families also lose health insurance coverage, for which Medicaid eligibility may prove vital protection. Previous research also emphasizes the consumption smoothing benefits of food stamps (Gundersen and Ziliak, 2003) and ubiquity of food stamp receipt among the AFDC population (Hoynes, 1996). In Section IV I estimate that the combined value represented by the three programs is far above that from AFDC-UP alone. With regard to the "outside" option of separation as a means for mothers and children to receive benefits under the AFDC-basic program, the program interactions once again strengthen the theoretical connection between AFDC-UP and two-parent family stability, since AFDC-basic recipients were also eligible for Medicaid and food stamps.

[24] The literature has taken an interesting arc. Earlier research found largely mixed associations between welfare generosity and single motherhood (Groeneveld et al., 1983), while the literature in the 1980s found more of a consistent impact (Moffitt, 1992). Subsequent critiques and advocation for fixed effects to account for state-specific factors influencing both welfare generosity and family structure (Ellwood and Bane, 1985; Moffitt, 1994; Hoynes, 1997) led again to no consensus, in part because the state-level relationship between public assistance generosity and family structure long-preceded the Post-1960s period (Moehling, 2007).

[25] Schram and Wiseman (1988) and Winkler (1995) studied the impact of AFDC-UP on single-motherhood in 1980 and 1987 cross-sections, respectively, and found either null or positive effects on single motherhood. Using 1980 census data, Schultz (1994) found no effect of AFDC-UP on either single motherhood or fertility. In the closest study to the present, Hoynes (1997) utilized individual-level data from the Panel Study of Income Dynamics (PSID) from 1968-1989 and finds the expected negative relationship between AFDC-UP availability and single motherhood for white women. For black women, the sign reverses, but when AFDC-UP is interacted with state benefit amounts, the sign is as expected.

In sum, economic theory suggests that AFDC-UP should protect against two-parent family dissolution. While most of the aforementioned empirical evidence focuses on marital status, this paper's primary focus is on cohabitation and separation because the eligibility conditions for AFDC-UP relating to family structure considered whether both natural or adoptive parents of the child were in the home, not whether the parents were married (Winkler, 1995; Moffitt et al., 1998).[26] At the same time, because couples with children are very likely to be married and many states required AFDC-basic applicants to file for divorce to receive benefits (Finman, 1971), I also consider the impact of AFDC-UP in promoting marital stability.

II Data

The primary analysis sample is based on CPS-ASEC repeated cross-sections between 1977 and 1988. An additional sample described in Section VI extends the period back to 1968 for a restricted set of states.[27] A chief advantage of the CPS-ASEC is its availability much earlier than most longitudinal surveys, allowing me to trace the impact of AFDC-UP on family structure for up to two decades. Because both prior unemployment and previous earnings are crucial for assigning AFDC-UP eligibility, the cross-sectional nature of the CPS-ASEC requires me to focus analyses on family structure responses among men.

To construct the primary sample, I first keep men who are aged 24-58, are the head of their household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks and report wage income in excess of

[26] In the early years of the program, this distinction may have been weaker, but was clarified after the King V. Smith (1968) decision struck down so called "Man in the House Laws" (Bell, 1965) that restricted AFDC benefits for families if a man was in the house, regardless of their biological or adoptive status with regard to the children, and which existed almost exclusively in the South (Fuller, 2022).

[27] The sample begins in 1977 as that is the first year after 1967 that all states are identifiable in the CPS, while it ends in 1988 because the FSA of 1988 mandated all states adopt AFDC-UP. Details of individually identifiable and grouped states in the CPS between 1967-1976, as well as the creation of the restricted sample ranging from 1968-1988, are described in Section VI and detailed in Appendix B.

$1,000 in the prior year, and who report zero farm or business income in the prior year.[28] The restriction on positive weeks worked and wage income is intended to proxy for having sufficient work history to qualify for AFDC-UP. I further restrict the sample to men who also live in households where no other adult is in the labor force (if they live with an adult), which is largely motivated by the fact that mothers in AFDC-UP families work just 6.2 percent of the time (Hoynes, 1996, Table 1).[29] My preferred sample also restricts consideration to ever-married men, to proxy for having children. On the other hand, because AFDC-UP impacted the relative costs and benefits of cohabitation among natural or adoptive parents regardless of marital status, I show that results are quite similar when including never-married men in Appendix A.

II.A Defining unemployment spells

My preferred unemployment measure indicates whether individuals report being unemployed for at least 6 months in the previous year or report current unemployment of at least 6 months. I focus on "long-term" unemployment because these families should be the most likely to qualify for and receive AFDC-UP. I also focus on long-term unemployment because these spells are associated with larger declines in income and consumption, various measures of health deterioration, and larger increases in measures of stress (see Nichols et al. (2013) for a review of this literature). I also include those currently unemployed longer than half a year

[28] I choose age 23 rather than 18 to focus on men who are more likely to have children, while I keep those younger than 59 to reduce the likelihood of "planned" unemployment as a bridge between work and early retirement under Social Security (available at age 62). I do not consider those with alternative sources of income because it is not clear what type of income shock is represented by unemployment spells for these families.

[29] This restriction is also motivated by the fact that families in which women are not working would be more likely to view AFDC-basic as a plausible outside option, thus increasing the incentive to separate absent AFDC-UP. In 1973, for example, less than 30 percent of mothers among all AFDC cases were in the labor force, and less than 1/6 were currently employed (U.S. NCSS, 1975a). The restriction may also imply lower "added worker" effects (Lundberg, 1985; Cullen and Gruber, 2000; Stephens Jr., 2002), which may crowdout welfare recipiency.

because long-term unemployment is a relatively rare event in the sample (Table 1).[30] On the other hand, comparisons of the gradients between unemployment duration and welfare receipt, family structure stability, and the protectiveness of AFDC-UP serve as additional checks on validity of the research design, and I therefore investigate heterogeneity by unemployment duration in Section IV. In the primary analysis, I further drop men who are not classified as long-term unemployed but who report some current or former unemployment, to more clearly delineate those experiencing an unemployment spell in the window of observation from those who do not. Henceforth I refer to the preferred long-term unemployment measure as "unemployment" unless otherwise noted.

II.B Estimating income terciles

Given variation in specific income and asset means tests both across states and over time (e.g., see Hoynes (1996)), I take a straightforward approach and conduct analyses separately by tercile of the previous year's personal wage and salary income.[31] Because the interest lies in comparing men who have similar *full time earnings*, I predict "potential" earnings from men who worked 52 weeks in the previous year (87.3 percent of the sample) using the Least Absolute Shrinkage and Selection Operator (LASSO; Tibshirani, 1996), with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry.[32] To avoid comparing relatively wealthier individuals in poor states to poorer

[30]Because the CPS-ASEC occurs in March, these men spent some time unemployed in the previous year, hence it is reasonable to expect questions regarding transfers in the previous year to reflect such spells.

[31]I use personal wage and salary income rather than household income because the former tracks individuals regardless of the household they are in, whereas the latter does not include income of people who no-longer live in the household. Because I examine the heads of household, condition on no farm or business income, and also condition on no other adults in the labor market, these measures are closely related (correlation of 0.83). I choose to partition into terciles because both unemployment and family dissolution are rare events. Finer quantiles tends to produce consistent but less precise and monotonic results over the income distribution.

[32]I use the square-root LASSO (Belloni et al., 2011), which does not require knowledge of the standard deviation of the error term, but also show that results using the Extended Bayesian Information Criteria (Chen and Chen, 2008) or linear regression are similar in Appendix A. Pairwise correlations across the

individuals in wealthier states and to account for differential price changes across space, I estimate terciles separately by state and year.

II.C State AFDC-UP programs

Data on AFDC-UP caseloads by state and month come from the Department of Health and Human Services (DHHS).[33] States varied in their timing of adoption (Figure 2) and some dropped and re-started their programs, so I define whether a state has an AFDC-UP program in a given calendar year by whether there were positive AFDC-UP caseloads in any month. Figure B.1 shows the years that each state had the program in effect for states that ever adopted by 1988.

Rather than parameterize the staggered initial roll out of the program, my preferred AFDC-UP treatment measure ($\text{UP}^{76}_{s(i)}$) indicates whether men reside in states that did or did not have the program by 1977, and is therefore time-invariant. There are several reasons to focus on a later period than the early 1960s. Perhaps most importantly, the absence of intercensal microdata containing information on family structure, labor force participation and work, and income and transfers in the 1950s complicates analyses in a window around 1961.[34]

I drop Alaska, Hawaii, and the District of Columbia, as well as Nevada (due to extreme outlier rates of marriage and divorce). I further drop five additional states (Colorado, Maine,

predicted income measures range between 0.94 and 0.98.

[33]These data are available at: `https://www.acf.hhs.gov/ofa/resource/tanf-and-afdc-historical-case-data-pre-2012`.

[34]Other reasons include the King V. Smith (1968) ruling (see footnote 26), which ensured that AFDC-UP is the only AFDC policy relevant for two-parent families after that period; the temporary provision of AFDC-UP through 1967, which may have impacted state adoption or otherwise affected legislative or individual expectations regarding stability of the program; that the correlation between unemployment and caseloads is not strong in the 1960s, with caseloads rising even as unemployment is falling; and that there is mostly clustered early adoption, with not many states responding to the high unemployment environment of the 1970s with programs, indicating the presence of program is more likely to be exogenous in later years. I do not attempt to leverage the FSA of 1988 because the legislation allowed for states that did not already have AFDC-UP to offer benefits on a reduced basis, and caseloads did not increase proportionately. The 1990s also saw much experimentation with state welfare policies, which further complicates analyzing that period.

Oregon, South Carolina, and Utah), due to long periods in which they dropped the program during the sample (three or more years). Finally, I treat an additional 5 states as always or never having the program because of short-periods (two or less years) in which they did not or did have provide AFDC-UP benefits, respectively. Further details are described in Appendix B. The final sample consists of 42 states including 22 treated and 20 control states and is mapped in Figure 2, Panel B.[35] Although control states are disproportionately in the South, I show in Section V that most of the central conclusions are unchanged by omitting all states in the region.

II.D Program participation and family structure outcomes

The CPS began collecting information on welfare receipt as of the previous calendar year in 1966, and the question pertains to the respondent themselves. AFDC-UP often, but not always, payed benefits to support the father themselves, but household recipiency is the relevant margin for household decisions such as separation. As such, I use as my preferred welfare receipt measure whether the householder or spouse report welfare receipt in the prior year, if there is a spouse present in the household.[36] The CPS also collected information on prior year receipt of other cash and in-kind government transfer programs of interest, including UI (beginning in 1976) and Medicaid and food stamps (beginning in 1980). Supplemental analyses for Medicaid and food stamps are therefore conducted on a restricted sample ranging from 1980-1988. Food stamp recipiency is at the household level, and I similarly construct a household-level Medicaid coverage variable indicating whether the head of household or spouse was covered by Medicaid.[37]

[35] This definition also circumvents issues with timing of adoption in difference in differences designs (Callaway and Sant'Anna, 2021; Goodman-Bacon, 2021). I define treatment as the first time a state adopts AFDC-UP because of difficulties arising in interpreting coefficients where units are treated multiple times (Sandler and Sandler, 2014).

[36] There are also some cases in which only children were eligible for benefits. Unfortunately, the universe for this question is either age 14 or 15 and over, so I choose not to augment this measure with child recipiency.

[37] The use of a household Medicaid variable is largely irrelevant, as over 99 percent of the sample is classified equivalently as the original variable.

The primary family structure outcomes are currently separated, divorced, or reporting spousal absence, currently married with spouse present, living with one's own child(ren) under age 18, and heterosexual cohabitation (defined by living with one's own minor child(ren) and the mother of at least one of these children). The analysis effectively examines how family structure outcomes are related to recent long-term unemployment (measured within the previous 15 months). There is good reason to think separation would respond within this time frame, and thus that any protective effects of AFDC-UP would also be evident. Lindo et al. (2022) find that divorces and separations respond within the first quarter after job displacement, with little evidence of growing magnitudes over time, and Charles and Stephens (2004) find that layoffs increase the divorce risk within 1-3 years, with *smaller* effects looking at a longer horizon. As an additional validity check, I also test for differences in the probability of being married with no children, families for whom AFDC-UP should have no effect. The within-state comparisons ensure that any state-year-specific differences in fertility rates are captured, while the inclusion of age dummies in some analyses are intended to account for differences in fertility by age.

II.E Summary statistics and balance

Table 1 shows summary statistics for the overall sample in column 1 and for each tercile of the (predicted) income distribution in columns 2-4 (all monetary variables are in 2019 dollars). Family structure characteristics are largely similar across terciles, whereas there is an expected negative gradient between income and educational attainment and home ownership. Poorer individuals are also more likely to experience long-term unemployment and to head families that receive welfare, receive food stamps, or are covered by Medicaid. They are also more likely to receive UI, but less likely conditional on unemployment, and are younger and less likely to be white.

While notable changes to family structure occurred between the introduction of AFDC-UP in 1961 and the start of the primary sample period in 1977, I investigate balance between

AFDC-UP and non-AFDC-UP states using the 1960 census 5 percent sample (Ruggles et al., 2023) and the same restrictions as the primary analysis sample. Table 2 shows these tests provide little evidence of differences in family structure, including no differences in probabilities of being married, divorced, living with children under 18, or cohabiting with children under 18. Men in AFDC-UP states marry for the first time around 1 year later, and are roughly 3 percent less likely to have been married multiple times. Unsurprisingly given that many control states are in the South, AFDC-UP states are whiter, richer, and more educated, although expenses are also higher, with meaningful differences in annual rent and a "cost of living" measure that includes the cost of electricity, gas, water, and fuel. Finally, AFDC-UP states had higher UI recipiency (FIU) — largely due to differential industrial composition — had higher UI maximums, higher ADC benefits per case, and similar ADC recipiency rates per 1,000 women ages 15-54.

Most of these characteristics would be expected to mean a disproportionately *larger* family structure response to unemployment in non-AFDC-UP states, biasing down the protectiveness of AFDC-UP. Nevertheless, because men in these states are fundamentally different on characteristics shown to affect or correlate with family structure, I investigate robustness to propensity score weighting methods in Section IV and also show in Section V that results are quite stable when dropping all men in the South.

III Empirical Strategy

The empirical analysis is motivated by a "no-timing" difference in differences research design that compares unemployed men to those who are employed, in AFDC-UP states versus non-AFDC-UP states, and separately by tercile of the income distribution. Due to the means-tested nature of AFDC-UP, public assistance recipiency and measures of family stability should respond primarily in the first tercile. I use this feature as a test of the identifying assumptions underlying these models, which I describe further below.

Consider the following simple difference in differences linear probability specification estimated via Ordinary Least Squares:

$$y_{ist} = \beta_1 \times \text{UN}_{ist} + \beta_2 \times \text{UP}^{76}_{s(i)} + \beta_3 \times \text{UN}_{ist} \times \text{UP}^{76}_{s(i)} + \varepsilon_{ist} \tag{1}$$

Where y_{ist} is the program participation or family structure outcome of interest for individual i in state s and year t (program participation asked in t regards receipt in $t-1$). There are two coefficients of interest per regression: $\widehat{\beta}_1$ measures the association between unemployment and outcomes in non-AFDC-UP-adopting states, while $\widehat{\beta}_3$ measures the effect of unemployment in AFDC-UP states relative to this benchmark association. In other words, $\widehat{\beta}_3$ represents the additional program participation and the protectiveness of AFDC-UP against unemployment-associated family dissolution among recently or currently unemployed men and their families in AFDC-UP states.

To understand the identifying assumptions necessary for $\widehat{\beta}_3$ to estimate a causal effect and what treatment effect it estimates under these assumptions, let $y_{ist}(\text{UP}^{76}_{s(i)}, \text{UN}_{ist})$ denote the potential outcomes for being exposed to AFDC-UP if unemployed ($y_{ist}(1,1)$), exposed to AFDC-UP if employed ($y_{ist}(1,0)$), not exposed to AFDC-UP if unemployed ($y_{ist}(0,1)$), and not exposed to AFDC-UP if employed ($y_{ist}(0,0)$). $\widehat{\beta}_3$ mechanically estimates:

$$\begin{aligned}&\mathbf{E}[y_{ist}|\text{UP}^{76}_{s(i)} = 1, \text{UN}_{ist} = 1] - \mathbf{E}[y_{ist}|\text{UP}^{76}_{s(i)} = 1, \text{UN}_{ist} = 0] \\ &\quad - \big(\mathbf{E}[y_{ist}|\text{UP}^{76}_{s(i)} = 0, \text{UN}_{ist} = 1] - \mathbf{E}[y_{ist}|\text{UP}^{76}_{s(i)} = 0, \text{UN}_{ist} = 0]\big)\end{aligned} \tag{2}$$

In Appendix D I show that the above estimand is equivalent to the following expression:[38]

$$\widehat{\beta}_3 \xrightarrow{p} ATT - ATT^{ind} + \Delta ATT^{ue} + BS \tag{3}$$

The first term represents one parameter of interest; the Average Treatment Effect on the Treated (ATT) for unemployed individuals in AFDC-UP states. The second term represents an indirect treatment effect that AFDC-UP may have on the families of currently employed

[38]This derivation draws heavily on Goodman-Bacon (2023).

men — for example, through its insurance value — and is another parameter of interest. ΔATT^{ue} represents the difference in causal effects of unemployment on family structure (not including the effect of AFDC-UP) for actually unemployed men in AFDC-UP and non-AFDC-UP states, and BS is the mean difference in non-AFDC-UP, employed potential outcomes between actually unemployed and employed workers in AFDC-UP states net the same measure in non-AFDC-UP states, and is analogous to parallel trends in more canonical difference in differences research designs that leverage timing.

The empirical analysis does not separately identify ATT and ATT^{ind}. Assuming that $ATT^{ind} = 0$ would rule out any "spillover" effects of AFDC-UP on the families of household heads observed employed. As mentioned above, this assumption may not be tenable in the situation where households know they will have access to an income stream were the head to become unemployed. However, given that the insurance value is one component of the *effect* of AFDC-UP provision, the policy relevant parameter should include both direct (ATT) and indirect (ATT^{ind}) effects.

Identification of $ATT - ATT^{ind}$ requires that ΔATT^{ue} and BS are both zero or sum to zero. Focusing on the latter term first, I term the assumption that $BS = 0$ as the "bias stability" assumption, which states that average un-treated potential outcomes can vary across observed-unemployed and employed men, but this difference must be stable across treated and control states. The assumption that $\Delta ATT^{ue} = 0$ states that the direct effect of unemployment (i.e., excluding the effect of AFDC-UP) on outcomes must be the same among unemployed men in AFDC-UP and non-AFDC-UP states. These assumptions are not directly testable, but the assumption that they sum to zero can be indirectly tested by examining whether unemployment impacts family structure differentially in AFDC-UP states in higher terciles of the income distribution, where AFDC-UP is not meaningfully present (i.e., where we expect $ATT = ATT^{ind} = 0$). I also consider results using re-weighting methods in Section IV that are identified off of different, conditional assumptions regarding untreated potential outcomes.

In practice, the primary analysis is based instead on variants of the following augmented model:

$$y_{ist} = \delta_{st} + \gamma_1 \times \text{UN}_{ist} + \gamma_2 \times \text{UN}_{ist} \times \text{UP}^{76}_{s(i)} + \mathbf{X}'_{ist}\boldsymbol{\beta} + \varepsilon_{ist} \tag{4}$$

Where the second term in specification (1) has been replaced with state-by-year fixed effects (δ_{st}) and the model now allows for the inclusion of controls ($\mathbf{X}_{ist}$). This specification imposes weaker identifying assumptions than specification (1), which is comprised in part of comparisons between unemployed and employed individuals across different states (within treatment and control groups) and over time. In contrast, specification (4) limits these comparisons to be within state-year. In practice, I show in Appendix A that results using models (1) and (4) are quite similar.

While my preferred specification omits additional controls, various robustness checks include in $\mathbf{X}_{ist}$ the individual's race, fixed effects for educational attainment bins (less than high school, high school, some college, and college or more), age fixed effects, and occupation and industry fixed effects. In some specifications, $\mathbf{X}_{ist}$ also includes the interaction of the state-year UI maximum and unemployment status, or the interaction of the 1980 FIU and unemployment status, as attempts to control for underlying differences in UI generosity or eligibility. All regressions use household survey weights and cluster standard errors at the state level.

IV Results

The primary results from specification (4) are shown separately by tercile in Table 3. Consistent with prior research, unemployment is associated with increased rates of separation and lower rates of marriage, while the results also provide novel evidence of a decline in cohabitation as well. These effects are all large relative to sample means, are statistically significant at conventional levels in the lowest income tercile, and are generally also present and similar in magnitude at higher terciles (although less precise for some outcomes).[39] Taking the coefficient on separation in the lowest income tercile as illustrative of the magnitudes, the association of 11.7 percentage points (p.p.; p-value< 0.01) relative to the (weighted) mean among employed men in the control group (20.4 p.p.) implies a 57.1 percent increase in separation probabilities (60.2 percent relative to the sample mean). The separation associations are similar to results from Charles and Stephens (2004), who find effects on unemployment divorce hazards between roughly 18 and 30 percentage points.

Column 1 shows that welfare receipt is strongly related to the interaction of previous unemployment and living in AFDC-UP states (18.5 p.p.; p-value< 0.01), constituting a more than 300 percent increase in the probability of welfare receipt relative to unemployed men in non-AFDC-UP states.[40] While there are statistically significant impacts on welfare participation in the middle and upper terciles, the coefficients are around 40 percent and 20 percent as large as in the lowest tercile, respectively.[41]

Turning to estimates of the Intent to Treat (ITT) effects, the wealth of evidence suggests that the role AFDC-UP provided in mitigating unemployment-associated family dissolution is quite large and only present in the lowest income tercile, strongly supporting the identifying assumptions. Column 2 shows that AFDC-UP decreases the likelihood of separation by 6 p.p. (p-value$= 0.051$), counteracting 0.061/0.117 or 52.1 percent of unemployment-associated separation (p-value< 0.01 using the Delta Method). Results for married with

[39]In the middle tercile, the coefficient on cohabitation is statistically significant, but the coefficients on separation and marriage are not. In the upper tercile, the coefficients on separation and marriage are statistically significant, but the coefficient on cohabitation is not.

[40]The significant coefficient on welfare for unemployed men in non-AFDC-UP states likely reflects General Assistance (see footnote 23), which is commonly referred to as welfare but did not have its own question. It may also reflect some spouses (and children) who received AFDC-basic benefits if the couple had temporarily separated but were re-united by the time of the interview.

