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John T. Addison
Paulino Teixeira
Alex Bryson
André Pahnke

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John T. Addison
University of South Carolina, GEMF and IZA

Paulino Teixeira
University of Coimbra, GEMF and IZA

Alex Bryson
NIESR and CEP

André Pahnke
IfM Bonn

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IZA

P.O. Box 7240
53072 Bonn
Germany

Phone: +49-228-3894-0
Fax: +49-228-3894-180
E-mail: iza@iza.org

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ABSTRACT

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This paper depicts and examines the decline in collective bargaining coverage in Germany. Using repeat cross-section and longitudinal data from the IAB Establishment Panel, we show the overwhelming importance of behavioral as opposed to compositional change and, for the first time, document workplace transitions into and out of collective agreements via survival analysis. We provide estimates of the median duration of coverage, and report that the factors generating entry and exit are distinct and symmetric.

JEL Classification: J50, J53

Keywords: sectoral and firm agreements, changes in collective bargaining/works council coverage, shift-share analysis, bargaining transitions, survivability

Contact:

Paulino Teixeira
Faculdade de Economia
Universidade de Coimbra
Av. Dias da Silva, 165
3004-512 Coimbra
Portugal
E-mail: pteixeira@fe.uc.pt

I. Introduction

Use of the characteristics of collective bargaining to help motivate analysis of wage and employment outcomes occupies an important position in contemporary treatments of the covariation of institutions and macroeconomic outcomes. Thus, notions of the importance of the centralization of collective bargaining (or its absence) to wage and unemployment development figured heavily in policy discussions in the 1980s (Calmfors and Driffill, 1993). More recently, the importance of centralization has been supplemented if not supplanted by notions of coordination (OECD, 2004, Chapter 3). *Vulgo*: more centralized bargaining regimes – and latterly more coordinated ones – have been held out as offering scope for improved economic performance.

One important issue that has arisen is the stability of the underlying relationships. After all, it was the failure of the Swedish model that spawned the coordination thesis. Might not coordinated systems for their part also be subject to a possibly pre-set cycle of emaciation and decay? In any event, bargaining structures – centralized, coordinated, or otherwise – are typically observed at discrete points in time. Insufficient attention has been paid to within-country changes in the degree of centralization/coordination in collective bargaining regimes and correspondingly perhaps too much attention given over to shocks per se. In this sense, the literature on the role of bargaining structure is no different from that on the contribution of some other 'key' institutions such as employment protection and labor standards where time variation in regressors is at best sporadic.

Yet we live in a time in which systems are said to be increasingly under stress. If so, they might be expected to evolve or fail. Nevertheless, there is in general very little discussion of the change in institutions outside of studies of the decline in union density (which phenomenon has tended to be associated in the Anglo-Saxon literature at least with the notion of a decline in the 'disadvantages' of unionism; see, for example, Addison and Belfield, 2004). Although the change in German institutions has received some attention in the wake of a precipitous decline in

unionization, the research has proceeded in a patchwork fashion and remains controversial (see below). One aim of the present treatment, therefore, is to offer a comprehensive and updated examination of the course of collective bargaining *and* worker representation in Germany since 2000.

In the present paper, we will first chart the extent of erosion in the twin pillars of the dual system of industrial relations in that nation. Distinctions will be made between western and eastern Germany, between large and small firms, between manufacturing and services, and between surviving, newly-founded, and failing establishments. We also model changes in collective bargaining using shift-share analysis, providing points of contact with a German literature examining the determinants of union density (e.g. Fitzenberger, Kohn, and Wang, 2006) and the emerging consensus that changes in the composition of the workforce have played a minor role in the decline in union density (in our case, sectoral collective bargaining).

We will also update an altogether sparser and typically cross-sectional German literature on the determinants of the structure of bargaining covering both the application of sectoral agreements and their abandonment.[1] The novelty of our analysis stems from the longer observation window during which plants and their collective bargaining status are being observed consecutively. More concretely, collective bargaining 'membership' is analyzed within the framework of an unobserved (random) effects probit model, while empirical discussion of establishment transitions into and out of collective bargaining is tackled in the context of a survival model. We view these innovations as the principal contributions of the present study.

II. A Brief Thematic Survey of Past Research

There has been considerable discussion of the future of the German 'model' in recent years despite the continued institutional predominance of industry-wide or sectoral collective bargaining. In particular, the practical locus of collective bargaining has shifted to lower levels, leading observers to question whether this development represents an ongoing process of erosion

or is instead indicative of the natural accommodation to changed circumstances of a flexible system.

