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Unions increase job satisfaction in the United States

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ABSTRACT

We revisit the well-known negative association between unionization and workers' job satisfaction in the United States, first identified over forty years ago. We find the association has disappeared since the Great Recession. The job satisfaction of both younger and older union workers in the National Longitudinal Surveys of 1979 and 1997 no longer differs compared to that of their non-union counterparts. When controlling for person fixed effects with panel data unionization is associated with *greater* job satisfaction throughout, suggesting that when one accounts for worker sorting into unionization, becoming unionized has always been associated with improvements in job satisfaction. We find a diminution in unions' ability to lower quit rates which is consistent with declining union effectiveness as a 'voice' mechanism for unionized workers. We also find unions are able to minimize covered workers' exposure to underemployment, a phenomenon that has increasingly negatively impacted non-union workers since the Great Recession.

1. Introduction

It was in the late 1970s that economists first identified that unionized workers were less satisfied with their jobs than their non-unionized counterparts. The empirical work conducted by Freeman (1978) and Borjas (1979) was for the United States. The finding appeared to be at odds with unions' ability to raise wages above those set in the market (Lewis, 1963; 1986) – something which surely would have raised union workers' satisfaction? However, the association persisted even when Freeman incorporated wages as a control variable. Freeman offered an explanation:

> *"At the 1975 [NBER] meetings, I suggested that the inverse relation might reflect the role of unions as a 'voice' institution, encouraging workers to express discontent during contract negotiations and to make formal grievances rather than to quit, which would keep the dissatisfied from leaving the employer. If this view is correct, the satisfaction relation lends some support to the exit-voice model of the union" (Freeman, 1978: 139-140)*

Freeman was referring to a model, originally devised by Hirschman (1970) to explain the behavior of consumers when faced with a defective product or deficient service. They could choose to voice their concerns in the hope of rectifying the problem or else leave the provider of the good or service (exit) and seek an improved product or service elsewhere. Applying this model to employment relations, Freeman (1978, 1980) (and subsequently Freeman and Medoff (1984)) suggested that

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unions, by helping to solve employees' problems at work, made them less likely to quit the workplace than similarly dissatisfied non-union workers so that unions appeared to increase the 'stock' of dissatisfied workers. Freeman (1980) attributed the lower quit rates among unionized workers to the availability of an effective grievance system in the union sector.

In recent decades the effectiveness of unions in securing workplace improvements has been called into question due to falling union density and the increased ability of employers to replace union labor with non-union labor both domestically and globally (Rosenfeld, 2014; Milkman, 2013; Schnabel, 2013). Poor productivity growth since the Great Recession may have further limited unions' ability to extract concessions from employers (Addison, 2020). Thus, we hypothesize that dissatisfied union workers may no longer believe their grievances can be resolved effectively, and instead choose to leave. This behavior would result in higher union worker job satisfaction and a reduced ability of unions to reduce quit rates.

Yet, post-Great Recession union workers still enjoyed the relative stability in work hours and job security offered by collective bargaining contracts. These contracts offer unionized workers insurance against the vicissitudes of the market, which tend to more adversely impact non-unionized workers, especially during periods of prolonged economic stagnation. The Great Recession offers a useful and novel circumstance to investigate whether unions preserve work hours relative to non-union workers in periods of economic upheaval. Consequently, we also test whether an improvement in job satisfaction of unionized workers relative to their non-unionized counterparts post-Great Recession may be due, in part, to unions' ability to maintain covered workers' hours and limit their under-employment, compared to those in the non-union sector.

In this paper we return to the National Longitudinal Survey (NLS) initially used by Freeman (1978) to see whether the relationship between unionization and job satisfaction has switched signs in the United States. Moreover, we examine two mechanisms influencing any such switch, namely whether union voice may be eroding and whether unions offered work hour preservation and protection from underemployment post-Great Recession. We run OLS estimates, as Freeman did, but we also run person fixed effects models to assess whether becoming unionized raises job satisfaction, having accounted for fixed unobserved differences between union and non-union workers which might potentially affect selection into unionization and workers' propensity for job satisfaction. This is a valuable addition to what Freeman (1978) did originally because, as many have argued, those who have a tendency towards dissatisfaction with life, and perhaps also with their jobs, may be more inclined to unionize than those who feel less dissatisfied. Again, following Freeman (1978), estimates are reported, for older and younger workers to see whether effects are robust across cohorts. The NLS79 includes workers born between 1957 and 1964, and those from the NLS97, workers born between 1980 and 1984. [1]

Using OLS estimation we confirm the negative association between unionization and job satisfaction in the years prior to the Great Recession but find the sign on the union coefficient switches positive after the Great Recession, although it is statistically non-significant. In person fixed effects panel estimates for the older NLS1979 cohort the association between union status and job satisfaction is positive both pre- and post-Great Recession, but the coefficient almost doubles in size (becoming more positive) post-Great Recession. For the younger NLS1997 cohort the positive union coefficient on job satisfaction becomes statistically significant for the first time in the post-Great Recession period. These fixed effects models confirm that switching union status significantly and positively impacts job satisfaction, especially in the post-Great Recession period. The fact that union coefficients in the fixed effects models are more positive than those in the OLS models, regardless of period, confirms workers with a greater propensity for dissatisfaction also have a greater propensity to be unionized.

The union effects on job satisfaction are quite sizeable and economically meaningful. Drawing from job satisfaction estimates in the literature that use NLSY data, our union coefficient is approximately equal in size to the positive impact on job satisfaction of having three employer-provided fringe benefits (Artz, 2010a), having a job in the public versus the private sector (Artz, 2017; Heywood and Wei, 2006), taking part in an employer-provided profit-sharing scheme (Heywood and Wei, 2006) and roughly half the effect of getting a promotion (Kosteas, 2011).

We then consider two reasons why the job satisfaction of unionized workers may have risen, relative to that of non-unionized workers, since the Great Recession. First, we test for a potential reduction in their ability to lower quit rates, a change that would be consistent with a decline in unions' ability to operate as an effective 'voice' for union workers. We find strong support for this proposition among the older workers in the NLS79 cohort, and more limited support in the younger NLS97 cohort.

Second, we explore the role that underemployment may have played in workers' job satisfaction since the Great Recession. Underemployment denotes circumstances in which workers' actual hours are below their desired hours. Here we define underemployment as part-time for economic reasons divided by employment. As Bell and Blanchflower (2019) have shown, the Great Recession led to a substantial rise in underemployment around the world, whereby workers desired far more hours of work than they were able to find in the labor market. To our knowledge, there has been no investigation into the role played by trade unions in protecting unionized workers from underemployment, despite the fact that maintaining one's hours of paid employment is an important aspect of income maintenance, especially in a period of sluggish wage growth.

If unions are successful at reducing unionized workers' exposure to underemployment post-recession, this may help explain an increase in their job satisfaction relative to that of non-unionized workers. We uncover evidence that unionized workers are less likely to be underemployed than non-unionized workers, and that the gap is counter-cyclical, rising in

[1] Later in the paper (Table 10) we also run probits using the US Gallup Daily Tracker data, 2009-2013 when job satisfaction is measured as a Yes/No variable.

periods of recession. Moreover, NLS79 data suggest that union workers enjoyed some insurance from work hour reductions post-Great Recession.

The remainder of the paper is set out as follows. Section Two reviews the existing literature, focusing on studies examining links between job satisfaction and unionization in the United States. Section Three presents our data and approach to estimation. Section Four presents our results, Section 5 discusses potential mechanisms that explain our results, and Section Six concludes.