[41]This is not surprising, since AFDC-UP recipients were often higher income relative to participants in the AFDC-basic program (Hoynes, 1996). The presence of relative increases in welfare receipt in higher income terciles may also be due to systematic over-estimation of potential wages for the unemployed in light of evidence that wages begin declining before displacement (Jacobson et al., 1993). Indeed, if I assume that wages are 20 percent lower than their predicted value among the unemployed (roughly the magnitude found in Jacobson et al. (1993)), the coefficient on welfare receipt in the middle tercile is roughly half the size and no longer significant and is largely unchanged in the upper tercile.

spouse present (column 3) are similar.[42] The separation results are remarkably close to those of Lindo et al. (2022), whose estimates imply protectiveness of 52.4 percent against heightened divorce and separation risk were average UI benefits to equal average UI maximums.[43] Reassuringly, column 4 shows that results for being married with no children are small and insignificant, suggesting that changes in marriage, separation, and cohabitation are driven by couples with children.[44] The impacts of AFDC-UP on cohabitation and living with one's own children (columns 5 and 6) are *larger* relative to the corresponding unemployment associations, indicating AFDC-UP fully counteracts the roughly 7 and 8 p.p reductions associated with unemployment, respectively.

The large degree of protectiveness provided by AFDC-UP is consistent with descriptive evidence contained in reports from six states that terminated AFDC-UP and tracked participants, which found 12-28 percent of families become single-parent AFDC families (U.S. General Accounting Office, 1988 p. 4). The Utah report further reveals that separation rates were roughly 5 percentage points higher among former AFDC-UP participants after the program was terminated (a roughly 70 percent increase), while around 40 percent of respondents in Washington who separated stated that the unavailability of AFDC-UP contributed to their decision (U.S. General Accounting Office, 1988 pp. 20-21).

While men in higher terciles also received AFDC-UP to a lesser extent, it appears that family structure did not respond systematically to these benefits. This provides suggestive evidence that AFDC-UP provided larger welfare gains to families who are poorer. Given that AFDC-UP benefits were not subject to maximum durations and it is plausible that more needy families received benefits for longer, the effective targeting of high-marginal-

[42] Because I condition on ever-married individuals, the only difference between these measures is widowhood, although Figure A.4 shows that both separation and marriage results are similar when including never-married men.

[43] I calculate this value by taking their implied effect of a $1 increase in UI maximums on divorce and separation for men (0.000040), multiply by the average UI maximum ($334), and divide by the coefficient on layoffs (0.0255).

[44] Interestingly, there also does not seem to be an effect of unemployment on marriage with no children across all terciles, somewhat at odds with the findings from Lindo et al. (2022).

utility states *among* poor households is likely a primary channel through which AFDC-UP provided protection. I turn next to documenting this relationship explicitly through an analysis of heterogeneity by unemployment duration.

IV.A Heterogeneity by unemployment duration

The analysis thus far has focused on long-term unemployment, which corresponds roughly to the highest quartile of unemployment duration in the sample. The available evidence indicates that AFDC-UP should be more prevalent for longer-term unemployed individuals, while theory and evidence also indicates that longer-term unemployment corresponds with greater declines in household welfare (Nichols et al., 2013; Ganong and Noel, 2019). To test these hypotheses, I re-estimate the primary outcomes separately by quartile of the unemployment-duration distribution (for the lower income tercile).[45]

The results are plotted in Figure 5. Panel A shows the expected monotonic relationship between family welfare receipt and unemployment duration for families of men in AFDC-UP states, with little evidence of any trend among those in non-adopting states. Panel B shows heightened separation risk is associated with unemployment across quartiles of duration, but is also largely increasing in duration (strictly so between quartiles two and four). In contrast, the mitigating effects of AFDC-UP only begin to appear in the third quartile (not significant), with the significant coefficient in the fourth quartile largely reproducing the primary results of protectiveness. Panel C shows results for marriage are similar. Panel D shows that the patterns for cohabitation are also monotonic in the expected direction, save for the 2nd quartile, in which there is a puzzling *decline* in cohabitation. This anomaly is

[45] The quartiles of prior-year unemployment are 1-6, 7-12, 13-23, and 24-51 weeks. They are the same weighted and unweighted. The quartiles of current unemployment duration are similar (1-5, 6-13, 14-21, and 22 and above). For consistency, I use the bins defined by prior year unemployment, since these men constitute the majority of all unemployment in the sample. These regressions are from samples that are close, but not identical, to the primary sample. The only difference is that each drops those reporting unemployment duration outside of the range corresponding to the relevant quartile, which is analogous to dropping shorter unemployment durations in the primary sample and ensures the employed portion of the control group does not change across samples.

not present for living with minor children (Panel E).

These results provide additional evidence that AFDC-UP disproportionately targeted the most needy households, and that family structure was also more likely to respond to such benefits among these households. The literature also documents higher UI replacement rates lead to higher consumption smoothing benefits (Gruber and Madrian, 1997; East and Kuka, 2015). I turn next to showing that average AFDC-UP monthly replacement rates (total value of benefits as a share of income) were also in excess of average UI replacement rates when accounting for transfer program interactions.

IV.B Categorical eligibility for Medicaid and food stamps

One reason that AFDC-UP provides large protection may be because it conferred categorical eligibility for food stamps and Medicaid, which increases the total value of benefits substantially. Table 4 shows results for health insurance coverage, individual and joint receipt across transfer programs, and additional outcomes (the sample begins in 1980, when these variables are first recorded).[46] Consistent with prior evidence (Gruber and Madrian, 1997; Schaller and Stevens, 2015), column 1 shows that unemployment is associated with large and precise losses in employer-provided group coverage across income terciles. Column 2 shows that the families of unemployed men in AFDC-UP states are much more likely to be covered by medicaid, with a close magnitude to that of welfare receipt (21 p.p.; p-value< 0.01), undoing roughly 80 percent of the 27 p.p. coverage loss in the lowest tercile.[47] These are largely the same families; column 3 shows the coefficient on joint welfare receipt and Medicaid coverage is 18 p.p.[48] The coefficient on food stamps (column 4) is also positive, although

[46] Table A.2 shows results for the primary outcomes are similar when estimated only using data from 1980-1988.

[47] In line with the primary results for welfare participation, AFDC-UP also precisely predicts lower increases in Medicaid receipt in the middle tercile.

[48] It is unlikely these results are driven by differential state Medicaid policies per se, since AFDC receipt is the most important determinant of Medicaid coverage in this period. The roughly 14 million Medicaid recipients in AFDC families (basic and unemployed) in 1980 comprised over 90 percent of all Medicaid recipients not categorically eligible due to age, disability, or blindness (Gornick et al., 1985). Furthermore, administrative information from 1973 indicates that 90 percent of AFDC families lost Medicaid

much smaller (3 p.p.) and not precise. However, the great majority of welfare recipients jointly received food stamps (column 5; 16 p.p.; p-value< 0.01) and the coefficient on jointly receiving or being covered by the three programs is virtually identical (column 6; 16 p.p.; p-value< 0.01).[49]

To quantify the total value of categorical eligibility, I consider monthly replacement rates of previous earnings. While in-kind benefits may be valued above or below their equivalent dollar cost (Currie and Gahvari, 2008), estimates of the willingness to pay for Medicaid for childless adults cover a range including one (Finkelstein et al., 2019) and parents may value each dollar of health expenditure for their children more. Because I do not observe full-time earnings for unemployed men and the value of transfers in the CPS-ASEC is likely attenuated (see Section IV.C), I consider average replacement rates using 1979 wages for full-time earners with maximum educational attainment either less than or equal to a high school degree using the 1980 5 percent census (Ruggles et al., 2023) combined with administrative information on program spending per case or recipient in 1979 or close years, dividing annual statistics by 12 (all below values are in 1979 dollars).[50]

Median 1979 full time monthly earnings among married men with children and less than a high school degree was \$1,150 (\$1,478 for those with a high school degree). The 1979 AFDC-UP monthly benefit per family was \$398 (Bureau of Public Assistance, 1979), the value of food stamps per person was \$31 (U.S. Department of Agriculture Food and Nutrition Service , USDA), and the 1975 Medicaid benefit for each child was \$26 and for each adult

when their cases were terminated (U.S. NCSS, 1975b p. 19.) even though family income may actually *fall* after AFDC-UP termination (Hoynes, 1996). I find no evidence that income effects lead to higher private health insurance as is the case for UI (Kuka, 2020), which is consistent both with broad access to Medicaid crowding out some health insurance investment and with the fact that these families have very low "discretionary" income.

[49]The much smaller effect for food stamp receipt alone is likely because this is the only program considered where "need" is determined at the national level and non-AFDC-UP states were poorer (Table 2), which should imply lower participation in AFDC-UP states absent differential welfare policy.

[50]I consider these education groups because, as Table 1 shows, almost 50 percent of men in the lowest income tercile had less than a high school degree and around 90 percent had at most a high school degree.

\$51 (\$24 and \$47 in 1984, respectively; Gornick et al. (1985)).[51] The average number of kids per AFDC-UP case was roughly 2.5 (Bureau of Public Assistance, 1979). Summing the AFDC-UP benefit, the food stamp benefit multiplied by 4.5, and Medicaid for 2 adults and 2.5 children implies a total value of \$702, which is over 75 percent larger than the AFDC-UP benefit alone and represents a replacement rate of between 47-61 percent depending on the median wage. The replacement rate could be much higher for very low wage workers and may also be a low estimate because AFDC recipients were also prioritized for housing assistance (see Moffitt (1998) p. 52).[52] These estimates are generally larger than for UI, which are below 40 percent for individuals in some states (Kuka, 2020; Dube, 2021).[53] Thus, apart from disproportionate AFDC-UP receipt among families of the long-term unemployed, the high benefit value relative to earnings is another primary explanation for the large degree of protection provided by AFDC-UP.

IV.C Protectiveness per AFDC-UP case in the presence of measurement error

I next consider implied effects per AFDC-UP case by scaling the reduced form family structure outcomes by the welfare participation coefficients (Wald estimates). Dividing the estimate on separation by the probability of welfare receipt (-0.061/0.185) implies roughly 33 percent of individuals would have separated in the absence of AFDC-UP. Since the coefficient on unemployment is 11.7 percent, this suggests a level of protectiveness per case far in excess of 100 percent. For cohabitation, the impact is even larger. How can we rationalize these magnitudes? One plausible explanation is the presence of significant measurement error in the CPS-ASEC, which functions to attenuate recipiency estimates leading to positively

[51] Although Medicaid coverage is not equivalent to medical care usage, surveys in the 1960s and 1970s found that 90 percent of families on welfare with Medicaid coverage also reported using Medicaid (Goodman-Bacon, 2018).

[52] Downward trends in real transfers during this period suggest the replacement rates may have been higher in earlier years, and lower towards the end of the sample period.

[53] High benefits relative to previous earnings for low income workers may be one reason why Hoynes (1996) finds a sharp labor supply response to AFDC-UP.

biased estimates per case.[54]

In Appendix C, I conduct a series of exercises to bound recipiency effects under plausible assumptions regarding three sources of measurement error: systematic underreporting of welfare receipt, missed welfare receipt in the first quarter of the current year, and error in prior-year unemployment recall. The results suggest that each source of measurement error functions to attenuate the estimated recipiency effects, biasing up estimates per case. A simplistic aggregation across sources of measurement error implies that recipiency rates may be upwards of 4 times larger than the estimates (around 90 p.p.) among the families of the long-term unemployed in the lowest income tercile in AFDC-UP states, which is broadly consistent with the actual size of the program. These results would be consistent with levels of protectiveness per case that are around 60 percent for separation and indistinguishable from 100 percent for cohabitation.

IV.D The interaction of AFDC-UP and UI

What does the empirical evidence reveal about the the degree of substitutability between AFDC-UP and UI? Table 4, column 8 shows differentially more UI recipiency in AFDC-UP states across income terciles. In the lowest tercile, the effect of 12 p.p. is roughly 60 percent the size of welfare participation. This may be because AFDC-UP states provide more generous UI benefits and higher UI benefit levels increase take up (Blank and Card, 1991; Anderson and Meyer, 1997), although differential UI receipt is likely in part also due to differences in the composition of the workforce (see the comparisons of UI maximums and the 1960 FIU in Table 2). The ratios of the coefficients on welfare and UI participation relative to those who are unemployed in non-AFDC-UP states, over 300 percent and roughly 25 percent, respectively, strongly suggest that the results are not being driven by

[54] Because the derivation of Section III shows that coefficient estimates are comprised of direct effects on the unemployed in AFDC-UP states net any indirect effects on those currently employed, this point is strengthened if such indirect effects exist. At the same time, the presence of such indirect effects suggests that measures of protectiveness per case may not be a policy-relevant treatment effect, particularly since the number of families headed by employed men for whom such indirect effects exist is not measured.

differential UI receipt. Perhaps more importantly, the majority of these recipients also received welfare (column 9), so that AFDC-UP functioned to "extend" UI benefits.[55] Indeed, column 11 shows that UI receipt conditional on no welfare receipt is much less common and not statistically distinguishable from zero. Because I do not model the length of AFDC-UP recipiency, I do not attempt to separate joint or consecutive recipiency of AFDC-UP and UI in interpreting the magnitude of the results, but only note that it was part of the suite of benefits provided. Moreover, I show in Figures A.3-A.6 that none of the central conclusions are changed when controlling directly for either the UI maximum or the 1980 FIU interacted with unemployment status.

IV.E Doubly Robust estimates

The unconditional identifying assumption in Section III regarding ATT^{ue} and BS summing to zero may not be tenable *a priori*, given evidence of imbalance between AFDC-UP and non-AFDC-UP states in pre-treatment characteristics (Table 2). Furthermore, the geographic concentration of the early 1980s recession in regions more likely to have AFDC-UP (the Midwest, East South Central, and Western census divisions (Stuart, 2022)), which may confound estimates because layoffs are better predictors of divorce than plant closings (Charles and Stephens, 2004). While I present evidence in Section V that controlling for fixed effects for every occupation and industry in the sample does not change any of the central conclusions presented thus far, I investigate robustness to a different estimation strategy that uses re-weighting methods to assign higher weights to unemployed men in control states who more closely resemble unemployed men in AFDC-UP states. I then compare this estimate to a similarly re-weighted estimate of the effect of unemployment by weighting employed men in control states to more closely resemble unemployed men in these states. These estimates are identified under two unconfoundedness assumptions in place of $BS = 0$, but still require

[55] These findings complement those of Rothstein and Valletta (2017), who study a period after AFDC-UP ceased to exit and find that public assistance receipt does not increase significantly after UI benefits are exhausted.

$ATT^{ue} = 0$ (i.e., that the (re-weighted) effect of unemployment on family structure is the same in treated and control states).[56]

I leverage recently developed Doubly Robust (DR) estimators, which yield consistent estimates if either the Outcome Regression (OR) or the propensity score is miss-specified (but not both). Following the notation in Section III, I employ the DR ATT estimator proposed by Mercatanti and Li (2014), estimating the following *among unemployed men*:

$$\begin{aligned}\widehat{ATT}^{\mathrm{UP}} =& \frac{1}{N_1^{\mathrm{UN}}}\sum_{i=1}^{N} y_{ist}\mathrm{UP}^{76}_{s(i)} \\ &- \frac{1}{N_1^{\mathrm{UN}}}\sum_{i=1}^{N} \frac{y_{ist}(1-\mathrm{UP}^{76}_{s(i)})\widehat{p}^{\mathrm{UP}}(\mathbf{X}_{ist}) + \widehat{\mu}_0^{\mathrm{UN}}(\mathbf{X}_{ist})(\mathrm{UP}^{76}_{s(i)} - \widehat{p}^{\mathrm{UP}}(\mathbf{X}_{ist}))}{1-\widehat{p}^{\mathrm{UP}}(\mathbf{X}_{ist})}\end{aligned} \tag{5}$$

Where N_1^{UN} denotes the number of unemployed men in AFDC-UP states in the sample, $\widehat{p}^{\mathrm{UP}}(\mathbf{X}_{ist})$ are the propensity score estimates for receiving treatment (living in an AFDC-UP state) among unemployed men, and $\widehat{\mu}_0^{\mathrm{UN}}(\mathbf{X}_{ist})$ are the predicted values from an OR estimated only among unemployed men in control states of the form: $y_{ist} = \mathbf{X}'_{ist}\boldsymbol{\beta} + \varepsilon_{ist}$. For both the OR and the propensity score, covariates $\mathbf{X}'_{ist}$ include year fixed effects, race, industry fixed effects, age fixed effects, and a metro dummy. I estimate the propensity score using a logistic regression (Figure A.2, Panel A shows propensity score overlap for this group). Assuming either the OR or the propensity score is correct,

$$\widehat{ATT}^{\mathrm{UP}} \xrightarrow{p} ATT^{\mathrm{UP}} \equiv \mathbf{E}[y_{ist}(1,1) - y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1] \tag{6}$$

I also consider a similar formula for estimating the effect of unemployment in control states, re-weighting employed men to resemble unemployed men. The estimator is the following, applied *among men in control states*:

[56] Formally, for the DR estimator given by (5) the unconfoundedness assumption is $y_{ist}(1,1), y_{ist}(0,1) \perp\!\!\!\perp \mathrm{UP}^{76}_{s(i)}|\mathbf{X}_{ist}$ and for the DR estimator given by (7) the unconfoundedness assumption is $y_{ist}(0,1), y_{ist}(0,0) \perp\!\!\!\perp \mathrm{UN}_{ist}|\mathbf{X}_{ist}$.

$$\widehat{ATT}^{\mathrm{UN}} = \frac{1}{N_0^{\mathrm{UN}}}\sum_{i=1}^{N} y_{ist}\mathrm{UN}_{ist} - \frac{1}{N_0^{\mathrm{UN}}}\sum_{i=1}^{N} \frac{y_{ist}(1-\mathrm{UN}_{ist})\widehat{p}^{\mathrm{UN}}(\mathbf{X}_{ist}) + \widehat{\mu}_0^{\mathrm{EMP}}(\mathbf{X}_{ist})(\mathrm{UN}_{ist} - \widehat{p}^{\mathrm{UN}}(\mathbf{X}_{ist}))}{1-\widehat{p}^{\mathrm{UN}}(\mathbf{X}_{ist})} \tag{7}$$

Where N_0^{UN} denotes the number of unemployed men in non-AFDC-UP states, $\widehat{p}^{\mathrm{UN}}(\mathbf{X}_{ist})$ are the propensity score estimates for being unemployed among men living in control states, and $\widehat{\mu}_0^{\mathrm{EMP}}(\mathbf{X}_{ist})$ are the predicted values from OR models estimated only among employed men in control states (of the same form as above). Figure A.2, Panel B shows propensity score overlap for this group. Again assuming either the OR or the propensity score is correct,

$$\widehat{ATT}^{\mathrm{UN}} \xrightarrow{p} ATT^{\mathrm{UN}} \equiv \mathbf{E}[y_{ist}(0,1) - y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1] \tag{8}$$

To mirror the prior, fully parametric analysis, I consider the impact of AFDC-UP in mitigating unemployment-associated family dissolution as $\widehat{ATT}^{\mathrm{UP}}$ and the protectiveness of AFDC-UP as $\widehat{ATT}^{\mathrm{UP}}/\widehat{ATT}^{\mathrm{UN}}$.

The results from this procedure are shown in Table 5, where terciles 2 and 3 are pooled.[57] The estimates of $\widehat{ATT}^{\mathrm{UP}}$ in the lower tercile show precise impacts on welfare receipt, separation, marriage, and living with children of comparable magnitudes to the prior results, and an effect on cohabitation of roughly half the magnitude (4 p.p.; p-value=0.124). The magnitudes for the effect of unemployment are all precise and relatively reduced, particularly for separation and marriage, so that measures of protectiveness are higher for these outcomes (not distinguishable from 100 percent). Again, all results are driven by couples with children. There is little systematic evidence of any family structure response to AFDC-UP in the middle and upper income terciles.

Taken together, these results indicate that the general patterns of protectiveness found for AFDC-UP are not due to differential pre-treatment characteristics of men residing in treated

[57] I pool these terciles due to low rates of long-term unemployment and many empty cells when attempting to estimate the propensity score in the upper tercile alone.

versus non-treated states or to any systematic differences in the type of unemployment (e.g., voluntary versus mass layoff). If anything, the DR estimates strengthen the relationship between AFDC-UP and two-parent family structure stability among unemployed men. I turn next to providing a host of additional sensitivity checks, which all confirm these findings.

V Additional sensitivity checks, robustness, and ruling out alternative explanations

V.A Dropping the South

A notable geographic pattern evident from Figure 2, Panel B is the concentration of control states in the Southern Census region. In the lowest income tercile, about 70 percent of the control group is in the South, compared with just 7 percent of the treatment group. The 1960s saw stark convergence in general fertility among Southern black women (Thompson, 2019), diverging trends in nonmarital fertility by mothers' educational attainment (Bailey et al., 2013), and strong regional variation in access to contraceptive technology (Bailey, 2010). While these trends largely began earlier than the primary analysis period, a natural question is whether families in the South form a valid control group.

Reassuringly, Table A.3 shows that results dropping the South are quite similar. All unemployment coefficients for family structure are close in magnitude to the primary results and are significant at the 1 percent. The effect of AFDC-UP on welfare receipt is slightly smaller (14 p.p.) and AFDC-UP is similarly protective against cohabitation (p-value$<$0.01), living with children (p-value=0.038), and marriage (p-value=0.067). For separation, the degree of protectiveness is slightly lower and not precise (p-value=0.252). Taken together, there is little evidence that results are driven by differential trends in family structure or differential responsiveness of family structure to unemployment in the South.

V.B Pre-AFDC-UP comparisons in the 1960 Census

I also consider whether the preferred AFDC-UP variation implies protectiveness in the 1960 Census, before the implementation of AFDC-UP. While measures of family structure in 1960 are balanced across treated and non-treated states (Table 2), these results may mask heterogeneity by prior unemployment status. I use the sample for Table 2 (see Section II) and the primary treatment definition. Because there is no question regarding prior unemployment, I proxy for my preferred unemployment measure with having worked 26 weeks or less in the previous year. To mirror the primary analysis, I also drop those currently or previously unemployed for shorter durations.