Unambiguously the German system has been decentralizing. Apart from embracing 'individual' as opposed to collective bargaining, firms were initially to switch from sectoral to firm-level collective bargaining (Hassel, 1999). But sectoral agreements were also evolving to permit greater flexibility. The means included *opening clauses* and latterly *pacts for employment and competitiveness* (see, respectively, Bispinck, 2004; Seifert and Massa-Wirth, 2005). The issue has been whether the working out of such contractual innovations – particularly the latter – reflects a *coordinated* or managed decentralization or, in conjunction with declining collective bargaining coverage, a distinct change in model?

Observers such as Massa-Wirth and Niechoj (2004: 22-23) speak of a process of increasingly uncontrolled decentralization associated in particular with pacts – even in those cases where they are not in actual contravention of sectoral labor contracts. Other observers also see the seeds of ultimate destabilization in otherwise organized decentralization (i.e. where issues have been formally delegated from central level to the plant level) by virtue of the effects on the disparate interest membership of employers' associations, chiefly large versus small firms (Hassell, 1999).[2]

For its part, orthodoxy has tended to stress the notion of transformation without disruption. Specifically, it has been argued that German employers have a vested interest in maintaining the dual system, that the system possesses powerful flexibility, permitting adjustment to outsourcing and other major changes without conflict, that pervasive cooperation is the order of the day, and that the appearance of institutional instability is a response to the business cycle (see, respectively, Thelen and Van Wijnbergen; Streeck, 2001; Frege, 2003; Klikauer, 2002).

Nevertheless, information on the facts of the case as reflected in the dual system as a whole is sparse. Much of the extant literature referred to earlier tends to focus on sectoral bargaining alone (see, for example, Kohaut and Schnabel, 2003a, 2003b, 2007). Wider-ranging

analyses include the study by Addison et al. (2009), which covers the interval 1998-2004, and upon which the present treatment builds and the partial updates provided in German-language studies by Ellguth and Kohaut (2008, 2010). As noted, one important goal of the present treatment is to modernize and extend the focus of previous research, even if the issue of performance of the full range of institutions in question raised by this thematic review necessarily is the task of future research.

III. Data

Our data is extracted from the IAB Establishment Panel (or *Betriebspanel*). The Panel is based on a stratified random sample of plants from the population of all establishments with at least one employee covered by social insurance. The basis for sampling is the Federal Employment Agency establishment file, containing information on some 2 million establishments. Since good detailed descriptions of the Panel, which is conducted annually and now contains information on a little over 16,000 plants, are now widely available (e.g. Fischer et al., 2009), we choose here to confine our remarks to outlining the procedures used to generate our various estimation samples.

First, given that we seek to offer a complete picture of the course of collective bargaining coverage over a sufficiently long period of time, we took the most recent survey available at the time of writing and appended all the previous surveys back to 2000. We decided not to range further back in time primarily to avoid having to deal with material changes in industry classification in 2000 (from a 3- to a 5-digit system).

Second, we focus on establishments from the private, profit-oriented sector of the economy. For reasons connected with the need to include works councils in our sample, we also excluded establishments having less than 5 employees – the legal size threshold for the establishment of works councils. In total, we have some 82,000 observations on approximately 24,000 establishments in the whole of Germany.

Third, the selected covariates – data-driven in the main and largely self-explanatory – are presented in Table 1. The principal covariates comprise two measures of workforce composition based on skill and gender, foreign ownership, single versus multi-site firm status, establishment age, establishment size, and an indicator of the state of technology in use. They are augmented by a total of thirty seven 2-digit industry dummies plus sixteen regional dummies. Although somewhat sparse, our choice of regressors is guided by the literature (notably, Schnabel, Zagelmeyer, and Kohaut, 2006) *and* the need to minimize the loss of establishments occasioned by missing observations.

(Table 1 near here)

Fourth, the (nine) surveys selected are used in cross-section fashion to chart the main developments in collective bargaining and worker representation coverage (in section IV of the paper). For its part, the constructed longitudinal dataset – namely, the panel in which establishments are followed over time for a maximum period of nine years (in the case of those plants populating all surveys from 2000 through 2008) – is used initially to examine the determinants of collective bargaining (in sections V and VI) and thence the duration of collective bargaining status as either a covered or uncovered institution (section VII).

