Union Membership and Perceived Job Insecurity: 30 Years of Evidence from the American General Social Survey

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Abstract

Using the American General Social Survey, we explore the link between union membership and perceived job insecurity. We find that union members are more likely to fear for their current (and future) job. This finding is mainly attributed to the primary and secondary sectors and for recessionary periods. Instrumental-variables estimation and the use of attitudinal proxy variables suggest that the positive correlation between union membership and perceived job insecurity is not due to self-selection.

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1 Introduction

It is widely believed that unions allow their members to benefit from a rent (Bennett & Kaufman 2007). As such, it is not surprising to observe that union members are more likely than their non-unionized counterparts to feel that it would be hard to find a job with similar wage and fringe benefits in the event that they were to become unemployed. What is puzzling, however, is that the same union members are also more likely to feel insecure about their current job (Aaronson & Sullivan 1998, Brochu & Zhou 2009).[1] For instance, the American General Social Survey (1978-2008) reveals that 14.2 percent of union workers are insecure about their current job, while only 10.5 percent of non-union workers report the same concern. These figures are surprising as standard labor economics textbooks suggest that one of the principle purposes of a union is to protect the jobs of its members. The main goal of our paper is to investigate the causes of the positive correlation between union membership and current job insecurity.

Understanding the link between union membership and perceived job insecurity is important as perceptions can have a direct impact on economic outcomes. Within a search framework, for example, current job insecurity can affect (on-the-job) search intensity. More generally, increased fear of job loss may have an effect on bargaining, and ultimately on wage outcomes (Campbell et al. 2007); heightened insecurity regarding *future job*[2] prospects can affect the wage determination process as it reduces the worker's outside option. Finally, perceived job insecurity can influence workers' consumption/investment patterns as insecure workers could delay making important purchases or investments.

The fact that union workers are more insecure about future jobs is intuitive. If unions succeed in raising wages (Card 1996, Blanchflower & Bryson 2004) and non-wage benefits (Budd 2005), and the number of union jobs is constrained (Farber 1990), it would be harder for a displaced union worker to find a job with similar wages and non-wage benefits (Kuhn & Sweetman 1998). But, the link between unions and current job insecurity is not as straightforward as it seems. The positive correlation could simply be due to self selection—'worriers' self-selecting into unionized jobs (Aaronson & Sullivan 1998). Unions could also affect job insecurity directly. If job insecurity enters the union objective function, then a successful bargaining process would decrease the job insecurity of its members. On the other hand, a union-induced wage increase would have a disemployment effect within a simple neo-classical

[1] Aaronson & Sullivan (1998) relied on American General Social Survey data (1978-1996), while Brochu & Zhou (2009) used Canadian Gallup Poll data (1979-2000). Both controlled for confounding factors, and both found an economically and statistically significant union effect.

[2] In this paper, we refer to 'future job' as the job a worker would expect to find if she were to lose her current job.

framework, thereby increasing job insecurity. Finally, if unions reduce the productivity and profitability of firms, one could see more plant closures or firm failures which would also increase job insecurity.[3]

There is a theoretical literature suggesting that, in a dynamic setting, the presence of a union can directly impact layoffs. For example, Trejo (1993) proposes a simple dynamic framework where layoffs increase because the union tries to stabilize working hours and union membership using overtime pay. The exit-voice model predicts that unions reduce quits, which implies that firms cannot simply rely on normal attrition to reduce their workforce, and as such, they may be forced to rely more on layoffs. There is some empirical evidence (Medoff 1979, Blau & Kahn 1983, Pearce 1983, Montgomery 1991) suggesting that firms with unionized workers may rely more on layoffs than their non-unionized counterparts.

This paper investigates the causes of the positive correlation between union membership and current job insecurity using American General Social Survey (GSS) data over the 1978-2008 period. In particular, we explore whether the positive correlation arises from specific sectors, specific time periods, or whether this correlation is simply due to self-selection into unionized jobs (based on unobserved ability or inherent fear).

To our knowledge, Bender & Sloane (1999) is the only paper whose focus has been on the causal link between union status and job insecurity perceptions.[4] Our paper differs from that of Bender & Sloane (1999) in two important ways. First, our dataset consists of 17 repeated cross-sections that cover a 30-year period. This allows us to use time varying instruments (e.g. changes in labor laws, and unionization rates at the state level).[5] We are also able to take full advantage of the richness of the attitudinal (i.e. opinion) information in the GSS. We can, for example, proxy for an individual's personal traits and perceived employer-employer relations, and thus deal with aspects of worker/firm heterogeneity. The many cross sections also allow us the see whether the union/non-union differential in perceived job insecurity varies across the business cycle as some of the objective data would suggest (Pearce 1983).

Second, our paper is the first to explore the causal link between union membership and

[3]Freeman & Kleiner (1999) find that unions reduce firm profitability, but do not lead to more plant closures and firm failures. They interpret this as meaning that the union is a rational agent that takes the economic situation into account when bargaining.

[4]The main focus of the existing job insecurity perception literature has been on determining whether perceptions of job insecurity have changed over time (Aaronson & Sullivan 1998, Schmidt 1999, Green et al. 2000, Linz & Semykina 2008, Brochu & Zhou 2009). There also exists a related (and extensive) literature that looks at the link between union membership and job satisfaction (Borjas 1979, Bender & Sloane 1998, Bryson et al. 2004).

[5]Bender & Sloane (1999) rely on one cross section, i.e. the 1986 British Social Change and Economic Life Initiative, and on opinion-based variables as instruments which can be problematic when looking at a subjective dependent variable (Bryson et al. 2004). The (subjective) instruments are likely not to satisfy the exclusion restriction.

perceived job insecurity for the United States—a country that experienced a dramatic decline in unionization rates over the last forty years (Farber 1990, Hirsch 2008). We can therefore examine the extent to which perceived job insecurity is linked to the union environment. In particular, we can see if the union/non-union perception differential changed with the decline of unions in the United States.

We find, as in previous North American studies (Aaronson & Sullivan 1998, Brochu & Zhou 2009), that union members are more likely to be insecure about both their current and future job prospects after controlling for confounding factors. Over the 1978-2008 period, union members are about 3.5 percentage points more likely to feel insecure about their current job. This difference is significant as the average level of insecurity was only 10.5 percent in the non-union sector. The difference is also significant for the insecurity about future jobs.

Our evidence—both with respect to our instrumental variables approach and proxy variables approach—suggest that self-selection is not driving our finding that union workers are relatively more insecure about their current job. We find that the positive correlation is localized in the primary and secondary sectors, and is counter-cyclical in nature. Interestingly, import penetration does not explain the perception differential found in the manufacturing sector. Finally, we do not find any evidence to support the claim that job insecurity perceptions changed with the decline of unions in the United States. There continued to be a counter-cyclical relationship in the manufacturing sector well after the dramatic decline in unionization of the 1970s and 1980s.