The results for this exercise are shown in Table A.4. Consistent with the fact that no states had no-fault divorce laws in 1960 (Wolfers, 2006) and divorce may not always be the preference of both partners, the results indicate that unemployment is associated with much smaller increases (decreases) in separation (marriage) but is similarly associated with cohabitation and living with one's own children. The results also suggest that families of the unemployed in AFDC-UP states were *more likely* to dissolve in response to job loss. The AFDC-UP coefficients on separation and marriage are statistically significant and positive (negative), whereas the coefficients on cohabitation and living with one's children are not statistically distinguishable from zero (but are also opposite signed of the main results). This may be explained by the fact that AFDC-UP states had more generous welfare benefits (Table 2) and the long-standing correlation between single motherhood and welfare generosity (Moehling, 2007); however, this cannot explain why coefficients at higher terciles are also opposite signed and significant. If anything, these results reaffirm the importance of AFDC-UP in contributing to diverging trends in family structure responsiveness to unemployment across states that would and would not adopt AFDC-UP in the years to come.

V.C Robustness to treatment definition, covariates, and specification choice

Figures A.3-A.6 show that the coefficient magnitudes for welfare participation and the central family structure outcomes are all similar when including controls for race, fixed effects for age, and fixed effects for educational attainment bins (less than high school, high school, some college, and college or more), and are all significant at conventional levels. Results are also similar when including fixed effects for every occupation and industry represented in the sample and are generally *more precise*, complementing the stability of the DR estimates presented above. The figures also indicate that estimates of the more parsimonious model given in specification (1) — which only includes as right hand side variables dummies for AFDC-UP, unemployment, and their interaction — yields consistent results that are all significant at conventional levels. Figures A.3-A.6 also show that results are robust to instead using all 48 contiguous states and a treatment measure that reflects whether each state had the program in each year (e.g., defined by Figure B.1). Finally, these figures show that estimates using alternative methods to predict wage income for unemployed individuals, and stratified by the corresponding income terciles formulated from these predicted values, are comparable to the primary results.

V.D No-fault divorce laws

As a final validity check, I consider heterogeneity according to whether states had adopted no-fault divorce laws, which allow individuals to file for divorce without the consent of their spouse. Of the 42 states included in the primary sample, 23 had adopted no-fault divorce laws before 1977 and 17 had yet to adopt such laws by 1988 (Friedberg, 1998; Wolfers, 2006). Close to half (11) of the early no-fault divorce adopters are AFDC-UP states, whereas roughly two thirds (11) of the late-adopters are AFDC-UP states.

The results are shown in Table A.5 separately for early and late no-fault divorce adopters. The associations between unemployment and family structure are largely similar across both groups of states, with the marriage and divorce coefficients somewhat larger in early no-fault states and the cohabitation coefficient larger among later adopters. The magnitudes of AFDC-UP tend to mirror these differences, so that measures of protectiveness are similar. Many of the coefficients on unemployment and the interaction of unemployment and AFDC-UP are statistically significant, while many others are not. There is some evidence of uniformly more precise effects of AFDC-UP among earlier adopters, although small numbers of clusters in each sample complicates inference. Higher terciles show similar patterns as before of some evidence of heightened welfare receipt but little evidence of differential family structure responses. Taken together, there is perhaps some evidence that the presence of no-fault divorce laws led to a greater association between two-parent family dissolution and unemployment, but these laws cannot explain the degree of protectiveness found for AFDC-UP.

VI Discussion: AFDC-UP and trends in family structure by employment status

Figure 1 shows the long-run trend has been towards increasing rates of separation and declining rates of cohabitation, while also showing steeper trends in age-adjusted rates of separation among unemployed men. It is therefore reasonable to ask whether the increasing association between unemployment and separation would have been even larger in the absence of AFDC-UP, and whether the association of unemployment and noncohabitation would not have stabilized around 1960. The primary motivation for beginning the sample in 1977 is the availability of state identifiers; however, some states are identifiable in earlier years — typically those with larger populations — and for some groups of states treatment status is uniform within group. In Appendix B I describe the 12 states that adopted by 1968 or never adopted and which are also uniform in their adoption if they are part of a group.

Figure B.2, Panel C maps the sample.[58]

I consider whether the effects of AFDC-UP are stable when examining this restricted set of states between 1977-1988, and then test for stability of effects back to 1968. Because each regression is conducted on small numbers of clusters (10), I provide bootstrapped 2-sided and 1-sided p-values using the Wild Cluster Restricted (WCR) bootstrap (Cameron et al., 2008; MacKinnon et al., 2023).[59] Table 6, Panel A shows that results for 1977-1988 are largely similar as the primary results in Table 3, providing additional evidence of treatment effect homogeneity. Panel B shows that results extending the sample back to 1968 are also consistent. Across both periods, estimates are statistically significant at conventional levels for separation and marriage, but less precise for cohabitation and living with children.

The evidence thus suggests AFDC-UP provided protection against familial dissolution for at least two decades. In Panels A and B of Figure 6, I show age-adjusted rates of separation and cohabitation between 1940-2000 among employed and unemployed men in states that adopt AFDC-UP in the 1960s and states that never adopt through 1988, normalized by their 1960 value.[60] While trends among employed men look quite similar in both groups, significant differences exist among the unemployed. Panels C and D illustrate this, showing that the normalized difference in rates of cohabitation between unemployed and employed men in AFDC-UP states rise in the 1960s (after AFDC-UP is implemented) and fall only in the 1990s (after AFDC-UP was eliminated). In contrast, the central tendency is towards larger differences among control states. Panel D shows a tendency towards larger differences in

[58]There are two eras of different groupings in the CPS pre-1977: the first is between 1968-1972 (18 states individually identified) and the second is between 1973-1977 (12 states individually identified). The states includes as part of groups include Michigan and Wisconsin (treated) and Alabama and Mississippi (control). Panels A and B show the states in each era of CPS state groupings that satisfy the above criteria.

[59]I conduct 1,024 replications, which is the maximum number of draws with 10 clusters using the preferred Rademacher distribution to generate the "wild weights". I use the `boottest` command to implement the bootstrap procedure (Roodman et al., 2019).

[60]Rates are age-adjusted to resemble all men in labor force. Figure B.1 shows the years that each state had the program in effect for states that ever adopted by 1988. I include states that adopted the program in the 1960s and which dropped the program for at most 2 years, which mirrors the primary treatment definition for the main results.

separation among both groups of states, but also shows disproportionately larger differences in control states. The difference in these differences is somewhat smaller in magnitude than cohabitation, which is consistent with the smaller coefficients found for separation in the econometric analysis.

Estimating counterfactual rates is complicated by the fact that, while the results of Section IV shows the families of long-term unemployed men are very responsive to AFDC-UP, they also indicate that the families of men with higher incomes and shorter unemployment durations — who also differentially reported welfare receipt of a smaller magnitude — exhibit little response. The lack of detailed unemployment duration information in the Census complicates attempts to differentiate these groups in the aggregate series'. Nevertheless, the divergence in trends around the introduction of AFDC-UP, reversal in treated states after the removal of AFDC-UP, and the primary econometric results together suggest that AFDC-UP helped mitigate an even higher aggregate association between unemployment and family dissolution, and also helps to explain differentials in such associations across the states.

VII Conclusion

This paper estimates effects of an understudied public assistance program that provided a form of unemployment insurance for poor two-parent families on family structure stability. AFDC-UP operated in roughly half of the states for three and a half decades, forming the basis of the difference in differences research design that compares program participation and family structure among lower income men recently or currently experiencing unemployment versus those not and in AFDC-UP versus non-AFDC-UP states. Results show that AFDC-UP provided meaningful protection against long-term unemployment-associated separation and reductions in cohabitation that would otherwise have occurred among poor families. The large degree of responsiveness found is consistent with theoretical predictions regarding high marginal utility of the long-term unemployed, high replacement rates of AFDC-UP when accounting for categorical eligibility, long-standing geographic differences in AFDC-UP's availability, and with the removal of the outside option of separation as a means to receive AFDC-basic benefits, although the results do not distinguish between these factors.

The short-term effects of AFDC-UP on separation for men must translate into similar effects for women (Moffitt, 1992), which indirectly implies that female partners were more likely to head households with minor children in non-AFDC-UP states after their spouse became unemployed. Future research could explore the impact AFDC-UP had in mitigating even higher rates of single motherhood than observed since the 1960s, perhaps through the use of restricted data linkages to follow mothers during a period in which public longitudinal surveys are largely unavailable. This topic is of interest given the close connection between U.S. public assistance policy, public discourse, and rising rates of single motherhood over the last 60 years. Additionally, the reduced economic opportunities often attributed to children raised in single-parent families (e.g., McLanahan, 1988; Astone and McLanahan, 1991; Stevens, 1994; Hoynes, 1997; Heckman et al., 2006; Lee and McLanahan, 2015; Gould et al., 2020; Kearney, 2023) make this an important area of further investigation.

Policy proposals calling for federalizing UI benefits point out that even if earnings and work requirements are relaxed, efficacious social insurance against unemployment would still require a supplementary assistance program for the most needy unemployed (e.g., Dube (2021)). Lingering long-term unemployment in the wake of the Great Recession (Aaronson et al., 2010) suggests the importance of such a program may grow moving forward. The results in this paper suggest that the welfare gains of extending UI to these types of families may be quite large and have far reaching implications for how U.S. families respond to unemployment spells, particularly those of longer duration.

References

Aaronson, Daniel, Bhashkar Mazumder, and Shani Schechter, "What is Behind the Rise in Long-Term Unemployment?," *Economic Perspectives*, 2010, *34* (2).

Akerlof, George A., Janet L. Yellen, and Michael L Katz, "An Analysis of Out-of-Wedlock Childbearing in the United States," *The Quarterly Journal of Economics*, 1996, *111* (2), 277–317.

Anderson, Patricia M. and Bruce D Meyer, "Unemployment Insurance Takeup Rates and the After-Tax Value of Benefits," *The Quarterly Journal of Economics*, 1997, *112* (3), 913–937.

Armour, Philip, "The Role of Information in Disability Insurance Application: An Analysis of the Social Security Statement Phase-In," *American Economic Journal: Economic Policy*, 2018, *10* (3), 1–41.

Astone, Nan Marie and Sara S McLanahan, "Family structure, parental practices and high school completion," *American sociological review*, 1991, pp. 309–320.

Autor, David, David Dorn, and Gordon Hanson, "When Work Disappears: Manufacturing Decline and the Falling Marriage Market Value of Young Men," *American Economic Review: Insights*, 2019, *1* (2), 161–178.

Baicker, Katherine, Claudia Goldin, and Lawrence F Katz, "A Distinctive System: Origins and Impact of US Unemployment Compensation," in "The Defining Moment: The Great Depression and the American Economy in the Twentieth Century," University of Chicago Press, 1998, pp. 227–264.

Bailey, Martha J, "Momma's Got the Pill: How Anthony Comstock and Griswold v. Connecticut Shaped US Childbearing," *American Economic Review*, 2010, *100* (1), 98–129.

Bailey, Martha J., Melanie Guldi, and Brad J Hershbein, "Is There a Case for a "Second Demographic Transition"?," *Human Capital in History*, 2013, p. 273.

Becker, Gary S, *A Treatise on the Family*, Cambridge, MA: Harvard University Press, 1981.

Becker, Gary S., Elisabeth M. Landes, and Robert T Michael, "An Economic Analysis of Marital Instability," *Journal of Political Economy*, 1977, *85* (6), 1141–1187.

Bell, Winifred, *Aid to Dependent Children*, New York, Columbia University Press, 1965.

Belloni, Alexandre, Victor Chernozhukov, and Lie Wang, "Square-Root Lasso: Pivotal Recovery of Sparse Signals Via Conic Programming," *Biometrika*, 2011, *98* (4), 791–806.

Bertrand, Marianne, Emir Kamenica, and Jessica Pan, "Gender Identity and Relative Income Within Households," *The Quarterly Journal of Economics*, 2015, *130* (2), 571–614.

Black, Dan, Terra McKinnish, and Seth Sanders, "The Economic Impact of the Coal Boom and Bust," *The Economic Journal*, 2005, *115* (503), 449–476.

Black, Sandra E., Paul J. Devereux, and Kjell G Salvanes, "Losing Heart? The Effect of Job Displacement on Health," *ILR Review*, 2015, *68* (4), 833–861.

Blank, Rebecca M. and David E Card, "Recent Trends in Insured and Uninsured Unemployment: Is There an Explanation?," *The Quarterly Journal of Economics*, 1991, *106* (4), 1157–1189.

Blau, Francine D., Lawrence M. Kahn, and Jane Waldfogel, "Understanding Young Women's Marriage Decisions: The Role of Labor and Marriage Market Conditions," *ILR Review*, 2000, *53* (4), 624–647.

Blaustein, Saul J., Wilbur J. Cohen, and William Haber, *Unemployment Insurance in the United States: the First Half Century*, WE Upjohn Institute for Employment Research, 1993.

Buckles, Kasey, Melanie Guldi, and Joseph Price, "Changing the Price of Marriage Evidence from Blood Test Requirements," *Journal of Human Resources*, 2011, *46* (3), 539–567.

Bureau of Family Services, "Aid to Dependent Children of Unemployed Parents: The First Seven Months of Operation," 1962.

Bureau of Public Assistance, *Characteristics of General Assistance in the United States*, Security Administration. Washington D.C., 1959.

_ , *Illegitimacy and its Impact on the Aid to Dependent Children Program*, Social Security Administration. Washington D.C., 1960.

_ , *Advanced Release of Statistics of Public Assistance*, Social Security Administration Division of Statistics and Analysis. Washington D.C., 1961-1988.

_ , "Characteristics of General Assistance in the United States," 1969.

Callaway, Brantly and Pedro HC Sant'Anna, "Difference-in-Differences with Multiple Time Periods," *Journal of Econometrics*, 2021, *225* (2), 200–230.

Cameron, A. Colin, Jonah B. Gelbach, and Douglas L Miller, "Bootstrap-Based Improvements for Inference with Clustered Errors," *The Review of Economics and Statistics*, 2008, *90* (3), 414–427.

Carter, Susan B., Scott Sigmund Gartner, Michael R. Haines, Alan L. Olmstead, Richard Sutch, Gavin Wright, Price V. Fishback, and Melissa A. Thomasson, *Historical Statistics of the United States Millennial Edition Online*, New York: Cambridge University Press, 01 2006.

Charles, Kerwin Kofi and Melvin Stephens Jr, "Job Displacement, Disability, and Divorce," *Journal of Labor Economics*, 2004, *22* (2), 489–522.

_ **and Ming Ching Luoh**, "Male Incarceration, the Marriage Market, and Female Outcomes," *The Review of Economics and Statistics*, 2010, *92* (3), 614–627.

Chen, Jiahua and Zehua Chen, "Extended Bayesian Information Criteria for Model Selection with Large Model Spaces," *Biometrika*, 2008, *95* (3), 759–771.

Classen, Timothy J. and Richard A Dunn, "The Effect of Job Loss and Unemployment Duration on Suicide Risk in the United States: A New Look Using Mass-Layoffs and Unemployment Duration," *Health Economics*, 2012, *21* (3), 338–350.

Coe, Norma and April Yanyuan Wu, "What Impact Does Social Security Have on the Use of Public Assistance Programs Among the Elderly?," *CRR WP*, 2014, *5.*

Cohen, Wilbur J., *Some Major Policy Issues in Unemployment Insurance and General Assistance*, US Government Printing Office, 1960.

_ **, William Haber, and Eva Mueller**, *The Impact of Unemployment in the 1958 Recession: A Report on Nationwide Surveys of Unemployment, Unemployment Insurance, and Attitudes of the Unemployed.*, Washington D.C.: US Government Printing Office, 1960.

Cullen, Julie B. and Jonathan Gruber, "Does Unemployment Insurance Crowd out Spousal Labor supply?," *Journal of Labor Economics*, 2000, *18* (3), 546–572.

Currie, Janet and Firouz Gahvari, "Transfers in Cash and in-Kind: Theory Meets the Data," *Journal of Economic Literature*, 2008, *46* (2), 333–383.

Curtin, Sally C. and Gladys M Martinez, "Recent Declines in Nonmarital Childbearing in the United States," 2014, (162).

Deb, Partha, William T. Gallo, Padmaja Ayyagari, Jason M. Fletcher, and Jody L Sindelar, "The Effect of Job Loss on Overweight and Drinking," *Journal of Health Economics*, 2011, *30* (2), 317–327.

Deshpande, Manasi and Yue Li, "Who is Screened Out? Application Costs and the Targeting of Disability Programs," *American Economic Journal: Economic Policy*, 2019, *11* (4), 213–248.

Doiron, Denise and Silvia Mendolia, "The Impact of Job Loss on Family Dissolution," *Journal of Population Economics*, 2012, *25*, 367–398.

Dube, Arindrajit, "A Plan to Reform the Unemployment Insurance System in the United States," *The Hamilton Project*, 2021.

Duggan, Mark, Perry Singleton, and Jae Song, "Aching to Retire? The Rise in the Full Retirement Age and its Impact on the Social Security Disability Rolls," *Journal of Public Economics*, 2007.

East, Chloe N. and Elira Kuka, "Reexamining the Consumption Smoothing Benefits of Unemployment Insurance," *Journal of Public Economics*, 2015, *132*, 32–50.

Edwards, Kathryn Anne, "Who Helps the Unemployed? Workers' Receipt of Public and Private Transfers," *IZA Journal of Labor Economics*, 2020, *9* (1).

Edwards, Roxanna and Sean M Smith, "Job Market Remains Tight in 2019, as the Unemployment Rate Falls to its Lowest Level Since 1969," *Monthly Labor Review*, 2020, *143*, 1.

Eliason, Marcus, "Lost Jobs, Broken Marriages," *Journal of Population Economics*, 2012, *25* (4), 1365–1397.

_ **and Donald Storrie**, "Does Job Loss Shorten Life?," *Journal of Human Resources*, 2009, *44* (2), 277–302.

Ellwood, David T. and Mary J. Bane, "The Impact of AFDC on Family Structure and Living Arrangements," in "Research in Labor Economics," Greenwich, CT: JAI Press, 1985.

Fetter, Daniel and Matthew Pesner, "The Expansion of Social Security and the Decline of Elderly Public Assistance," Technical Report, University of Michigan 2021.

Finkelstein, Amy and Matthew J Notowidigdo, "Take-Up and Targeting: Experimental Evidence from SNAP," *The Quarterly Journal of Economics*, 2019, *134* (3), 1505–1556.

_ **and Robin McKnight**, "What did Medicare do? The Initial Impact of Medicare on Mortality and out of Pocket Medical Spending," *Journal of Public Economics*, 2008, *92* (7), 1644–1668.

_ **, Nathaniel Hendren, and Erzo FP Luttmer**, "The Value of Medicaid: Interpreting Results from the Oregon Health Insurance Experiment," *Journal of Political Economy*, 2019, *127* (6), 2836–2874.

Finman, Ted, "OEO Legal Service Programs and the Pursuit of Social Change: The Relationship Between Program Ideology and Program Performance," *Wisconsin Law Review*, 1971, p. 1001.

Flood, Sarah, Miriam King, Steven Ruggles, Robert J. Warren, and Michael Westberry, "Ipums CPS: Version 10.0 [dataset]," *Minneapolis, MN: IPUMS*, 2022, *10*, D010.

Friedberg, Leora, "Did Unilateral Divorce Raise Divorce Rates? Evidence from Panel Data," *American Economic Review*, 1998.

Fuller, Sheridan, "New Deal, Same Compromise? Measuring the Effect of Racially Linked Welfare Policies on Program Participation and Children's Development," *Unpublished Manuscript*, 2022.

Ganong, Peter and Pascal Noel, "Consumer Spending During Unemployment: Positive and Normative Implications," *American Economic Review*, 2019, *109* (7), 2383–2424.

_ , Fiona Greig, Pascal Noel, Daniel M. Sullivan, and Joseph Vavra, "Lessons Learned from Expanded Unemployment Insurance during COVID-19," in Wendy Edelberg, Louise Sheiner, and David Wessel, eds., *Recession Remedies: Lessons Learned from the US Economic Policy Response to COVID-19*, Hutchins Center on Fiscal and Monetary Policy at Brookings, The Hamilton Project, 2022.

Goldin, Claudia and Lawrence F Katz, "The Power of the Pill: Oral Contraceptives and Women's Career and Marriage Decisions," *Journal of Political Economy*, 2002, *110* (4), 730–770.

Goodman-Bacon, Andrew, "Public Insurance and Mortality: Evidence from Medicaid Implementation," *Journal of Political Economy*, 2018, *126* (1), 216–262.

_ , "Difference-in-Differences with Variation in Treatment Timing," *Journal of Econometrics*, 2021, *225* (2), 254–277.

_ , "Difference-in-Differences without Time," Technical Report, Federal Reserve Bank of Minneapolis 2023.

_ **and Jamein P Cunningham**, "Changes in Family Structure and Welfare Participation Since the 1960s: The Role of Legal Services," Technical Report, National Bureau of Economic Research 2019.

_ **and Lucie Schmidt**, "Federalizing benefits: The introduction of Supplemental Security Income and the size of the safety net," *Journal of Public Economics*, 2020, *185*, 104174.

Gornick, Marian, Jay N. Greenberg, Paul W. Eggers, and Allen Dobson, "Twenty Years of Medicare and Medicaid: Covered Populations, Use of Benefits, and Program Expenditures," *Health Care Financing Review*, 1985, *1985* (Suppl), 13.

Gould, Eric D., Avi Simhon, and Bruce A Weinberg, "Does parental quality matter? Evidence on the transmission of human capital using variation in parental influence from death, divorce, and family size," *Journal of Labor Economics*, 2020, *38* (2), 569–610.

Groeneveld, Lyle, Michael Hannan, and Nancy Tuma, "Income and Marital Events: Review of Previous Research," in "Final Report of the Seattle-Denver Income Maintenance Experiment," Vol. 1 1983.

Gruber, Jonathan and Brigitte C Madrian, "Employment Separation and Health Insurance Coverage," *Journal of Public Economics*, 1997, *66* (3), 349–382.

Guldi, Melanie, "Fertility Effects of Abortion and Birth Control Pill Access for Minors," *Demography*, 2008, *45* (4), 817–827.

Gundersen, Craig and James P Ziliak, "The Role of Food Stamps in Consumption Stabilization," *Journal of Human Resources*, 2003, pp. 1051–1079.

Heckman, James J., Jora Stixrud, and Sergio Urzua, "The effects of cognitive and noncognitive abilities on labor market outcomes and social behavior," *Journal of Labor Economics*, 2006, *24* (3), 411–482.

Herd, Pamela and Donald P Moynihan, *Administrative Burden: Policymaking by Other Means*, Russell Sage Foundation, 2019.

Horvath, Francis W, "Forgotten Unemployment: Recall Bias in Retrospective Data," *Monthly Labor Review*, 1982, *105* (3), 40–43.

Hoynes, Hilary W, "Welfare Transfers in Two-Parent Families: Labor Supply and Welfare Participation under AFDC-UP," *Econometrica*, 1996, *64* (2), 295–332.