2. Previous literature

Although trade unions as institutions vary somewhat across the countries of the world, there is a common acceptance that Freeman and Medoff's (1984) characterization of their 'voice' and 'monopoly' faces captures a large part of what they do at the workplace. These two faces of unionization have potentially countervailing effects on how workers feel about their jobs. On the one hand, unions' ability to monopolize the supply of labor to the employer and thus bargain for wage and non-wage conditions above those set in the absence of unions, should positively impact how covered workers feel about those jobs. This, in turn, should reduce their desire to quit unionized jobs since their outside options set in the market are likely to be inferior. Of course, there is the possibility for dissatisfaction to arise if workers' expectations about bargained outcomes are not met. But the literature is quite clear on this issue: unions continue to procure premia on wages and related conditions such as paid leave, despite some debate about decline in their bargaining power (Blanchflower and Bryson, 2004).

On the other hand, unions' 'voice' face might conceivably generate job dissatisfaction via what Freeman and Medoff (1984) termed 'voice-induced complaining', part of the process by which unions foster support from covered workers strengthening their hand in dealings with the employer, or through the increased flow of information that comes to unionized workers as a result of the two-way communication between management and employees which is a pre-requisite for unions in representing their members. This, together with the increased likelihood of dissatisfied workers remaining in the presence of union voice – as predicted under the exit, voice, loyalty model discussed above – results in cross-sectional estimates of the partial correlation between unionization and job dissatisfaction originally identified in the work of Freeman (1978) and Borjas (1979).[2]

In his original paper Freeman (1978) partly motivated the importance of job satisfaction as a variable that economists should be interested in by showing that it was a strong predictor of labor mobility. Using panel data on older men in the NLS Older Men and NLS Younger Men samples and the Michigan Panel Survey of Income Dynamics (PSID) Freeman showed job dissatisfaction predicted higher quit rates in the late 1960s and early 1970s (1978: 137). However, unionization was associated with both lower job satisfaction and lower quit rates in the Michigan PSID and NLS Older Male samples which, he argued, was consistent with unions providing effective voice for workers, thus reducing their quit probabilities for a given level of job dissatisfaction. He revisits the issue in Freeman (1980) paper using the same data sets plus the Current Population Survey (CPS), confirming reduced separation rates among unionized workers, particularly among the least satisfied workers. The reduction in quits among unionized workers is apparent having conditioned on wages (p. 666) which are intended to net out potential monopoly face benefits of unionization, thus isolating voice effects.[3]

Borjas (1979) comes to similar conclusions analyzing data from the National Longitudinal Survey of Mature Men aged 50-64. His estimates of union negative associations with job satisfaction are robust to his efforts to account for potential simultaneity using an instrumental variables approach. But he goes a stage further in testing the voice hypothesis for union dissatisfaction effects by showing the effects are strongest among workers with high tenure.[4] However, he adds to Freeman's analysis of quits by showing the union effect in reducing quits is largely confined to low tenured workers, a finding he attributes to the flatter wage profile faced by older workers arising from the seniority wage system promoted by unions.

Recently though, Blanchflower et al. (2022) found that the partial correlation between unionization and job dissatisfaction no longer holds in the United States and in Europe. Their estimates for the United States, based on cross-sectional data from the General Social Survey for the period 1972-2018, confirm Freeman (1978) and Borjas' (1979) finding of a negative correlation between unionization and job satisfaction in the early years. However, the correlation turned positive and statistically non-significant between 1998 and 2008 and became positive and statistically significant after the Great Recession. They confirm this positive, significant correlation for the period after the Great Recession in data from the Gallup Daily Tracker. Investigating why the change may have occurred, they present evidence from analyses of the General Social Survey indicating that unionized workers were more likely than their non-union counterparts to expect job loss in the period prior to the Great Recession, but that this was no longer the case after 2008. At the same time, union workers continued

[2] Borjas (1979: 21) expressly refers to the "politicization of the unionized labor force" as a contributory factor. In a study for the UK, Bender and Sloane (1998) confirm that unions negatively impact the climate of employment relations at the workplace and that, when one accounts for this, the association between unionization and job dissatisfaction is no longer significant. They conclude that "*union workers' relative dissatisfaction….stems from poor industrial relations or from unions forming where satisfaction would be low anyway*" (p. 222).

[3] A parallel literature using establishment-level data for Britain confirms a partial correlation between the presence of union voice and lower quit rates (Bryson et al., 2013a). The effect did not vary significantly over the period 1990-2004 despite declining within-establishment union density, something which may have impaired the effectiveness of union voice.

[4] Borjas' contention is that low tenure workers provide information to the firm about the problems workers face via exit whereas high tenure workers provide that information through expressed dissatisfaction (p. 30).

to benefit from a wage premium. They conclude: “This likely helps to explain the positive coefficient in the job satisfaction equations: union workers are less fearful of job loss than previously, yet they continue to receive the substantial wage premium they have always received” (p. 13).

Blanchflower et al. (2022) also consider the possibility that changes in the composition of union and non-union workers over time may have contributed to changes in the correlation between job satisfaction and union status. They do so by examining the partial correlation between job satisfaction and union status in the Gallup Daily Tracker data for the period 2009-2013 for members of different birth cohorts. They find early birth cohorts who would have made up most of the sample in Freeman and Borjas' studies in the 1970s continued to exhibit a negative union partial correlation with job satisfaction, whereas subsequent birth cohorts (born after 1959) exhibited a positive partial correlation.[5]

There is, as yet, no obvious explanation for the role of birth cohorts in the change in the partial correlation between job satisfaction and unionization, although the persistent negative correlation between job satisfaction and unionization among earlier birth cohorts may be related to Artz's (2010b; 2012) findings based on panel data for the period 1979–2004 from the National Longitudinal Survey of Youth which showed the negative association between unionization and job satisfaction increased with union experience. In any event, the finding points to the potential importance of accounting for compositional change in those becoming unionized.

One way to do this is to undertake panel estimation which focuses on changes in union status within individuals over time, thus accounting for fixed unobserved differences across union and non-union workers over time. There are a number of such estimates for the United Kingdom, and these tend to show the negative partial correlation between job satisfaction and union status is ameliorated and, in some cases, even switches sign with the inclusion of person fixed effects.

The most recent example is the only panel analysis presented in Blanchflower et al. (2022). These estimates, based on data from the British Household Panel Survey and its successor Understanding Society for the period 1996-2018, show the union partial correlation with job satisfaction is negative and statistically significant in OLS estimates but becomes positive and statistically significant when introducing person fixed effects which net out fixed unobserved differences between union and non-union workers.[6] However, the authors do not present estimates for early and later periods. There is no other recent panel evidence on the association between unionization and job satisfaction for the United States, so it is unclear what their incorporation might imply for the switch in the union partial correlation with job satisfaction apparent in OLS estimates. We address this issue below.

If the switch in the partial correlation between job satisfaction and unionization is linked to a diminution in union effectiveness as a ‘voice’ for unionized workers, a corollary might be a reduction in unions' capacity to lower quit rates. We investigate this issue below.

Finally, when considering how union effects on job satisfaction may have changed since the Great Recession it is important to recall the importance of changes in labor market trends since 2008. The discussion above noted the potential role played by unions in insuring against job loss, but another potential role unions might play relates to their ability to guarantee income security through the avoidance of underemployment. The issue of underemployment has come to the fore in the United States and elsewhere since the Great Recession (Bell and Blanchflower, 2019). In a period characterized by wage stagnation and recession-induced unemployment, it might not be surprising to find unions are bargaining to maintain covered workers' hours to ensure income security. Although the existing literature points to a positive partial correlation between unionization and satisfaction with hours worked (Bryson and White, 2016a, 2016b) nobody has investigated this issue to date.

