Union dissolution and income inequality among separating women

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Abstract

Union dissolution has large consequences for women's standard of living. The consequences may be particularly serious for women from low-income unions, raising concerns about income inequality. Hence, this study describes the consequences of union dissolution for income inequality among separating women. I started from the idea that dissolution may drive cumulative disadvantage. To test this idea, I used administrative data from the Netherlands, following women in coresidential unions from 2003 to 2015 (N = 38,059). Using fixed-effects individual-slopes regressions and recentered influence functions, I compared women's household incomes to a counterfactual scenario in which incomes continued along their predissolution trajectories. The results showed that dissolution prompted income convergence, as women from high-income unions experienced sizeable losses yet women from low-income unions actually gained. At the same time, aggregate inequality increased somewhat, due to a combination of downward mobility by most women and strong upward mobility by some. These results demonstrate that union dissolution increased inequality among separating women, but this inequality did not accumulate over women's life courses. They also demonstrate how demographic events at the micro level can be connected to outcomes at the macro level.

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Introduction

Women experience sizeable income losses following union dissolution. Average income losses range from 13% to 40% (for the United States, see Avellar & Smock, 2005; Tach & Eads, 2015; Holden & Smock, 1991; for other countries, see Andreß et al., 2006; De Vaus et al., 2017). Especially compared to men, who experience little change, these figures affirm that dissolution is an event with dramatic consequences for women’s standard of living (Leopold, 2018).

The consequences of dissolution might be particularly harmful for women from low-income unions. Women from low-income unions face more difficulties in expanding labor supply, benefit less from repartnering, and are more likely to have physical custody over children (Cancian et al., 2014; Dewilde & Uunk, 2008). As a result, they might be less able to compensate the loss of household income. This has raised the concern that union dissolution could feed into a process of cumulative disadvantage, potentially increasing inequality among all separated women (Hogendoorn et al., in press).

Despite these concerns, little is known about the consequences of union dissolution for income inequality. Several reasons could be at play. First of all, declines in household income often precede dissolution (Boheim & Ermisch, 2001). This renders it difficult to disentangle ongoing income declines from those following dissolution. In addition, the links between differential consequences at the micro level and aggregate inequality at the macro level are ill understood. Group differentials need not carry over to the aggregate, as many women lose more or less than their income group does on average (DiPrete & Eirich, 2006). Methods to address these issues have proliferated only recently.

This study therefore describes the consequences of union dissolution for income inequality among separating women. I started from the idea that dissolution could drive cumulative disadvantage. To test this idea, I used administrative data from the Netherlands, following women in coresidential unions from 2003 to 2015. These data were uniquely suited to the analysis as they covered almost the entire population of separating women and their household incomes. Fixed-effects individual-slopes regressions were used to disentangle dissolution from other characteristics that affected household income trajectories, showing how the consequences of dissolution differed by prior income. Recentered influence functions were used to estimate distributional effects (Firpo et al., 2009), showing the extent to which differential consequences were accompanied by changes in aggregate income inequality.

The results demonstrate complex dynamics in women’s standard of living. Union dissolution may differentially affect groups of women, but these group differentials need not translate to inequality among separated women in the aggregate. The reason is that many women experience income mobility away from their group average. In other words, the consequences of union dissolution are described by the interplay between income divergence, mobility, and inequality.

In this way, this study contributes to debates on demographic events and stratification. It shows that events play an important role for individual women as well as for inequality at large. Understanding such links between the micro level and the macro level is fundamental to sociodemographic analysis. Furthermore, its focus on the Netherlands may prove an interesting case. The Dutch administrative registers cover nearly all coresidential unions, which is important because of strong legal harmonization between marriage and cohabitation and widespread acceptance of unmarried cohabitation (Hiekel & Keizer, 2015; Perelli-Harris & Gassen, 2012). Yet, full coverage may also be informative for other countries, where marriage concerns an increasingly smaller and more selective section of the population.

Theory

An influential idea in demography is that of cumulative disadvantage. Many formulations of this idea exist, including the compensatory advantage model (Bernardi, 2014; DiPrete & Eirich, 2006). The model holds that individuals from advantaged backgrounds are better able to compensate for the negative consequences of life events, resulting in growing inequality over the life course. A clear example has been provided by Crystal and Waehrer (1996). Their study has shown how higher-income individuals are better able to secure private pensions upon retirement, resulting in stark increases in income inequality among retired individuals.

The idea of cumulative disadvantage can be applied to union dissolution as well. Women stand to lose substantially from dissolution, as they earn less than their partners across all household incomes (Winslow-Bowe, 2006). Yet, women from high-income unions may be able to compensate some immediate losses. They are less likely to take physical custody of children after dissolution (Cancian et al., 2014). They also tend to be better educated, active in the labor force, and able to increase working hours (Jansen et al., 2009). They further repartner as quickly as women from low-income unions, but their new partners tend to have higher incomes (Dewilde & Uunk, 2008; Shafer & James, 2013). Other factors might go in the opposite direction, but their importance is difficult to determine a priori. For example, the role of taxes and transfers depends on the extent to which separated women become eligible (Tach & Eads, 2015). Similarly, child support and partner alimony could compensate some losses, but amounts are typically small and noncompliance widespread (De Vaus et al., 2017; Huang et al., 2005). These arguments suggest a divergence in women’s standard of living.

Besides income divergence, union dissolution may induce further mobility. That is, even if average household incomes diverge between women from low- and high-income unions, there may be considerable variation within each groups. Previous work confirms this. Whereas the majority of women lose, a minority actually gains (Ananat & Michaels, 2008; Bratberg & Tjøtta, 2008). Moreover, the variation in income losses might differ by predissolution household income. Women from low-income unions might have little scope for mobility, since taxes and transfers bound the very bottom and dampen upward mobility, although repartnering might occasionally lift them up. Women from middle- and high-income unions might be more mobile, since they are less subject to hikes in taxes and transfers and since they have more latitude to adjust labor supply (Bradbury & Katz, 2002). Hence, the consequences of dissolution cannot be fully described by group divergence between women from low- and high-income unions. Doing so would mask any underlying mobility within these groups.

As a result, income divergence need not carry over to aggregate inequality. To illustrate, consider that mobility is a form of "noise". This noise alters the income distribution as a whole. The distribution could become more or less unequal and this could depend on where in the distribution inequality is measured. For example, greater average losses of low-income women could coexist with a relatively large fraction of high-income women moving to the middle. Union dissolution would then differentiate group averages while reducing aggregate inequality. Conversely, greater average losses of high-income women could coexist with a relatively large fraction retaining their incomes. Union dissolution would then equalize group averages while increasing aggregate inequality. These examples illustrate that income divergence between groups does not imply increased inequality at the macro level (for a technical discussion, see Cheng, 2014). Therefore, an assessment of cumulative disadvantage requires attention to divergence, mobility, and inequality as connected but distinct processes.