Hoynes, Hilary W., "Does Welfare Play Any Role in Female Headship Decisions?," *Journal of Public Economics*, 1997, *65* (2), 89–117.

Hsu, Joanne W., David A. Matsa, and Brian T Melzer, "Unemployment Insurance as a Housing Market Stabilizer," *American Economic Review*, 2018, *108* (1), 49–81.

Jacobson, Louis S., Robert J. LaLonde, and Daniel G Sullivan, "Earnings Losses of Displaced Workers," *The American Economic Review*, 1993, pp. 685–709.

Jr., Melvin Stephens, "Worker Displacement and the Added Worker Effect," *Journal of Labor Economics*, 2002, *20* (3), 504–537.

Kearney, Melissa S, *The Two-Parent Privilege: How Americans Stopped Getting Married and Started Falling Behind*, University of Chicago Press, 2023.

Kearney, Melissa S. and Riley Wilson, "Male Earnings, Marriageable Men, and Nonmarital Fertility: Evidence from the Fracking Boom," *Review of Economics and Statistics*, 2018, *100* (4), 678–690.

Kearney, Melissa Schettini, "Is There an Effect of Incremental Welfare Benefits on Fertility Behavior? A Look at the Family Cap," *Journal of Human Resources*, 2004, *39* (2), 295–325.

Kennedy, John F., "Special Message to the Congress: Program for Economic Recovery and Growth, February 2, 1961," in "The Cumulated indexes to the public papers of the Presidents of the United States, John F. Kennedy, 1961-1963," 156 p. Millwood, N.Y.: KTO Press, 1977.

Kuhn, Andreas, Rafael Lalive, and Josef Zweimüller, "The Public Health Costs of Job Loss," *Journal of Health Economics*, 2009, *28* (6), 1099–1115.

Kuka, Elira, "Quantifying the Benefits of Social Insurance: Unemployment Insurance and Health," *Review of Economics and Statistics*, 2020, *102* (3), 490–505.

_ **and Bryan A Stuart**, "Racial Inequality in Unemployment Insurance Receipt and Take-Up," Technical Report, National Bureau of Economic Research 2021.

Lansdale, Robert T, "The Unemployed Parent Segment of AFDC: Category Within a Category," *Social Work*, 1967, *12* (1), 40–50.

Lee, Dohoon and Sara McLanahan, "Family Structure Transitions and Child Development: Instability, Selection, and Population Heterogeneity," *American sociological review*, 2015, *80* (4), 738–763.

Leung, Pauline and Christopher O'leary, "Unemployment Insurance and Means-Tested Program Interactions: Evidence from Administrative data," *American Economic Journal: Economic Policy*, 2020, *12* (2), 159–192.

Lindo, Jason M., Krishna Regmi, and Isaac Swensen, "Stable Income, Stable Family," Technical Report, National Bureau of Economic Research 2022.

Lundberg, Shelly, "The Added Worker Effect," *Journal of Labor Economics*, 1985, *3* (1, Part 1), 11–37.

_ **and Robert A Pollak**, "Bargaining and Distribution in Marriage," *Journal of Economic Perspectives*, 1996, *10* (4), 139–158.

MacKinnon, James G., Morten Ørregaard Nielsen, and Matthew D Webb, "Cluster-Robust Inference: A Guide to Empirical Practice," *Journal of Econometrics*, 2023, *232* (2), 272–299.

McLanahan, Sara, Glen Cain, Michael Olneck, Irving Piliavin, Sheldon Danziger, and Peter Gottschalk, "Losing Ground: A Critique," 1985.

McLanahan, Sara S, "Family structure and dependency: Early transitions to female household headship," *Demography*, 1988, *25* (1), 1–16.

Mercatanti, Andrea and Fan Li, "Do Debit Cards Increase Household Spending? Evidence from a Semiparametric Causal Analysis of a Survey," *The Annals of Applied Statistics*, 2014, pp. 2485–2508.

Meyer, Bruce D., Wallace KC Mok, and James X Sullivan, "Household Surveys in Crisis," *Journal of Economic Perspectives*, 2015, *29* (4), 199–226.

Michael, R Haines, "Inter-University Consortium for Political and Social Research. Historical, Demographic, Economic, and Social Data: The United States, 1790–2002," *Ann Arbor, MI: Inter-university Consortium for Political and Social Research [distributor]*, 2010, pp. 05–21.

Moehling, Carolyn M, "The American Welfare System and Family Structure An Historical Perspective," *Journal of Human Resources*, 2007, *42* (1), 117–155.

Moffitt, Robert A, "Incentive Effects of the US Welfare System: A Review," *Journal of Economic Literature*, 1992, *30* (1), 1–61.

_ , "Welfare Effects on Female Headship with Area Effects," *Journal of Human Resources*, 1994, pp. 621–636.

_ , "Welfare, the Family, and Reproductive Behavior: Research Perspectives," 1998.

_ , "The Temporary Assistance for Needy Families Program," *Means-tested transfer programs in the United States*, 2003, pp. 291–364.

Moffitt, Robert A., Robert Reville, and Anne E Winkler, "Beyond Single Mothers: Cohabitation and Marriage in the AFDC Program," *Demography*, 1998, *35* (3), 259–278.

Murray, Charles, "Losing Ground: American Social Policy, 1950–1980," 1984.

Nichols, Austin, Josh Mitchell, and Stephan Lindner, "Consequences of Long-Term Unemployment," *Washington, DC: The Urban Institute*, 2013.

of Health, U.S. Department and Human Services (U.S. DHHS), "Aid to Families with Dependent Children: the Baseline," *Washington, DC: Office of the Assistant Secretary for Planning and Evaluation*, 1998.

Price, Daniel N, "Unemployment Insurance, Then and Now, 1935-85," *Social Security Bulletin*, 1985, *48*, 22.

Roodman, David, Morten Ørregaard Nielsen, James G. MacKinnon, and Matthew D Webb, "Fast and Wild: Bootstrap Inference in Stata Using boottest," *The Stata Journal*, 2019, *19* (1), 4–60.

Rosenzweig, Mark R, "Welfare, Marital Prospects, and Nonmarital Childbearing," *Journal of Political Economy*, 1999, *107* (S6), S3–S32.

Ross, Heather L., Isabel V. Sawhill, and Anita R MacIntosh, *Time of Transition The Growth of Families Headed by Women*, The Urban Insitute, 1975.

Rossin-Slater, Maya, "WIC in your Neighborhood: New Evidence on the Impacts of Geographic Access to Clinics," *Journal of Public Economics*, 2013, *102*, 51–69.

Rothstein, Jesse and Robert G Valletta, "Scraping by: Income and program participation after the loss of extended unemployment benefits," *Journal of Policy Analysis and Management*, 2017, *36* (4), 880–908.

Ruggles, Steven, Sarah Flood, Matthew Sobek, Danika Brockman, Grace Cooper, Stephanie Richards, and Megan Schouweiler, "Ipums usa: Version 13.0 [dataset]," *Minneapolis, MN: IPUMS*, 2023, *10*, D010.

Sandler, Danielle H. and Ryan Sandler, "Multiple Event Studies in Public Finance and Labor Economics: A Simulation Study with Applications," *Journal of Economic and Social Measurement*, 2014, *39* (1-2), 31–57.

Schaller, Jessamyn and Ann Huff Stevens, "Short-Run Effects of Job Loss on Health Conditions, Health Insurance, and Health Care Utilization," *Journal of Health Economics*, 2015, *43*, 190–203.

Schmidt, Lucie and Purvi Sevak, "AFDC, SSI, and Welfare Reform Aggressiveness: Caseload Reductions versus Caseload Shifting," *Journal of Human resources*, 2004, *39* (3), 792–812.

Schram, Sanford and Michael Wiseman, "Should Families be Protected from AFDC-UP," *University of Wisconsin-Madison Institute for Research on Poverty Discussion Papers*, 1988.

Schultz, T. Paul, "Marital Status and Fertility in the United States: Welfare and Labor Market Effects," *Journal of Human Resources*, 1994, pp. 637–669.

Shenhav, Na'ama, "Lowering Standards to Wed? Spouse Quality, Marriage, and Labor Market Responses to the Gender Wage Gap," *Review of Economics and Statistics*, 2021, *103* (2), 265–279.

Skandalis, Daphné, Ioana Marinescu, and Maxim N Massenkoff, "Racial Inequality in the US Unemployment Insurance System," Technical Report, National Bureau of Economic Research 2022.

Stevens, Ann Huff, "The Dynamics of Poverty Spells: Updating Bane and Ellwood," *The American Economic Review*, 1994, *84* (2), 34–37.

Stuart, Bryan A, "The Long-Run Effects of Recessions on Education and Income," *American Economic Journal: Applied Economics*, 2022, *14* (1), 42–74.

Sullivan, Daniel and Till Von Wachter, "Job Displacement and Mortality: An Analysis Using Administrative Data," *The Quarterly Journal of Economics*, 2009, *124* (3), 1265–1306.

Thompson, Owen, "Fertility Decline in the Civil Rights Era," Technical Report, National Bureau of Economic Research 2019.

Tibshirani, Robert, "Regression Shrinkage and Selection Via the Lasso," *Journal of the Royal Statistical Society: Series B. (Methodological)*, 1996, *58* (1), 267–288.

U.S. Advisory Council on Public Assistance, *Public Assistance. A Report of the Findings and Recommendations of the Advisory Council on Public Assistance*, Washington D.C.: U.S. Government Printing Office, 1960.

U.S. Department of Agriculture Food and Nutrition Service (USDA), *Supplemental Nutrition Assistance Program Participation and Costs*, Washington D.C., 2023.

U.S. Department of Health and Human Services (DHHS), *Unemployment Insurance*, Social Security Administration. Washington D.C., 2012.

U.S. Department of Labor, "ET Financial Data Handbook 394 [dataset]," https://oui.doleta.gov/unemploy/hb394.asp 2019.

U.S. General Accounting Office, "Welfare Reform: Projected Effects of Requiring AFDC for Unemployed Parents Nationwide," 1988.

U.S. National Center for Social Statistics (U.S. NCSS), "Findings of the 1973 AFDC study. Part I. Demographic and program characteristics.," 1975.

_ , "Findings of the 1973 AFDC study. Part IV. Discontinuances for AFDC money payments during 1973.," 1975.

U.S. Office of the Assistant Secretary for Planning and Evaluation (U.S. ASPE), "Background Material and Data on Programs Within the Jurisdiction of the Committee on Ways and Means. Section 8. Aid to Families with Dependent Children and Related Programs (Title IV-A)," 1985.

_ , *1996 Green Book. Section 8. Aid to Families with Dependent Children and Related Programs (Title IV-A)*, U.S. Government Printing Office, Washington D.C., 1996.

Ventura, Stephanie J. and Christine A Bachrach, "Nonmarital Childbearing in the United States, 1940-99.," *National Vital Statistics Reports*, 2000, *48* (16), n16.

Weiss, Yoram and Robert J Willis, "Match Quality, New Information, and Marital Dissolution," *Journal of Labor Economics*, 1997, *15* (1, Part 2), S293–S329.

Wexler, Sandra and Rafael J Engel, "Historical Trends in State-Level ADC/AFDC Benefits: Living on Less and Less," *Journal of Sociology and Social Welfare*, 1999, *26*, 37.

Willis, Robert J, "A Theory of Out-of-Wedlock Childbearing," *Journal of Political Economy*, 1999, *107* (S6), S33–S64.

Wilson, William J., *When Work Disappears: The World of the New Urban Poor*, New York: Knopf, 1996.

_ , *The Truly Disadvantaged: The Inner City, the Underclass, and Public Policy*, University of Chicago Press, 2012.

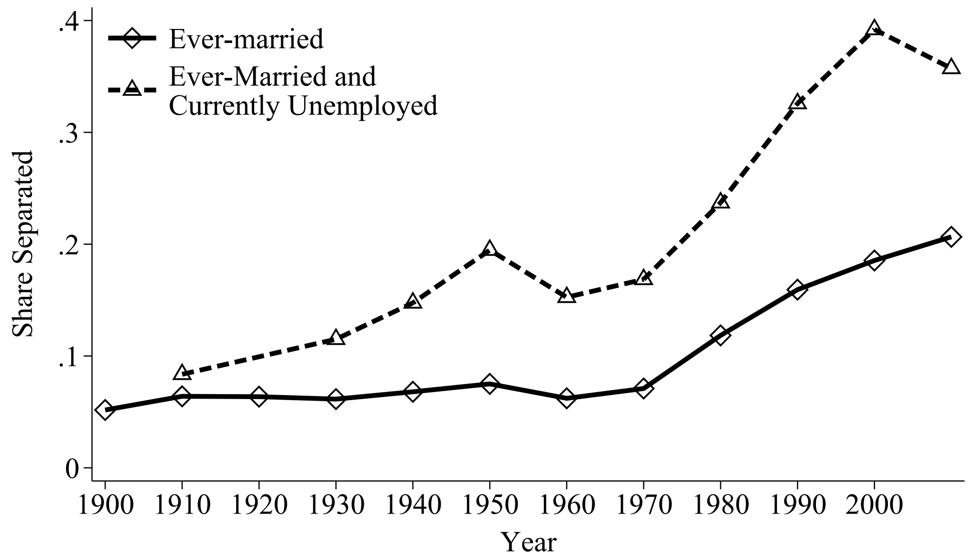
Winkler, Anne E, "Does AFDC-UP Encourage Two-Parent Families?," *Journal of Policy Analysis and Management*, 1995, *14* (1), 4–24.

Wolfers, Justin, "Did Unilateral Divorce Laws Raise Divorce Rates? A Reconciliation and New Results," *American Economic Review*, 2006, *96* (5), 1802–1820.

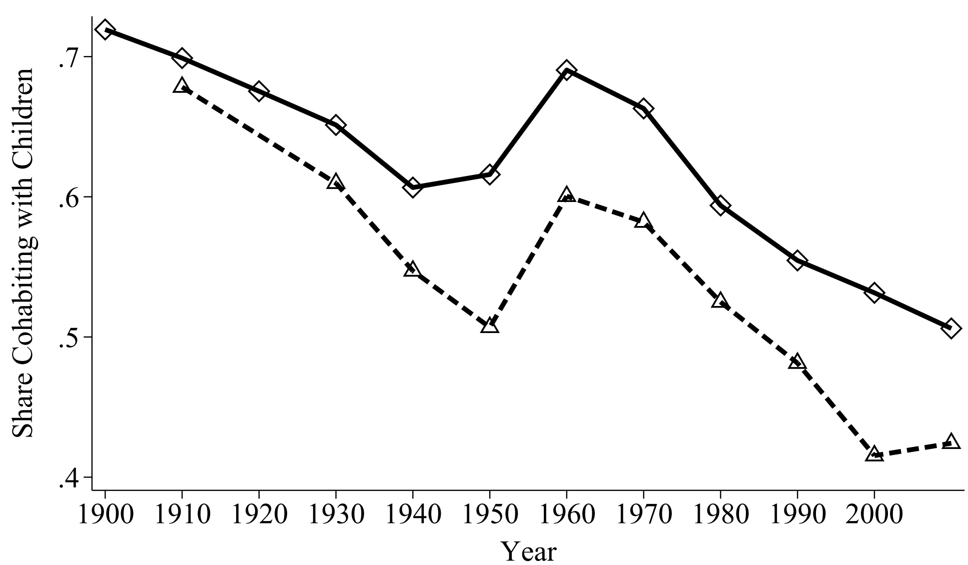
Figures and Tables

Figure 1: Trends in separation and cohabitation among men, by current employment status

A. Fraction divorced, separated, or spouse absent



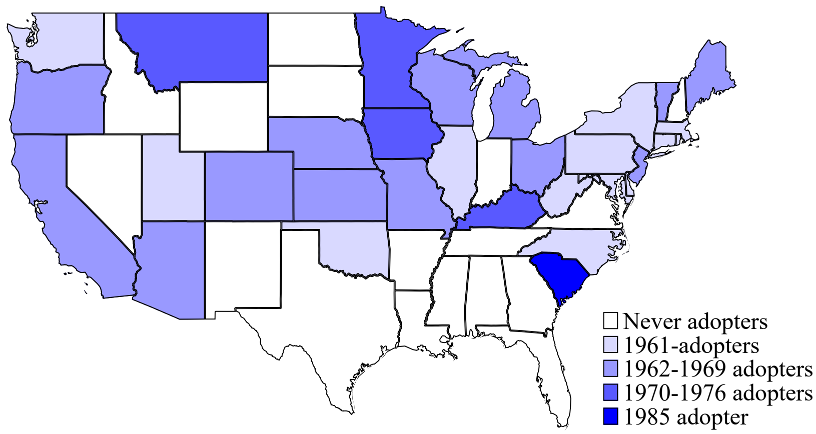
B. Fraction cohabiting with minor children and their mother



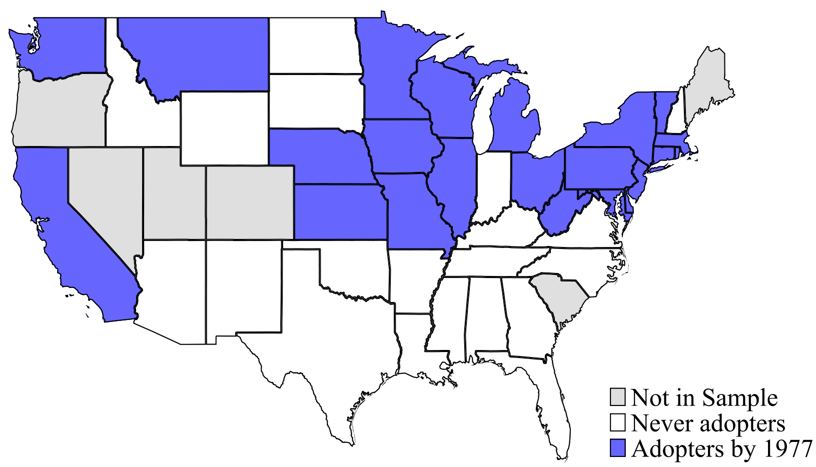
Notes: Data come from 1 percent decennial census samples (Ruggles et al., 2023) and are comprised of ever-married men ages 24-58 who are in the labor force. Panel A plots the share who are divorced, separated, or report spousal absence (black solid line with diamonds) and the age-adjusted share who are currently unemployed and divorced, separated, or report spousal absence (black dashed line with triangles). Panel B plots the share who are cohabiting with children, defined as living with their own children under age 18 and the mother of at least one of those children (black solid line with diamonds); and the age-adjusted share who are currently unemployed and cohabiting (black dashed line with triangles). Age-adjustment is performed using the age distribution in each year among ever-married men ages 24-58 who are in the labor force.

Figure 2: Geographic and timing variation in first-time AFDC-UP adoption

A. All pre-1990 first-time adoption

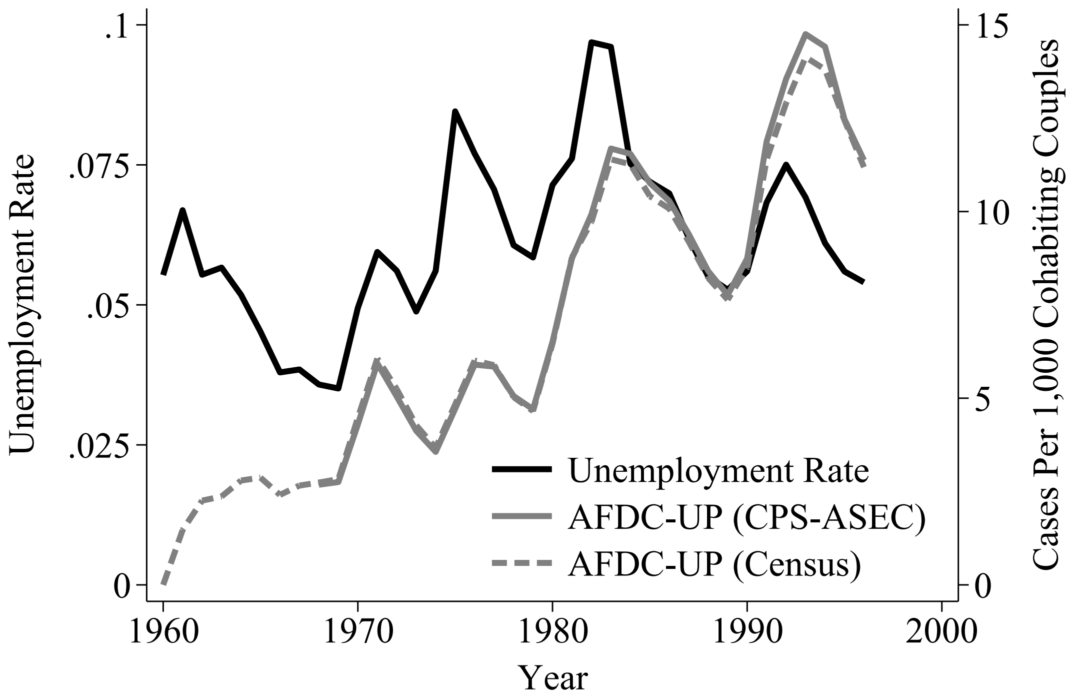


B. CPS-ASEC definition



Notes: Panel A maps states in the contiguous United States that ever adopted AFDC-UP by 1988, by the following eras of first-time adoption: 1961, the first year of AFDC-UP; 1962-1969; 1970-1976; and after (the only state is South Carolina, in 1985). First-time adoption is defined as the first year in which caseloads were positive in any month. Data on AFDC-UP caseloads by state and month are from (Bureau of Public Assistance, various years). Panel B plots the 42 states that are classified as always or never treated (see Section II and Appendix B more details).