The next section describes the GSS data used in this paper while Section 3 presents the empirical strategies used to address the potential endogeneity issue related to union membership. Sections 4 presents our results. Finally, Section 5 concludes.

2 Data

We use the American General Social Survey over the 1978-2008 period. There are two advantages to using the GSS: 1) there are many years of job perception data which makes it possible to use time-variant state-level characteristics as instruments; and 2) the GSS asks respondents their views on a variety of topics, from very specific (e.g. perceptions of employee-employer relations) to more general (e.g. political views).[6] The presence of these

[6]There are two parts to the GSS: the core questionnaire and a supplement (i.e. a topical module). The core questionnaire is made up of permanent questions (e.g. age, gender and job insecurity perceptions) that are asked on a regular basis. The topical modules, however, are only asked at irregular intervals, i.e. in select years. By relying on so many cross sections, we can take advantage of the wide range of attitudinal information that come from the many topical modules.

variables will allow us to address the endogeneity of union status in ways that are not possible with other more traditional labor data sources like the Current Population Survey (CPS).

The GSS is a repeated cross section dataset; the same survey is repeated over time, but with different respondents in each survey year. Most importantly, the GSS consistently ask respondents two job insecurity perception questions (Davis et al. 2009): 1) *"Thinking about the next 12 months, how likely do you think it is that you will lose your job or be laid off: very likely, fairly likely, not too likely, or not at all likely?"* and 2), *"About how easy would it be for you to find a job with another employer with approximately the same income and fringe benefits you now have? Would you say very easy, somewhat easy, or not easy at all?"* The GSS also gathers a broad range of social and demographic information. Note, however, that it identifies union membership, and not the broader measure of union coverage. Fortunately, the difference is very small in the U.S. (Card et al. 2003).

We restrict our sample to private-sector workers who are 18 to 64 years of age and are not self-employed. The self-employed were excluded because the process that determines job insecurity for both subjective and objective measures, is very different for them. Public sector employees were also removed because unions play a different role in the public sector, with union activity restrictions, for instances on wage determination and on the right to strike, that are not present in the private sector (Ehrenberg & Smith 2006).[7] We impose an upper age restriction of 64 years of age so as to abstract from retirement issues. Finally, individuals with missing information on gender, age, education, part-time status, union status, and industry classification are also excluded.[8]

Our restricted sample consists of 7,519 individual observations over the 1978-2008 period. We have data for the following 17 years: 1978, 1983, 1985, 1986, 1988, 1989, 1990, 1991, 1993, 1994, 1996, 1998, 2000, 2002, 2004, 2006, and 2008. The yearly sample range in size from 225 in 2002, to 703 in 1983.[9]

We provide some summary statistics in Table 1.[10] Most variables are self-explanatory,

[7] The union literature has tended to focus on the private sector. As a robustness check, we carried out regressions that included public sector employees, and it does not affect our findings. We present results including public-sector workers in the Appendix. We also tried specifications where we excluded those under the age of 25—because some may not have finished school. It does not materially affect our findings.

[8] Missing information due to non-response (don't know or refuse to answer) is relatively small, representing only 1.77 percent of the sample.

[9] In order to increase content, but not overburden respondents, the GSS has used a rotation design. Prior to 1988, the rotation was across time with the regular questions appearing two thirds of the time. The job insecurity perception questions, for example, were asked in 1985 and 1986, but not in 1987. Starting in 1988, the GSS switched to a split-ballot design. Under this design, a regular question would be asked in only one of three random sub-samples. This explains why the 2002 sample is much smaller than, say, the 1983 sample.

[10] In Table 1 and all subsequent regressions, we rely on normalized GSS weights; where the weights sum to the number of observations in each survey. We repeated our analysis using a different weighting strategy where the weights sum to 1 in each survey. We also repeated the analysis without using weights. Our results

except for the industry variables. We divided industries into four broad categories: the primary sector includes agriculture, fisheries, forestry, mining, and construction; the secondary sector includes manufacturing of both durable and non-durable goods; and the tertiary sector which is divided into two categories: 'Transport & Communication', and 'Other Services'. 'Transport & Communication' includes transportation, communication and other public utilities, and wholesale trade and retail trade. 'Other Services' represents finance, insurance, business and repair services, personal services, entertainment and recreation services, and professional and related services. Note that the unionization rates across various groups are in line with other studies: For example, the unionization rate is higher for men, and blacks. The positive relationship between age and unionization rates is consistent with the decline of unions over the 1978-2008 period.

3 Empirical Approach

The benchmark econometric model takes the form of a probit model

$$insecure^*_{it} = \beta_0 + \beta_1 union_{it} + \boldsymbol{X_{it}\beta} + \epsilon_{it} \tag{1}$$

where $insecure^*_{it}$ is the latent perceived job insecurity for individual i in period t, and ϵ_{it} is the normally distributed error term. We observe the binary variable $insecure_{it}$ which equals one if the latent variable is greater than zero (i.e. the job is perceived to be insecure), and zero otherwise. $union_{it}$ is a dummy variable that equals one if the worker is a union member and zero otherwise. Finally, $\boldsymbol{X_{it}}$ is a vector of individual and job characteristics, and region and time dummies.

We explore two measures of perceived insecurity: insecurity with respect to the current job, and insecurity with respect to the future job. For the current job, the insecurity dummy equals one if the respondent thinks that it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). For the future job, the insecurity dummy equals one if the respondent thinks that finding a job with the same income and fringe benefits that he/she now has would not be easy at all (and zero otherwise).[11]

are very similar irrespective of the weighting strategy. See the Appendix for more details.

[11] We also estimated ordered probits for both current and future job insecurities and got very similar findings. We decided to focus on the probit model for a few reasons: One, we believe that for current job insecurity (the main focus of this paper) going from (very or fairly) likely to (not too, or not at all) likely is the margin of interest; two, those answering not too likely or not at all likely are already a small proportion of all workers, so making a further distinction within this group will (and does) affect the precision of the marginal effects.