Empirical evidence on divergence, mobility, and inequality is limited. Regarding divergence, studies have found that income losses differ little by predissolution household income or are still greater for women from high-income unions (Fisher & Low, 2016; Jarvis & Jenkins, 1999; Uunk, 2004; Weiss, 1984). Regarding mobility, studies have suggested that most separated women move to a lower income quintile, though a moderate fraction move up to the middle and a small fraction to the top (Bradbury & Katz, 2002; Gittleman & Joyce, 1999). Regarding inequality, a study has shown that income inequality is larger among separated than partnered women and that increases in separation entail small increases in aggregate inequality (Martin, 2006). However, these studies did not account for unobserved heterogeneity in income trajectories. Dissolution may be selective of households at a downward income trajectory, and this trajectory should not be conflated with the losses from dissolution (Boheim & Ermisch, 2001). In addition, they did not directly link individual dissolutions to aggregate inequality among separated women. Direct linking would provide better insight in the effects of micro-level events on macro-level outcomes. Lastly, they focused on different aspects of cumulative disadvantage. The present study aims to bring these different aspects together. In the next sections, I describe the consequences of union dissolution for income divergence, mobility patterns, and aggregate inequality in women's standard of living.

Data and method

Data

I used longitudinal data from the Dutch administrative registers. These data covered all individuals with legal residence in the Netherlands and combine information from the population register, education registers, social insurance bank, and revenue service. I examined unions formed between 2003 and 2005 and followed their incomes for ten years, the period during which incomes were available without break in definition. Access to the data can be requested via Statistics Netherlands (https://www.cbs.nl/en-gb/our-services/customised-services-microdata).

The study population was selected from the cohabitation file. The cohabitation file is a unique administrative file with information of all married and unmarried cohabiting couples in the Netherlands. Couples are identified by joint residential moves, as well as by marriage, civil union, children, joint welfare transfers, or joint taxation, which is available to all cohabiting couples. In this way, the cohabitation file achieves very high coverage and low misclassification. The only unions that are not covered are those that never moved, are not registered, have no children, and have not used any income provisions.

I selected all women aged 21 to 35 at first union formation (N = 222,114). The lower bound represents the age until which individuals can legally claim maintenance from their parents, the upper bound the age at which most first unions in the Netherlands have formed (Mulder et al., 2006). Experiments with higher age bounds did not change the results. I restricted the population to women outside of full-time education, because the incomes of students are little indicative of economic well-being (N = 183,790). I further restricted the population to women whose complete union spell took place in the Netherlands, because union episodes abroad were not registered (N = 171,568). Finally, I selected women who separated between two and ten years after union formation (N = 38,829). The lower bound was necessary because the estimator required at least two time points before dissolution. The upper bound ensured that the dissolution effect was estimated among ever-separating women only, a group more comparable than ever- and never-separating women combined. These women were followed from union formation through union dissolution. This yielded a total sample size of 234,901 person-year observations nested in 38,829 persons. After list-wise deletion of observations with missing values, the analytic sample consisted of 208,308 person-year observations nested in 38,059 persons.

Table 1 describes the sample at union formation. The average woman was aged in their mid-twenties (26.00), one in eight were born abroad (12%), completed education was fairly average (14.91), and most were in employment (88%). Few unions started by marriage (17%), confirming the acceptance of cohabitation in the Netherlands. Households had relatively few members (2.12) and few resident children (0.18). Household disposable income was quite average (25,000 EUR; 31,000 USD in 2015 after inflation and purchasing power correction).

Measures

Union dissolution was measured as separation from the household by a single move or by a dual move into two different households. Dissolutions by death (0.7%) or widowhood (1.2%) were censored. *Household disposable income* was measured as the annual sum of earnings, business income, and property income of all household members, after taxes and transfers. This included partner alimony, which is registered for by the Dutch revenue service, though not child support. Income was top-coded at one million euros (0.1%) and bottom-coded at zero (0.6%). To account for economies of scale, it was adjusted using the square root equivalence scale. This equivalence scale assigned each household member the total household disposable income divided by the square root of household size, and is widely used in income research (e.g. Atkinson et al., 1995; Solt, 2016). All incomes were inflated to their 2015 values. *Time since union formation* was measured as the number of years since the start of cohabitation.

Time intervals were specified in years because income taxes were filed annually. All variables were measured by the end of each calendar year.

Table 1: Descriptive statistics at the end of the union formation year

	M	*SD*	*min*	*max*	*N*
Individual characteristics					
Age	26.00	3.70	21	35.92	31,816
Foreign-born	0.12		0	1	31,816
Education years	14.91	3.31	2	22	24,066
Employment status					
Employed	0.88		0	1	31,816
Unemployed with benefits	0.06		0	1	31,816
Other non-employed	0.06		0	1	31,816
Union characteristics					
Cohort	2003.97		2003	2005	31,816
Married	0.17		0	1	31,816
Household characteristics					
Size	2.12	0.82	1	17	31,816
Children	0.18	0.50	0	6	31,816
Disposable income	24,922	13,932	0	707,107	31,816
Distribution characteristics					
Gini coefficient	0.23				31,816
P50/P10 quantile ratio	1.90				31,816
P90/P50 quantile ratio	1.50				31,816

Notes: Variables measured by the end of the calendar year. The number of observations is less than the total number of women (N = 38,059), because some women had missing values at union formation.

Income inequality was measured using three variables. The *Gini coefficient* measured the (rescaled) average income difference between all pairs of women. This variable indicated income inequality overall. *P50/P10 quantile ratio* measured the ratio between median income and income at the 10th percentile. This variable indicated inequality in the lower half of the income distribution. The *P90/P50 quantile ratio* measured the ratio between income at the 90th percentile and median income. This variable indicated inequality in the upper half of the distribution. All variables were based on household disposable income.

I included the following control variables. *Marital status* was a binary indicator of formal marriage. *Children* was measured as the number of resident children by the end of the year. *Labor market status* was measured as employed, unemployed with benefits, and other non-employed. Unemployed women who did not request or were not entitled to benefits were categorized as other non-employed, since they were not registered at the social insurance bank. All three variables were lagged one year to reduce simultaneity with income. *Unemployment rate* was measured as the annual rate of unemployment in the total working population. It adjusted the correlation between income and dissolution for periodic fluctuations due to the economic cycle. Period effects proper were not included because of their linear dependency on union cohort and union duration. Time-invariant and time-linear variables, such as union cohort, ethnicity, or age, were also not included as they were captured by the fixed-effects individual-slopes model (see next section).