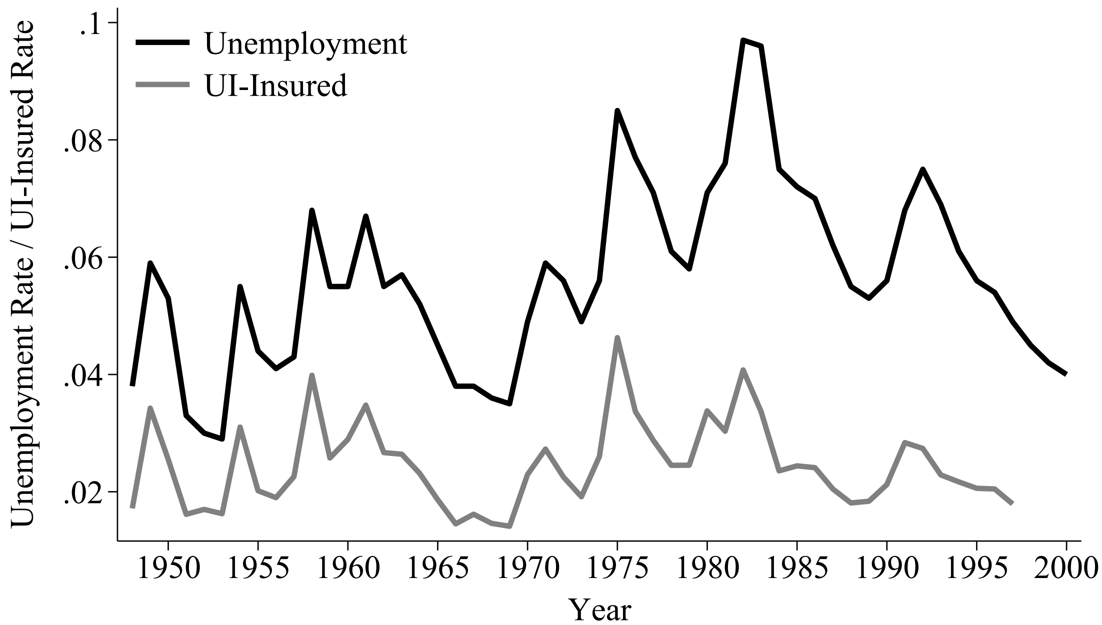
Figure 3: Trends in Unemployment and AFDC-UP Caseloads



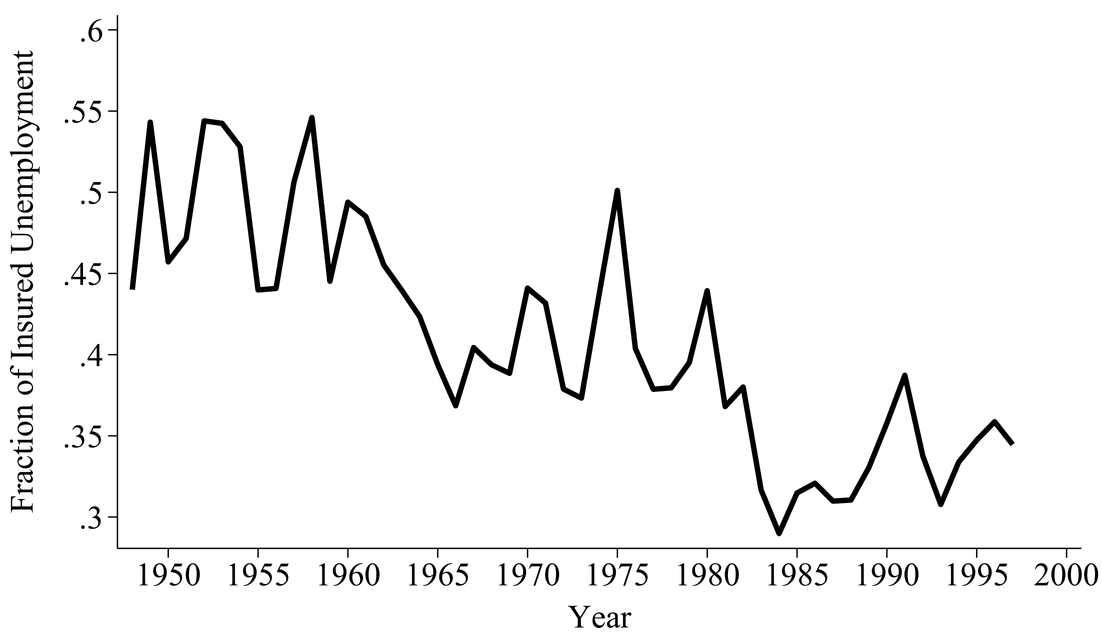
Notes: Plots the national unemployment rate (left axis, black solid line), the number of caseloads per thousand cohabiting couples as measured in the CPS-ASEC (right axis, gray solid line), and the number of caseloads per thousand cohabiting couples as measured in the decennial census (right axis, gray dashed line). Cohabiting couples are defined as heterosexual couples who both report living with their own children with at least one of their children under 18, further restricted to households in which there is only one such couple. Data on the national unemployment rate are from Carter et al. (2006) Series Ba478-486. Monthly data on AFDC-UP caseloads come from (Bureau of Public Assistance, various years) and are collapsed by year. CPS-ASEC data are from Flood et al. (2022) and decennial census data are from Ruggles et al. (2023). The CPS-ASEC series begins in 1968, as this was the first year that the age of the youngest child in the household was collected. Cohabiting couple counts in the decennial census series are linearly interpolated in intercensal years. Census samples are 5 percent for 1960 and 1980-2000, and 1 percent for 1970.

Figure 4: Trends in unemployment and UI

A. The unemployment rate and UI-insured rate

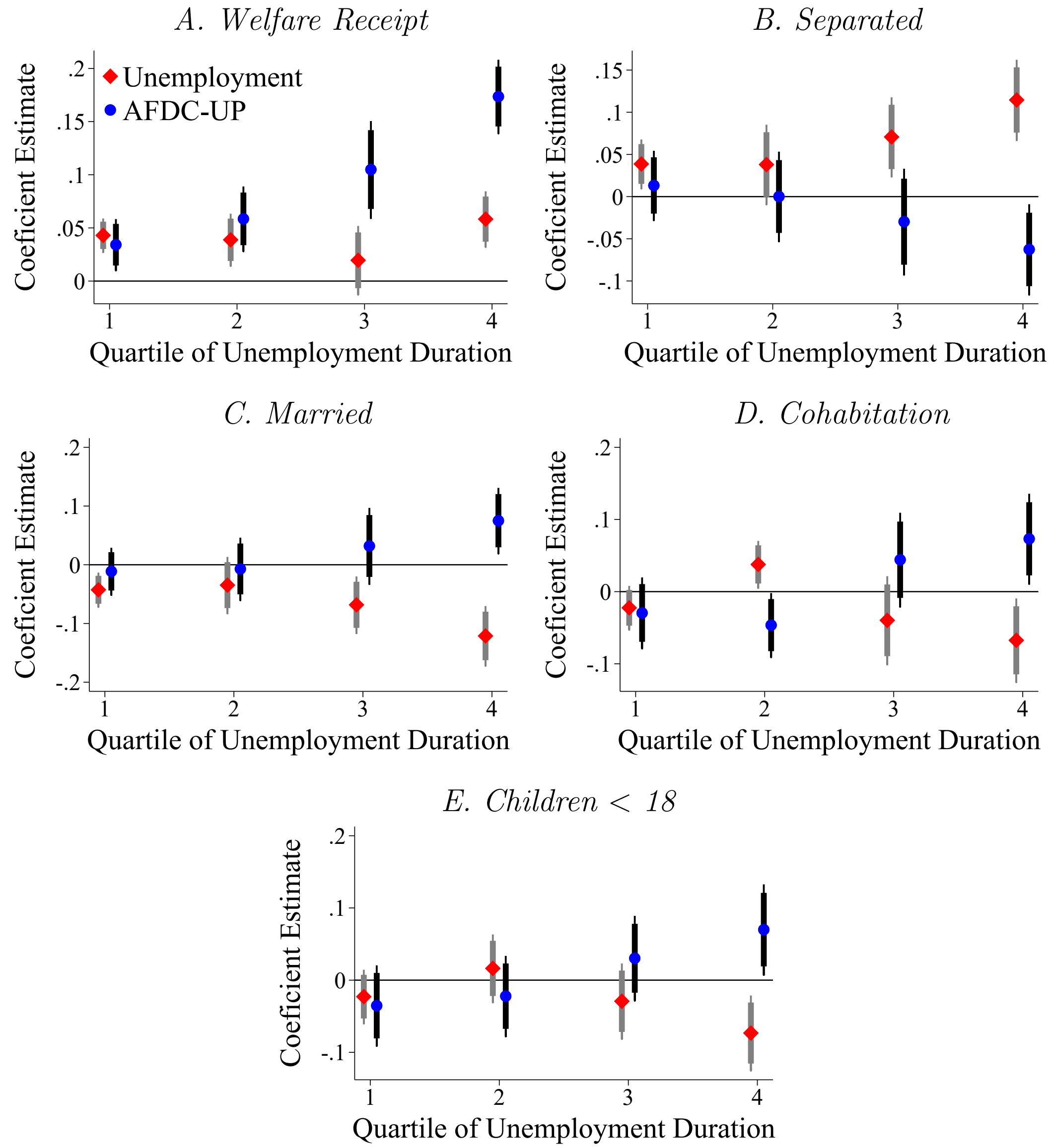


B. Fraction of Insured Unemployment



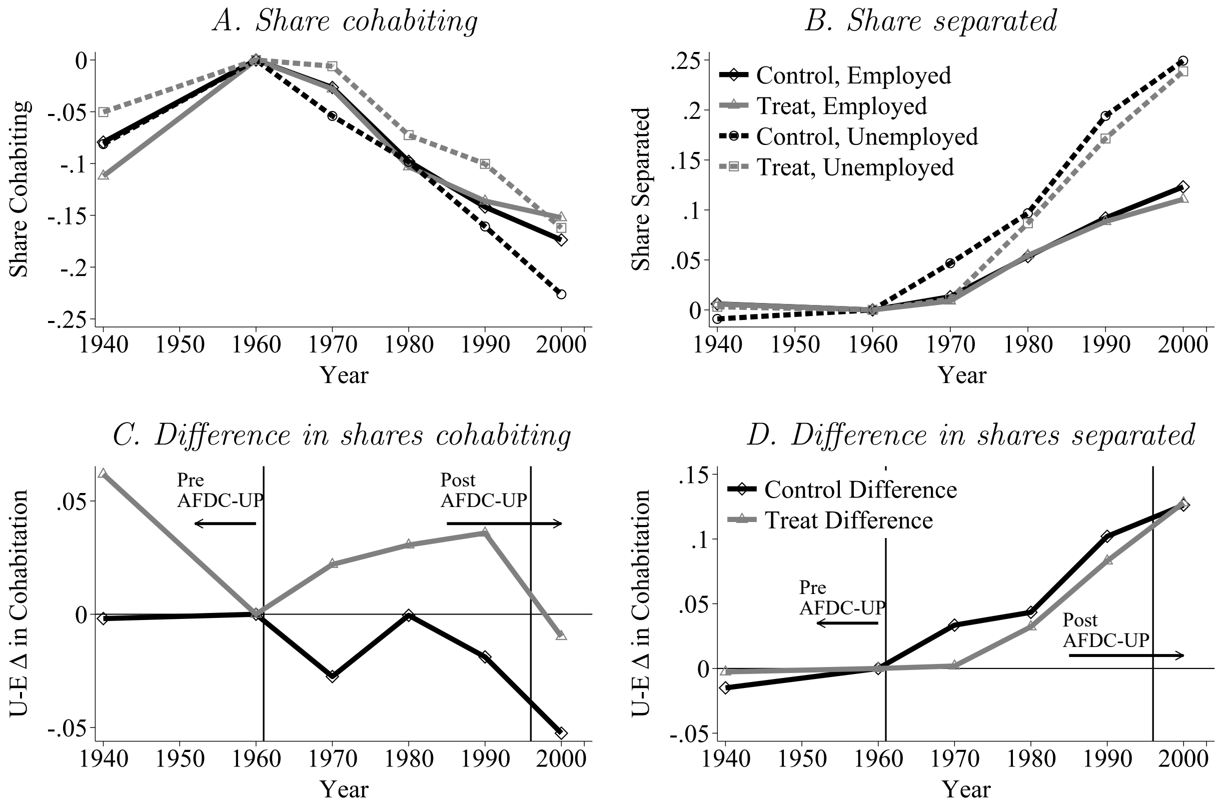
Notes: Panel A plots the national unemployment rate (black line; Carter et al., 2006, series Ba485) and the national UI-insured rate (gray line; Carter et al., 2006, series Bf485), defined as the average number of people weekly insured for UI divided by the number of people in the labor force. Panel B plots the Fraction of Insured Unemployment (FIU; Blank and Card, 1991), which is the ratio of the national UI-insured rate to the unemployment rate.

Figure 5: Heterogeneity by unemployment duration



Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$, red diamonds with gray bars) and unemployment ($\widehat{\gamma}_1$, blue circles with black bars) from specification (4) among men in the lowest income tercile, separately by outcome (given in panels A-E) and by quartile of unemployment duration, with 90 percent and 95 percent cluster robust (CR) confidence intervals at the state level represented by thick and thin bars, respectively. Samples pool March CPS-ASEC data between 1977-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. The quartiles of prior-year unemployment are 1-6, 7-12, 13-23, and 24-51 weeks. For consistency, these bins are also applied to current unemployment duration. For the four quartiles of duration, each sample drops those reporting duration outside of the corresponding range of included unemployment duration (as with the primary sample), so all regression coefficients are relative to the same group of employed workers. Terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. All specifications use CPS-ASEC survey weights.

Figure 6: Trends in separation and cohabitation among men, by current employment and by treatment status



Notes: Data come from 1 percent decennial census samples (Ruggles et al., 2023) between 1940-2000 and are comprised of ever-married men ages 24-58 who are in the labor force. Trends are presented separately for employed and unemployed men in treated and control states. All trends are normalized by their 1960 value. Employed men are defined as those not currently employed and who also did not report any unemployment in the prior year. Unemployed men are those reporting current unemployment. Data in 1950 for prior unemployment are largely missing, so that census is not included. Treated states are those that adopted AFDC-UP in the 1960s and dropped the program for at most two years through 1988 (See Figure B.1) and control states are those that never adopt the program by 1988. All trends are age-adjusted using the age distribution in each year among ever-married men ages 24-58 who are in the labor force. Panel A plots the share who are cohabiting with children by treatment and employment status. Panel B plots these shares for separation. Panel C plots the differences between unemployed and employed trends in Panel A, and Panel D plots the corresponding differences for Panel B.

Table 1: Summary statistics

	(1) Overall	(2) Tercile 1	(3) Tercile 2	(4) Tercile 3
Demographic				
Share married	.808	.794	.802	.828
Share divorced, separated, or spouse absent	.182	.194	.186	.165
Age	39	36.7	40.2	40.2
Share white	.912	.82	.957	.96
Share with children under 18	.635	.645	.609	.649
Share cohabiting with children under 18	.676	.684	.654	.691
Share completed less than highschool	.194	.459	.105	.013
Share completed highschool	.362	.439	.521	.123
Share completed some college	.175	.095	.281	.149
Share completed 4 year college or more	.269	.007	.093	.715
Economic				
Wage income (52-week workers)	\$72,658	\$50,385	\$69,114	\$99,122
Wage income (predicted)	\$73,149	\$53,716	\$69,597	\$96,702
Household income	\$78,801	\$53,414	\$74,358	\$109,371
Share unemployed $\geq$ 26 weeks	.037	.068	.032	.011
Share unemployed $\geq$ 26 weeks last year	.034	.062	.029	.01
Share currently unemployed $\geq$ 26 weeks	.008	.013	.007	.003
Duration $\mid$ Unemployed $\geq$ 26 weeks last year	32.4	32.3	32.5	32.7
Duration $\mid$ Currently unemployed $\geq$ 26 weeks	40.5	39.9	41.3	41.3
Share below the poverty line	.047	.104	.025	.01
Share own home	.699	.56	.738	.804
Share with employer-based group coverage	.877	.795	.908	.929
Program participation				
Share receiving welfare (family)	.013	.029	.008	.002
Welfare income $\mid$ welfare receipt	\$5,106	\$5,483	\$3,947	\$4,452
Share receiving UI	.042	.068	.043	.013
Share receiving UI $\mid$ unemployed $\geq$ 26 weeks	.574	.557	.615	.559
Share receiving foodstamps (family)	.036	.082	.022	.005
foodstamp value $\mid$ foodstamp receipt	\$2,809	\$2,967	\$2,297	\$2,400
Share covered by Medicaid (family)	.023	.048	.014	.005
Observations (1977–1988)	66,518	23,361	22,044	21,113
Observations (1980–1988)	47,906	16,792	15,893	15,221

Notes: Sample pools March CPS-ASEC data between 1977-1988 (Flood et al., 2022) and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of \$1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Column 1 shows statistics for the full sample, while columns 2-4 show statistics separately by tercile of the previous year's predicted personal wage and salary income. Earnings are estimated using personal wage and salary income among men who worked 52 weeks in the previous year and predicted for those who worked less than 52 weeks, using the square-root Least Absolute Shrinkage and Selection Operator (LASSO; Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. Statistics on on health insurance, food stamps, and Medicaid are from 1980-1988. Monetary variables are in 2019 dollars. All statistics are calculated using CPS-ASEC survey weights.

Table 2: 1960 balance

	(1) AFDC-UP States	(2) Non-AFDC-UP States	(3) Difference	(4) p-value
Demographic				
Age	38.9	38.2	.638	< .01
Share white	.946	.902	.044	< .01
Share married	.951	.957	-.006	.239
Age at first marriage	24.9	23.9	1.1	< .01
Share married more than once	.102	.135	-.033	.035
Share divorced, separated, or spouse absent	.039	.035	.005	.353
Share with children under 18	.789	.79	-.001	.878
Share cohabiting with children under 18	.78	.781	-.001	.888
Share completed less than highschool	.464	.523	-.059	< .01
Share completed highschool	.291	.255	.036	< .01
Share completed some college	.107	.098	.009	.338
Share completed 4 year college or more	.134	.114	.02	< .01
Economic				
Share in metro area	.784	.527	.256	< .01
Share own home	.64	.626	.014	.62
Home value — own home	$125,574	$93,796	$31,778	< .01
Family income	$58,662	$49,298	$9,364	< .01
Wage income	$54,388	$45,502	$8,885	< .01
Other income| have other income	$5,499	$6,184	-$685	.021
Annual rent| renter	$6,933	$5,586	$1,347	< .01
Annual cost of living| renter	$7,378	$5,967	$1,410	< .01
State-level benefit				
FIU	.603	.429	.174	< .01
ADC recipiency per 1,000 women aged 15–54	14.2	17.5	–3.3	.227
UI maximum (4 weeks)	$1,604	$1,096	$508	< .01
ADC benefit per case	$1,306	$718	$588	< .01
N	456,137	191,058		

Notes: Sample for demographic and economic variables is drawn from the 1960 5 percent census (Ruggles et al., 2023) and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. State-level data on the 1960 FIU is calculated using the overall unemployment rate from the 1980 5 percent census and Department of Labor information on the average number weekly insured for UI (available via the following link: `https://oui.doleta.gov/unemploy/claimssum.asp`). State level data on 1960 ADC payments per caseload and ADC caseloads per 1,000 women come from Goodman-Bacon and Schmidt (2020) and Haines (2010). Column 1 shows means for treated states according to the primary definition (see Section II and Appendix B, column 2 shows means in non-AFDC-UP adopting states, Column 3 shows the difference, and Column 4 the p-value from tests of equality clustering at the state level.

Table 3: Difference in differences estimates of the effect of AFDC-UP on welfare participation and family structure

	(1) Welfare	(2) Div/Sep /Sp.Abs.	(3) Married	(4) Married No Child	(5) Cohabit-ation	(6) Child < 18
A. Lower Tercile						
AFDC-UP	.19	-.06	.07	-.02	.08	.07
	(.02)	(.03)	(.03)	(.02)	(.04)	(.03)
CR p-value	[<.01]	[.051]	[.021]	[.498]	[.042]	[.033]
Unemployment	.06	.12	-.12	-.02	-.07	-.08
	(.01)	(.03)	(.03)	(.02)	(.03)	(.03)
CR p-value	[<.01]	[<.01]	[<.01]	[.254]	[.043]	[<.01]
N	23,361	23,361	23,361	23,361	23,361	23,361
B. Middle Tercile						
AFDC-UP	.07	.04	-.04	-.01	0.00	-.01
	(.02)	(.05)	(.05)	(.05)	(.05)	(.05)
CR p-value	[<.01]	[.373]	[.458]	[.812]	[.95]	[.807]
Unemployment	.04	.05	-.07	.04	-.09	-.10
	(.01)	(.04)	(.04)	(.04)	(.05)	(.05)
CR p-value	[<.01]	[.264]	[.12]	[.333]	[.053]	[.035]
N	22,044	22,044	22,044	22,044	22,044	22,044
C. Upper Tercile						
AFDC-UP	.04	.07	-.06	.01	-.09	-.08
	(.02)	(.05)	(.04)	(.07)	(.1)	(.09)
CR p-value	[.082]	[.141]	[.192]	[.908]	[.363]	[.411]
Unemployment	.01	.13	-.14	-.01	-.10	-.08
	(.02)	(.04)	(.04)	(.06)	(.09)	(.08)
CR p-value	[.347]	[<.01]	[<.01]	[.905]	[.32]	[.367]
N	21,113	21,113	21,113	21,113	21,113	21,113

Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from specification (4), separately by outcome (given in Columns 1-6) and by tercile of the previous year's predicted personal wage and salary income (given in panels), with cluster robust (CR) standard errors at the state level given in parentheses and associated p-values in brackets. Sample pools March CPS-ASEC data between 1977-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. All specifications use CPS-ASEC survey weights.