3.1 Econometric Issues

Two important econometric issues need to be addressed. The first difficulty arises from the fact that our dependent variable is subjective, while the second comes from the potential selection into unionized jobs. We discuss these issues below. Bertrand & Mullainathan (2001) argue that subjective measures can be affected by the wording and the order of the survey questions, as well as by the potential 'instability' of attitudes. In theory, these problems could be serious enough to render such measures unreliable for regression analysis. We do not believe that these potential problems are of serious concern in our case. The wording of our questions of interest has stayed the same since the GSS started asking them in 1978. The choice of answers offered to respondents has also been stable over time. The job insecurity perception questions tend to be asked early in the survey, prior to other perception questions. As such, one does not need to worry that the respondent will try to be consistent with answers given to other perception questions. Bertrand & Mullainathan (2001) also warn against the use of subjective measures when the respondents could have unstable—changing drastically over a short period of time—or no attitudes toward the question of interest. Our empirical results suggest that the perceived job insecurity is stable to the extent that our results are in line with those looking at actual job insecurity. For example, perceived job insecurity is inversely related to education, and positively correlated with the unemployment rate.[12]

The second issue is related to self-selection. It is possible that the union effect observed in the data is spurious, i.e. it is due to endogeneity problems. Aaronson & Sullivan (1998) state that "workers who are more insecure about their future employment are more likely to join a union". Workers may also choose to join/form a union based on their ability. In addition, Union Determination models (e.g. the Supply and Demand model (Schnabel 2003), and the Queuing model (Farber 1983), emphasize the endogenous aspect of union status. An increase in perceived job insecurity at the firm level, (e.g. due to mismanagement, unfair labor practices) could also make the non-union worker prefer a unionized environment.

Given the repeated cross sectional nature of our data, we cannot include an individual fixed effect term in our regression model.[13] We have to use a different approach to cope with the self-selection issue. The richness of the GSS allows for two empirical strategies to deal with potential self-selection: a proxy approach and an instrumental variables (IV) approach.

[12] We address these concerns in more details in the Appendix.

[13] It should be noted, however, that panel data sets are not a panacea when exploring for union effects. Few workers switch between union and non-union jobs. Some respondents of the 2006 GSS were re-interviewed in 2008 to construct a panel data set. We investigated the possibility of using this panel to address the self-selection issue, but, of all respondents that were asked the job security questions, only nine had either switched from a unionized to non-unionized job, or the opposite. Given that the measurement error problem is amplified in a fixed effect model (Card 1996), we are more confident using the GSS as repeated cross-sections rather than as a panel.

The time-series aspect of the GSS allows for the use of instruments that rely on fluctuations in state-level characteristics. In particular, we used both (together and separately) right-to-works laws and state-level union prevalence as instruments. Another key strength of the GSS data is the attitudinal (i.e. opinion) information. The availability of such information allows us to complement our IV-approach with an alternative strategy for dealing with potential differences between unionized and non-unionized workers. We used the attitudinal information (e.g. importance of job security) to proxy for an individual's personal traits, and thus deal with some aspects of worker heterogeneity. Combining these two fundamentally different strategies will shed light on the likelihood that the positive correlation between union membership and perceived job insecurity is due to worker self-selection.

4 Results

In this section, we start by presenting the results from estimating equation (1), and then present the results from our IV and proxy approaches—two strategies attempting to deal with potential endogeneity issues.

Tables 2 and 3 present the probit marginal effects from estimating equation (1) with current and future job insecurity measures as dependent variables, respectively.[14] Controls are added sequentially, from columns (1) to (4) to see whether the union coefficient estimate is robust to changes in specification. Columns (1) to (4) all suggest that the union effect is economically and statistically significant. Union members are between 3.1 and 3.9 percentage points more likely to feel insecure about their current job than their non-unionized counterparts, which is significant given that the average level of insecurity was only 10.5 percent in the non-union sector. Finally, while the estimates presented in columns (1) to (4) are based on a broad industry classification, column (5) presents results from an identical specification to column (4) with the exception that we used a much narrower industry classification (3-digit industry classification).[15] The narrower classification does not affect our

[14]Using a probit model or a linear probability model gives very similar results. Table A.1 presents the results found in Table 2 when a linear probability model is used, instead of a probit. For Table 2 and all subsequent tables, we rely on weighted standard errors. We also tried clustering at the region-year level. The standard errors are of similar magnitude.

[15]Note that the probit estimation drops observations from industries in which all, or none of the workers fear for their job. This is why the sample size drops from 7,436 to 6,659. Again, note that the union coefficient estimate is very similar to the one obtained from a linear probability model that retains all 7,436 observations. The 3-digit industry classification in the GSS is based on the 1970 and 1980 Census Industry Classifications. As some classes have changed when moving from the 1970 to the 1980 classification, we interact industry with a dummy equal to 1 for observations for which we only have the 1980 classification so that the 1970 and 1980 classifications with the same identifier (i.e. numerical code) are allowed to represent different industries.

results.

Although we do not show the coefficients estimates for all control variables, their signs are in line with expectations and with other studies that have looked at job security perception. For example, while more educated individuals and older individuals fear less for their job, part time workers and blacks tend to fear more.

In Table 3, we investigate whether union workers think that it would be hard for them to find a job with similar pay and benefits if they were to be laid off. If unions conferred a rent to their members, then we could imagine that they would be more likely to be insecure about their future job than similar non-unionized workers. Table 3 confirms this hypothesis: unionized workers are generally more than 15 percentage points more likely to fear for their future job. This difference is large considering that 35.4 percent of non-union workers are insecure about their future job prospects.

Since the more surprising finding is that unionized workers are more likely to fear for their *present* job, we now investigate whether this finding could be due to self-selection.

4.1 Instrumental-Variables Approach

We estimated equation (1) using a set of instruments to tackle the potential endogeneity problem coming from the union variable. Table 4 presents the IV-estimation results (based on the specification of column (4) in Table 2) using two instruments.[16] The state unionization rate should not directly affect an individual's job insecurity once we control for her/his job industry and labor market conditions (controlled for using year effects), but should be correlated with the probability that an employee is unionized.[17] The estimated effect of union membership on insecurity is still positive when we use state unionization rate as an instrument. However, as in many cases where the instrument does not vary at the individual level, the estimate is imprecise. The estimate of the union parameter is 0.202 with standard errors equal to 0.126—the p-value being 0.111. Not surprisingly, the instrument passes the first-stage (rule-of-thumb) test as the F-statistic is clearly above 10 (Staiger & Stock 1997). The estimated union effect is, in our opinion, too large to be credible. However, one should take into account that the standard errors are relatively large. If the main endogeneity problem comes from self-selection (from 'worriers') into unionized jobs, we believe that our

[16]We present results from two-stage least squares for linear probability models. We also estimated a bivariate probit with very similar results. These results are available upon request.

[17]The state unionization rates are from Hirsch et al. (2001). We tested for whether the union variable is endogenous in the linear probability model using a Hausman endogeneity test, and in the probit model using a Rivers-Vuong endogeneity test. In both cases, we assumed that state unionization rate is a valid instrument. For both tests, the null hypothesis that union membership is not endogenous is only rejected at the 10 percent confidence level; the p-values are 0.098 and 0.101, respectively.

instrument is valid as it should not be correlated with the error term once we control for the state of the labor market (which is done here using time fixed-effects). We obtain similar results when we control for the region- or state-level unemployment rates.