Finally, observe that in general we do not know the elapsed duration of the observed spells. That is to say, we do not know the number of years in which a given establishment has been either covered or uncovered at the point it is first observed in the survey. As a result, all establishments are left-truncated, with the notable exception of the newly-founded establishments (i.e. births) that we were able to follow from the outset. One of our tasks therefore was to ensure that the year of birth coded in the survey panel was correct. To this end, we used the establishment register (or *Betriebsdatei*) and the fact that establishments in the two raw datasets (i.e. *Betriebsdatei* and *Betriebspanel*) share exactly the same identification code (or *Betriebsnummer*). Further information on the construction of the different estimation samples is provided below.

IV. The Course of Collective Bargaining

The extent of collective bargaining and worker representation is described in Table 2 and Figure 1, where we distinguish between industry-wide and firm-level collective bargaining, an absence of collective bargaining, and works council presence. Coverage is given by employment and by establishment for Germany as a whole and its western and eastern halves. The most notable feature of Table 2 is the increase in the prevalence of no collective bargaining. This is largely the result of a fall in industry-wide or sectoral bargaining. Note that there has been little change in firm collective bargaining, while works council coverage has fallen over the sample period, despite the passage of legislation in 2001 designed to increase their coverage. These trends are graphed annually in panels (a) and (b) of Figure 1 for employment and establishment shares, respectively. Bargaining coverage, especially sectoral bargaining, is much lower in eastern than western Germany throughout the period, but the rate of decline in bargaining coverage is more pronounced in the latter region.

(Table 2 and Figure 1 near here)

Another important distinction to be made in addressing the decline in traditional bargaining and the growth of bargaining-free regimes is establishment size. Figures 2 and 3 graph coverage by employment and establishment for plants with less than and greater than 250 employees. First, it is clear that *levels* of sectoral bargaining are considerably higher in larger establishments and absence of collective bargaining correspondingly lower. Even more striking is the disparate coverage of works councils in establishments of different sizes. Whereas the vast majority of larger firms have councils, only a minority of smaller ones do so. Larger plants are also more likely to have firm-level collective agreements than their smaller counterparts, although the disparities here are very much smaller.

(Figures 2 and 3 near here)

In terms of changes in levels, however, the growth in bargaining-free regimes has been somewhat more pronounced in larger plants. The figures are reversed in respect of the declines in

sectoral bargaining and works council coverage. For sectoral bargaining, declines in coverage by employment and establishment are substantially higher among smaller plants. In the case of works council coverage, rather small declines are observed in the case of larger plants as compared with major declines in smaller plants. Finally, if anything modest upward trends in firm-level collective bargaining characterize both large and small establishments in terms of their employment coverage.

(Figures 4 and 5 near here)

Another disaggregation worth pursuing is coverage in manufacturing versus that in services. The situation is depicted in Figures 4 and 5. Again, the principal distinction is more one of levels than first differences. Thus, services are clearly differentiated from manufacturing by their lower incidence of traditional bargaining and correspondingly higher shares of bargaining-free regimes, but over the period in question the decline in collective bargaining and the growth in no collective bargaining was fairly similar as between the two sectors. And while the decline in works council coverage by employment was much more sizeable in services than in manufacturing broadly similar declines in establishment coverage were observed in the two sectors. Finally, the employment coverage of firm-level agreements grew in both sectors, although establishment shares hardly budged.[3]

To determine whether the observed changes in collective bargaining between 2000 and 2008 are the result of behavioral or compositional factors we next turn to a shift-share analysis. (A parallel treatment of workplace representation is available from the authors upon request.)

V. Shift-share Analysis

The percentage point change in collective agreements (mean) coverage between 2000 and 2008 can be decomposed into its Oaxaca-Blinder components: the between or compositional effect, and the within or behavioral effect. The between effect, or the ‘explained component,’ is that part of the observed change that can be attributed to differences in observable characteristics. The within

effect, or unexplained component, measures the change in coverage arising from differences in propensities (or coefficients).

More formally, let $x_{2008}b_{2000}$ be the 2008 (predicted) coverage based on year 2000 coefficients, where x denotes the mean vector of observed (establishment) characteristics and b indicates the vector of estimated coefficients. Then, the between effect is given by $(x_{2008} - x_{2000})b_{2000}$ and the within effect by $x_{2008}(b_{2008} - b_{2000})$, where the reference groups are the year 2000 coefficients and the year 2008 characteristics, respectively. (A different choice of reference groups would yield $(x_{2008} - x_{2000})b_{2008}$ and $x_{2000}(b_{2008} - b_{2000})$ for the between and within effects, respectively.)

For expositional convenience, our decompositions rely on linear estimates.[4] Following on the data description given in section III, our selected vector of covariates x includes establishment size, the proportion of skilled and female workers, and dummies for single-establishment status, foreign ownership, establishment age, state of technology, industry and region.