Table 4: The effect of AFDC-UP on additional measures of program participation and family structure

	(1) Group coverage	(2) Medicaid	(3) Welfare + Medi-caid	(4) Food stamps	(5) Welfare + food stamps	(6) Welfare + Medicaid + food stamps	(7) UI	(8) Welfare + UI	(9) Welfare (No UI)	(10) UI (No Welfare)
A. Lower Tercile										
AFDC-UP	-.02	.21	.18	.03	.16	.16	.12	.07	.11	.05
	(.04)	(.02)	(.02)	(.03)	(.02)	(.02)	(.04)	(.02)	(.02)	(.04)
CR p-value	[.595]	[<.01]	[<.01]	[.452]	[<.01]	[<.01]	[<.01]	[<.01]	[<.01]	[.271]
Unemployment	-.27	.08	.05	.31	.06	.05	.44	.03	.04	.40
	(.03)	(.01)	(.01)	(.03)	(.01)	(.01)	(.04)	(.01)	(.01)	(.04)
CR p-value	[<.01]	[<.01]	[<.01]	[<.01]	[<.01]	[<.01]	[<.01]	[.021]	[<.01]	[<.01]
N	15,420	16,792	16,792	16,792	16,792	16,792	23,361	16,792	16,792	16,792
B. Middle Tercile										
AFDC-UP	.08	.10	.08	0.00	.06	.06	.10	.05	.04	.03
	(.06)	(.03)	(.02)	(.03)	(.02)	(.02)	(.05)	(.02)	(.02)	(.05)
CR p-value	[.152]	[<.01]	[<.01]	[.987]	[<.01]	[.02]	[.042]	[<.01]	[.067]	[.518]
Unemployment	-.33	.07	.04	.22	.04	.04	.52	.01	.03	.51
	(.04)	(.02)	(.01)	(.03)	(.02)	(.02)	(.04)	(.01)	(.01)	(.03)
CR p-value	[<.01]	[<.01]	[<.01]	[<.01]	[.015]	[.015]	[<.01]	[.024]	[.037]	[<.01]
N	14,581	15,893	15,893	15,893	15,893	15,893	22,044	15,893	15,893	15,893
C. Upper Tercile										
AFDC-UP	-.05	.03	.03	-.05	.04	.03	.19	.02	.02	.18
	(.08)	(.04)	(.03)	(.04)	(.03)	(.03)	(.08)	(.01)	(.02)	(.09)
CR p-value	[.497]	[.342]	[.195]	[.241]	[.096]	[.185]	[.022]	[.074]	[.416]	[.059]
Unemployment	-.21	.05	.02	.15	.02	.02	.43	0.00	.02	.39
	(.06)	(.03)	(.02)	(.04)	(.02)	(.02)	(.06)	(0.00)	(.02)	(.07)
CR p-value	[<.01]	[.08]	[.349]	[<.01]	[.349]	[.349]	[<.01]	[.012]	[.349]	[<.01]
N	13,964	15,221	15,221	15,221	15,221	15,221	21,113	15,221	15,221	15,221

Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from specification (4), separately by outcome (given in Columns 1-6) and by tercile of the previous year's predicted personal wage and salary income (given in panels), with cluster robust (CR) standard errors at the state level given in parentheses and associated p-values in brackets. Sample pools March CPS-ASEC data between 1977-1988, with Medicaid, food stamp, and group coverage outcomes available beginning in 1980 and group coverage unavailable in 1988. Sample is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. All specifications use CPS-ASEC survey weights.

Table 5: Doubly robust estimates of the effect of AFDC-UP on welfare participation and family structure

	(1) Welfare	(2) Div/Sep /Sp.Abs.	(3) Married	(4) Married No Child	(5) Cohabit-ation	(6) Child < 18
A. Lower Tercile						
AFDC-UP	.18	-.07	.07	-.01	.04	.06
	(.02)	(.02)	(.02)	(.02)	(.02)	(.02)
p-value	[<.01]	[<.01]	[<.01]	[.778]	[.124]	[.01]
N	1,518	1,518	1,518	1,518	1,518	1,518
Unemployment	.05	.07	-.08	-.01	-.05	-.05
	(.01)	(.02)	(.02)	(.01)	(.02)	(.02)
p-value	[<.01]	[<.01]	[<.01]	[.594]	[<.01]	[<.01]
N	8,276	8,276	8,276	8,276	8,276	8,276
B. Middle & Upper Terciles						
AFDC-UP	.07	.02	-.03	-.05	.07	.02
	(.02)	(.03)	(.03)	(.03)	(.04)	(.04)
p-value	[<.01]	[.44]	[.404]	[.098]	[.08]	[.621]
N	841	841	841	841	841	841
Unemployment	.03	.09	-.10	.01	-.10	-.09
	(.01)	(.02)	(.02)	(.02)	(.03)	(.02)
p-value	[<.01]	[<.01]	[<.01]	[.799]	[<.01]	[<.01]
N	14,292	14,292	14,292	14,292	14,292	14,292

Notes: Displays Doubly Robust (DR) estimates of the effect of AFDC-UP and of unemployment, where the effect of AFDC-UP is estimated according to equation (5) and the effect of unemployment is estimated according to equation (7). Sample pools March CPS-ASEC data between 1977-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Estimates are for men in the lower income tercile, where terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year.

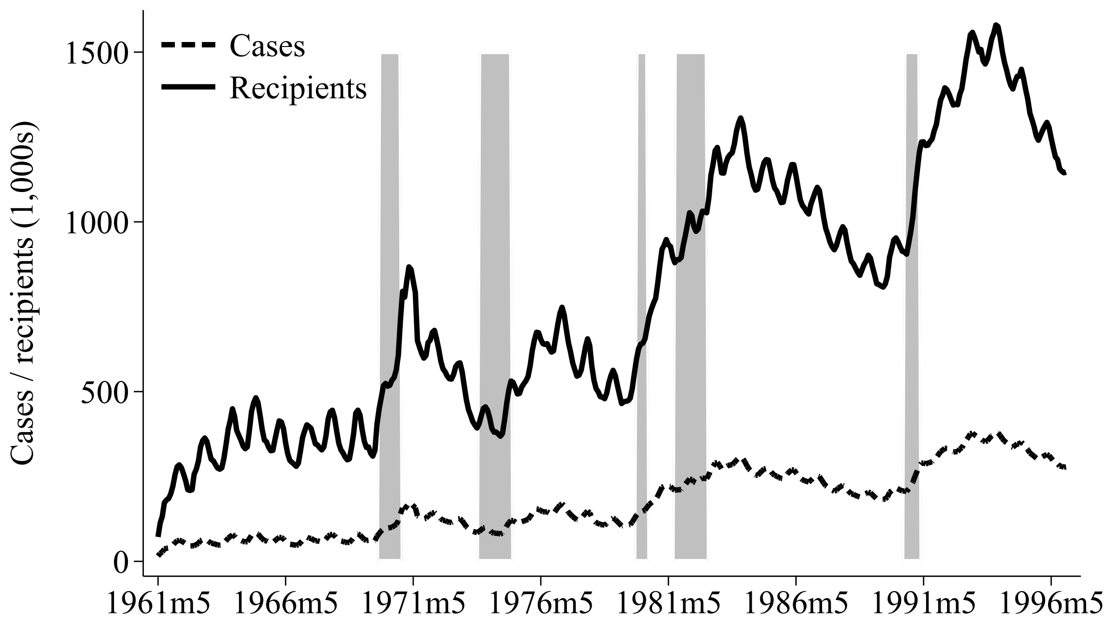
Table 6: Difference in differences estimates of the effect of AFDC-UP between 1968-1988 for a restricted set of states

	(1) Welfare	(2) Div/Sep /Sp.Abs.	(3) Married	(4) Married No Child	(5) Cohabit-ation	(6) Child < 18
A. 1977–1988						
AFDC-UP	.19	-.07	.09	0.00	.08	.05
CR p-value	[<.01]	[.117]	[.049]	[.892]	[.215]	[.322]
2-sided WCR p-value	[<.01]	[.193]	[<.01]	[.895]	[.216]	[.396]
1-sided WCR p-value	[<.01]	[.097]	[<.01]	[.447]	[.108]	[.198]
Unemployment	.07	.13	-.15	-.02	-.09	-.08
CR p-value	[.05]	[<.01]	[<.01]	[.336]	[.154]	[.122]
2-sided WCR p-value	[.171]	[<.01]	[.032]	[.434]	[.184]	[.063]
1-sided WCR p-value	[.086]	[<.01]	[.017]	[.217]	[.092]	[.031]
N	12,891	12,891	12,891	12,891	12,891	12,891
B. 1968–1988						
AFDC-UP	.19	-.07	.09	0.00	.07	.05
CR p-value	[<.01]	[.078]	[.036]	[.931]	[.248]	[.387]
2-sided WCR p-value	[<.01]	[.066]	[<.01]	[.926]	[.283]	[.529]
1-sided WCR p-value	[<.01]	[.033]	[<.01]	[.463]	[.142]	[.265]
Unemployment	.06	.12	-.14	-.02	-.08	-.07
CR p-value	[.021]	[<.01]	[<.01]	[.299]	[.187]	[.194]
2-sided WCR p-value	[.154]	[.016]	[<.01]	[.424]	[.184]	[.176]
1-sided WCR p-value	[.077]	[<.01]	[<.01]	[.212]	[.092]	[.088]
N	27,825	27,825	27,825	27,825	27,825	27,825
Number of states	12	12	12	12	12	12
Number of clusters	10	10	10	10	10	10

Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from specification (4), separately by outcome (given in Columns 1-6) and for two samples periods; 1977-1988 and 1968-1988 (given in panels), with cluster robust (CR) p-values given in the first set of brackets. The second and third sets of brackets provide bootstrapped 2 and 1-sided p-values, respectively, using the Wild Cluster Restricted (WCR) bootstrap (Cameron et al., 2008; MacKinnon et al., 2023) with 1,024 replications implemented via the `boottest` command (Roodman et al., 2019). The sample is restricted to 10 clusters comprised of 12 states that are identifiable in the CPS before 1977 (see Appendix B for further details). Sample pools March CPS-ASEC data between 1977-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. All specifications use CPS-ASEC survey weights.

A Additional Results

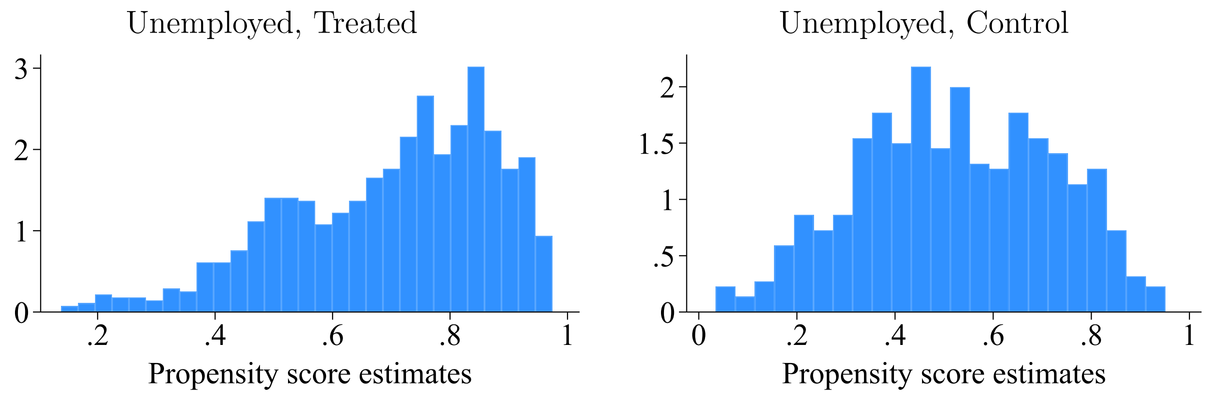
Figure A.1: AFDC-UP caseloads and recipiency, by NBER recession dates



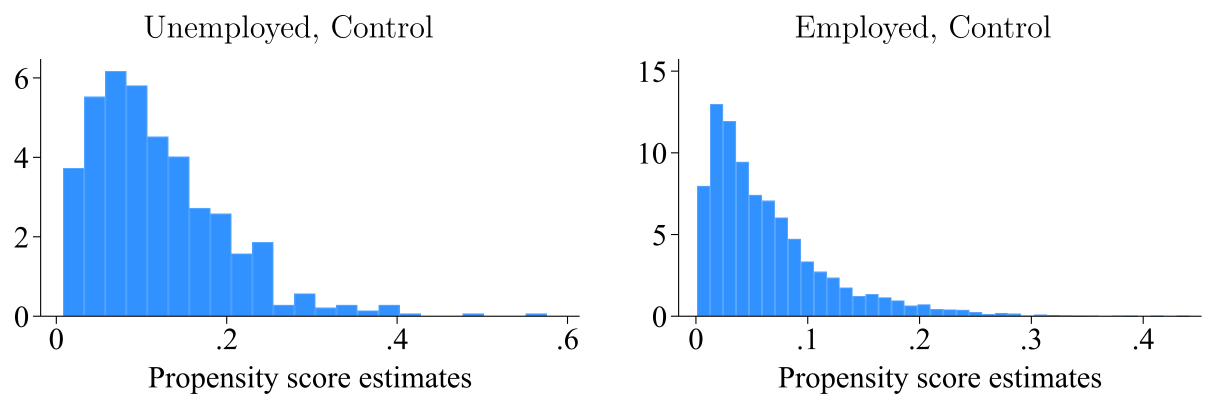
Notes: Plots monthly AFDC-UP caseloads between March, 1961 and December, 1996, with shaded areas representing recessions. Data on caseloads by state and month are from Bureau of Public Assistance (Various Years) and recession dating is taken from the National Bureau of Economic Research (NBER).

Figure A.2: Propensity score overlap in the lower income tercile

A. Propensity score overlap for DR estimates of AFDC-UP

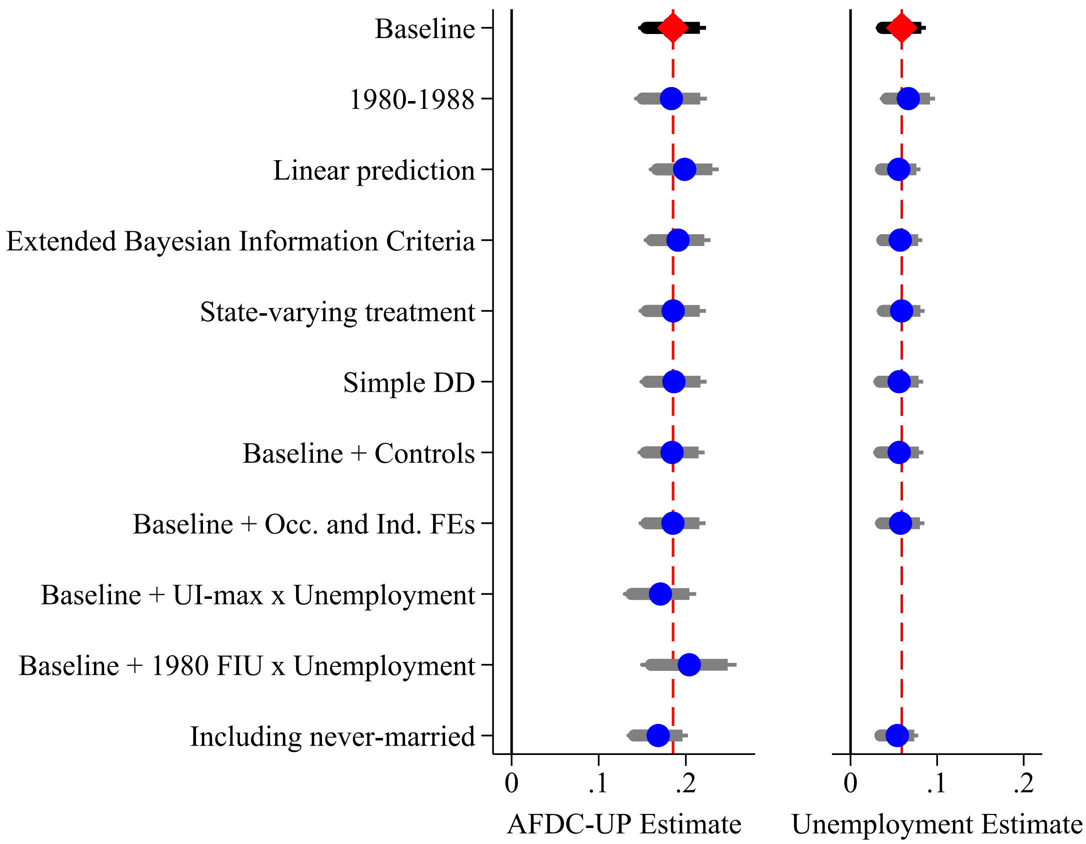


B. Propensity score overlap for DR estimates of Unemployment



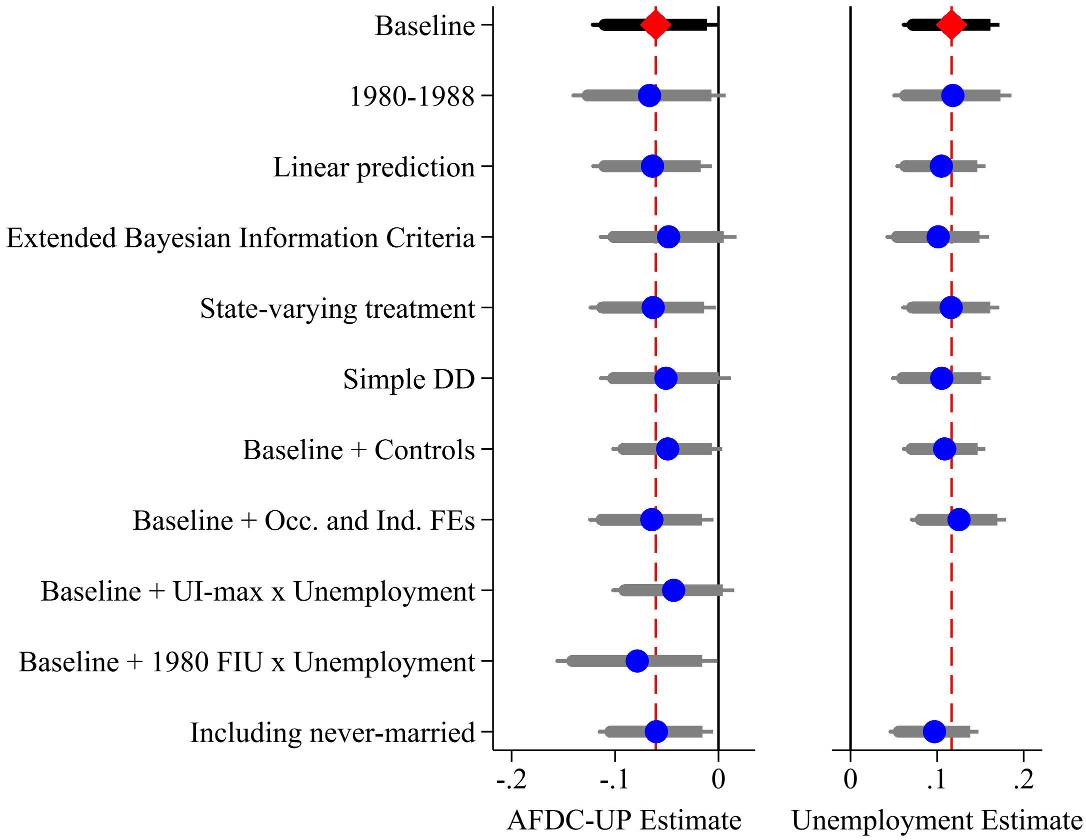
Notes: Displays densities of propensity score estimates predicted using logistic regressions to predict group status. Group status is living in AFDC-UP states among unemployed men in Panel A (estimates used for the Doubly Robust (DR) estimate given in equation (5)) and is experiencing unemployment among men living in control states in Panel B (estimates used for the DR estimate given in equation (7)). Covariates include year fixed effects, race, industry fixed effects, age fixed effects, and a metro dummy. Sample pools March CPS-ASEC data between 1977-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Only men in the lower income tercile are included, where terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year.

Figure A.3: Robustness: Welfare receipt estimates



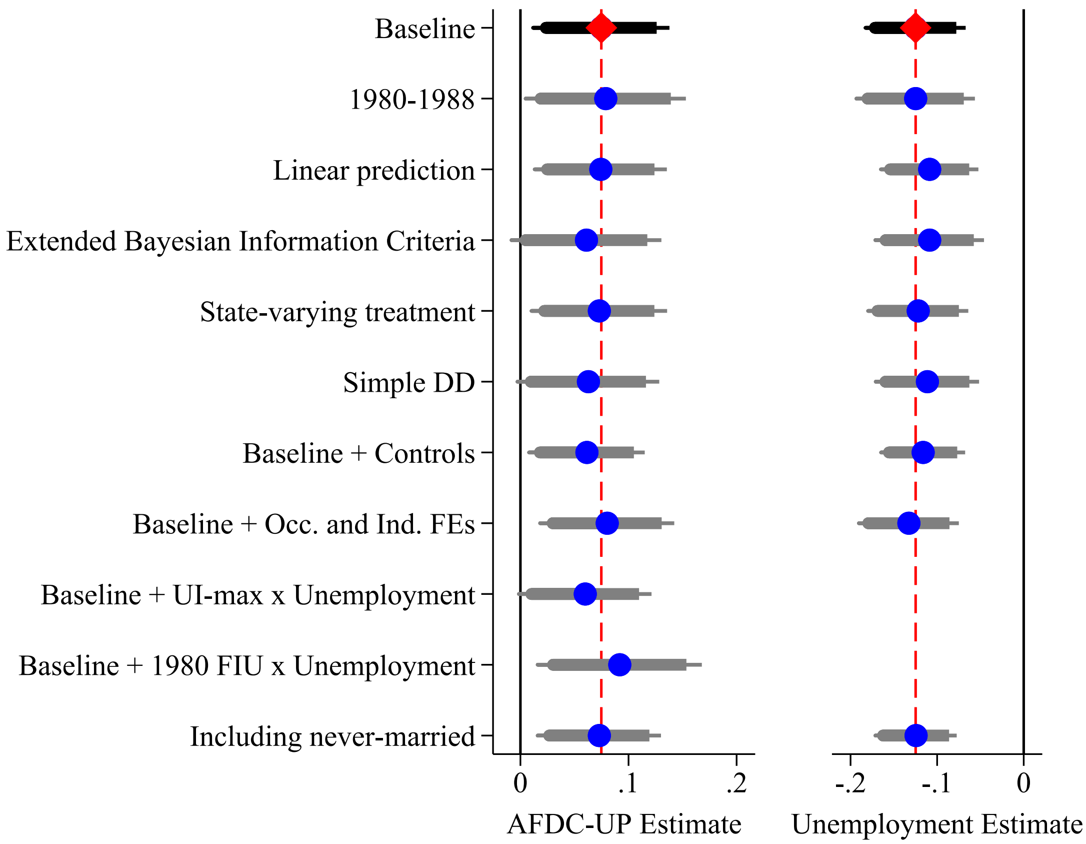
Notes: The outcome indicates household welfare receipt. Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from variants of specification (4) for the lowest tercile of predicted full-time earnings, where terciles are estimated separately by state and year using full-time earnings among men who worked 52 weeks in the previous year. Samples pool March CPS-ASEC data between 1977-1988 or 1980-1988 and are comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and, in all but one specification, are ever-married. The 1st row recreates the baseline results (Table 3); the 2nd shows results restricted to 1980-1988; the 3rd and 4th show estimates where terciles are formed instead using linear regression or the LASSO Extended Bayesian Information Criteria (Tibshirani, 1996; Chen and Chen, 2008) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry; the 5th uses as AFDC-UP treatment every contiguous state and whether they had positive caseloads in a given year (see Appendix B); the 6th uses a dummy for AFDC-UP in place of state-by-year fixed effects (specification (1)); the 7th is the baseline model plus individual-level controls for race, educational attainment-bin fixed effects, and age fixed effects; the 8th is the baseline model plus fixed effects for 1950 occupation and industry codes; the 9th is the baseline model plus an interaction term between the state-year UI maximum and unemployment status (Edwards, 2020), using the maximum for two children in states that varied UI benefits for dependents; the 10th is the baseline model plus an interaction term for the 1980 Fraction of Insured Unemployment (FIU) and unemployment, where the FIU is calculated using the overall unemployment rate from the 1980 5 percent census (Ruggles et al., 2023) and Department of Labor information on the average number weekly insured for UI (available via the following link: `https://oui.doleta.gov/unemploy/claimssum.asp`); and the 11th includes never-married men in the sample. All specifications use CPS-ASEC survey weights and cluster standard errors at the state level, with 90 percent and 95 percent confidence intervals represented by thick and thin bars, respectively.