Analytic strategy

Throughout the analysis, I modeled incomes using fixed effects individual slopes (FEIS). FEIS models are an extension to fixed-effects models, whereby each person receives not only an individual intercept but also an individual slope (Ludwig & Brüderl, 2018). The individual intercepts and slopes account for the fact that time-invariant unobserved heterogeneity may manifest itself not only in outcome levels but also in outcome trajectories. For example, dissolution might be selective of unions who experienced a downward income trajectory, and this downward trajectory should not be conflated with the losses caused by the dissolution event itself. FEIS estimates the dissolution effect as the average deviation from all individual trajectories of household income.

The FEIS estimates should be interpreted as average treatment effects on the treated. They tell us what would have happened to separating women if they had not separated. That is, these models compared women's observed household incomes upon dissolution to a counterfactual situation in which their household incomes would have continued as before. Note that this counterfactual required at least two years before dissolution. For this reason, the analytic sample comprised women who separated after two years or later, followed from union formation through dissolution.

The analysis proceeded in three steps. In the first step, I examined the consequences of union dissolution for income divergence. To do so, I regressed household incomes on dissolution and its interaction with predissolution income. Income was logged to obtain relative changes and one added to include women without income. The regression model was as follows:

$$\ln(Y_{it} + 1) = \alpha_{1i} + \alpha_{2i}T_{it} + \boldsymbol{\beta} D_{it}\boldsymbol{P_i^k} + \boldsymbol{\gamma X_{it}} + \varepsilon_{it}$$

where Y_{it} represented household disposable income, α_{1i} represented the individual intercept, α_{2i} the individual slope, T_{it} linear time since union formation, D_{it} a dummy for the dissolution event, $\boldsymbol{P_i}$ a set of dummies for predissolution income quintiles k, $\boldsymbol{X_{it}}$ a set of time-varying control variables, and ε_{it} an idiosyncratic error term. In this model, each woman followed her own household income trajectory. The trajectory was disruption by dissolution, and this disruption was allowed to vary according to the woman's initial income position. Hence, the coefficients $\boldsymbol{\beta}$ estimated the differential consequences of union dissolution for income.

In the second step, I examined the consequences of union dissolution for mobility at large. To do so, I regressed women's income positions on dissolution and its interaction with predissolution income. Income positions were bracketed in quintiles. The regression model was as follows:

$$I[Q^{k-1} < Y_{it} \leq Q^k] = \alpha_{1i} + \alpha_{2i}T_{it} + \boldsymbol{\beta D_{it}^y P_i^k} + \boldsymbol{\gamma X_{it}} + \varepsilon_{it}$$

The model was similar to that in the previous step, except that it substituted income by indicators for each income quintile Q^k. It estimated changes in women's income positions while taking account of changes that were already underway. Hence, the coefficients $\boldsymbol{\beta}$ estimated the consequences of union dissolution for mobility.

In the third step, I examined the consequences of union dissolution for aggregate inequality. To do so, I regressed recentered influence functions (RIFs) on dissolution and its interaction with predissolution income. Influence functions give the influence of each individual observation on a distributional measure. Adding back the sample average ("recentering") renders a variable whose expected value equals the expected value of that distributional measure (Firpo et al., 2009). For example, the expected value of the Gini coefficient's RIF is the expected value of the Gini coefficient itself. RIFs can be regressed on a set of covariates, including individual intercepts and slopes, to estimate the effects of those covariates on the distributional measure (Killewald & Bearak, 2014). The entire procedure is then bootstrapped to estimate the standard errors. Here, the regression model was as follows:

$$RIF^m(Y_{it}) = \alpha_{1i} + \alpha_{2i}T_{it} + \beta D_{it} + \boldsymbol{\gamma X_{it}} + \varepsilon_{it}$$

This model was similar to that in the previous steps, except that it substituted income by the RIF of each inequality measure m and omitted the interactions with predissolution income. It estimated the change in income inequality while taking account of changes that were already underway. Hence, the coefficient β estimated the consequences of union dissolution for income inequality in the aggregate.

Income quintiles and RIFs were measures of relative household income. They should measure women's relative positions even if the income distribution changed over time. Therefore, I temporarily expanded the sample to include all person-year observations after dissolution, with each woman's last observation being the tenth year after union formation. Next, I defined the income quintiles and the RIFs within each union duration. This implies that a woman's reference group consisted of other women who started cohabiting just as long ago. I then classified women into predissolution income groups according to their average income position across the years before dissolution. This was preferred over the income position in a single year, because annual incomes fluctuated substantially.

Results

Descriptive results

Figure 1 describes household income trajectories. The curves show similar trajectories prior to dissolution, with women from different income groups experiencing more or less the same growth. These trajectories were disrupted upon dissolution, with women from higher groups losing more. For instance, the average income loss was 6% for women in the middle quintile, compared to 12% for the fourth quintile, and 14% for the fifth quintile. Women in the lower quintiles lost little or even gained. This convergence suggests that, instead of accumulating disadvantage, union dissolution strongly equalized the averages of different income groups.

Figure 2 describes the underlying mobility patterns. The heat map shows considerable variation in the transition from predissolution to postdissolution income position, with the majority of women moving downward. Downward moves increased toward higher-income unions, yet a substantial share of women at the top retained their position. For instance, downward moves of at least twenty percentiles occurred among 16% of women from the second quintile, 26% from the third quintile, 32% from the fourth quintile, yet only 25% from the top quintile. These downward moves were partially compensated by (large) upward moves from the lower and middle quintiles. This suggests that union dissolution altered the income distribution, raising the possibility that group convergence did not carry over to aggregate inequality.