We have also used right-to-work (RTW) laws—alone and combined with the unionization rate—as instruments since they are expected to decrease the likelihood of union membership. Again, the estimated effect of union membership on insecurity stays positive. In the case where RTW is used alone, although the estimate is large (0.412), it is only statistically significant at a 10 percent confidence level (s.e. = 0.243).[18] When combining the unionization rate and the RTW laws as instruments, the estimated effect is 0.221 with standard errors equal to 0.125. Overall, the instrumental-variables estimation does not suggest that self-selection is a main factor behind our findings in Table 2.

4.2 Proxy Approach

As previously discussed, we do not believe that opinions of respondents make good instruments, but they can, however, be used as proxy variables. As is the case with ability in a Mincerian wage equation, we recognize that there are no perfect proxies; there will probably remain some unobserved heterogeneity that may be correlated with union status. If the evidence, when using a series of proxies, points systematically in the same direction (and it does) then the proxy approach, albeit imperfect, still provides useful evidence in determining whether (or not) the union effect is an artifact of self selection.[19]

The 1985, 1991 and 1993 surveys ask respondents to rank what matters most in a job. The choices include 'no danger of being fired', 'high income', 'chance of advancement', 'short working hours', and 'work gives a sense of accomplishment'. By including a variable that accounts for whether an individual values job security, one can directly tackle the issue raised by Aaronson & Sullivan (1998) who believe that insecure workers self-select into union jobs. Table 5 presents the effect of controlling for the importance of job security on the union coefficient estimate. We tried two different ways of controlling for the importance of job security.[20] The sign of the importance of job security is positive, but it does not affect the union finding.

We tried a variety of other controls to account for whether the respondent was a 'worrier',

[18]The F-statistic for the RTW dummy is 16.12 in the first-stage regression. We also tried as instruments the proportion of Republicans in the house of representatives at the state level, and also a dummy variable equal to one if the house (at the state level) is controlled by the Republicans; neither passed the first stage.

[19]Although the GSS asks a variety of opinion questions, many of them are only asked in select years. As such, our proxy approach will consist of estimating the effect of each proxy separately.

[20]'Importance of Job Security 1' equals one if job security is ranked most important, and zero otherwise. 'Importance of Job Security 2' equals one if it is most or next most important, and zero otherwise.

including whether the respondent tended to worry about little things, whether he/she felt safe at home, and whether he/she would be afraid to walk alone at night. Again, the inclusion of these variables did not affect our findings. Although not an opinion variable, we also tried to account for whether the respondent had been unemployed in the recent past. This could affect the workers perception of job security (the individual could become more fearful) which in turn could affect his/her decision to join the union sector. The variable that accounted for past unemployment spells was statistically significant in our regression, but the union coefficient did not change.[21]

Finally, firm-specific characteristics could affect job insecurity levels, which could also influence workers' decisions to form (or join) a union. The quality of a firm's management team is a prime example. Recent GSS surveys ask respondents about the quality of employer-employee relations in their workplace, allowing us to see whether poor employer-employee relations could explain the observed correlation between union membership and perceived job insecurity. As shown in Table 5, we control for poor relations by including a binary variable that equals one if the respondent says that the relations are quite bad or very bad, and zero otherwise. The relations variable was economically and statistically significant, but as was the case with our other proxy variables, its inclusion did not affect our findings with respect to the impact of union membership on insecurity perceptions.

Table 5 does reveal, however, that the union effect is sensitive to the choice of sample years. More precisely, the effect appears to be counter-cyclical; when we focus more on expansionary periods, as is the case when controlling for employer-employee relations (i.e. columns (4) and (5)), the union effect is muted.[22] We further investigate the cyclicality of the 'union membership effect' in the next sub-section.

4.3 Potential Sources of the Differential

Empirical evidence suggests that the presence of unions can result in more layoffs. Medoff (1979) and Blau & Kahn (1983), for example, found a significant union effect in the manufacturing sector.[23] Pearce (1983) examined the link between unionism and cyclical behavior,

[21] The results based on the proxies discussed in this paragraph are available upon request.

[22] For the columns (1) through (3) sub-sample—where one sees a strong union effect—the unemployment rate was 7.0 percent, while it was only 5.4 percent for the sub-sample used in columns (4) and (5).

[23] Montgomery (1991), on the other hand, found that the union effect was larger in the non-manufacturing sector. He also found that controlling for establishment size (in a non-linear way) muted the manufacturing sector union effect. Establishment size is available in the GSS, but only for 1991 onward (except for 1993 when it was not asked). We tried specifications where we controlled for establishment size by including binary variables for six of the seven response intervals, for samples that included all industries, and also for manufacturing and non-manufacturing separately. In all cases, controlling for establishment size did not affect our union finding.

and found the sensitivities of employment to excess demand were relatively greater for union members. Table 6 examines whether these sectorial and cyclical differences are also present in the subjective data.

The first set of results in Table 6 (columns (1) through (3)) is for the full sample, the second (columns (4) through (6)) and third (columns (7) through (9)) sets are for the non-recessionary and recessionary years, respectively. Within each set, we investigate whether the union effect is homogeneous across industries by including interaction terms. For these specifications, the union coefficient estimate (e.g. 0.070 under column (3)) represents the union effect for the secondary sector industry. The interaction terms (e.g. Primary Sector*Union) capture differences in the union effect as compared to the effect in the secondary sector.[24]

Focussing on the full sample, one can see that the positive correlation between union membership and perceived job insecurity comes from the primary and the secondary sectors. Union members in the manufacturing sector are between 7.0 to 7.1 percentage points more likely to feel insecure about their present job than non-union members—a difference that is economically very significant. For the tertiary sector (transportation and communications, and other services) there is essentially no difference between union and non-union workers. These sectoral differences remain essentially unchanged when one splits the sample into recession and non-recessionary years. What does change, however, is that the (union) insecurity differential in the manufacturing sector rises dramatically in recessionary periods.[25]

We also tried controlling for cyclicality by using the full sample, but adding the regional unemployment rate (UR) as a control (instead of year fixed effects).[26] The UR coefficient was both economically and statistically insignificant, and did not affect the union findings. The UR coefficient estimate became significant, however, when the year dummies were excluded. As such, job insecurity perceptions are correlated with the unemployment rate, but the link disappears once year dummies are included in the regression model, possibly because of insufficient variation in the unemployment rate within regions. Adding an unemployment rate-union interaction term, however, does make a difference. The interaction term is positive and statistically significant. As was the case when we split our sample into recessionary and

[24]The estimation of the interaction coefficients is done using the methodology proposed by Ai & Norton (2003). These estimates are very similar to the ones obtained using a linear probability model (see columns (5) and (6) in Table A.1).