3. Data and estimation

In accordance with the previous literature reviewed in Sections One and Two we estimate job satisfaction equations using Ordinary Least Squares (OLS) estimation supplemented with models incorporating person fixed effects. The fixed effects estimator holds constant all time-invariant unobserved individual characteristics that may affect selection into union status and job satisfaction, thus measuring changes in job satisfaction as an individual's union status changes. This approach requires longitudinal panel data tracking workers and their job satisfaction over time.

In the United States, the National Longitudinal Survey of Youth (NLSY) is the most widely used and respected panel containing both job satisfaction and union status measures, and it is one of two data sets Freeman (1978) used in his examination of unionization and job satisfaction. We analyze two separate cohorts from the NLSY. The first cohort (NLS79), born between 1957 and 1964, were first interviewed as teenagers in 1979, and were most recently reinterviewed in 2018. The second cohort (NLS97), born between 1980 and 1984, were first interviewed in 1997 and were most recently surveyed in 2017.

Both surveys initially interviewed individuals annually but shifted to a biennial survey (from 1994 in the case of the NLS79 and from 2011 in the case of the NLS97). In addition to union status and job satisfaction, both cohorts contain information on demographic and job characteristics which we include as covariates in our estimations. These include year,

[5] They subsequently confirmed the persistence in the negative partial correlation between job satisfaction and unionization in a UK birth cohort born in 1958 (Blanchflower and Bryson, 2022).

[6] For earlier similar results for the UK see Bryson and White (2016a, 2016b).

gender, race, age, region of residence, education level, and industry. Summary statistics separately for the 1979 and 1997 NLSY cohorts are presented in Table 1.

The way union status is identified in the NLS79 differs over time. The survey records union membership status in the period 1988 to 2018. Consequently, we limit our analysis involving union membership to those years alone. The NLS79 also records coverage by unions using two other questions that together span all the NLS79 waves. Between 1979 and 1993 workers were asked whether their wages were set by a collective bargaining agreement, and between 1994 and 2018 the survey asked whether workers were covered by union contracts. In creating the coverage indicator, we count both union members and those covered by collective bargaining agreements and union contracts as covered workers. The NLS97 survey, however, only asks whether respondents are covered by a contract negotiated by a union. Thus, coverage alone is used in the NLS97.[7] We confine our analyses to overall job satisfaction because satisfaction with facets of the job is only available in a sub-set of waves.[8]

The self-employed and those working in the military and private households are removed from the estimation sample to increase the comparability of employees entering the union and non-union samples. We also restrict the estimation sample to those in the private sector. Although unions loom large in government settings, their role and bargaining power are somewhat different to those in private sector settings. Moreover, in many instances public and private sector jobs consist of distinct tasks and employment relationships, making job satisfaction comparisons difficult. Finally, NLS respondents sometimes fail to acknowledge their region of residence, occupation or industry. The final job satisfaction estimation samples are approximately 159,000 and 65,000 in the NLS79 and NLS97 respectively.

In examining the mechanisms influencing the switch in union coefficient signs, we first measure the partial correlation between union status and the propensity to quit using NLS79 and NLS97. We run linear estimation models on the (0,1) voluntary quit outcome where 1 denotes all worker-initiated (voluntary) job separations except for family or pregnancy reasons. These separations specifically identify workers quitting to look for another job, take another job, or for other reasons: they exclude employer-initiated job loss.

Second, to measure whether union status protects workers from underemployment post-Great Recession, we leverage Gallup Daily Tracker data available after 2008 that contain large sample sizes including measures of underemployment concerns among workers. These data allow us to test the proposition that some of the higher job satisfaction among union workers relative to non-union workers may be their increased ability to avoid underemployment.

4. Results

We make use of two questions on job satisfaction that vary slightly between the NLS1979 and NLS1997. They are as follows

Q1. 1979 - How (do/did) you feel about (the job you have now/your most recent job)? (do/did) you like it very much (41%), like it fairly well (48%), dislike it somewhat (9%), or dislike it very much (3%)?

Q2. 1997 Which of the following best describes how you [feel/felt] about your {job_assignment}] [as/with your employer? = like it very much (35%), like it fairly well (31%), think it is ok (27%), dislike it somewhat (5%), dislike it very much (3%)

The numbers in parentheses are the overall distribution of responses averaged over the years. Of note is that in both sweeps small minorities of workers say they are dissatisfied with their jobs 12% in the NLS1979 and 8% in the NLS1997.

Fig. 1 presents mean job satisfaction for private sector covered and non-covered workers in the NLS79 over the period 1979-2018. Covered workers include all union members and those covered by collective bargaining agreements or union contracts. Among these workers, born between 1957 and 1964, mean job satisfaction is lower for the union than the non-union workers, apart from the period since 2008.

4.1. Main results

To assess whether the switch in the union association with a 4-step job satisfaction variable is statistically significant and robust to the introduction of controls, we present regression-adjusted estimates in Table 2. We code the dependent variables such that a positive coefficient means higher job satisfaction. Our union status variable is coded 1 if an individual is a union member. Due to limited availability of the union membership status variable, we restrict our sample to the years 1988–2018.

The first four columns of the table present pooled OLS cross-sectional estimates for the periods 1988–2004 and 2006–2018 respectively, with and without controls for year, education, race, gender, industry and region. Then in the final four columns OLS fixed effect estimates are reported with the same four specifications, with and without controls.

[7] Throughout the union status measures also encompass membership of “employee associations”. We cannot disentangle the two. We remove the 1994 wave of the NLS79 from the estimation sample due to inaccurate or incomplete union data in the wave arising due to recording errors in the union coverage variables.

[8] NLS79 records all workers' job satisfaction. However, NLS97 limits the question to those workers who have been at their employer for more than 12 weeks, thus reducing the estimation sample in the NLS97.

Table 1
Summary statistics – mean (standard deviation in parentheses).

NLSY variable descriptions	1979 cohort	1997 cohort
Job satisfaction: "How do you feel about the job you have now?" from 1 "dislike it very much" to 4 or 5 "like it very much"[$]	3.260 (0.743)	3.900 (1.040)
Quit: = 1 if respondent voluntarily quits their job, excluding for pregnancy or family reasons; 0 otherwise [*]	0.061 (0.239)	0.090 (0.286)
Union member: = 1 if respondent is a member of a union and 0 otherwise. [%]	0.104 (0.305)	N/a
Covered: = 1 if respondent is covered by a union contract or is a union member, and 0 otherwise[%]	0.141 (0.348)	0.084 (0.277)
Years / panel waves [@]	27 years	18 years
Observations [#]	158,081	66,181

Notes:
[$]Job satisfaction: 4 categories in the 1979 cohort and 5 categories in the 1997 cohort.
[*]We omit waves 1979–1982 as voluntary quits were particularly turbulent during these initial years in the 1979 cohort. We note the manuscript's results, and this summary statistic, are relatively unchanged when including these years.
[%]We omit the 1994 wave from the 1979 NLSY panel due to recording errors when collecting information regarding union status. Union status reflects union membership alone, and is only available in waves 1979 and 1988–2018. Covered combines union membership, collectively bargained wages, and coverage under a union bargained contract.
[&]The NLSY over-samples racial minorities.
[@]1979 cohort: annual waves from 1979 to 1993 and biennial thereafter until 2018; 12,395 total individuals in the unbalanced sample; average of 13 observations per individual.
1997 cohort: annual waves from 1997 to 2011 and biennial thereafter until 2017; 8402 individuals in the unbalanced sample; average of 8 observations per individual.
[#]We omit all private household employees, self-employed workers, military and public sector employees from the working sample.