The results from the shift-share exercise are presented in Table 3 for Germany as a whole and for eastern and western Germany separately. Panel (a) of the table refers to collective agreements of any type, while panel (b) refers to sectoral agreements.

(Table 3 near here)

Three main findings emerge from the table. First, the within effect is overwhelmingly dominant, accounting for at least 90 percent of the observed change in coverage in either panel. Second, the decompositions in the two panels are very similar, which of course reflects the fact that the share of firm-level agreements is relatively small. Third, the declining coverage observed in eastern Germany, while less pronounced is again dominated by a within effect of approximately the same proportion as in western Germany. Evidently, changes in the propensity of being covered lie at the root of the decline in collective agreements irrespective of the magnitude of that decline.[5]

We also note that given that the percentage point change over 2000-20008 is close to zero, the decompositions with respect to the changes in firm-level coverage – not reported in the table – are something of a *curiosum*: the between and within effects become very large in percentage terms even if they are actually very small in absolute size.[6] Finally, we found no evidence that any particular variable (or set of variables) is driving the results of the decomposition described in Table 3. All individual composition (or characteristics) effects are small, and no individual within effect (attributed to any observable characteristic) is statistically significant, with the sole exception of the industry dummies.

These findings would seem to suggest that unobserved establishment traits play a role in the observed decline in collective bargaining coverage in Germany over the last decade. We now turn to a closer examination of this issue.

VI. Collective Bargaining Coverage Propensity

We now take full advantage of the longitudinal nature of our panel to ascertain the determinants of coverage propensity. We will consider in particular the extent to which observations within an individual establishment are correlated over time. Since the outcome variable is a binary variable, we shall deploy an unobserved (random) effects probit model. It will be recalled that the maximum length of any individual time series in our panel is nine years (in the case of those establishments observed consecutively from 2000 to 2008).

Let Y_{it} represent the coverage outcome for the tth observation in the ith establishment. Given the random effect u_i which represents the establishment's persistent *unobserved traits* – its unobserved propensity to be covered – the random-effects probit model can be specified as

$$\Pr(Y_{it} = 1 \mid u_i, X_{it}) = \Phi(X_{it}\beta + u_i)\,, \tag{1}$$

where Φ is the standard cumulative distribution function and $u_i \sim N(0, \sigma_u^2)$, with u_i uncorrelated with X_{it}; X includes all observed establishment characteristics that have an impact on the binary response probability; and β denotes the set of parameters to be estimated.[7]

Conditional on (u_i, X_{it}), outcomes Y_{i1}, Y_{i2}, ..., Y_{iT} are independent, with probabilities depending on u_i and X_{it}. This means that, conditioning only on X_{it}, Y_{i1}, Y_{i2}, ..., Y_{iT} will be dependent across *t*. A useful statistic therefore is the (latent) intra-class (establishment) correlation, given by $\rho = \frac{\sigma_u^2}{\sigma_u^2 + 1}$, which indicates the relative importance of the unobserved effect u_i or the correlation between $u_i + e_{it}$ across any two time periods (see, for example, Rodríguez and Elo, 2003). We will also exploit an additional measure of (manifest) association based on the actual binary outcomes Y_{it}, rather than on the latent variable Y_{it}^*, namely Pearson's *r* coefficient. Along with these measures, we will use other indicators evaluated with the linear predictor set at various percentiles, the goal being to have different measures of status persistence.

Using the model in equation (1), the determinants of being covered by type of collective agreement are presented in Table 4. We retain in the sample all plants surveyed in the 2000-2008 observation window, including those switching collective bargaining status more than once. As a practical matter, however, dropping the latter produced virtually no change in the results. Our set of covariates is unchanged from section V, and for expositional convenience we focus exclusively on Germany as a whole. The broad rationale for inclusion of these covariates can be found, for example, in Willman, Bryson, and Gomez's (2007) modeling of employer voice-choice decisions. Based on the argument that firms face non-trivial switching costs (i.e. costs connected with uncertainty surrounding the benefits from moving from coverage to non-coverage, and vice-versa), one would expect the returns to being covered by collective agreements to be higher in large establishments and in plants integrated in multi-site establishments. Establishments with a

higher proportion of low-skill employers are also likely to rely less on voice mechanisms and therefore expected to be associated with a lower presence of collective agreements. By the same token, older establishments are more likely to be covered given that the incidence of collective tended to be higher in the past.