[25]We included 1983 as part of the recessionary years despite the fact that the recession officially ended in November 1982 (according to the NBER). We did so because the GSS interviews for 1983 were carried early in the year (February, March and April), and the unemployment rate was still very high over these three months. As a robustness check, we restricted our 'recessionary' years sample to only include 1991 and 2008, and we got essentially the same union results.

[26]We also estimated specifications where we included the state monthly unemployment rate instead of the regional rate, and we got similar results.

non-recessionary years, we find that the manufacturing sector drives the union effect.[27]

Import penetration can have an impact on plant survival and employment growth (Bernard et al. 2006, Khandelwal 2010), so we explored whether it could explain why we find a union/non-union difference in job insecurity in the manufacturing sector.[28] Table 7 presents the results when an import penetration variable and its interaction it with union status are included as additional regressors.[29] Given that the trade data were only available up to 2005, and there are no GSS survey in this year, we restricted our analysis to the 1978-2004 period. The import penetration coefficient estimate is statistically insignificant. The same applies for the interaction term. We also tried a specification where we included the change in import penetration from period $t-k$ and period t for $k = 1, 2, 3, 4$ (and its interaction with union status), and found very similar results as in Table 7.[30] These findings would seem to indicate that competitive pressure (as measured by import penetration) cannot explain why union workers feel more insecure about their jobs. As a final robustness check, we allowed low-wage and high-wage countries (as defined in Kandilov 2010) to have their own 'import-penetration effect'—again, the union effect is unchanged.

Finally, if unions represent the preferences of the median 'voter', the difference in perceived job security between unionized and non-unionized workers could be concentrated among more juniors workers (Blau & Kahn 1983). Unfortunately, the GSS has information about tenure only for a few select years (1991, 2002, and 2006). We did investigate whether the 'union effect' changes with (a quadratic function of) tenure using the available information, but the results were inconclusive. There seems to be a U-shape relationship between tenure and the 'union effect' but the coefficient estimates are statistically insignificant—not surprisingly given that there are only 122 unionized workers in this restricted sample. As an alternative, we focus on age (as opposed to tenure), allowing us to use all 7,436 observations. There again, we find a statistically insignificant U-shape relationship between age and the 'union effect'. Interestingly, the union effect (i.e. $\hat{\beta}_1 + \hat{\beta}_{union*age} age + \hat{\beta}_{union*age^2} age^2$) is

[27] These results are available upon request.

[28] Import penetration is defined as in Bertrand (2004):

$$\text{Import Penetration} = \frac{\text{imports}}{\text{shipment} - \text{exports} + \text{imports}}$$

where imports are cost cif (cost, insurance, freight) valued.

[29] Import data is available at the Standard Industry Classification levels. The GSS, however, used the Census classification. The early GSS surveys relied on the 3-digit 1970 Census Industry Classification (CIC70), and starting in 1988, the GSS moved to the 3-digit 1980 Census Industrial Classification (CIC80). To construct our combined dataset, we relied on Schott's Standard Industry Classification (SIC87) trade data (1976-2005) and the following two crosswalks: a SIC87 to CIC80 crosswalk to make the trade data compatible with the GSS data; and a CIC70 to CIC80 crosswalk to create a uniform industry classification (i.e. CIC80) across GSS surveys. More details on both the trade data and the crosswalks can be found in the Appendix.

[30] These results are available upon requests.

positive and statistically significant for all age groups.

5 Concluding Remarks

This paper seeks to investigate why we observe a positive correlation between union membership and job insecurity. In particular, we try to see whether this correlation is simply due to self-selection into unionized jobs. Using the American General Social Survey over the 1978-2008 period, we find that union members are about 3.5 percentage points more likely to feel insecure about their current job relative to their non-unionized counterparts. Our evidence—using both IV and proxy approaches—suggests that this effect is not due to self-selection.

The positive correlation appears to localized in the primary and secondary (manufacturing) sectors. Interestingly, this latter finding is not driven by import penetration, but is counter-cyclical in nature; the union/non-union perception differential increases as the economy worsens.

Finally, there is no evidence of a structural change in job security perception differential accompanying the well-documented decline in unionization rates in the United States. There continued to be a counter-cyclical relationship in the manufacturing sector well after the dramatic decline in unionization of the 1970s and 1980s—as evident by the positive and economically significant union/non-union perception differential of the recessionary periods of the early 1990s and late 2000s.

A potential limitation of our study is that we cannot make a distinction between fears of temporary versus permanent layoffs. As discussed in the introduction, unions may be more willing to accept layoffs as opposed to, say, a reduction in hours—as long as they are temporary in nature. Therefore, one may see differences not just in aggregate layoffs but also in composition, i.e. temporary versus permanent (Blau & Kahn 1983). Hence, it is possible that our results are, in part, due to the type of job layoff that is more likely to affect unionized versus non-unionized workers. However, the identification of layoffs—whether temporary or permanent—is typically done ex-post. If there is some uncertainty in the mind of workers about the exact nature of the layoff, our measure of insecurity is still appropriate. As an extension of this work, it would be interesting to see whether the difference in perceived job insecurity (and its cyclicality) is reflected in differences in consumption/investment patterns between unionized and non-unionized workers.

A Appendix

Public Sector

As discussed in Section 2, we excluded the public sector because unions may play a different role (e.g. on wage negotiation and the right to strike). In this subsection, we examine whether the inclusion of the public sector changes any of our main findings. Table A.2 presents the probit marginal effects from estimating equation (1)—using present job insecurity as dependent variable—on our sample of respondents including public-sector workers. As with Table 2, we add controls sequentially.

One can draw two conclusions when expanding the sample to include public-sector workers: 1) The key findings about union membership are still present. The union coefficient remains positive (although slightly dampened), and both economically and statistically significant. When we allow for the union effect to vary by sector (columns (5) and (6)), we still find that the primary and secondary sectors drive our results; and 2), the union effect is negative in the public sector. If it were the case that insecure workers self-select into union jobs, one should also have seen a negative, and not positive, correlation between unionization and job insecurity perception. As such, we believe this to be one further piece of evidence to support our claim that self-selection is not the driving force behind our results.

Weights

In this subsection we explore alternative weighting strategies. We compare three approaches: 1) weighting the data so that the weights in survey year t add up to number of observations in year t ('Weight 1'). This weighting approach ensures that larger surveys play a larger role in the estimation; 2) re-weighting the data so that the weights sum to one in each survey year ('Weight 2'); this means that each survey, irrespective of its sample size, plays an equal role in the estimation; and 3) using equally weighted observations ('Unweighted').