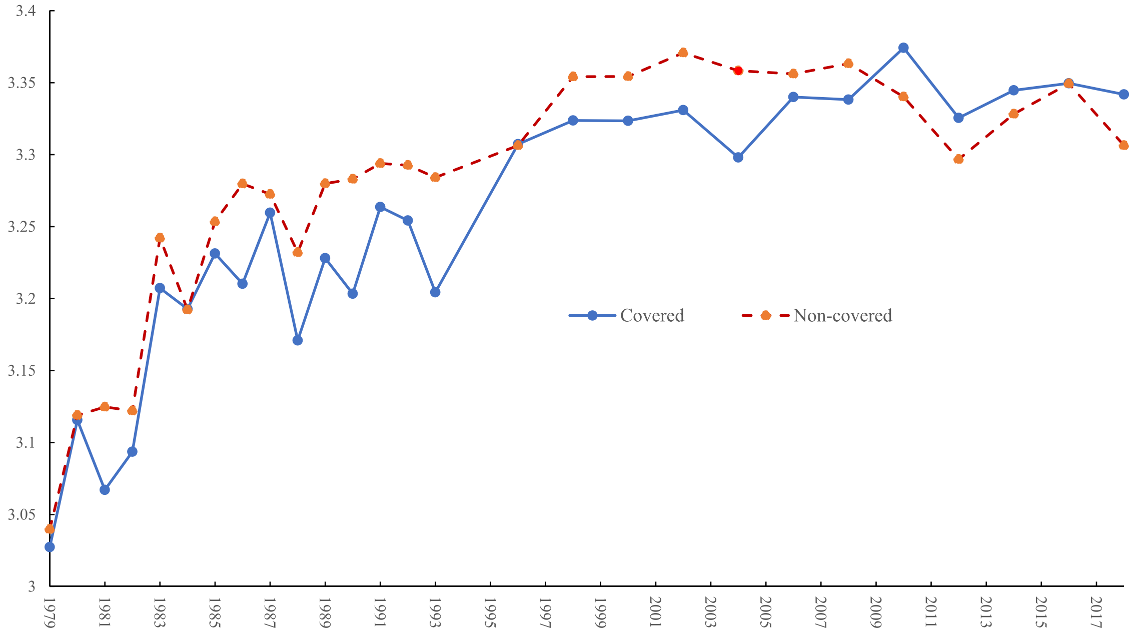

Fig. 1. NLS79 job satisfaction means.

Table 2
Job satisfaction. OLS, pooled cross-sections; OLS person fixed effects using union membership, NLS1979, 1988–2018.

	OLS				Fixed effects			
	1988–2004	1988–2004	2006–2018	2006–2018	1988–2004	1988–2004	2006–2018	2006–2018
Union	-.0481	-.0233	-.0112	.0203	+.0565	+.0542	+.0806	+.0808
	(3.32)	(1.62)	(0.51)	(0.91)	(3.63)	(3.52)	(2.51)	(2.52)
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Personal controls	No	Yes	No	Yes	No	Yes	No	Yes
Constant	3.2286	3.1579	3.355	3.3707	3.2303	3.1421	3.3634	3.5141
Observations	67,265	67,265	28,204	28,204	67,265	67,265	28,204	28,204
Number of groups					10,096	10,096	6,112	6,112
R-squared	.0038	.0201	.0009	.0191	.0020	.0139	.0003	.0111
F-statistic	19.10	20.63	4.54	9.02	15.35	10.39	8.25	4.49

Notes: controls gender, race (3), education (3), regions (4) and industries (15). Fixed effects drop gender and race. T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector.

Using OLS, the coefficient of the union status variable is significantly negative without controls (column 1) and only weakly so ($t = 1.6$) with them (column 2). Columns 3 and 4 report the same analysis, but for the period since 2004, the union variable becomes statistically non-significant. Columns 5 through 8 do the same using person fixed effects models. Switching into union status, because that is what the fixed effects identifies is positively and significantly correlated with improvements in job satisfaction in columns 6 and 7 in both the pre- and post-Great Recession periods, with and without control variables.

Table 3 replaces the union membership variable with a union coverage variable which is set to one if a worker was covered by a union bargaining agreement or was a union member. Coverage is available for a longer run of years from 1979 to 2018. With the additional years we test whether the union effect is different in the second period than the first by interacting union coverage with an additional indicator set to one for the years from 2006 onwards and zero in the earlier years. We find that the union partial correlation with job satisfaction is indeed different post-Great Recession.

Table 3 has the same structure as Table 2 with estimates using OLS in the first five columns and FE in the last three. Columns 5 and 8 pool the years and add union*year interaction terms. As in the case of Table 2, the OLS estimates in columns 1 and 2 return negative partial correlations between union status and job satisfaction in the early period, but the sign switches and is even weakly statistically significant and positive in the later period (columns 3 and 4). Coverage status is positive and weakly statistically significant throughout in the person fixed effects models, with the coefficient doubling, post-Great Recession. These effects on job satisfaction are large and roughly equivalent to the satisfaction of having three employer-provided fringe benefits (Artz, 2010a), having a job in the public versus the private sector (Artz, 2017;

Table 3
Job satisfaction. OLS pooled cross-sections; OLS person fixed effects using union coverage, NLS1979, 1979–2018.

OLS						Fixed effects		
	1979-2004	1979-2004	2006-2018	2006-2018	1979-2018	1979-2004	2006-2018	1979-2018
Covered	-.0367	-.0259	.0093	.0283	-.0217	.0146	+.0377	.0150
	(3.98)	(2.89)	(0.51)	(1.56)	(2.43)	(1.86)	(1.84)	(1.95)
Covered*2006-18					.0513			.0394
	(2.80)	(2.40)						
Year dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Personal controls	No	Yes	No	Yes	Yes	Yes	Yes	Yes
Constant	3.0421	3.0572	3.3523	3.3701	3.0562	3.0298	3.5157	3.0208
Observations	129,840	129,840	28,241	28,241	158,081	129,840	28,241	158,081
Number of groups						12,375	6,119	12,406
R-squared	.0127	.0321	.0009	.0190	.0301	.0269	.0115	.0254
F-statistic	56.43	49.17	4.43	9.05	42.66	35.49	4.45	32.28

Notes: controls gender, race (3) and education (3), regions (4), and industries (15). Fixed effects drop gender and race. T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector.

Table 4
Job satisfaction. OLS pooled cross-sections; OLS person fixed effects using union coverage, NLS1997, 1997–2017.

	OLS				Fixed effects			
	1997-2007	1997-2007	2008-2017	1997-2017	1997-2007	2008-2017	2008-2017	1997-2017
Covered	-.0561	-.0048	.0202	-.0132	.0068	+.0786	+.0704	-.0111
	(2.21)	(0.19)	(0.67)	(0.54)	(0.27)	(2.48)	(2.22)	(0.48)
Covered*2008-2017				.0433				.0675
				(1.23)				(2.13)
Personal controls	No	Yes	Yes	Yes	Yes	No	Yes	Yes
Constant	3.7263	3.9662	3.8733	3.9259	3.7928	3.9042	3.9342	3.7635
Observations	37,751	37,751	28,430	66,181	37,751	28,430	28,430	66,181
Number of groups					8,110	6,771	6,771	8,428
R-squared	.0023	.0350	.0332	.0336	.0123	.0047	.0181	.0153
F-statistic	6.58	22.54	18.02	25.41	11.26	18.09	6.68	16.68

Notes: controls are gender, race (3), education (3), regions (4), and industries (15). Fixed effects drop gender and race. T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector.