Figure 1: Income trajectories through union dissolution

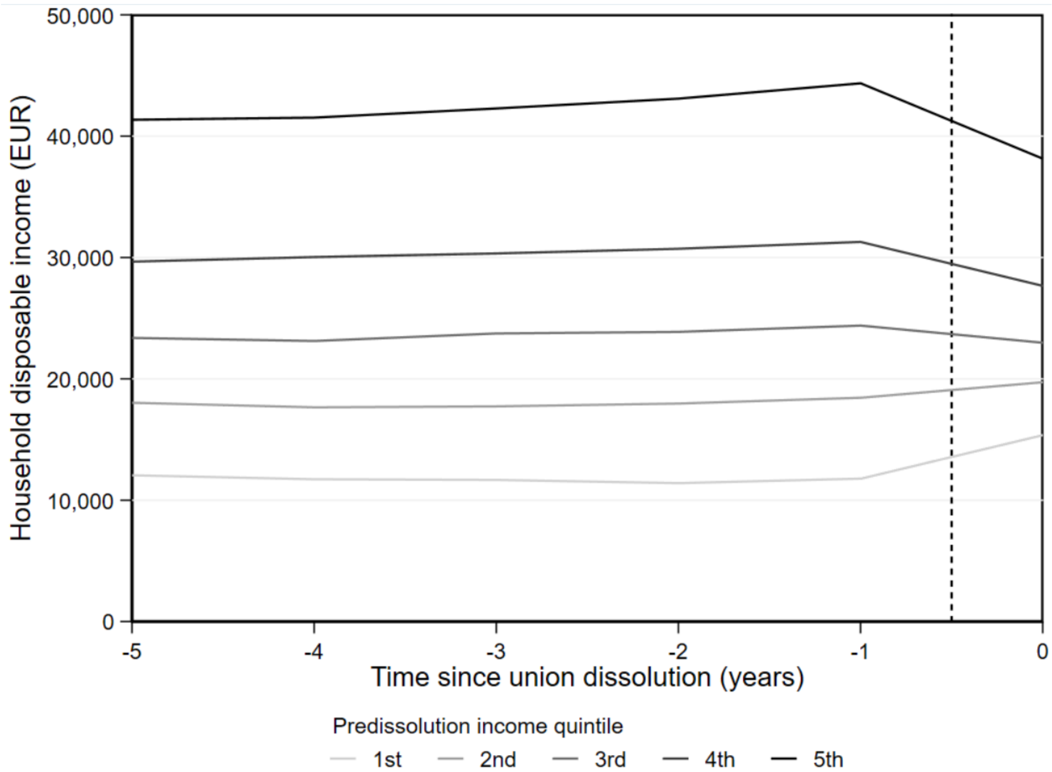
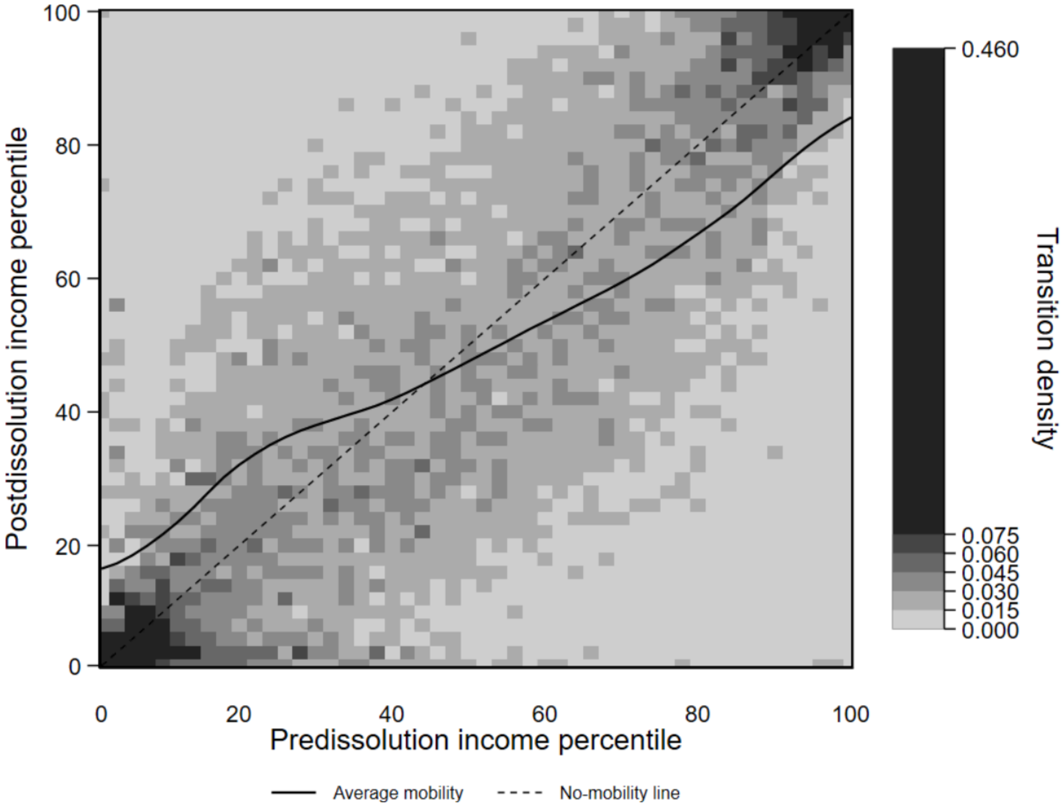
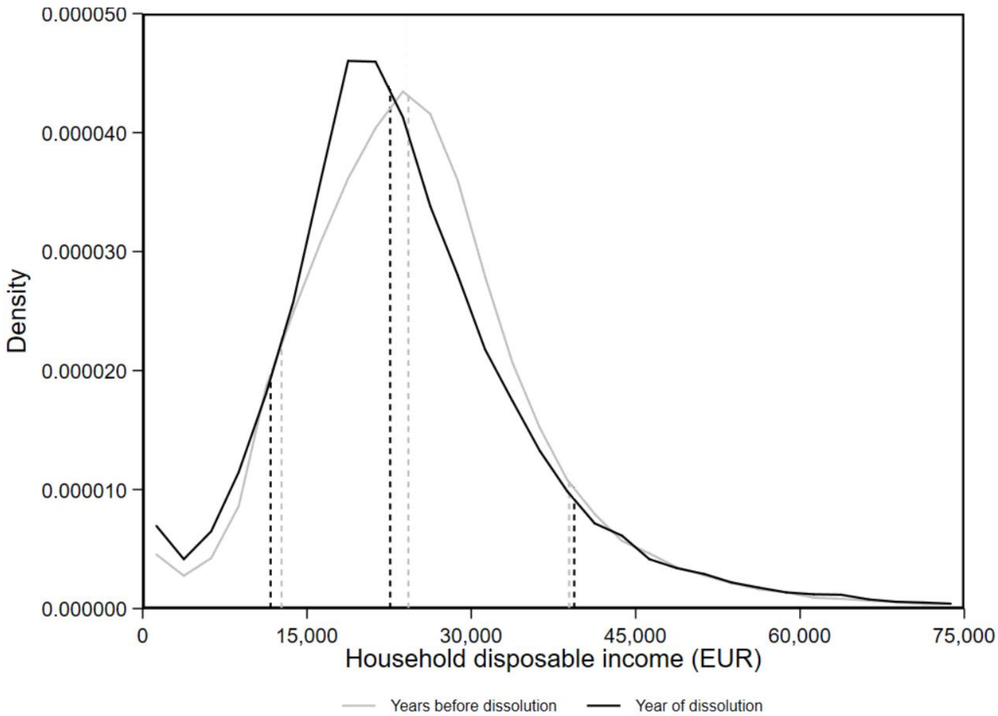


Figure 2: Income mobility upon union dissolution



Notes: Transition densities conditional on predissolution income percentiles.