(Table 4 near here)

From the first column of the table, which refers to coverage propensity by any type of collective agreement, it can be seen that establishment size and establishment age are positively and single-establishment firm status negatively associated with coverage. This propensity is also increasing in the skill composition of the workforce. The sectoral agreements case, given in the second column of the table, mirrors the results for all collective agreements. The principal exception is the state of technology variable: more modern plants now evince a higher propensity to be covered by a collective agreement. Finally, from the third column of the table, we observe that the sign of the state of technology variable is reversed and that the association between plant age and coverage by a local, firm agreement turns negative. That said, given the statistical insignificance of the latter coefficient estimate, we have not uncovered evidence to favor the proposition that newly-founded firms are attracted by firm-level agreements, while the negative sign of the technology argument might suggest that firms facing more competition by reason of outdated technology may be those opting out of sectoral agreements.

With a few exceptions, the industry and region dummies are statistically significant. However, other than the lower propensity of eastern Germany establishments to be covered by a collective agreement, there are no obvious patterns in the data in this regard.

Of interest is the high value of ρ throughout, indicating considerable inertia in collective bargaining status. In short, there is strong evidence that, controlling for X_i, the probabilities of an establishment being covered in any t_0 and t_1 are highly correlated. (The presence of non-trivial switching costs may of course lie at the root of this outcome.) Equivalently, the size of σ_u

(ranging from 2.6 to 3.9) implies that a small difference in unobserved traits entails a quite different propensity of being covered by a collective agreement. We also note that since the significance test for ρ is itself a test for the presence of the unobserved (random) effect, we can reject the simple pooled probit as an appropriate model description of the data.

(Table 5 near here)

The manifest interclass correlation across distinct percentiles in given in Table 5.[8] We again focus on the any collective agreement case in panel (a) and on the median percentile. For an establishment with a median probability of being covered by any type of collective agreement (the 0.50 column), the inter-class correlation is 0.76, flagging a substantive within group persistence. Note also that for the median percentile, the corresponding joint probability in the second row (viz. the probability of being covered in two given years) is equal to 0.47. In turn, the corresponding marginal probability of being covered by any type of agreement in any given year is 0.53 (first row), which is not too far away from the mean coverage rate observed in the sample of 52.7 percent (see Table 1). Finally, the odds ratio in the fourth row indicates that the odds of an establishment being covered in t_0 and t_1 *versus* not being covered in t_0 but covered in t_1 are 145 times higher for the same observed characteristics. Since the odds ratio contrasts the (same) behavior of two individuals (viz. establishments) in t_1, given that in t_0 they may have behaved differently, the conclusion is that it is considerably more likely that establishments that are covered will stay covered than non-covered establishments will join. Inertia in non-coverage is therefore very strong as well. That said, there is much less persistence in firm-level bargaining.

Finally, by squaring the Pearson's r coefficient, we obtain the interesting result that collective bargaining coverage in a given year explains about 57 percent of the variation in collective bargaining behavior in another year. The inference is that there is no terminal inertia in collective bargaining status, which result offers more than sufficient justification for an analysis of transitions into and out of collective bargaining.

VII. Transitions and Collective Bargaining 'Survivability'

We have seen that certain characteristics are associated with collective bargaining coverage. But can we say for example that the longer lasting is its coverage, the less likely an establishment will be to change bargaining status? Our concern is now with the specific factors that induce *failure*, that is, transitions into or out of a collective agreement. The proper context for such analysis is survival modeling.

In our observation window, we have a maximum of nine annual observations which is insufficient to allow us to follow all production units from outset (birth) to death. The typical unit in our panel is indeed one that was born before 2000 and surveyed over a certain number of years within the observation interval. Figure 6 illustrates the array of possibilities. Establishment A, for example, was born before 2000 and is observed consecutively from 2000 up to point *e* (exit from a given state or point of 'failure'). Establishment A has therefore a left-truncation point as it is not possible to recover its bargaining status prior to 2000. Establishment B is not only left-truncated but also right censored as well since it rotates out of the panel at point c. For their part, establishments C, D and E are observed for a number of years up to (a) 'failure', (b) self-rotation, and (c) right censoring (in 2008), respectively. Establishments F and G were born after 2000 and are, respectively, right censored and exiting a given state before 2008. Finally, there are those 'permanent' establishments, represented by case H, which are both left- and right-censored (in 2000 and 2008, respectively). In general, we will not be able to know the exact length of all spells because it is simply not possible to recover the 'missing' information. On the other hand, newly-founded establishments – and, to some extent, permanent establishments – are a special case and they will be used to explain the survivability of collective bargaining. Again in the interests of expositional convenience, we focus on the aggregate category of collective agreements of any type.