Heywood and Wei, 2006), taking part in an employer-provided profit-sharing scheme (Heywood and Wei, 2006) and roughly half the effect of getting a promotion (Kosteas, 2011). The effects hold when conditioning on wages and tenure too.[9]

In column 5 for OLS and column 8 using fixed effects we pool the years together and include a coverage interaction term between the coverage variable and if the data was drawn in the years from 2006 onwards. In column 5 union coverage is significantly negative while the union interaction term is significantly positive. In the person fixed effects specification, both the union main effect and interaction terms are positive suggesting a bigger union impact in the later period. The impact of union is significantly higher in the later period.

One potential reason for this switch in the union partial correlation with job satisfaction could be a change in the composition of union and non-union workers over time. If those workers sorting into (out of) unionization have a greater (lower) propensity for job satisfaction over time, this might explain the increasing propensity for unionized workers to be more satisfied with their jobs than non-unionized workers. We can discount this possibility by focusing on the person fixed effects estimates. These estimates are based on changes in job satisfaction within worker over time, and how they relate to switches in union status. These models, which avoid making comparisons across workers which may be confounded by unobserved differences across workers, confirm that unionization was associated with bigger job satisfaction effects over time. That is to say, the change is apparent within workers over time, so cannot be driven by time-variance in the fixed unobserved traits of union and non-union workers.

Fig. 2 and Table 4 repeat the exercise for covered workers, but this time for the younger NLS97 cohort whose job satisfaction is recorded over the period 1997–2017. This time we split the sample before and after 2008 rather than at 2006 as we did with the NLSY1979 sample. We follow the same structure as before with five specifications in the first five columns and three fixed effects equations in the last three columns. The figure indicates that the raw difference in mean job satisfaction favors non-union workers until the onset of the Great Recession where, for several years, the union differential appears largely positive until 2013 where the differential once more turns negative. Table 4 offers a more thorough analysis using regression to control for potential confounders. The pattern of results is similar to those presented with the NLS1979. Using

[9] In Appendix Tables 1 we explore our results' robustness using both the NLSY1979 and NLSY1997 by also conditioning on wages and tenure, following Freeman (1978) and Borjas (1979). Even after including wages and tenure in the specifications, the effects of union coverage on job satisfaction are at minimum no longer negative and even significantly positive in fixed effects estimates, particularly in the advent of the Great Recession.

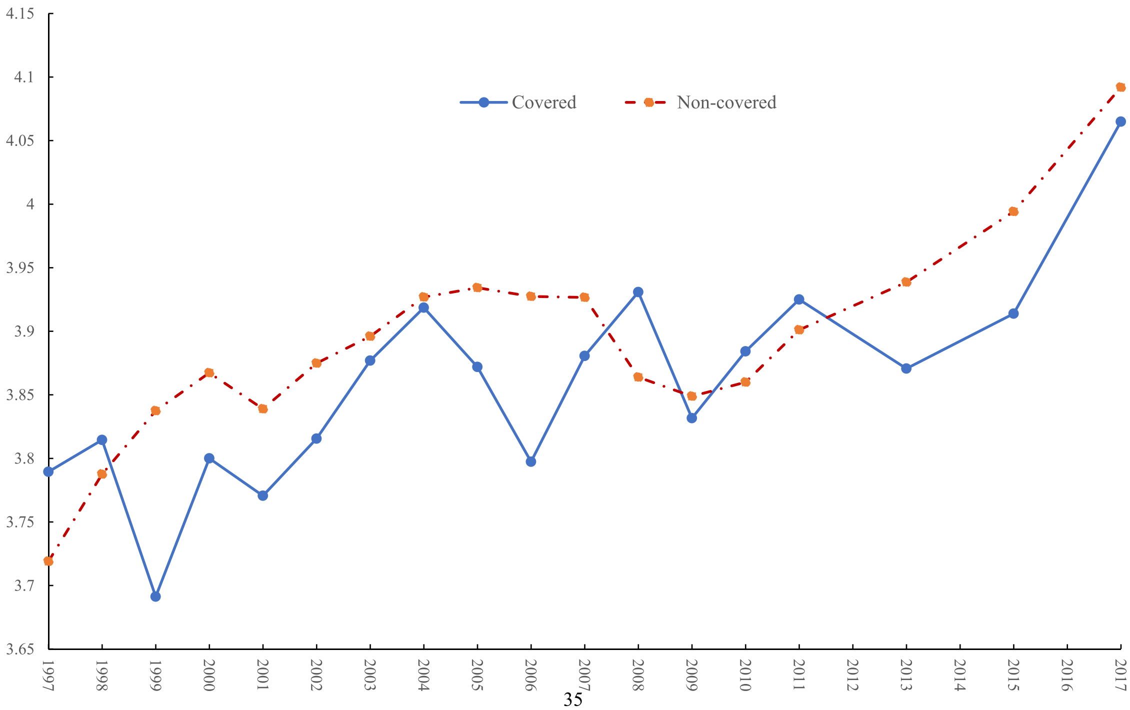

Fig. 2. Job satisfaction NLS97.

OLS the negative and statistically significant partial correlation between union coverage and job satisfaction prior to the Great Recession in column 1 without controls except year becomes insignificant adding controls in the first period (column 2) and is statistically non-significant in the post-Great Recession period (column 3). The person fixed effects estimates in columns 5 to 8 are particularly striking: a non-significant coverage coefficient prior to the Great Recession (column 5) becomes positive and statistically significant post-Great Recession without (column 6) and with controls (column 7). Column 8 shows that the union interaction term is significantly positive for the later period in the presence of person fixed effects.

Thus, results appear robust for the younger as well as the older cohort of workers.[10]

4.2. Potential mechanisms

We begin by considering whether the positive partial correlation between unionization and job satisfaction since the Great Recession indicates a decline in union effectiveness as a voice mechanism for workers. Freeman (1978: pp. 139-140) had suggested that union voice effects may explain the otherwise paradoxical finding that unions lower quit rates despite raising job dissatisfaction, something which is otherwise known to raise quit rates. Therefore, a corollary of unionization being associated with higher job satisfaction since the Great Recession might be a diminution in unions' ability to lower quit rates by providing effective voice for workers.

Figs. 3 and 4 plot quit rates for covered and non-covered workers separately for the NLS79 and NLS97, respectively. Fig. 3 is suggestive of a closure in quit rates between covered and non-covered workers, but only at the height of the Great Recession after 2006. In the NLS97 non-union quit rates have been persistently higher than those for union workers since the early 2000s, though they do close between 2015–2017 (Fig. 4).

We test more formally for a change in relative quit rates among union and non-union workers by running regressions similar to those in Tables 2 to 4 for job satisfaction. We restrict our sample to those who are in work and report job satisfaction. Table 5 presents OLS linear estimation models for the (0,1) probability of quitting the job over the period 1988-2018 among those born between 1957 and 1964 in the NLS79. The table uses the union membership variable, while Table 6 does the same with the union coverage variable. In Table 6 we restrict the sample to start in 1983 due to wide variations in reported quits before 1983, which seems to be due to differences in the survey methodology and with the cohort's youth at the time.[11]

[10] In Appendix Table 2 we distinguish between movements into and out of unionization. Gaining union coverage has a large positive effect on job satisfaction that increases between the pre- and post-Great recession periods, particularly in the younger cohort. Leaving coverage is statistically non-significant pre- and post-Great Recession in both cohorts.