Figure 3: Income distributions upon union dissolution



Notes: The dashed lines indicate the 10^{th}, 50^{th}, and 90^{th} percentile of each distribution.

Figure 3 describes the resulting income distribution. The histograms show a fairly symmetric distribution prior to dissolution, with most women having incomes close to the sample median. The distribution shifted to the left upon dissolution, and this shift was more pronounced in the left tail and the middle than in the right tail. For instance, the Gini coefficient increased from 0.24 to 0.27, the ratio between the 50th and the 10th percentile increased from 1.89 to 1.92, and the ratio between the 90th and the 50th percentile increased from 1.60 to 1.74. These figures suggest that, even though group averages converged, aggregate inequality increased. This might seem counterintuitive, but it must be borne in mind that union dissolution was a noisy event that induced mobility across the income distribution. I formally analyzed these descriptive results in the next section.

Regression results

In the first step of the analysis, I examined the differential consequences of union dissolution. The results confirmed that dissolution had negative consequences. This is illustrated by the left column of Table 2, which shows the coefficient of dissolution on household disposable income. On average, incomes dropped by 13% ($e^b = 0.87$). This means that women's standard of living declined compared to a scenario in which they would not have separated.

Nonetheless, union dissolution did not result in income divergence. It rather benefited women from the poorest households and disadvantaged women from richer households. This is illustrated by the right column of Table 2, which shows the coefficients of dissolution interacted with predissolution income quintiles on household disposable income. Incomes improved by 20% among women in the bottom quintile, remained stable in the second quintile, dropped by 15% in the third quintile, dropped by 21% in the fourth quintile, and dropped by 24% in the top quintile. This implies that women's standards of living converged upon union dissolution.

In the second step of the analysis, I examined mobility patterns beyond group averages. The results reaffirmed that union dissolution induced more downward mobility among women from higher-income unions. This is illustrated by the cells below the diagonal in Table 3, which show the coefficients of dissolution interacted with predissolution income quintiles on lower income quintiles. The probability of moving down at least one quintile increased by 7 percent points for women from the second quintile, compared to 21 points from the third quintile, 34 points from the fourth quintile, and 34 points from the top quintile. Nevertheless, women from the top quintile rarely fell below the median. This indicates that downward moves mainly befell women in the middle part of the income distribution.

These downward moves were partially offset by upward moves. This is illustrated by the cells above the diagonal in Table 3, which show the coefficients of dissolution interacted with predissolution income quintiles on higher income quintiles. The probability of moving up one quintile or more increased by 35 percent points for women from the bottom quintile, by 19 points from the second quintile, by 1 points from the third quintile, but decreased by 5 points from the fourth quintile. Moreover, a small portion of women from the lower three quintiles landed in the top quintile. This suggests that the combination of mobility patterns resulted in a reshuffling of the distribution's lower half, a downgrading of women in the upper half, and their partial replacement by women from the lower half.

Table 2: Regressions of log income on union dissolution

	General		Differential	
	b	**SE**	**b**	**SE**
Dissolution event	0.87***	0.01		
x predissolution Q1			1.20***	0.05
x predissolution Q2			1.00	0.02
x predissolution Q3			0.85***	0.01
x predissolution Q4			0.79***	0.01
x predissolution Q5			0.76***	0.01
Time-varying controls	✓		✓	
Individual intercepts	✓		✓	
Individual slopes	✓		✓	
χ^2 model fit	533		1,274	
N observations	208,308		208,308	
N individuals	38,059		38,059	

Notes: Coefficients were exponentiated to show the relative change in income. Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Standard errors accounted for clustering at the individual level. *** $p < .001$

Table 3: Regressions of income position on union dissolution

	Q1	**Q2**	**Q3**	**Q4**	**Q5**
Dissolution event					
x predissolution Q1	-0.35	+0.14	+0.14	+0.06	+0.02
x predissolution Q2	+0.07	-0.26	+0.04	+0.10	+0.05
x predissolution Q3	+0.12	+0.09	-0.22	-0.05	+0.06
x predissolution Q4	+0.07	+0.14	+0.13	-0.28	-0.05
x predissolution Q5	+0.04	+0.06	+0.11	+0.13	-0.33
Time-varying controls	✓	✓	✓	✓	✓
Individual intercepts	✓	✓	✓	✓	✓
Individual slopes	✓	✓	✓	✓	✓
χ^2 model fit	2,386	2,004	1,598	1,561	1,927
N observations	208,308	208,308	208,308	208,308	208,308
N individuals	38,059	38,059	38,059	38,059	38,059

Notes: Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Standard errors accounted for clustering at the individual level. All standard errors smaller than 0.012 (not shown for readability). All coefficients statistically significant at $p < .001$.

In the third step of the analysis, I examined how these mobility patterns played out for the aggregate inequality. The results revealed that aggregate inequality increased. This is illustrated by the left column of Table 4, which shows the coefficient of dissolution on the RIF of the Gini coefficient. The Gini coefficient increased by 0.02. The increase was moderately small, given the predissolution value of 0.24. This demonstrates that, despite the strong convergence between different income groups, overall inequality among separating women increased somewhat.

The increase in inequality was limited to the upper half of the income distribution. This is illustrated by the middle and right columns of Table 4, which show the coefficients of dissolution on the RIFs of the quantile ratios. The P50/P10 quantile ratio decreased by 0.03, whereas the P90/P50 quantile ratio increased by 0.06. In other words, the distance from the 10th income percentile to the median decreased whereas the distance from the median to the 90th percentile increased. The changes were small, given the predissolution values of 1.91 and 1.60, respectively. This shows that the increase in aggregate inequality was notable only in the upper half of the income distribution.

All in all, the links between union dissolution and income inequality can be described as follows. Women experienced moderately large income losses upon union dissolution. These losses followed a pattern of convergence rather than divergence, as they were concentrated in women from middle- and high-income unions. At the same time, aggregate inequality increased, albeit only in the upper half of the income distribution. This paradoxical situation occurred because most women moved down yet some climbed to the top. Put differently, union dissolution increased inequality among separating women, but this inequality did not accumulate over women's life courses.