Figure A.4: Robustness: Separation estimates



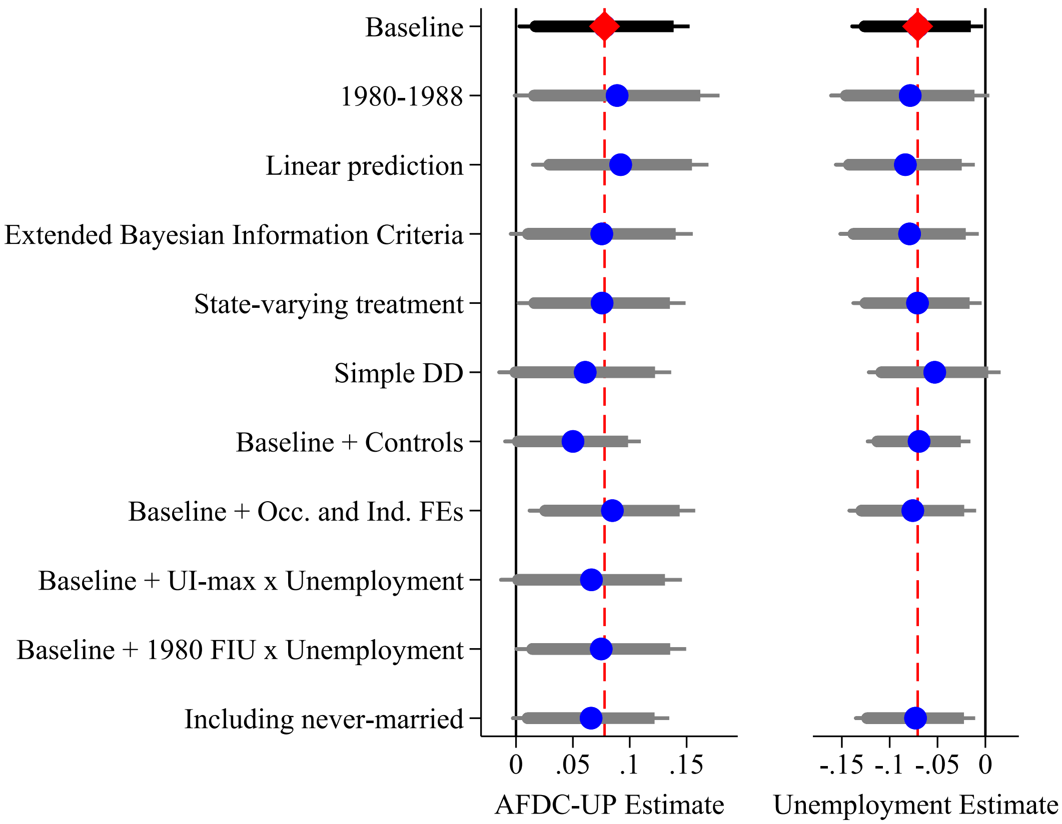
Notes: The outcome indicates currently divorced, separated, or reporting spousal absence. Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from variants of specification (4) for the lowest tercile of predicted full-time earnings, where terciles are estimated separately by state and year using full-time earnings among men who worked 52 weeks in the previous year. Samples pool March CPS-ASEC data between 1977-1988 or 1980-1988 and are comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and, in all but one specification, are ever-married. The 1st row recreates the baseline results (Table 3); the 2nd shows results restricted to 1980-1988; the 3rd and 4th show estimates where terciles are formed instead using linear regression or the LASSO Extended Bayesian Information Criteria (Tibshirani, 1996; Chen and Chen, 2008) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry; the 5th uses as AFDC-UP treatment every contiguous state and whether they had positive caseloads in a given year (see Appendix B); the 6th uses a dummy for AFDC-UP in place of state-by-year fixed effects (specification (1)); the 7th is the baseline model plus individual-level controls for race, educational attainment-bin fixed effects, and age fixed effects; the 8th is the baseline model plus fixed effects for 1950 occupation and industry codes; the 9th is the baseline model plus an interaction term between the state-year UI maximum and unemployment status (Edwards, 2020), using the maximum for two children in states that varied UI benefits for dependents; the 10th is the baseline model plus an interaction term for the 1980 Fraction of Insured Unemployment (FIU) and unemployment, where the FIU is calculated using the overall unemployment rate from the 1980 5 percent census (Ruggles et al., 2023) and Department of Labor information on the average number weekly insured for UI (available via the following link: `https://oui.doleta.gov/unemploy/claimssum.asp`); and the 11th includes never-married men in the sample. All specifications use CPS-ASEC survey weights and cluster standard errors at the state level, with 90 percent and 95 percent confidence intervals represented by thick and thin bars, respectively.

Figure A.5: Robustness: Married estimates



Notes: The outcome indicates currently married. Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from variants of specification (4) for the lowest tercile of predicted full-time earnings, where terciles are estimated separately by state and year using full-time earnings among men who worked 52 weeks in the previous year. Samples pool March CPS-ASEC data between 1977-1988 or 1980-1988 and are comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and, in all but one specification, are ever-married. The 1st row recreates the baseline results (Table 3); the 2nd shows results restricted to 1980-1988; the 3rd and 4th show estimates where terciles are formed instead using linear regression or the LASSO Extended Bayesian Information Criteria (Tibshirani, 1996; Chen and Chen, 2008) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry; the 5th uses as AFDC-UP treatment every contiguous state and whether they had positive caseloads in a given year (see Appendix B); the 6th uses a dummy for AFDC-UP in place of state-by-year fixed effects (specification (1)); the 7th is the baseline model plus individual-level controls for race, educational attainment-bin fixed effects, and age fixed effects; the 8th is the baseline model plus fixed effects for 1950 occupation and industry codes; the 9th is the baseline model plus an interaction term between the state-year UI maximum and unemployment status (Edwards, 2020), using the maximum for two children in states that varied UI benefits for dependents; the 10th is the baseline model plus an interaction term for the 1980 Fraction of Insured Unemployment (FIU) and unemployment, where the FIU is calculated using the overall unemployment rate from the 1980 5 percent census (Ruggles et al., 2023) and Department of Labor information on the average number weekly insured for UI (available via the following link: `https://oui.doleta.gov/unemploy/claimssum.asp`); and the 11th includes never-married men in the sample. All specifications use CPS-ASEC survey weights and cluster standard errors at the state level, with 90 percent and 95 percent confidence intervals represented by thick and thin bars, respectively.

Figure A.6: Robustness: Cohabitation estimates



Notes:
The outcome indicates cohabiting with children, defined by living with at least 1 child under age 18 and their mother. Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from variants of specification (4) for the lowest tercile of predicted full-time earnings, where terciles are estimated separately by state and year using full-time earnings among men who worked 52 weeks in the previous year. Samples pool March CPS-ASEC data between 1977-1988 or 1980-1988 and are comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and, in all but one specification, are ever-married. The 1st row recreates the baseline results (Table 3); the 2nd shows results restricted to 1980-1988; the 3rd and 4th show estimates where terciles are formed instead using linear regression or the LASSO Extended Bayesian Information Criteria (Tibshirani, 1996; Chen and Chen, 2008) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry; the 5th uses as AFDC-UP treatment every contiguous state and whether they had positive caseloads in a given year (see Appendix B); the 6th uses a dummy for AFDC-UP in place of state-by-year fixed effects (specification (1)); the 7th is the baseline model plus individual-level controls for race, educational attainment-bin fixed effects, and age fixed effects; the 8th is the baseline model plus fixed effects for 1950 occupation and industry codes; the 9th is the baseline model plus an interaction term between the state-year UI maximum and unemployment status (Edwards, 2020), using the maximum for two children in states that varied UI benefits for dependents; the 10th is the baseline model plus an interaction term for the 1980 Fraction of Insured Unemployment (FIU) and unemployment, where the FIU is calculated using the overall unemployment rate from the 1980 5 percent census (Ruggles et al., 2023) and Department of Labor information on the average number weekly insured for UI (available via the following link: `https://oui.doleta.gov/unemploy/claimssum.asp`); and the 11th includes never-married men in the sample. All specifications use CPS-ASEC survey weights and cluster standard errors at the state level, with 90 percent and 95 percent confidence intervals represented by thick and thin bars, respectively.

Table A.1: Summary statistics (unemployed men only)

	(1) Overall	(2) Tercile 1	(3) Tercile 2	(4) Tercile 3
Demographic				
Share married	.717	.73	.714	.642
Share divorced, separated, or spouse absent	.27	.26	.265	.348
Age	37.8	36	40.7	41.2
Share white	.849	.792	.957	.906
Share with children under 18	.577	.623	.5	.508
Share cohabiting with children under 18	.63	.674	.568	.529
Share completed less than highschool	.376	.542	.139	.016
Share completed highschool	.432	.392	.597	.21
Share completed some college	.121	.065	.231	.158
Share completed 4 year college or more	.071	0	.034	.616
Economic				
Wage income (predicted)	$60,197	$51,964	$67,797	$89,942
Share unemployed $\geq$ 26 weeks last year	.899	.903	.907	.85
Share currently unemployed $\geq$ 26 weeks	.21	.195	.235	.227
Duration $\mid$ Unemployed $\geq$ 26 weeks last year	32.4	32.3	32.5	32.7
Duration $\mid$ Currently unemployed $\geq$ 26 weeks	40.5	39.9	41.3	41.3
Share below the poverty line	.377	.453	.271	.207
Share own home	.468	.408	.58	.526
Share with employer-based group coverage	.58	.536	.637	.686
Program participation				
Share receiving welfare (family)	.153	.198	.093	.042
Welfare income $\mid$ welfare receipt	$5,485	$5,892	$3,747	$5,565
Share receiving UI	.574	.557	.615	.559
Share receiving foodstamps (family)	.315	.383	.236	.116
foodstamp value $\mid$ foodstamp receipt	$2,699	$2,805	$2,322	$2,706
Share covered by Medicaid (family)	.201	.252	.136	.076
Observations (1977–1988)	2,521	1,599	687	235
Observations (1980–1988)	2,022	1,277	554	191

Notes: Shows sample statistics only for men unemployed according to the primary definition (see Section II). Sample pools March CPS-ASEC data between 1977-1988 (Flood et al., 2022) and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Column 1 shows statistics for the full sample, while columns 2-4 show statistics separately by tercile of the previous year's predicted personal wage and salary income. Earnings are estimated using personal wage and salary income among men who worked 52 weeks in the previous year and predicted for those who worked less than 52 weeks, using the square-root Least Absolute Shrinkage and Selection Operator (LASSO; Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. Statistics on on health insurance, food stamps, and Medicaid are from 1980-1988. Monetary variables are in 2019 dollars.

Table A.2: The effect of AFDC-UP on welfare participation and family structure, 1980-1988 only

	(1) Welfare	(2) Div/Sep /Sp.Abs.	(3) Married	(4) Married No Child	(5) Cohabit-ation	(6) Child < 18
A. Lower Tercile						
AFDC-UP	.18	-.07	.08	-.02	.09	.08
	(.02)	(.04)	(.04)	(.03)	(.04)	(.03)
CR p-value	[<.01]	[.075]	[.037]	[.438]	[.053]	[.036]
Unemployment	.07	.12	-.12	-.02	-.08	-.07
	(.02)	(.03)	(.03)	(.02)	(.04)	(.03)
CR p-value	[<.01]	[<.01]	[<.01]	[.378]	[.062]	[.017]
N	16,792	16,792	16,792	16,792	16,792	16,792
B. Middle Tercile						
AFDC-UP	.09	.05	-.05	-.05	0.00	.02
	(.02)	(.05)	(.05)	(.05)	(.05)	(.05)
CR p-value	[<.01]	[.399]	[.31]	[.334]	[.936]	[.783]
Unemployment	.04	.03	-.04	.05	-.07	-.10
	(.01)	(.05)	(.04)	(.04)	(.04)	(.05)
CR p-value	[.01]	[.5]	[.343]	[.282]	[.137]	[.049]
N	15,893	15,893	15,893	15,893	15,893	15,893
C. Upper Tercile						
AFDC-UP	.04	.07	-.06	.01	-.10	-.08
	(.03)	(.06)	(.06)	(.08)	(.1)	(.09)
CR p-value	[.102]	[.25]	[.338]	[.945]	[.347]	[.402]
Unemployment	.02	.14	-.15	.01	-.11	-.10
	(.02)	(.05)	(.05)	(.07)	(.1)	(.08)
CR p-value	[.349]	[.012]	[<.01]	[.854]	[.262]	[.253]
N	15,221	15,221	15,221	15,221	15,221	15,221

Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from specification (4), separately by outcome (given in Columns 1-6) and by tercile of the previous year's predicted personal wage and salary income (given in panels), with cluster robust (CR) standard errors at the state level given in parentheses and associated p-values in brackets. Sample pools March CPS-ASEC data between 1980-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. All specifications use CPS-ASEC survey weights.

Table A.3: The effect of AFDC-UP on welfare participation and family structure, dropping the South

	(1) Welfare	(2) Div/Sep /Sp.Abs.	(3) Married	(4) Married No Child	(5) Cohabit-ation	(6) Child < 18
A. Lower Tercile						
AFDC-UP	.14	-.03	.06	-.03	.11	.07
	(.03)	(.03)	(.03)	(.03)	(.04)	(.03)
CR p-value	[<.01]	[.252]	[.067]	[.372]	[<.01]	[.038]
Unemployment	.10	.09	-.11	0.00	-.11	-.07
	(.03)	(.03)	(.03)	(.03)	(.03)	(.02)
CR p-value	[<.01]	[<.01]	[<.01]	[.908]	[<.01]	[<.01]
N	16,039	16,039	16,039	16,039	16,039	16,039
B. Middle Tercile						
AFDC-UP	.08	-.02	.06	.07	0.00	-.09
	(.02)	(.02)	(.03)	(.05)	(.05)	(.06)
CR p-value	[<.01]	[.268]	[.068]	[.168]	[.984]	[.119]
Unemployment	.04	.12	-.16	-.03	-.11	-.04
	(.01)	(.01)	(.02)	(.04)	(.05)	(.05)
CR p-value	[<.01]	[<.01]	[<.01]	[.439]	[.042]	[.424]
N	15,112	15,112	15,112	15,112	15,112	15,112
C. Upper Tercile						
AFDC-UP	.05	-.11	.10	.04	.05	0.00
	(.02)	(.07)	(.07)	(.09)	(.16)	(.07)
CR p-value	[<.01]	[.139]	[.169]	[.646]	[.735]	[.977]
Unemployment	0.00	.30	-.29	-.02	-.25	-.16
	(0.00)	(.07)	(.07)	(.08)	(.15)	(.05)
CR p-value	[<.01]	[<.01]	[<.01]	[.749]	[.118]	[<.01]
N	14,379	14,379	14,379	14,379	14,379	14,379

Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from specification (4), separately by outcome (given in Columns 1-6) and by tercile of the previous year's predicted personal wage and salary income (given in panels), with cluster robust (CR) standard errors at the state level given in parentheses and associated p-values in brackets. The sample does not include any states in the southern census region. Sample pools March CPS-ASEC data between 1977-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. All specifications use CPS-ASEC survey weights.

Table A.4: Falsification test using AFDC-UP variation applied to the 1960 census

	(1) Div/Sep /Sp.Abs.	(2) Married	(3) Married No Child	(4) Cohabit-ation	(5) Child < 18
A. Lower Tercile					
AFDC-UP	.04	-.04	0.00	-.04	-.04
	(.02)	(.02)	(.01)	(.03)	(.03)
CR p-value	[.021]	[.044]	[.863]	[.169]	[.178]
Unemployment	.02	-.03	.06	-.09	-.09
	(.01)	(.01)	(.01)	(.02)	(.02)
CR p-value	[<.01]	[<.01]	[<.01]	[<.01]	[<.01]
N	171,939	171,939	171,939	171,939	171,939
B. Middle Tercile					
AFDC-UP	.03	-.04	0.00	-.03	-.03
	(.01)	(.02)	(.02)	(.04)	(.04)
CR p-value	[.079]	[.043]	[.879]	[.394]	[.421]
Unemployment	.03	-.04	.10	-.14	-.14
	(.01)	(.01)	(.02)	(.02)	(.02)
CR p-value	[<.01]	[<.01]	[<.01]	[<.01]	[<.01]
N	171,778	171,778	171,778	171,778	171,778
C. Upper Tercile					
AFDC-UP	.02	-.02	-.04	.02	.02
	(.02)	(.02)	(.02)	(.03)	(.03)
CR p-value	[.364]	[.308]	[.026]	[.484]	[.549]
Unemployment	.05	-.06	.13	-.19	-.18
	(.01)	(.01)	(.01)	(.02)	(.02)
CR p-value	[<.01]	[<.01]	[<.01]	[<.01]	[<.01]
N	171,813	171,813	171,813	171,813	171,813

Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from specification (4), separately by outcome (given in Columns 1-6) and by tercile of the previous year's predicted personal wage and salary income (given in panels), with cluster robust (CR) standard errors at the state level given in parentheses and associated p-values in brackets. Sample is drawn from the 1960 5 percent census (Ruggles et al., 2023) and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state.

Table A.5: The effect of AFDC-UP, Stratifying by Early/Late no-fault Divorce

	Pre-1977 No-Fault Divorce Law Adopters						Post-1988 (or Never) No-Fault Divorce Law Adopters					
	(1) Welfare	(2) Div/Sep /Sp.Abs.	(3) Married	(4) Married No Child	(5) Cohabit-ation	(6) Child < 18	(7) Welfare	(8) Div/Sep /Sp.Abs.	(9) Married	(10) Married No Child	(11) Cohabit-ation	(12) Child < 18
A. Lower Tercile												
AFDC-UP	.19	-.07	.09	0.00	.04	.06	.20	-.05	.05	-.03	.12	.08
	(.02)	(.04)	(.04)	(.03)	(.05)	(.04)	(.03)	(.07)	(.07)	(.05)	(.04)	(.06)
CR p-value	[<.01]	[.086]	[.034]	[.991]	[.373]	[.144]	[<.01]	[.501]	[.524]	[.508]	[.014]	[.179]
Unemployment	.07	.12	-.14	-.02	-.06	-.08	.03	.10	-.09	-.02	-.10	-.07
	(.02)	(.03)	(.03)	(.02)	(.05)	(.03)	(.02)	(.07)	(.07)	(.04)	(.04)	(.05)
CR p-value	[<.01]	[<.01]	[<.01]	[.277]	[.21]	[.036]	[.168]	[.145]	[.193]	[.648]	[.024]	[.163]
N	12,567	12,567	12,567	12,567	12,567	12,567	10,285	10,285	10,285	10,285	10,285	10,285
B. Middle Tercile												
AFDC-UP	.11	.04	-.05	.01	-.02	-.06	.03	.11	-.08	-.08	-.01	.05
	(.03)	(.05)	(.05)	(.04)	(.06)	(.06)	(.04)	(.05)	(.04)	(.08)	(.08)	(.1)
CR p-value	[<.01]	[.46]	[.358]	[.883]	[.733]	[.345]	[.484]	[.035]	[.052]	[.312]	[.937]	[.623]
Unemployment	.03	.09	-.11	0.00	-.10	-.08	.07	-.04	.02	.13	-.07	-.15
	(.01)	(.05)	(.05)	(.03)	(.06)	(.06)	(.03)	(.04)	(.04)	(.07)	(.07)	(.09)
CR p-value	[<.01]	[.072]	[.041]	[.911]	[.094]	[.153]	[.027]	[.35]	[.587]	[.091]	[.372]	[.11]
N	11,565	11,565	11,565	11,565	11,565	11,565	9,960	9,960	9,960	9,960	9,960	9,960
C. Upper Tercile												
AFDC-UP	.07	.09	-.07	-.03	-.02	-.05	0.00	.07	-.08	.02	-.22	-.09
	(.03)	(.05)	(.05)	(.09)	(.12)	(.13)	(.05)	(.08)	(.08)	(.08)	(.12)	(.13)
CR p-value	[.032]	[.103]	[.206]	[.749]	[.852]	[.689]	[.956]	[.449]	[.363]	[.856]	[.075]	[.463]
Unemployment	0.00	.14	-.16	-.03	-.15	-.07	.04	.09	-.09	.05	.03	-.09
	(0.00)	(.05)	(.04)	(.09)	(.12)	(.12)	(.05)	(.08)	(.08)	(.07)	(.11)	(.12)
CR p-value	[<.01]	[<.01]	[<.01]	[.725]	[.205]	[.563]	[.355]	[.242]	[.282]	[.5]	[.793]	[.438]
N	10,901	10,901	10,901	10,901	10,901	10,901	9,721	9,721	9,721	9,721	9,721	9,721
Number of states	23	23	23	23	23	23	17	17	17	17	17	17

Notes: Displays the coefficients on AFDC-UP ($\widehat{\gamma}_2$) and unemployment ($\widehat{\gamma}_1$) from specification (4), stratified by whether states had adopted no-fault divorce laws before 1977 (Columns 1-6) or had not adopted no-fault divorce laws by 1988 (columns 7-13) and by tercile of the previous year's predicted personal wage and salary income (given in panels), with cluster robust (CR) standard errors at the state level given in parentheses and associated p-values in brackets. Information on no-fault divorce laws is taken from Wolfers (2006) who draws in part from Friedberg (1998). 23 of the 42 states in the primary sample had adopted before 1977, while 17 had not adopted by 1988. For the sample description, see Section II or notes to Table 3. All specifications use CPS-ASEC survey weights.

B Data Appendix

B.1 Primary treatment definition

As noted in the main text, I begin by dropping AK, HI, and DC. I also drop NV, because of outlier rates of marriage and divorce. Figure B.1 shows which states had AFDC-UP in every year as determined by positive caseloads in any calendar month, for all states that ever adopted before the FSA of 1988 (excluding the above states and DC). 19 of the 31 states had the program every year between 1977-1988. Although AZ and OK had the program in earlier years, they never do between 1977-1988, so I include them as control states. KY and NC each have the program only for 1 year in the sample (1977 and 1988, respectively), so I also include these states in the control group. MO, MT, and WA only drop the program for one or two years in the sample (1982, 1984-1985, and 1982, respectively), so I include these states as treated. Finally, because CO, ME, OR, and UT all drop the program for three or more years, and SC adopts for the first time in 1985 (four years in the sample), I drop these 5 states. The final sample includes 42 states (Figure 2, Panel B).