[11] We note though that including waves 1979–1982 does not appreciably alter the results.

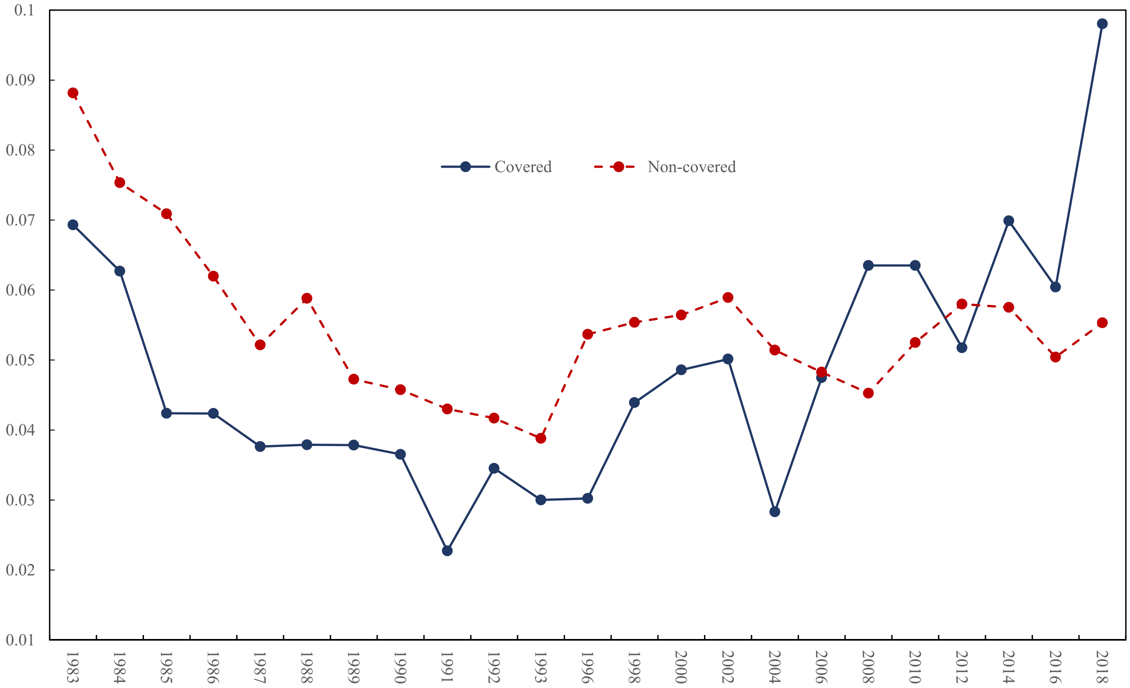

Fig. 3. Quit rates NLS79.

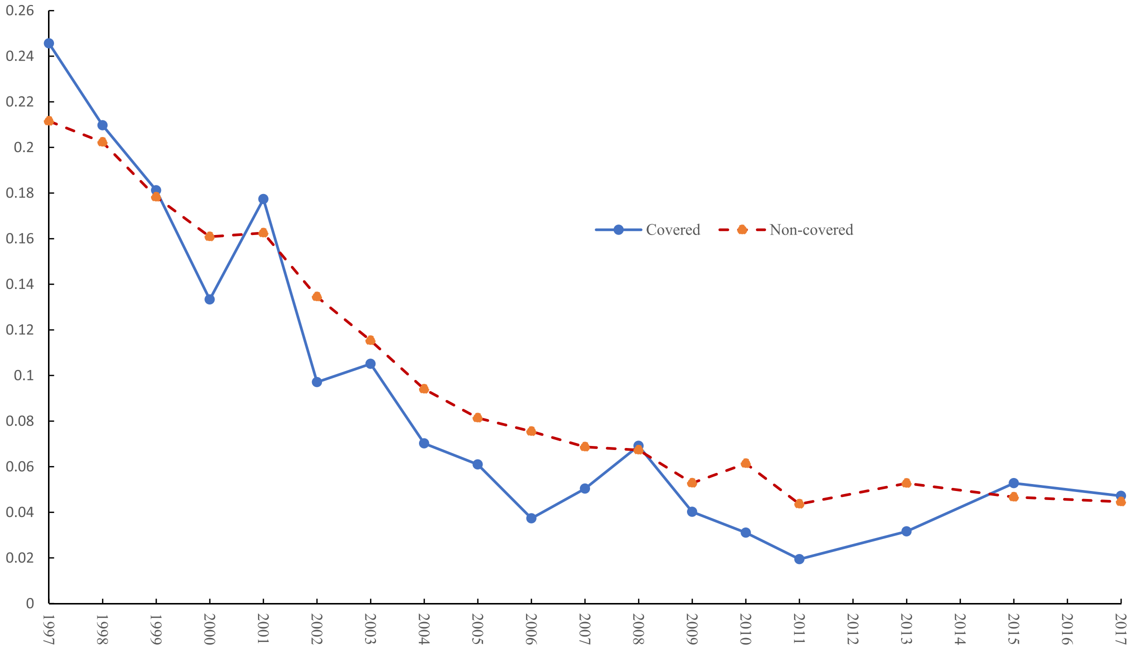

Fig. 4. Quit rates NLS97.

In the OLS estimates, there is a negative and statistically significant partial correlation between union membership and quits prior to 2006. Whilst it remains significant post 2006 the coefficient halves in the later period (compare columns 1 and 2). In the person fixed effects models the negative and significant coefficient in the early period becomes statistically non-significant from 2006.

Similar patterns are apparent in Table 6 using the union coverage metric: the negative coefficient in the early period drops substantially in the later period and even becomes statistically non-significant, a change which is itself statistically significant as indicated in the interaction effect in column 3. Again, similar effects are apparent in the person fixed effects

Table 5
Quit equations with union membership, NLS79 cohort, 1988–2018.

	OLS pooled cross-sections		OLS person fixed effects	
	1988-2004	2006-2018	1988-2004	2006-2018
Union	-.0248	-.0104	-.0118	-.0120
	(9.75)	(2.21)	(2.67)	(1.01)
Controls	Yes	Yes	Yes	Yes
Constant	.0984	.0762	.0519	-.1033
Observations	67,265	28,204	67,265	28,204
Number of groups	10,096	6,112		
R-squared	.0174	.0148	.0055	.0015
F-statistic	22.74	11.64	3.20	4.6

Notes: controls are gender, race (3), education (3), regions (4), years and industries (15). T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector. Sample is of workers who responded to the job satisfaction question.

Table 6
Quit equations with coverage, NLS79 cohort, 1983–2018.

	OLS pooled cross-sections			OLS person fixed effects		
	1983-2004	2006-2018	1983-2017	1983-2004	2006-2018	1983-2017
Covered	-.0207	-.0065	-.0210	-.0128	.0011	-.0154
	(10.38)	(1.52)	(10.67)	(4.69)	(0.15)	(5.98)
Covered*2008-2017	+.0146	.0128				
	(3.29)	(2.66)				
Controls	Yes	Yes	Yes	Yes	Yes	Yes
Constant	.0428	.0763	.0416	.0203	-.1015	.0122
Observations	104,460	28,241	132,701	104,460	28,241	132,701
Number of groups	11,844	6,119	11,898			
R-squared	.0189	.0148	.0178	.0071	.0015	.0090
F-statistic	39.51	11.70	39.32	17.91	13.07	22.62

Notes: controls are gender, race (3), education (3), regions (4), years and industries (15). T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector. Sample is of workers who responded to the job satisfaction question.