Additional analyses

I conducted three additional analyses. The results are available in the Appendix. The first analysis concerned income trajectories. In the previous sections, I examined the dissolution effect as the contrast between observed incomes and counterfactual incomes if predissolution trajectories had continued. This raises the question of what these trajectories looked like. Hence, I repeated the estimation of log income and recovered the individual intercepts and individual slopes (Table A1). This showed that income groups differed greatly in income level but hardly in income growth. Consequently, group differentials in income losses could not be attributed to group differentials in counterfactual income growth to which these losses were contrasted. Women from higher-income unions simply lost more upon dissolution.

Table 4: Regressions of recentered influence functions on union dissolution

	Gini		**P50/P10**		**P90/P50**	
	b	**SE**	**b**	**SE**	**b**	**SE**
Dissolution event	0.02***	0.00	-0.03*	0.01	0.06***	0.01
Time-varying controls	✓		✓		✓	
Individual intercepts	✓		✓		✓	
Individual slopes	✓		✓		✓	
χ^2 model fit	139		185		134	
N observations	208,308		208,308		208,308	
N individuals	38,059		38,059		38,059	

Notes: Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Standard errors were obtained using a cluster bootstrap at the individual level with 500 replications. *** $p < .001$, * $p < .05$

The second analysis concerned income convergence. In the previous sections, I showed that union dissolution was more consequential for women from higher-income unions. This raises the question of what explains these group differentials. Although a full explanation is outside the scope of this paper, I repeated the estimation of log income. Next, I fitted several models to capture the group differentials in income losses in one parameter. This showed that the differentials were well captured by the square root of the income quintile. Then, I estimated a mediation model to explain these differentials (Table A2). This showed that group differentials could not be explained by shared custody, reemployment, or repartnering, because these factors hardly compensated the effect of dissolution on disposable income. Therefore, I also repeated the estimation using different income concepts (Table A3). This showed that, before taxes and transfers, income losses were actually larger among women from lower-income unions. Income transfers turned the picture upside down, improving the situation of all women but especially that of women from low-income unions.

The third analysis concerned the medium-term consequences. In the previous sections, I examined the short-term consequences of union dissolution. This examination did not show whether group convergence and increased aggregate inequality persisted over time. Hence, I repeated the analysis through five years after dissolution (Figures A1-A2 and Table A4). This showed that dissolution had an instantaneous effect on income levels but not on subsequent income growth. Thus, the convergence brought about by dissolution persisted in the years after. This convergence was accompanied by mobility toward the middle and upper middle quintiles. Consequently, aggregate inequality reverted to its original trend. Inequality in the lower half of the distribution even fell more than it would have in the absence of dissolution.

Discussion

Union dissolution has severe consequences for women's standard of living. It could be particularly harmful for women from low-income unions, who are less well-equipped to compensate its consequences (Cancian et al., 2014; Dewilde & Uunk, 2008; Jansen et al., 2009). This raises the concern that union dissolution could feed into a process of cumulative disadvantage, potentially increasing inequality.

In this study, I described the consequences of union dissolution for income inequality among separating women. Using longitudinal data from the Netherlands (N = 38,059), I compared women's household incomes upon dissolution to a counterfactual scenario in which incomes continued along their predissolution trajectories. The results showed that the living standards of women from low-income unions improved, whereas those of women from high-income unions declined. At the same time, aggregate inequality increased somewhat. This was due to a combination of moderate downward mobility by most women and strong upward mobility by some women from low-income unions.

These results are surprising. First, they show that union dissolution does not harm the living standards of all women. Household incomes dropped by 13% on average but increased by 20% among women from low-income unions. Second, they show that dissolution does not feed into a process of cumulative disadvantage. Dissolution rather results in strong convergence between women from different income groups. Third, they show that convergence between groups can coexist with increased inequality in the aggregate. This demonstrates that care must be taken when formulating cumulative-disadvantage or "dissolution-as equalizer" arguments.

What explains the convergence between income groups? The answer probably regards the welfare state. Additional analyses showed that income losses before taxes and transfer followed a pattern of cumulative disadvantage. Public transfers, and to a lesser extent progressive taxation, turned the picture upside down. This reflects the Dutch one-and-a-half breadwinner model, whereby women in low-income unions often work part-time because of high childcare cost, means-tested childcare subsidies, and tax deductions for breadwinners with dependent partners (Evertsson et al., 2009). Union dissolution renders these women eligible to more childcare subsidies and fiscally rewards their labor supply, while social assistance cushions those with poor labor market prospects (Jansen et al., 2009).

This argument suggests that the consequences of union dissolution for inequality are context dependent. Incomes may converge in countries that promote a strong one-and-a-half breadwinner model, including Switzerland and the United Kingdom, or that provide generous minimum incomes for single-headed households, including Belgium and Denmark (Ciccia & Bleijenbergh, 2014; Wang & Van Vliet, 2016). Incomes may diverge in countries with limited employment opportunities and poor income provisions, including most countries in Southern Europe. A dual mode may exist in countries that rely heavily on in-work benefits, which protect women with low to moderate earnings but do not cover women without earnings, as is the case in Ireland, the United States, and to a lesser extent the United Kingdom (Immervoll & Pearson, 2009). Similarly, increases in aggregate inequality depend on the progressivity of the tax-benefit system. Welfare transfers and progressive taxation substantially compress the income distribution of the Netherlands, as they do in Belgium and the Nordic countries. The tax-benefit system is weaker the United Kingdom and the United States, where market inequality remains largely untouched (Smeeding, 2005).

A fuller picture of inequality also requires attention to women who did not separate. In this study, I examined the differential consequences of a given dissolution, showing that the consequences are greater for women from high-income unions. However, the risk of experiencing a dissolution in the first place is differentially distributed as well, as dissolution is more common among women from low-income unions (Ishizuka, 2018; Kalmijn et al., 2007). Risks and consequences may thus be countervailing forces. An assessment of dissolution and inequality among the entire population of partnered women, both separated and nonseparated, should therefore incorporate risks as well as consequences (Hogendoorn et al., in press).

Nonetheless, this study presents a step forward in the analysis of demographic events and social inequality. It provides an approach to link micro-level events to macro-level outcomes. Conceptually, these links consist in mobility. Statistically, they are modelled using influence functions. The approach easily extends to other events and outcomes, such as childbirth and earnings inequality among mothers, migration and income inequality among tied movers, or retirement and wealth inequality among people in old age. This could aid future research on the connections between demography and stratification.