Table A.3 compares the estimates presented in Table 2, obtained using our weighting strategy ('Weight 1'), to two alternative weighting strategies (i.e 'Weight 2' and 'Unweighted'). Table A.3 shows that the choice of weighting strategy does not materially affect our findings.

Subjective Measure and Measurement Error

As we are using a subjective measure as dependent variable, we need to address some potential concerns raised by some economists (Bertrand & Mullainathan 2001) on estimating regressions using such measure. We address a series of potential threats below.

Bertrand & Mullainathan (2001) mention that the wording and the ordering of the questions could affect how survey respondents answer them, which could make our job insecurity measure unreliable. This potential problem should not be of great concern in our case. The wording for the current-job question has stayed the same over the years, and the question clearly measures the outcome of interest, i.e. the perceived likelihood of job loss (or lay off) over the next twelve months. A similar argument holds true for the future job question. For this question, the type of new job is also precisely defined—a job with approximately the same income and fringe benefits as you now have. We believe that the job insecurity questions are less open to interpretation as compared to other perception measures such as happiness or job satisfaction. The respondents are offered four choices of answers: very likely, fairly likely, not too likely, and not likely at all. The recording of answers may be less precise than in the Survey of Economic Expectations (SEE) where probabilistic choices are given (Manski & Straub 2000), but it is more precise than the yes or no answers found in the Canadian Gallup (Brochu & Zhou 2009). Most importantly, the choice of answers has remained unchanged since 1977. Finally, the ordering of the questions should not create any problem here since both job insecurity questions are always asked early in the survey and typically follow standard employment questions. Hence, the GSS seems, in our opinion, to minimize the any potential 'cognitive problems' linked to the use of subjective measures.

A second set of problems mentioned in Bertrand & Mullainathan (2001) is related to the potential instability of the attitude (or the absence of an attitude) toward the question of interest. The questions we are interested in are not related to 'obscure' subjects; we expect workers to be familiar with the concept of job insecurity, and to have some opinion about their own job insecurity. In fact, many of the GSS respondents have 'strong' opinions about their perceived job insecurity: 66.3 percent of the respondents in our sample either answered 'very likely' or 'not at all likely', suggesting that the individuals in our sample do have an attitude toward job insecurity. Now, if perceptions were unstable it would have affected the estimated effect of other explanatory variables in the job insecurity perception regressions. However, other estimated coefficients (other than for union status) have the 'right' signs, i.e. they are intuitive and in-line with objective data. Perceptions of job insecurity, for example, are found to be inversely related with educational attainment (Brochu & Zhou 2009, Schmidt 1999).[31] That insecurities tend to mirror the business cycle, i.e. decrease in periods of expansion only to bounce back in periods of recession, would also go against the argument that perceptions are unstable over time. Finally, the correlation between

[31] Other intuitive and consistent findings include: part-time workers feeling relatively less secure about their job; white-collar workers feeling relatively more secure than blue collar workers; the tertiary sector workers feeling more secure than primary and secondary sector workers.

perceptions of insecurity and subsequent job loss should be low if perceptions were in fact unstable. Yet, researchers using panel data (Campbell et al. 2007, Stephens 2004) have found a strong correlation between perceptions of job insecurity and future unemployment spells.[32]

International Trade Data and Crosswalks

We constructed our import penetration variable using Schott's (2010) trade data. These data are available on Schott's International Economics Resource Page which can be accessed through the NBER website. Schott extends Feenstra's trade data up to 2005. The dataset contains information on U.S. manufacturing exports, imports and shipments at the year, country and 4-digit SIC87 level. A more detailed description of the data can be found in Schott (2010).

Crosswalks

- CIC70 to CIC80 crosswalk; based on U.S. Bureau of the Census (1989), we created a CIC70 to CIC80 crosswalk. The changes in industry classification across time were modest in nature, particularly in the manufacturing sector. As such, it was possible to find a compatible CIC80 match for all but seven CIC70 codes.

- SIC87 to CIC80 crosswalk; the 3-digit Census Industry classification is based on the 4-digit Standard Industry classification (U.S. Bureau of the Census 1989). We used an exact SIC87 to CIC80 correspondence that is available on the BLS website. The correspondence can be accessed through the following BLS link address: http://ferret.bls.census.gov/items/value/valu_59185.htm. Given that the Current Population Survey (CPS) also relied on the CIC80 for the 1983-1991 period, the same exact correspondence (with detailed accompanying notes) can also be found in the CPS codebooks (1983-1991) made available on the BLS website.

[32] Bertrand & Mullainathan (2001) also mention that 'social desirability' could affect they way respondents express their opinion in front of an interviewer. They could try not to look 'bad' (Bertrand & Mullainathan 2001). This could be a serious problem if we were interested in respondents' attitude toward sexual orientation for example, but we do not think it applies to job insecurity.

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Table 1: Summary Statistics, 1978-2008

	Mean	Standard Deviation	Fraction Unionized
A. Gender			
Male	0.503	(0.500)	0.199
Female	0.497	(0.500)	0.134
B. Race			
Non-Black	0.885	(0.319)	0.158
Black	0.115	(0.319)	0.232
C. Age			
Age 18 to 24	0.137	(0.343)	0.083
Age 25 to 34	0.292	(0.455)	0.138
Age 35 to 44	0.265	(0.442)	0.188
Age 45 to 54	0.203	(0.402)	0.212
Age 55 to 64	0.103	(0.304)	0.212
D. Educational Attainment			
Less than Hign School	0.127	(0.333)	0.184
Hign School	0.555	(0.497)	0.166
Associate/Junior College	0.073	(0.259)	0.142
Bachelor's and up	0.245	(0.430)	0.165
E. Job Characteristics			
Full-time	0.873	(0.333)	0.175
Part-time	0.127	(0.333)	0.106
F. Industry			
Primary Sector	0.076	(0.265)	0.203
Secondary Sector	0.210	(0.408)	0.214
Trans. & Comm.	0.281	(0.450)	0.158
Other Services	0.433	(0.495)	0.142
Observations	7,519		

Notes. The summary statistics are weighted. The weights are normalized to sum up to the number of observations in each survey.