(Figure 6 near here)

In the limit, the probability of failure, given by the hazard function, is constant and independent of any establishment attribute. This case is not particularly helpful in the present context since we believe that the selected covariates do have an impact on the hazard rate. Thus, we assume that leaving (or joining) a collective agreement of any type is a function of an observed set of time-constant (e.g. industry dummies) *and* time-varying (e.g. establishment size) covariates.[9]

Our hazard function belongs to the family of proportional hazard (PH) models

$$h(t; X) = k_1(X)k_2(t), \tag{2}$$

where k_1 and k_2 are the same functions for all individuals (establishments) and X is the vector of the selected covariates (see, for example, Lancaster, 1990, chapter 3). Setting $k_2(t) \equiv h_o(t)$ and $k_1(X) = \exp(X\beta)$, we have the standard proportional hazard Cox model

$$h(t; X) = h_0(t)\exp(X\beta), \tag{3}$$

where $h_0(t)$ is the baseline hazard (or the hazard rate when all covariates are set at zero).[10] Thus, $h(t)$ denotes, for covered (uncovered) establishments, the probability of an establishment leaving (joining) a collective agreement of any type, given that it has been covered (uncovered) up to time *t*. Given the longitudinal nature of our dataset, the standard errors of the estimated hazard coefficients are adjusted to account for the possible intra-group (establishment) correlation.

As mentioned earlier, we have both stock and flow sampling in our data, in the sense that we are able to observe *entrants* (newly-founded establishments) and non-entrants (i.e. establishments born at some point in the pre-observation period.[11] In the case of non-entrants, for whom left-censoring is the key problem, some further data manipulation will be required. For entrants, the survival analysis is straighforward since all spells for these units are either complete or right censored. In this context, the subsample of births turns out to be extremely useful, and we

will discuss below the extent to which inferences based on births can be carried forward, first, to the subset of permanent establishments and then the entire sample of surveyed units.

As shown in Table 6, we observe 2,679 births in the 1999-2007 period. Of the total number of births, there are 266 collective agreement transitions in the 2001-2008 interval, comprising 149 leavers and 117 joiners. In other words, 9.9 percent of all births either switched into or out of a collective agreement during the sample period.

(Table 6 near here)

Table 6 also gives the collective agreement status in the year of birth *and* in the year of exit for all births in the sample, as well as the average year of exit (i.e. self-rotation or transition into a different state) for each cohort. For example, an establishment born in 1999 is observed over an average period of 2.6 years before switching to a different regime or leaving the panel. Interestingly, the expected year of exit for our sample is virtually the same for covered and uncovered establishments. In any event, for establishments born later in the period, the average number of years prior to exit is necessarily smaller given that their number of years in the observation window becomes shorter.

From the total number of births in our dataset, and ignoring the 2007 cohort for which no transitions can be observed, in 52.2 percent of the cases establishments remain non-covered and 37.9 percent remain covered. This implies, as we have seen, that in 9.9 percent [100-(52.2+37.9)] of the cases we do observe establishments changing – either leaving or joining – their collective agreement status. Of those plants that are covered in the year of birth, some 12.8 percent do switch out of collective agreement within the observation window, while 7.7 percent of their non-covered counterparts will join a collective agreement. (Multiple failures – establishments with more than one transition over the observation period – are now dropped from our sample.)

The results of model (2) – the hazard function – are presented in Table 7 for the two possible failure events: leaving a collective agreement and joining one (first and second columns of the table, respectively). In the last row of the table, we also present the median duration of

coverage/‘uncoverage,’ based on a PH exponential model without covariates. According to our estimates, the median duration of coverage for newly-founded establishments is approximately two years, while the median duration of uncoverage is around three years.

(Table 7 near here)

As for the role of the selected covariates, greater establishment size decreases the probability of leaving a collective agreement, as does the use of modern technology. In contrast, foreign ownership and single-establishment status are associated with a higher failure rate. Note that the role of single-establishment status and foreign-owned variables are particularly strong; in particular, being a single establishment implies an 83 percent higher hazard rate, while foreign-ownership increases the hazard by 58 percent. In turn, a 1 percent increase in establishment size reduces the hazard by 0.35 percent. All other covariates included in the regression are poorly determined.

The results for joining a collective agreement of any sort (second column) look quite symmetric, such that where the probability of leaving a collective agreement is found in the first column of the table to be decreasing in employment size, it is now increasing in employment size when it comes to joining an agreement. But no other covariate is found to be statistically significant. *Vulgo*: establishment size is the major determinant of joining a collective agreement. The evidence is much weaker in the case of transitions into collective agreements, however, which is not altogether unexpected given the smaller number of establishments engaging in such switching behavior.