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Appendix

Table A1: Individual intercepts and slopes by predissolution income group

	Intercepts		**Slopes**	
	b	**SE**	**B**	**SE**
Predissolution Q1	9,882***	353	1.01	0.01
Predissolution Q2	17,596***	346	1.00	0.01
Predissolution Q3	23,870***	381	1.01*	0.00
Predissolution Q4	29,849***	426	1.02***	0.00
Predissolution Q5	41,331***	585	1.02***	0.00
χ^2 test of group equality	8,439***		12*	
N individuals	38,059		38,059	

Notes: Coefficients were exponentiated to show absolute intercepts and growth rates. Individual intercepts and slopes were recovered from the fixed-effects individual-slopes regression of log income on the interaction between dissolution and predissolution income quintile, marital status, children, labor market status, and unemployment rate. Standard errors were obtained using a cluster bootstrap at the individual level with 500 replications. *** $p < .001$, * $p < .05$.

Table A2: Mediation of group differentials upon union dissolution

	No mediators		Children		Reemployment		Repartnering		Full model	
	b	**SE**	**b**	**SE**	**b**	**SE**	**b**	**SE**	**b**	**SE**
Dissolution event	1.63***	0.08	1.66***	0.09	1.61***	0.08	1.62***	0.08	1.65***	0.09
x predissolution income group (square root)	0.70***	0.02	0.70***	0.02	0.70***	0.02	0.70***	0.02	0.71***	0.02
x custody share			0.97	0.03					0.96	0.01
x change in working hours					1.29***	0.06			1.29***	0.17
x repartnered							1.04	0.02	1.05*	0.04
Time-varying controls	✓		✓		✓		✓		✓	
Individual intercepts	✓		✓		✓		✓		✓	
Individual slopes	✓		✓		✓		✓		✓	
Mediation			0.61%		1.77%		0.10		2.75	
t test of mediation			0.79		3.77***		1.12*		3.05**	
$\chi 2$ model fit	1,222		1,222		1,251		1,226		1,260	
N observations	207,638		207,638		207,638		207,638		207,638	
N individuals	38,010		38,010		38,010		38,010		38,010	

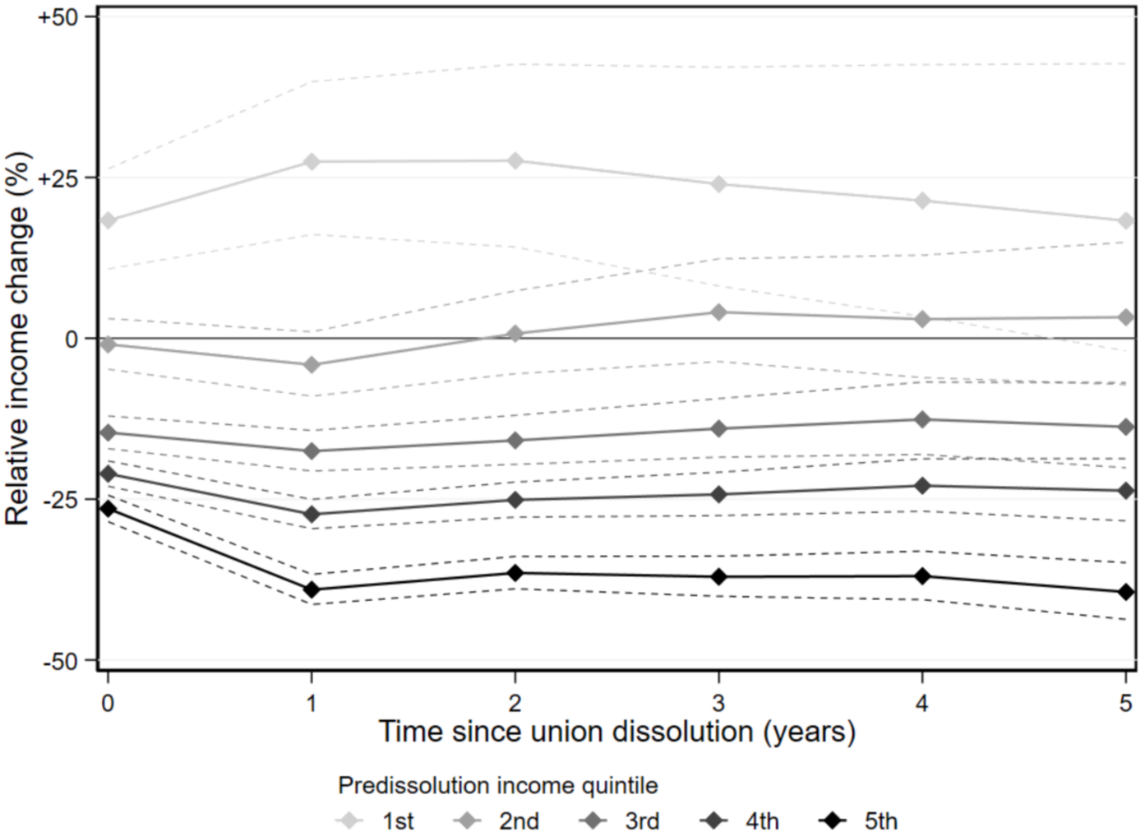
Notes: Coefficients were exponentiated to show the relative change in income. Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Mediation regards the change in the interaction effect between dissolution and predissolution income group. The number of observations is slightly smaller than in the main analysis because of list-wise deletion on more variables. *** $p < .001$.

Table A3: Regressions of log income on union dissolution under different income concepts

	Pretax pretransfer		Pretax posttransfer		Posttax posttransfer	
	b	**SE**	**b**	**SE**		
Dissolution event						
x predissolution Q1	0.62***	0.05	1.18***	0.04	1.20***	0.05
x predissolution Q2	0.74***	0.03	1.03	0.02	1.00	0.02
x predissolution Q3	0.77***	0.02	0.87***	0.01	0.85***	0.01
x predissolution Q4	0.78***	0.01	0.80***	0.01	0.79***	0.01
x predissolution Q5	0.75***	0.01	0.77***	0.01	0.76***	0.01
Time-varying controls	✓		✓		✓	
Individual intercepts	✓		✓		✓	
Individual slopes	✓		✓		✓	
χ^2 model fit	1,413		1,428		1,274	
N observations	208,308		208,308		208,308	
N individuals	38,059		38,059		38,059	

Notes: Coefficients were exponentiated to show the relative change in income. All outcomes were equivalized using the square root equivalence scale. Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Standard errors accounted for clustering at the individual level. *** $p < .001$