Table 7
Quit equations with coverage, NLS97 cohort, 1997–2017.

	OLS pooled cross-sections		OLS person fixed effects	
	1997–2007	2008–2017	1997–2007	2008–2017
Covered	-.0170	-.0069	-.0122	.0005
	(2.85)	(1.49)	(1.55)	(0.06)
Controls	Yes	Yes	Yes	Yes
Constant	.1436	.0887	.1964	-.0228
Observations	37,751	28,430	37,751	28,430
Number of groups	8,110	6,771		
R-squared	.0291	.0199	.0179	.0066
F-statistic	27.58	12.10	20.38	2.16

Notes: controls are gender, race (3), education (3), regions (4), years and industries (15). T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector. Sample is of workers who responded to the job satisfaction question.

models in columns 4 and 5, with the change in the union effect being statistically significant (column 6). The negative union impact on quits diminished over time.

The union effect on quits is reasonably similar for the younger people surveyed in the NLS97 born between 1980 and 1984 (Table 7). Union effects in the OLS model are again negative in the base period but the coefficient approaches zero in the second period and is not statistically significant. The coefficients are similar in the person fixed effects models although in this case the coefficients are not statistically significant.

Taken together, the evidence from Tables 5 to 7 is that the association between unionization and lower quit rates was only apparent prior to the Great Recession. Together with our finding that the partial correlation between job satisfaction and unionization turned positive post-Great Recession, there is some indication that the efficiency with which unions provide workers with voice may have declined somewhat recently.

Table 8
Unemployment and underemployment, 2000–2019.

PTFER % of employment			Unemployment
	Union	Non-union	rate %
2000	1.6	2.4	3.9
2001	1.9	2.8	4.7
2002	2.0	3.2	5.8
2003	2.1	3.4	6.0
2004	2.0	3.4	5.5
2005	2.0	3.2	5.1
2006	1.8	2.8	4.6
2007	2.0	2.9	4.6
2008	2.3	4.0	5.8
2009	3.8	6.5	9.3
2010	3.6	6.4	9.6
2011	3.5	6.3	9.0
2012	3.2	5.8	8.1
2013	3.3	5.7	7.4
2014	2.8	5.1	6.2
2015	2.6	4.4	5.3
2016	2.0	4.0	4.9
2017	2.0	3.5	4.4
2018	2.0	3.1	3.9
2019	1.9	2.8	3.7

Source: MORG files of the CPS and BLS (own calculations). PTFER is part-time for economic reasons available from the BLS.

Table 9
Well-being weighted means.

	Job satisfaction
FT employee union	.891
FT employee non-union	.877
FT self-employed union	.920
FT self-employed non-union	.920
PT wants PT union	.944
PT wants PT non-union	.944
PT wants FT union	.803
PT wants FT non-union	.772
Unemployed	n/a
OLF	n/a
N	704,925

Source: Gallup US Daily Tracker, 2009–2013. Weight=comb_weight.

Next we turn to the issue of underemployment and the possibility that positive union correlations with job satisfaction since the Great Recession may reflect unions' ability to shore up unionized workers' income by ensuring they continue to work sufficient hours. It is apparent from columns 1 and 2 of Table 8 and Appendix Fig. 1 that underemployment rates rose after the Great Recession, only returning to their pre-recession levels in 2016/17, but it should be noted they have still not returned to the lower rates observed at the start of the millennium. It is also apparent from columns 1 and 2 of Table 8 that underemployment rates were much higher among non-union workers than among union workers throughout the first two decades of the century. They were one-third higher among the non-union workers at the beginning and end of the period, but the underemployment rate was actually double for the non-employed in 2016.

Workers in search of additional paid hours of work may have sought second jobs or supplemented their income with self-employment, but Bureau of Labor Statistics figures indicate this did not happen. This suggests that workers were 'stuck' off their labor supply curves due to depressed labor demand. It is perhaps no surprise to discover, therefore, that the underemployed – whether unionized or not – were less happy with their jobs than other workers (Table 9). It is also notable that among the under-employed, non-union workers were less happy than union workers.

To establish whether underemployment may play a role in the job satisfaction of union workers relative to non-union workers post-Great Recession we ran estimates for job satisfaction using US Gallup Daily Tracker data for 2009-2013, with a sample size of just under 650,000 workers. These estimates, which are the first in the literature to consider the links between unionization, underemployment and employee wellbeing, are presented in Table 10. The exact question used is

Q3. Are you satisfied or dissatisfied with your job or the work you do, Yes = 1, No = 0?

In the absence of the underemployment variable the union variable, with year dummies only, in column 1 of Table 10 is significantly positive. Column 2 shows the importance of underemployment for job satisfaction, which is measured as a (0,1) dummy using probit; the results are very similar using OLS. Whereas those in part-time employment who do not want

Table 10
Job satisfaction and Underemployment, 2009–2013, workers only.

Union	.0202 (3.10)	.0347 (5.13)	.0287 (4.09)
Union*PT wants FT	.0767 (3.08)		
FT Self-employed	.1952 (22.73)	.1946 (22.65)	
PT doesn't want FT	.4232 (50.30)	.4227 (50.24)	
PT wants FT	-.4109 (61.33)	-.4167 (59.88)	
Year	Yes	Yes	Yes
Controls	No	Yes	Yes
Constant	1.2503	1.0501	1.0518
Pseudo R^2	.0001	.0314	.0336
N	624,863	624,861	624,861

Notes: All equations include year, state, gender, age and age squared. Probit estimates as job satisfaction is a (1,0) dummy. T-statistics in parentheses.
Source: US Gallup Daily Tracker Poll.

additional hours are more satisfied than full-time employees (the reference category), those part-timers who want additional hours are significantly less satisfied with their jobs than full-time employees. Column 2 also shows union workers are more satisfied with their jobs than their non-union counterparts, confirming earlier work (Blanchflower et al. (2022)).

Column 3 extends the analysis by interacting union status with underemployment. The interaction coefficient is positive and statistically significant, partly offsetting the large negative and statistically significant effects of underemployment these workers would have experienced if non-unionized. A significant, positive union effect persists even when introducing the interaction between unionization and underemployment, so unions' ability to combat the worst effects of underemployment are not the sole reason for the positive union association with worker wellbeing post-Great Recession. It would therefore appear that the switch in the union coefficient on job satisfaction from negative to positive is due, in part, both to the lower rates of underemployment in the union sector compared to the non-union sector, as well as unions' ability to ameliorate the negative effects of underemployment on workers' satisfaction.[12]

6. Conclusions

Although still controversial, policymakers and academics alike have grown increasingly accustomed to measuring utility with subjective wellbeing metrics such as life satisfaction and satisfaction with domains of one's life, including job satisfaction. Economic crises arising from the Great Recession and, more recently, the COVID pandemic, have raised concerns about citizens' wellbeing, and have heightened interest in the role of institutions in mitigating the worst effects of economic shocks. As democratic institutions answerable to their membership base, and with the avowed intention of improving workers' welfare through the monopoly, voice and insurance roles they perform, trade unions are one such institution. And yet the demise of trade unions, as indicated by falling union density over recent decades, raises questions of their ability to represent workers effectively. In addition, the existing evidence regarding the subjective wellbeing of unionized workers, relative to their non-union counterparts, is somewhat equivocal.