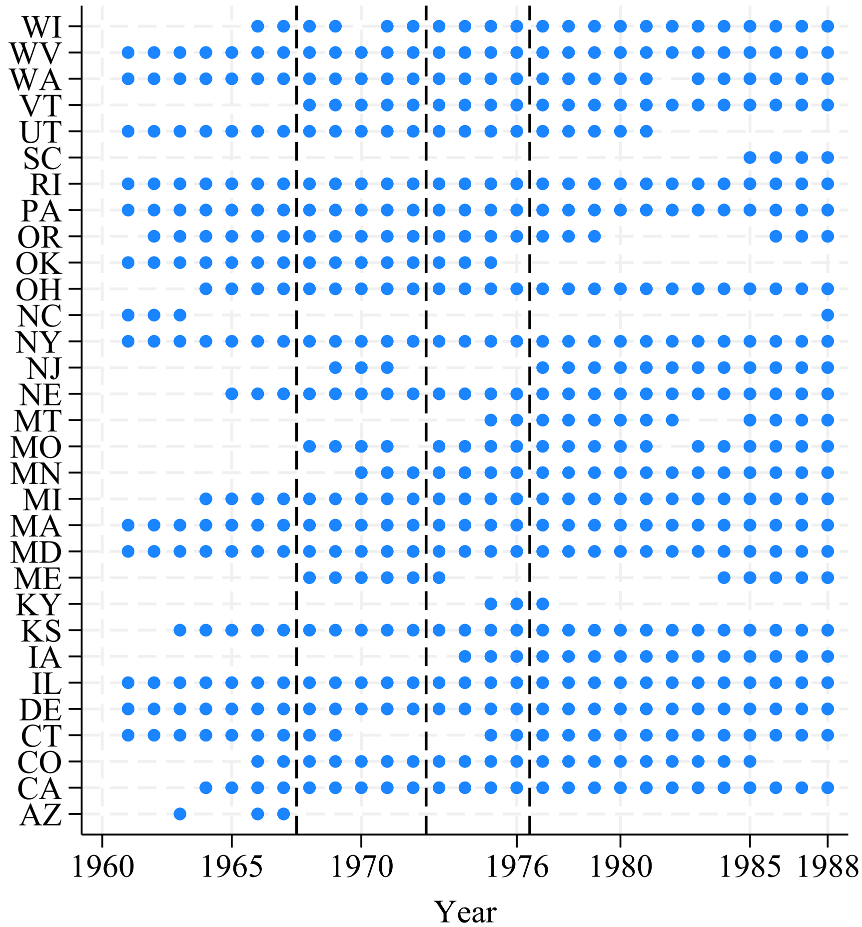
B.2 State groupings and AFDC-UP status in the CPS-ASEC between 1968-1977

There are two types of restrictions that must be made on the sample of states to extend analyses in the CPS-ASEC back before 1977. The first is that there are less states classified as either treated or untreated under the preferred definition that considers always versus never adopters, with always adoption defined at the beginning of the sample period examined. Given the years for which states are grouped in the CPS and when these groups change (see next paragraph), I consider two additional measures: always adopters from 1973-1988, and always adopters from 1968-1988. These are subsets of the primary treatment classification; that is, no state is coded as either having or not having a program if they were not included in the original sample. Beginning in 1973, I drop KY, OK, MT, NJ, leaving 38 states. Beginning in 1968, I further drop IA and CT (classifying both WI, which dropped the program for 1 year, and MN, which started the program in 1970, as always adopters). This leaves 36 states.

The second type of restriction stems from the fact that the CPS does not identify all individual states between 1968-1976.[1] There are two key periods of groupings: the first is between 1968-1972 in which 18 states are identified (14 of which are classified as treated or not), and the second between 1973-1977 in which 12 states are identified (11 of which are classified as treated or not). Table B.1 shows all groupings of states between 1973-1976 and Table B.2 shows the same for those between 1968-1972. The tables first list the states that are individually identified in each group of years, followed by the groupings of states, with pairwise columns listing the state and corresponding treatment classification. States in groupings are only included in the extended sample if their treatment classifications (always or never adopters) are uniform. In both periods, there are two groupings of states that satisfy these criteria. The MI and WI grouping are always adopters, and the AL and MS grouping are never adopters. Figure B.2 maps each set of included states. There are 18 states comprised of 16 clusters for 1968-1972 (Panel A), 15 states comprised of 13 clusters for 1973-1976 (Panel B), and 12 states comprised of 10 clusters that are included across both periods (Panel C).

[1] See https://cps.ipums.org/cps-action/variables/statefip#comparability_section for further details.

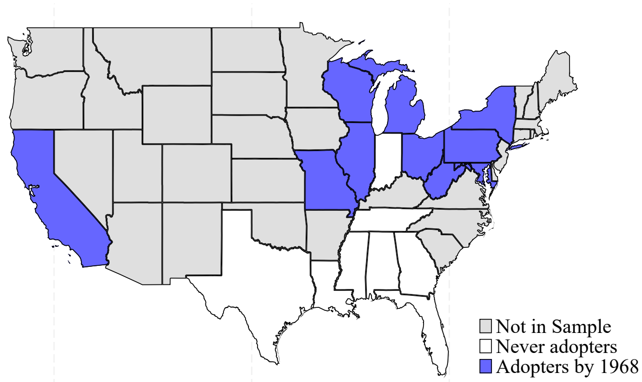
Figure B.1: All states adopting AFDC-UP before 1990, by years in which the program operated



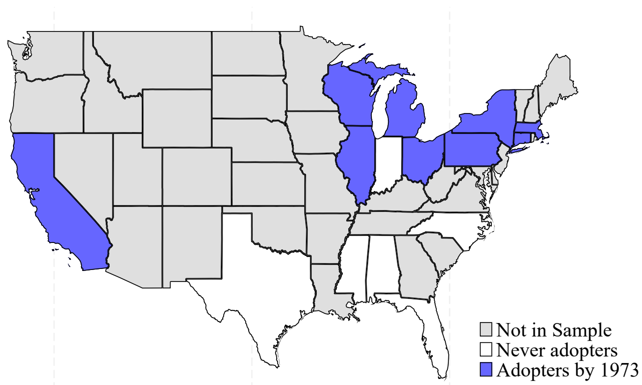
Notes: Plots each year a state had AFDC-UP brtween 1961-1988, for all states that ever adopted before 1990. Data on AFDC-UP caseloads by state and month come from the Department of Health and Human Services (DHHS) available at: https://www.acf.hhs.gov/ofa/resource/tanf-and-afdc-historical-case-data-pre-2012. A state is classified as having an AFDC-UP program in a given calendar year if there were positive AFDC-UP caseloads in any month during that year.

Figure B.2: States included in earlier CPS-ASEC samples

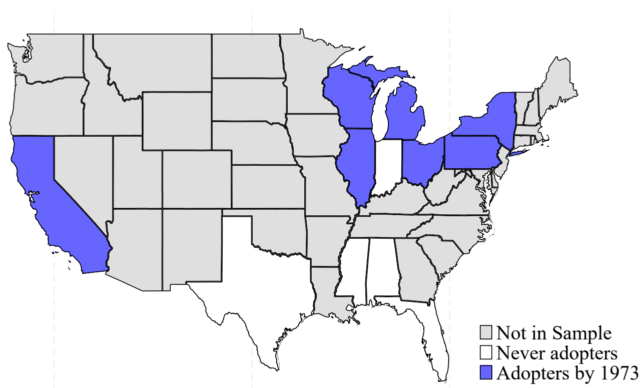
A. 1968-1972 and 1977-1988



B. 1973-1988



C. 1968-1988



Notes: Panel A maps states that are classified as treated or control from 1968-1988 and are either individually identifiable in the CPS-ASEC between 1968-1972 or are in groups with constant treatment status (see text and notes to Table B.1). Panel B maps states that are classified as treated or control from 1973-1988 and are either individually identifiable in the CPS-ASEC between 1973-1976 or are in groups with constant treatment status (see text and notes to Table B.2). Panel C maps the intersection of Panels A and B.

Table B.1: Sample 1968-1972 and 1977-1988

	State 1		State 2		State 3		State 4		State 5		
		AFDC-UP		AFDC-UP		AFDC-UP		AFDC-UP		AFDC-UP	Classified treated
Individual											
	CA	Y									Y
	CT	-									-
	FL	N									N
	GA	N									N
	IL	Y									Y
	IN	N									N
	KY	-									-
	LA	N									N
	MD	Y									Y
	MO	Y									Y
	NJ	-									-
	NY	Y									Y
	OH	Y									Y
	OR	-									-
	PA	Y									Y
	TN	N									N
	TX	N									N
	WV	Y									Y
Grouped											
	ME	-	MA	Y	NH	N	RI	Y	VT	Y	-
	MI	Y	WI	Y							Y
	MN	Y	IA	-							-
	NE	Y	ND	N	SD	N	KS	Y			-
	DE	Y	VA	N							-
	NC	N	SC	-							-
	AL	N	MS	N							N
	AR	N	OK	-							-
	AZ	N	NM	N	CO	-					-
	ID	N	WY	N	UT	-	MT	-	NV	-	-
	AK	-	WA	Y	HI	-					-

Notes: Tabulates all states that are individually identifiable in the CPS-ASEC as well as states that are grouped between 1968-1972. For groups of states, pairwise columns indicate the state abbreviation and their AFDC-UP treatment status. "Y" indicates whether a state always or almost always had an AFDC-UP program between 1968-1988, "N" indicates whether they never or almost never had an AFDC-UP program between 1968-1988, and "-" indicates whether they adopted or dropped the program for long periods, and also include HI and AK. The final column indicates whether all states in the group have identical treatment status of either "Y" or 'N", and is therefore a reproduction of column 1 for individual states. There are 18 of these states, including two groupings of two states, or 16 clusters in total.

Table B.2: Sample 1973-1988

	State 1		State 2		State 3		State 4		State 5		State 6		State 7		
		AFDC-UP		AFDC-UP		AFDC-UP		AFDC-UP		AFDC-UP		AFDC-UP		AFDC-UP	Classified treated
Individual															
	CA	Y													Y
	CT	Y													Y
	FL	N													N
	IL	Y													Y
	IN	N													N
	MA	Y													Y
	NJ	-													-
	NY	Y													Y
	NC	N													N
	OH	Y													Y
	PA	Y													Y
	TX	N													N
Grouped															
	MI	Y	WI	Y											Y
	AL	N	MS	N											N
	NH	N	ME	-	VT	Y	RI	Y							-
	SC	-	GA	N											-
	KY	-	TN	N											-
	AR	N	LA	N	OK	-									-
	IA	Y	ND	N	SD	N	NE	Y	KS	Y	MN	Y	MO	Y	-
	WA	Y	OR	-	AK	-	HI	-							-
	MT	-	WY	N	CO	-	NM	N	UT	-	NV	-	AZ	N	-
	DE	Y	MD	Y	VA	N	WV	Y							-

Notes: Tabulates all states that are individually identifiable in the CPS-ASEC as well as states that are grouped between 1973. For groups of states, pairwise columns indicate the state abbreviation and their AFDC-UP treatment status. "Y" indicates whether a state always or almost always had an AFDC-UP program between 1973-1988, "N" indicates whether they never or almost never had an AFDC-UP program between 1968-1988, and "-" indicates whether they adopted or dropped the program for long periods, and also include HI and AK. The final column indicates whether all states in the group have identical treatment status of either "Y" or 'N", and is therefore a reproduction of column 1 for individual states. There are 15 of these states, including two groupings of two states, or 13 clusters in total.

C Accounting for measurement error in the CPS-ASEC

The first source of measurement error is systematic under-reporting of welfare receipt in the CPS-ASEC. Meyer et al. (2015) advocate comparing administrative and survey information to estimate the "proportional bias" of transfer dollars reported received in common household surveys:

$$\text{Proportional Bias} = \left(\frac{\text{dollars reported in survey, population weighted}}{\text{dollars reported in administrative data}}\right) - 1 \tag{C.1}$$

Using newly entered information on AFDC-UP spending for some years in the 1970s and 1980s (Bureau of Public Assistance, Various Years), I estimate the proportional bias across four different samples, finding a range of underreporting between 22.5 percent and 53.9 percent, with the upper bound quite similar the 50 percent finding for the AFDC-basic program in Meyer et al. (2015).[1] While the bias is comprised both of underreporting and non-reporting, an extreme assumption is that it is all driven by non-reporting.[2] The highest proportional bias magnitude then implies scaling up the first stage by about a factor of 2, or scaling down the implied effect per case by a factor of 2.

The second explanation is that CPS-ASEC welfare participation refers to the prior year, which misses some families who began receiving benefits between January 1st and March. Given that the median AFDC-UP length is around 6 months or less (see discussion in Section I), all indications point to current receipt being disproportionately common among non-welfare reporting unemployed individuals in AFDC-UP states. As the interview date is sometime in March, a simple exercise is to scale up recipiency by 25 percent. Combined with the previous exercise, the cumulative factor

[1] Administrative information on AFDC-UP spending is available by state and month for 1968-1970 and 1978-1980. Across methodologies, I calculate AFDC-UP reported benefits in the CPS-ASEC (using survey weights) or in the 1980 5 percent census as all welfare income reported by men cohabiting with children under 18 who are themselves aged 18-64. The first procedure uses national spending across all states across the 6 available years (bias=-53.9 percent), the second only compares spending in treated states, which necessitates restricting attention to 1978-1980 (bias=-36.6 percent); the third instead compares administrative spending and welfare income in the 1980 census (-22.5 percent), and the fourth does the same only in AFDC-UP states (-49.5 percent). Using all men (regardless of cohabitation status) yields similar and slightly reduced estimates across methodologies.

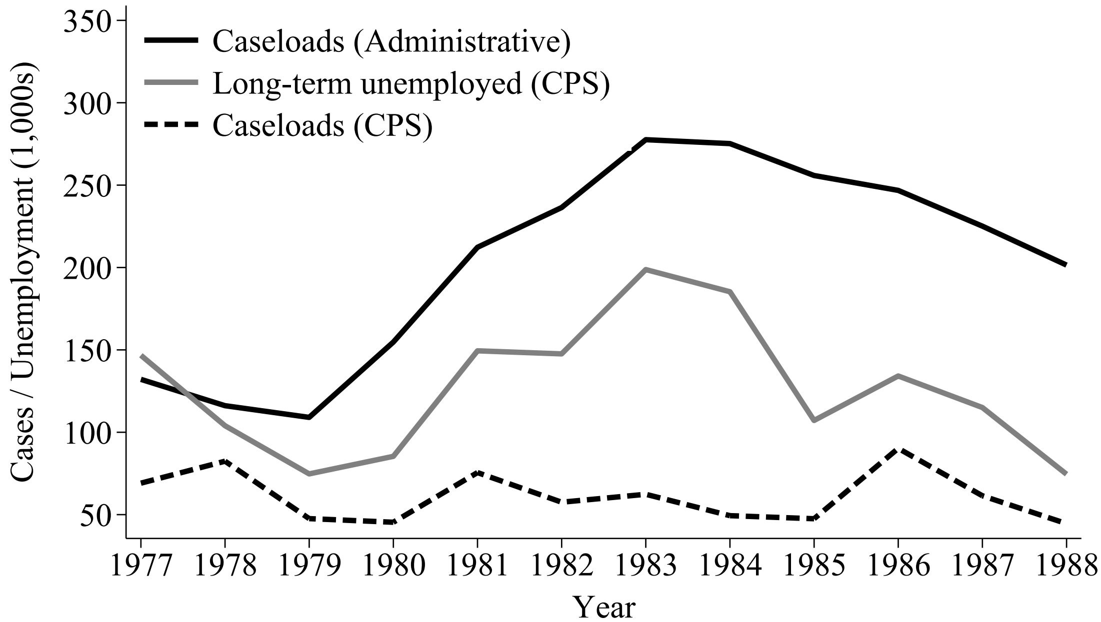
[2] Because there is no information on length of recipiency in the CPS-ASEC and administrative information on caseloads are stocks of recipients in any given month, it is not possible to do a similar analysis comparing recipiency rates.

is 2.5.

The third explanation centers on measurement error in unemployment recall, which Horvath (1982) documents in the CPS-ASEC between 1967-1979 to be on the order of 9-25 percent. Again taking an extreme assumption — that all employed welfare recipients (currently and last year) in AFDC-UP states were actually unemployed — and re-running specification (4) leads to a welfare coefficient about 90 percent higher (36.4 p.p.). Combined with the previous cumulative factor of 2.5 implies an upper bound on recipiency of roughly 90 p.p., which is broadly consistent with the actual size of AFDC-UP.[3] Using this estimate of welfare receipt suggests AFDC-UP prevented around 60 percent of *recipient families* from separating ((-0.061/0.9)/0.117). The corresponding estimate for cohabitation suggests AFDC-UP prevented around 120 percent of recipient families from non-cohabiting, although I cannot reject 100 percent (using the Delta Method and ignoring uncertainty due to measurement error yields a 95 percent confidence interval of [-1.70,-0.75]).

[3] To gauge the plausibility of this estimate, one useful metric is how the (survey estimated) population of unemployed men in the lowest tercile compares with recipiency over the sample period. Figure C.1 shows that caseloads stand well above the count of (long-term) unemployed in each year of the sample, indicating that the vast majority of these individuals could have been receiving welfare while some in higher terciles contributing to the remainder. For reference, I also include implied caseloads from the CPS-ASEC, which are significantly lower.

Figure C.1: AFDC-UP caseload comparisons between administrative records and the CPS-ASEC



Notes: Plots the number of AFDC-UP caseloads (black solid line), estimates of the long-term unemployed from the CPS-ASEC (gray solid line), and estimates caseloads in the CPS-ASEC (black dashed line), with each series measured in 1,000s. Data on monthly AFDC-UP caseloads come from Bureau of Public Assistance (Various Years) and are collapsed by year. Both CPS-ASEC series' are from the primary sample, which pools March CPS-ASEC data between 1977-1988 and is comprised of men ages 24-58 who are the head of the household, do not live in group quarters or multi-family households, are classified as wage and salary workers, worked non-zero weeks, report positive wage income in excess of $1,000, report zero farm or business income in the prior year, live in households where no other adult is in the labor force, and are ever-married. Series are presented only for men in the lower income tercile, where terciles are formed using predicted full-time earnings among men who worked 52 weeks in the previous year, where the prediction uses the square-root LASSO (Tibshirani, 1996; Belloni et al., 2011) with the set of possible predictors including fixed effects for age, race, educational attainment bins (less than high school, high school, some college, and college or more), state, year, occupation, and industry. Terciles are estimated separately by state and year. The two CPS-ASEC series' are estimated using survey weights.

D Derivation of no-timing difference in differences identification conditions

This section derives equation (3) in Section II, which relates the probability limit of $\widehat{\beta}_3$ from specification (1) to parameters of interest and measures of selection. I first re-create the estimand given in equation (2) here:

$$\begin{aligned}\widehat{\beta}_3 \xrightarrow{p} \mathbf{E}[y_{ist}|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 0] \\ - \Big(\mathbf{E}[y_{ist}|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 0]\Big)\end{aligned} \tag{D.1}$$

Substituting in the potential outcomes $y_{ist}(\mathrm{UP}^{76}_{s(i)}, \mathrm{UN}_{ist})$:

$$\begin{aligned}= \mathbf{E}[y_{ist}(1,1)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(1,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 0] \\ - \Big(\mathbf{E}[y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 0]\Big)\end{aligned} \tag{D.2}$$

Adding and subtracting the counterfactual mean potential outcome for unemployed men in AFDC-UP states were they not exposed to AFDC-UP ($\mathbf{E}[y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1]$) and re-arranging terms:

$$\begin{aligned}=&\underbrace{\mathbf{E}[y_{ist}(1,1) - y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1]}_{ATT} \\ &+ \Big(\mathbf{E}[y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(1,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 0]\Big) \\ &- \Big(\mathbf{E}[y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 0]\Big)\end{aligned} \tag{D.3}$$

The first term labeled ATT is the Average Treatment Effect on the Treated for workers who are unemployed in AFDC-UP states. Adding and subtracting the counterfactual mean potential outcome for employed men not exposed to AFDC-UP had they resided in AFDC-UP states ($\mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 0]$) and re-arranging again yields:

$$\begin{aligned}=&ATT - \underbrace{\mathbf{E}[y_{ist}(1,0) - y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 0]}_{ATT^{ind}} \\ &+ \Big(\mathbf{E}[y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 0]\Big) \\ &- \Big(\mathbf{E}[y_{ist}(0,1)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 0]\Big)\end{aligned} \tag{D.4}$$

The second term, labeled ATT^{ind}, represents any indirect treatment effect that AFDC-UP exposure has on employed individuals in AFDC-UP states.

Next, adding and subtracting the counterfactual mean untreated potential outcome if unemployed and in AFDC-UP states ($\mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1]$) and the counterfactual mean untreated potential outcome if unemployed and in non-AFDC-UP states ($\mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1]$), and re-arranging again yields:

$$\begin{aligned}
=&ATT - ATT^{ind} \\
&+ \underbrace{\mathbf{E}[y_{ist}(0,1) - y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1]}_{ATT^{ue}(1)} \\
&- \underbrace{\mathbf{E}[y_{ist}(0,1) - y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1]}_{ATT^{ue}(0)} \\
&+ \Big(\mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 1, \mathrm{UN}_{ist} = 0]\Big) \\
&- \Big(\mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 1] - \mathbf{E}[y_{ist}(0,0)|\mathrm{UP}^{76}_{s(i)} = 0, \mathrm{UN}_{ist} = 0]\Big)
\end{aligned} \tag{D.5}$$

$ATT^{ue}(1)$ represents the causal effect of unemployment for actually unemployed men in AFDC-UP states and $ATT^{ue}(0)$ represents the causal effect of unemployment for actually unemployed men in non-AFDC-UP states. Term the difference $\Delta ATT^{ue} \equiv ATT^{ue}(1) - ATT^{ue}(0)$. Finally, the last set of unlabeled terms together represent a "Bias-Stability" (BS) measure, which is the mean difference in non-AFDC-UP, employed potential outcomes between actually unemployed and employed workers in AFDC-UP states minus the same measure in non-AFDC-UP states and is analogous to parallel trends in more canonical difference in differences research designs leveraging timing. Plugging in BS yields the final probability limit derivation:

$$\widehat{\beta}_3 \xrightarrow{p} ATT - ATT^{ind} + \Delta ATT^{ue} + BS \tag{D.6}$$