Figure A1: Changes in average incomes over time since union dissolution



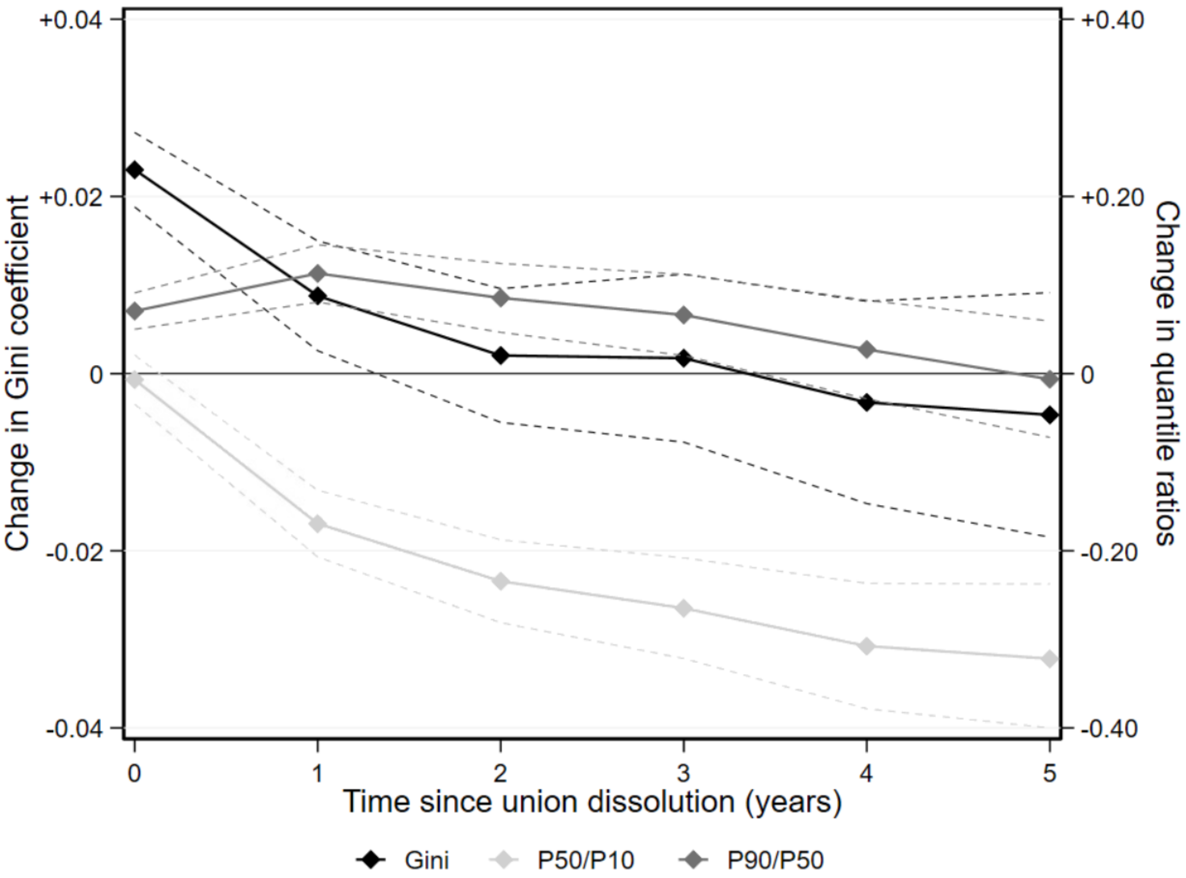
Notes: Coefficients were exponentiated to show the relative change in income. Dashed lines indicate 95% confidence intervals. Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Standard errors accounted for clustering at the individual level.

Table A4: Regressions of income position on union dissolution by time since dissolution

Time		Q1	Q2	Q3	Q4	Q5
0	Dissolution event					
	x predissolution Q1	-0.35	0.14	0.12	0.06	0.03
	x predissolution Q2	0.08	-0.28	0.04	0.10	0.06
	x predissolution Q3	0.12	0.10	-0.24	-0.04	0.07
	x predissolution Q4	0.07	0.16	0.11	-0.30	-0.05
	x predissolution Q5	0.04	0.08	0.12	0.12	-0.36
1	x predissolution Q1	-0.29	0.14	0.09	0.04	0.01[b]
	x predissolution Q2	0.19	-0.29	-0.02[a]	0.07	0.05
	x predissolution Q3	0.17	0.15	-0.29	-0.09	0.06
	x predissolution Q4	0.07	0.26	0.13	-0.36	-0.10
	x predissolution Q5	0.04	0.13	0.20	0.16	-0.52
2	x predissolution Q1	-0.34	0.15	0.12	0.04	0.03
	x predissolution Q2	0.15	-0.29	0.00[a]	0.07	0.07
	x predissolution Q3	0.14	0.15	-0.29	-0.09	0.08
	x predissolution Q4	0.05	0.23	0.16	-0.35	-0.09
	x predissolution Q5	0.01	0.11	0.19	0.18	-0.49
3	x predissolution Q1	-0.39	0.16	0.13	0.06	0.04
	x predissolution Q2	0.13	-0.29	-0.01[a]	0.09	0.08
	x predissolution Q3	0.12	0.15	-0.28	-0.09	0.10
	x predissolution Q4	0.03	0.21	0.16	-0.33	-0.08
	x predissolution Q5	0.00[a]	0.10	0.17	0.20	-0.46
4	x predissolution Q1	-0.44	0.19	0.14	0.06	0.05
	x predissolution Q2	0.11	-0.29	-0.01[a]	0.09	0.10
	x predissolution Q3	0.10	0.15	-0.26	-0.10	0.12
	x predissolution Q4	0.01[a]	0.21	0.19	-0.33	-0.08
	x predissolution Q5	-0.02[b]	0.07	0.17	0.24	-0.46
5	x predissolution Q1	-0.47	0.20	0.14	0.07	0.06
	x predissolution Q2	0.10	-0.31	-0.01[a]	0.10	0.12
	x predissolution Q3	0.09	0.16	-0.27	-0.11	0.12
	x predissolution Q4	-0.01	0.21	0.21	-0.32	-0.09
	x predissolution Q5	-0.04	0.07	0.17	0.26	-0.46
	Time-varying controls	✓	✓	✓	✓	✓
	Individual intercepts	✓	✓	✓	✓	✓
	Individual slopes	✓	✓	✓	✓	✓
	χ^2 model fit	5,145	4,673	3,587	2,998	4,321
	N observations	327,511	327,511	327,511	327,511	327,511
	N individuals	38,059	38,059	38,059	38,059	38,059

Notes: Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Standard errors accounted for clustering at the individual level. All standard errors were smaller than 0.02 (not shown for readability). [a] n.s., [b] p <.01, all other coefficients statistically significant at $p < .001$.

Figure A2: Changes in income inequality over time since union dissolution



Notes: Dashed lines indicate 95% confidence intervals. Models controlled for marital status, children, and labor market status at *t*-1 and for the national unemployment rate. Standard errors accounted for clustering at the individual level.