In this paper we revisit the literature which identifies a negative association between unionization and individuals' job satisfaction in the United States, first identified over forty years ago. We find the association has flipped since the Great Recession such that union workers are now more satisfied than their non-union counterparts. We show this to be the case for younger and older workers in the National Longitudinal Surveys of Youth of 1979 and 1997. In panel data unionized workers tend to have higher job satisfaction throughout, but the differential has risen over time. Since using the panel data accounts for fixed differences in those who are and are not unionized changes in worker sorting into union status are not the reason for the change.

The absence of substantial change in the union wage gap over time suggests the change is not associated with changes in unions' wage bargaining. Instead, we find some diminution in unions' ability to lower quit rates which is consistent with a decline in their effectiveness in operating as a 'voice' mechanism for unionized workers. We also present evidence suggestive of unions' ability to minimize covered workers' exposure to underemployment, a phenomenon that has been particularly detrimental to the wellbeing of non-union workers.

Unions play three important roles that may affect workers' job satisfaction: they bargain for better terms and conditions, such as higher wages (their monopoly face); they represent covered workers to the employer in grievance and other matters, helping to resolve problems at the workplace through worker 'voice'; and they offer insurance to workers, protecting them against fluctuations in their fortunes, against dismissal and against other threats to their job security such as hours reductions. We have shown in this paper that a diminution in the effectiveness of union voice, and a greater capacity to insure against hours insecurity – in the form of underemployment – are two of the mechanisms at play which have led

[12] This last point is somewhat reminiscent of recent work suggesting unions are able to ameliorate the adverse wellbeing effects of other aspects of labor market experience, such as workers' exposure to anxiety-inducing innovation at the workplace (Bryson et al., 2013b).

to improvements in the job satisfaction of unionized workers compared to their non-union counterparts. Our data do not permit us to examine why there appears to be a diminution in unions' voice capability at the same time as their bargaining (monopoly) face has remained in-tact and their effectiveness in providing insurance may have improved. It is possible that unions have reoriented themselves to face new challenges since the Great Recession, but this is conjecture and would be an issue worthy of investigation in future research.

Additional results and copies of the computer programs used to generate the results presented in the article are available from the corresponding author at artzb@uwosh.edu.

Declaration of Competing Interest

No funding or personal relationships inappropriately biased or influenced this manuscript.

Data availability

Data will be made available on request.

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Appendix A

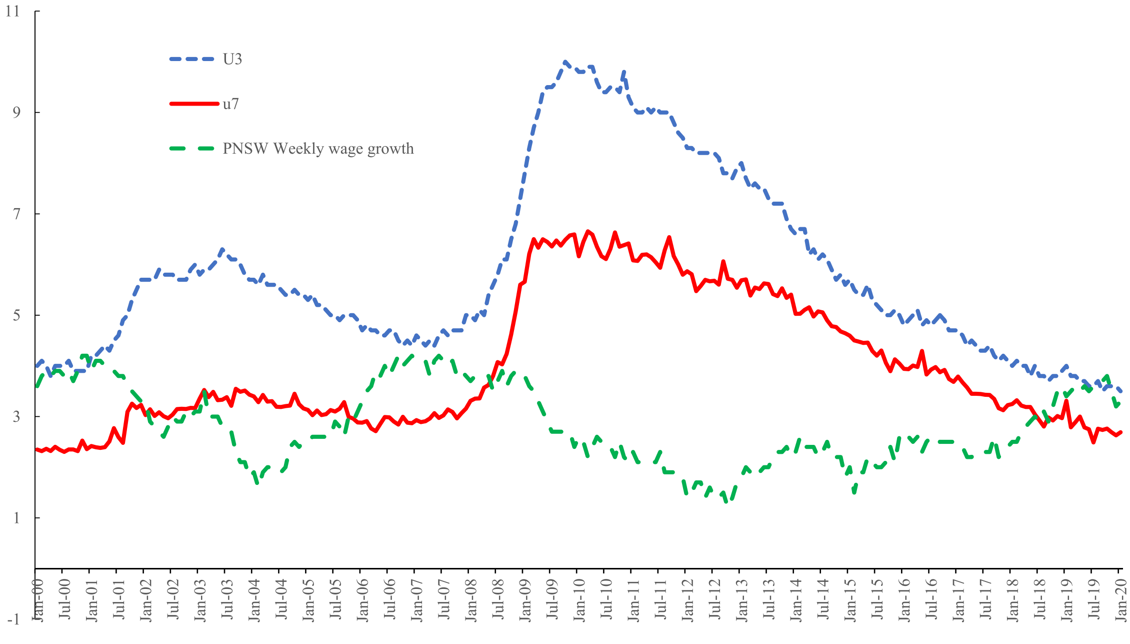

Fig. A1. Unemployment (u3), Underemployment (u7) and PNSW Wage growth, Source BLS.

Table A1
Job satisfaction including wages and tenure.

	NLS79 Cohort				NLS97 Cohort			
	OLS Pooled Cross-sections		OLS Person Fixed Effects		OLS Pooled Cross-sections		OLS Person Fixed Effects	
	1979-2004	2006-2008	1979-2004	2006-2008	1997-2007	2008-2017	1997-2007	2008-2017
Covered	-0.0505	0.0214	-0.0004	0.0368	-0.0227	-0.0055	-0.0036	0.0650
	(5.607)	(1.168)	(0.055)	(1.791)	(0.920)	(0.185)	(0.143)	(2.029)
Log wages	0.1241	0.1058	0.1134	0.0918	0.1762	0.2125	0.1086	0.1600
	(20.425)	(10.443)	(19.843)	(6.678)	(11.985)	(12.605)	(7.165)	(8.236)
Personal controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Constant	2.3778	2.6410	2.4492	2.8014	2.8242	2.3636	3.0820	2.7957
Observations	119,191	27,564	119,191	27,564	36,688	27,628e	36,688	27,628
Number of groups			11,883	6,035			8,068	6,706
R-squared	0.0329	0.0270	0.0276	0.0174	0.0413	0.0434	0.0179	0.0298
F-statistic	46.45	12.30	37.62	5.45	26.26	22.19	12.33	8.47

Notes: controls are gender, race (3), tenure, education (3), regions (4), years and industries (15). T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector.

Table A2
Job satisfaction equations including flows in and out of union coverage.

	OLS Person Fixed Effects			
	NLS79 Cohort		NLS97 Cohort	
	1979-2004	2006-2018	1997-2007	2008-2017
Entered coverage	0.0303	0.0381	0.0484	0.1568
	(3.390)	(1.626)	(1.780)	(4.367)
Exited coverage	0.0177	0.0045	-0.0085	0.0534
	(1.855)	(0.198)	(0.249)	(1.354)
Remained covered	0.0027	0.0400	-0.0984	-0.0410
	(0.238)	(1.329)	(-2.302)	(0.865)
Personal controls	Yes	Yes	Yes	Yes
Constant	3.0303	3.4835	3.7953	4.0573
Observations	129,284	28,241	37,759	28,430
Number of groups	12,375	6,119	8,110	6,771
R-squared	0.0269	0.0113	0.0123	0.0184
F-statistic	34.01	4.12	11.06	7.03

Notes: Omitted group are those who remained uncovered. Controls are gender, race (3), education (3), regions (4), years and industries (15). T-statistics are in parentheses. Heteroskedastic robust standard errors are clustered at the individual level. Private sector.

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