We recall that in our observation window all units are left-censored except for newly-founded plants. Since we cannot recover the entire record on collective bargaining participation in respect of the left-censored units, we can either ignore all transitions other than in the case of the sample of births or instead try to figure out an alternative procedure that avoids losing the valuable transition information we have on other types of establishments.

We chose the second route and therefore create a constructed pre-observation period in which collective agreement status is unchanged for all units included in the risk analysis. To this end, we (a) divide the 2000-2008 period into the two sub-periods 2000-2003 and 2004-2008, (b) use the set of permanent establishments (these units were observed for a reasonably long period of time anyway), and (c) impose the additional restriction of no change in status from 2000 to 2003. Transitions in the 2004-2008 interval will then be used to estimate the hazard. We will refer to this sample as the 'restricted sample of permanent establishments.' (Note that in enlarging the 'pre-observation' period from 2000-2003 to 2000-2004, for example, we reduced the risk period with no appreciable change in the results, other than a slight decrease in significance levels.)

In a second stage, and to test the role of left-censoring in our results – and ultimately evaluate whether the use of left-censored data in our survival analysis is legitimate – we added to the restricted sample of permanent establishments all those units in which collective bargaining status prior to 2004 is not fixed.[12] Taking, for example, the case of covered establishments this *counterfactual* exercise serves to compare the results from an experiment in which the left-censored units are necessarily covered with the case in which the presumed fixed coverage prior to 2004 is false for some units – and similarly for the case where the initial state is non-coverage. If the determinants of the hazard rate in the two counterfactual experiments are not too different (that is, where the hazard is not too sensitive to changes in the selected samples), we may conclude that left-censoring for permanent units of the panel is not really an issue, and that running the survival analysis on an 'unrestricted' set of permanent establishments is not too much of a stretch. In this vein, our third and final exercise applies the survival model to all permanent establishments observed in 2000-2008 period, without further restrictions. Again, in this case we are simply ignoring left-censoring, implicitly assuming that either there was no change in status in the past (i.e. before 2000) or, alternatively, that it occurred too long ago to be a matter of concern.

We have exactly 1,448 establishments in the restricted estimation sample of permanent establishments, of which 821 (627) were covered (not covered) in 2000-2003. Of those that were

covered (in 2000-2003), 93 switched out of collective bargaining between 2004 and 2008 – 93 out of 821, or 11 percent. Of those that were not covered, 35 switched into collective agreements after 2003 – 35 out of 627, or 6 percent.

The corresponding survival analysis, shown in the first column of Table 8, again indicates that establishment size is critical: the larger the establishment, the lower the probability that a covered establishment will leave a collective agreement. The single establishment variable is also well determined, and positively signed as expected. All the other variables are poorly determined. In turn, as shown in the second column of the table, joining collective agreements is a lot less common among permanents than among newly-founded establishments; recall that the number of observed failures is only one-third that of the number of transitions out of coverage. Not surprisingly, therefore, all variables in the second column are statistically insignificant, with the sole exception of the establishment age dummy. In this case, older establishments tend to have a lower exit rate (from non-coverage). Apparently, non-covered establishments tend to stay non-covered, while the considerable minority that join collective agreements do not seem to share any particularly visible characteristics.

(Table 8 near here)

The second experiment – the *counterfactual* – is given in Table 9. In this exercise, we added some 50 establishments to the sample in the first column of Table 8. The results are basically unchanged, so that we conclude that once we observe the state (coverage) of a permanent establishment, transition behavior tends to be quite predictable. The same obtains with respect to the transition behavior of initially uncovered establishments, shown in the second column, where some 100 establishments have been added to the sample. The main implication from the counterfactual is, again, that within the subsample of permanent establishments there seems to be no particular penalty in ignoring left-censoring.

(Tables 9 and 10 near here)

Given these findings, the final step is to present the survival analysis for the full set of permanent establishments. This procedure yields an enlarged estimation sample of 1,597 units, surveyed consecutively from 2000 to 2008. Of this total, we have exactly 922 (675) establishments that were covered (not covered) by any type of collective agreement in 2000, and 275 transitions comprising 193 leavers and 82 joiners. The results are presented in Table 10. As expected, the results reported in the table mimic those obtained earlier in Table 8. From this perspective, it appears legitimate to conclude that in the case of permanent panel members there is enough evidence to support the proposition that plant size and skill content of the workforce matter in terms of collective bargaining survivability, while single establishment status favors the abandonment of collective bargaining. The influence of the remaining covariates on survivability of collective agreements is statistically weak but nevertheless mildly visible, with the exception of the establishment age variable. However, it is more difficult to discern equally strong patterns in respect of transitions into collective agreements. Here, size and, to some limited extent, foreign ownership are the unique determinants, with again strikingly symmetric effects.