Table 2: Union Membership and Perceived (Current) Job Insecurity: Probit Models (Marginal Effects Reported)

	(1)	(2)	(3)	(4)	(5)
Union	0.039***	0.033***	0.035***	0.031***	0.039***
	(0.012)	(0.011)	(0.011)	(0.011)	(0.013)
Primary Sector	-	-	-	0.002	-
	-	-	-	(0.014)	-
Trans. & Comm.	-	-	-	-0.044***	-
	-	-	-	(0.009)	-
Other Services	-	-	-	-0.052***	-
	-	-	-	(0.010)	-
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Gender and Race Controls	No	Yes	Yes	Yes	Yes
Age and Education Controls	No	Yes	Yes	Yes	Yes
Part-time Status Control	No	No	Yes	Yes	Yes
Region Fixed Effects	No	No	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No	Yes
Observations	7,436	7,436	7,436	7,436	6,659

Notes. The dependent variable equals 1 if the respondent thinks it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 3: Union Membership and Perceived (Future) Job Insecurity: Probit Models (Marginal Effects Reported)

	(1)	(2)	(3)	(4)	(5)
Union	0.203***	0.176***	0.176***	0.173***	0.129***
	(0.017)	(0.018)	(0.018)	(0.018)	(0.020)
Primary Sector	-	-	-	-0.165***	-
	-	-	-	(0.023)	-
Trans. & Comm.	-	-	-	-0.124***	-
	-	-	-	(0.017)	-
Other Services	-	-	-	-0.168***	-
	-	-	-	(0.017)	-
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Gender and Race Controls	No	Yes	Yes	Yes	Yes
Age and Education Controls	No	Yes	Yes	Yes	Yes
Part-time Status Control	No	Yes	Yes	Yes	Yes
Region Fixed Effects	No	No	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No	Yes
Observations	7,430	7,430	7,430	7,430	7,288

Notes. The dependent variable equals 1 if the respondent thinks that finding a job with another employer with approximately the same income and fringe benefits he/she now has would not be easy at all (and zero otherwise). Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 4: Instrumental-Variables Estimation Union Membership and Perceived (Current) Job Insecurity

	(1)	(2)	(3)
A. First-Stage			
State Unionization Rate	0.010***	-	0.010***
	(0.002)	-	(0.002)
Right-to-Work Laws	-	-0.061***	-0.020
	-	(0.015)	(0.016)
F-Statistic	45.43	16.12	23.45
{p-value}	{0.0000}	{0.0001}	{0.0000}
B. Second-Stage			
Union	0.202	0.412*	0.221*
	(0.126)	(0.243)	(0.125)
Year Fixed Effects	Yes	Yes	Yes
Gender and Race Controls	Yes	Yes	Yes
Age and Education Controls	Yes	Yes	Yes
Part-time Status Control	Yes	Yes	Yes
Region Fixed Effects	Yes	Yes	Yes
Industry Fixed Effects	No	No	No
Observations	7,436	7,436	7,436

Notes. The dependent variable equals 1 if the respondent thinks it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). The estimated models are based on the specification of column (4) in Table 2. The reported F-statistics are from a joint test of significance on the excluded instruments. Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 5: Union Membership and Perceived (Current) Job Insecurity, Proxy Approach: Probit Models (Marginal Effects Reported)

	1985, 1991 and 1993 only			1989, 1991, 1998, 2002 2004 and 2006 only	
	(1)	(2)	(3)	(4)	(5)
Union	0.098***	0.095***	0.099***	0.001	-0.001
	(0.040)	(0.040)	(0.040)	(0.021)	(0.020)
Importance of Job Security 1	-	0.045	-	-	-
	-	(0.031)	-	-	
Importance of Job Security 2	-	-	0.037	-	-
	-	-	(0.047)	-	-
Bad Relations	-	-	-	-	0.087***
	-	-	-	-	(0.034)
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Gender and Race Controls	Yes	Yes	Yes	Yes	Yes
Age and Education Controls	Yes	Yes	Yes	Yes	Yes
Part-time Status Control	Yes	Yes	Yes	Yes	Yes
Region Fixed Effects	Yes	Yes	Yes	Yes	Yes
Industry Controls	Yes	Yes	Yes	Yes	Yes
Observations	790	790	790	1,987	1,987

Notes. The dependent variable equals 1 if the respondent thinks it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). We used three different measures to control for the importance of job security. 'Importance of Job Security 1' equals one if job security is ranked most important, and zero otherwise. 'Importance of Job Security 2' equals one if job security is most or next most important, and zero otherwise. To control for employer-employee relations, our bad relations variable equals one if the respondent considers that employer-employee relations are quite bad or very bad, and zero otherwise. Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 6: Union Membership and Perceived (Current) Job Insecurity, Heterogeneous Union Effect across Industries and Time: Probit Models (Marginal Effects Reported)

	All Years			Non-Recession Years (excluding 1983, 1991 and 2008)			Recession Years (1983, 1991 and 2008 only)		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
Union	0.031***	0.071***	0.070***	0.022*	0.050**	0.042*	0.068**	0.175***	0.213***
	(0.011)	(0.019)	(0.022)	(0.012)	(0.020)	(0.023)	(0.029)	(0.053)	(0.066)
Primary Sector	0.002	0.003	0.002	0.009	0.009	0.003	-0.002	0.005	0.032
	(0.014)	(0.015)	(0.016)	(0.017)	(0.017)	(0.018)	(0.035)	(0.035)	(0.036)
Trans. & Comm.	-0.044***	-0.033***	-0.034***	-0.039***	-0.032***	-0.033***	-0.065***	-0.041	-0.034
	(0.009)	(0.010)	(0.011)	(0.10)	(0.011)	(0.011)	(0.023)	(0.027)	(0.028)
Other Services	-0.052***	-0.041***	-0.041***	-0.062***	-0.055***	-0.056***	-0.078***	-0.042	-0.034
	(0.010)	(0.011)	(0.011)	(0.011)	(0.011)	(0.012)	(0.024)	(0.027)	(0.028)
Primary Sector*Union	-	-	-0.005	-	-	0.038	-	-	-0.133
	-	-	(0.044)	-	-	(0.048)	-	-	(0.105)
Trans. & Comm.*Union	-	-0.072***	-0.071***	-	-0.048*	-0.037	-	-0.137*	-0.174**
	-	(0.028)	(0.031)	-	(0.029)	(0.031)	-	(0.073)	(0.083)
Other Services*Union	-	-0.080***	-0.079***	-	-0.042*	-0.031	-	-0.222***	-0.260***
	-	(0.028)	(0.031)	-	(0.029)	(0.031)	-	(0.064)	(0.086)
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Gender and Race Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Age and Education Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Part-time Status Control	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Region Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Observations	7,436	7,436	7,436	5,957	5,957	5,957	1,479	1,479	1,479

Notes. The dependent variable equals 1 if the respondent thinks it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). The interaction term coefficient estimates were obtained using the methodology proposed in Ai and Norton (2003). Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table 7: Union Membership and Perceived (Current) Job Insecurity in the Manufacturing Sector: Probit Models (Marginal Effects Reported)

	(1)	(2)	(3)
Union	0.069***	0.069***	0.095***
	(0.026)	(0.026)	(0.041)
Import	-	0.031	0.060
	-	(0.073)	(0.080)
Union*Import	-	-	-0.161
	-	-	(0.204)
Year Fixed Effects	Yes	Yes	Yes
Gender and Race Controls	Yes	Yes	Yes
Age and Education Controls	Yes	Yes	Yes
Part-time Status Control	Yes	Yes	Yes
Region Fixed Effects	Yes	Yes	Yes
Observations	1,336	1,336	1,336

Notes. The dependent variable equals 1 if the respondent thinks it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). The interaction term coefficient estimates were obtained using the methodology proposed in Ai and Norton (2003). Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table A.1: Union Membership and Perceived (Current) Job Insecurity: Linear Probability Models

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Union	0.039***	0.033***	0.036***	0.034***	0.090***	0.092***	0.040***
	(0.012)	(0.011)	(0.011)	(0.011)	(0.023)	(0.027)	(0.012)
Primary Sector	-	-	-	0.003	0.004	0.005	-
	-	-	-	(0.019)	(0.019)	(0.020)	-
Trans. & Comm.	-	-	-	-0.056***	-0.039***	-0.039***	-
	-	-	-	(0.012)	(0.013)	(0.013)	-
Other Services	-	-	-	-0.060***	-0.044***	-0.044***	-
	-	-	-	(0.012)	(0.012)	(0.013)	-
Primary Sector*Union	-	-	-	-	-	-0.004	-
	-	-	-	-	-	(0.052)	-
Trans. & Comm.*Union	-	-	-	-	-0.085***	-0.086***	-
	-	-	-	-	(0.031)	(0.034)	-
Other Services*Union	-	-	-	-	-0.088***	-0.089***	-
	-	-	-	-	(0.027)	(0.031)	-
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Gender and Race Controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Age and Education Controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Part-time Status Control	No	No	Yes	Yes	Yes	Yes	Yes
Region Fixed Effects	No	No	Yes	Yes	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No	No	No	Yes
Observations	7,436	7,436	7,436	7,436	7,436	7,436	7,436

Notes. The dependent variable equals 1 if the respondent thinks it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table A.2: Union Membership and Perceived (Current) Job Insecurity When Including the Public Sector: Probit Models (Marginal Effects Reported)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Union	0.031***	0.025**	0.028***	0.024**	0.069***	0.069***	0.029**
	(0.011)	(0.010)	(0.011)	(0.010)	(0.019)	(0.022)	(0.012)
Primary Sector	-	-	-	0.002	0.003	0.002	-
	-	-	-	(0.014)	(0.014)	(0.016)	-
Trans. & Comm.	-	-	-	-0.043***	-0.033***	-0.033***	-
	-	-	-	(0.009)	(0.010)	(0.010)	-
Other Services	-	-	-	-0.051***	-0.040**	-0.040***	-
	-	-	-	(0.009)	(0.010)	(0.011)	-
Public Sector	-	-	-	-0.049***	-0.032**	-0.032**	-
	-	-	-	(0.011)	(0.014)	(0.014)	-
Primary Sector*Union	-	-	-	-	-	0.004	-
	-	-	-	-	-	(0.043)	-
Trans. & Comm.*Union	-	-	-	-	-0.070**	-0.069**	-
	-	-	-	-	(0.027)	(0.030)	-
Other Services*Union	-	-	-	-	-0.079***	-0.078***	-
	-	-	-	-	(0.028)	(0.030)	-
Public Sector*Union	-	-	-	-	-0.093***	-0.092***	-
	-	-	-	-	(0.032)	(0.034)	-
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Gender and Race Controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Age and Education Controls	No	Yes	Yes	Yes	Yes	Yes	Yes
Part-time Status Control	No	No	Yes	Yes	Yes	Yes	Yes
Region Fixed Effects	No	No	Yes	Yes	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No	No	No	Yes
Observations	8,068	8,068	8,068	8,068	8,068	8,068	7,287

Notes. The dependent variable equals 1 if the respondent thinks it is very likely or fairly likely that he/she will lose his/her job or be laid off over the next twelve months (and zero otherwise). The interaction term coefficient estimates were obtained using the methodology proposed in Ai and Norton (2003). Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.

Table A.3: Union Membership and Perceived (Current) Job Insecurity When Using Alternative Weighting Strategies: Probit Models (Marginal Effects Reported)

	Specification (4)			Specification (6)		
	Weight 1	Weight 2	Unweighted	Weight 1	Weight 2	Unweighted
Union	0.031***	0.033***	0.027***	0.070***	0.086***	0.061***
	(0.011)	(0.012)	(0.011)	(0.022)	(0.026)	(0.021)
Primary Sector	0.002	0.005	0.007	0.002	0.011	0.005
	(0.014)	(0.016)	(0.015)	(0.016)	(0.018)	(0.017)
Trans. & Comm.	-0.044***	-0.036***	-0.044***	-0.034***	-0.023*	-0.033***
	(0.009)	(0.010)	(0.009)	(0.011)	(0.012)	(0.010)
Other Services	-0.052***	-0.050***	-0.049***	-0.041***	-0.037***	-0.040***
	(0.010)	(0.010)	(0.009)	(0.011)	(0.011)	(0.010)
Primary Sector*Union	-	-	-	-0.005	-0.026	0.019
	-	-	-	(0.044)	(0.049)	(0.042)
Trans. & Comm.*Union	-	-	-	-0.071***	-0.089***	-0.069***
	-	-	-	(0.031)	(0.035)	(0.028)
Other Services*Union	-	-	-	-0.079***	-0.096***	-0.064***
	-	-	-	(0.031)	(0.035)	(0.028)
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Gender and Race Controls	Yes	Yes	Yes	Yes	Yes	Yes
Age and Education Controls	Yes	Yes	Yes	Yes	Yes	Yes
Part-time Status Control	Yes	Yes	Yes	Yes	Yes	Yes
Region Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Industry Fixed Effects	No	No	No	No	No	No
Observations	7,436	7,436	7,436	7,436	7,436	7,436

Notes. Table A.3 compares the estimates presented in Table 2, obtained using our weighting strategy (Weight 1), to two alternative weighting strategies (i.e 'Weight' 2 and 'Unweighted'). Estimates under 'Weight 1' are obtained using weights that sum to the sample size of each (yearly) survey. Estimates under 'Weight 2' are obtained using weights that sum to 1 for each year of the survey. The interaction term coefficient estimates were obtained using the methodology proposed in Ai and Norton (2003). Weighted standard errors are in parentheses. * significant at 10%; ** significant at 5%; *** significant at 1%.