VIII. Conclusions

The steady decline in collective bargaining coverage in Germany has been documented in a patchwork fashion in the extant literature. Based on a detailed analysis of its development over the last decade, this paper establishes that the downward trend identified in that literature has likely not come to a halt, although there is no real indication of any continued substitution *within* collective bargaining (i.e. of multi-employer, sectoral agreements being replaced by firm level agreements). That process seems to have been sidelined by the decentralization of sectoral bargaining, not that we can yet speak of a clear process of organized decentralization. Nor for that matter do plant births or deaths emerge as the main driving force behind the observed fall in collective bargaining coverage. Rather, the decline appears to be across the board, affecting

regions, sectors, small and large firms alike, and proceeding irrespective of the establishment's workforce composition.[13]

Our multivariate shift-share analysis suggests that changes in establishment characteristics play a small role on the course of collective bargaining over the observed period. The main source, therefore, is attributable to behavioral effects, even if no single factor can easily be identified as the chief suspect. What is clear is that economic circumstances are such that establishments in the late 2000s are definitely less prone to be covered than they were earlier in the decade. Globalization might be a good candidate: all else constant, increased product market competition is likely to stimulate a move away from 'sticky' collective agreements. This trend is also revealed by our analysis of 'membership,' or establishment coverage, where it is shown that unobserved establishment traits can explain much of the variation in coverage.

Despite the role played by unobserved heterogeneity, however, our duration analysis had shown that the set of regressors deployed here have non-negligible predictive power. That is to say, we are able to explain a material part of transitions into and out of collective agreements, especially in the case of the decision to leave collective bargaining agreements. Based primarily on a very careful coding of all births in the sample and then on a thorough modeling of left-censored *permanent* establishments, we were able to present – for the first time to our knowledge – the median duration of coverage and 'uncoverage' for newly-founded establishments, while at the same time offering an analysis of collective bargaining transitions for other types of establishments.

Although we cannot provide conclusive evidence that the decline in collective bargaining is irreversible, it is unquestionably the case that the German model is under stress. To be sure, our finding of considerable inertia in the process is not consonant with the claim that German collective bargaining is currently an endangered species on the U.S. (private-sector) pattern, but even here others have suggested that this inertia is undergirded by political support, without

which the erosion of the German system of industrial relations would be even more rapid and more pronounced (e.g. Hassel, 2002).

The *consequences* of changes in collective bargaining will form the next stage in our empirical inquiry, the first step of which will be to determine whether wages are lower in plants that abandon industry-level collective bargaining, and if not whether, say, organized decentralization (in the German terminology) allows sufficient adaptation to changing circumstances. Such work should assist in our understanding of the efficacy of existing broad-based classifications of collective bargaining systems used in macro treatments. It will also have a bearing on standard identification strategies used in tackling unobserved firm and worker characteristics.

Endnotes

1. See Kohaut and Schnabel (2001), Kohaut and Schnabel (2003b), Schnabel, Zagelmeyer, and Kohaut (2006).

2. Hassel (1999, 2002), in particular, is also concerned to stress the role of the works council because its purported decline necessarily limits what can be expected of organized decentralization; that is to say, the transfer of collective bargaining functions from the collective bargaining arena to plant level is only viable where works councils are actually in place.

3. A breakdown of collective agreement and worker representation coverage for plants that are observed in both 2000 and 2008 (i.e. *stayers*) offers the same broad pattern: a growth in the collective bargaining free zone; a certain decline in sectoral bargaining (albeit somewhat less obvious than earlier reported); a shrinking works council sector; and volatile levels of firm-level bargaining. Regarding *births* and *deaths*, however, there is indication that absence of collective bargaining is higher among newly-born firms for both coverage measures (with the growth rate being much higher for the employment measure). Sectoral bargaining is more common among dying establishments, again on either measure. Newly-born establishments are also less likely to have works councils than dying establishments, which serves to confirm the growth of a codetermination-free zone. Full details are available from the authors upon request.

4. Our findings are robust to probit estimation, with within- and between components of virtually the same size as those reported below. Full results of the probit exercise are available from the authors upon request.

5. We note that the decomposition is insensitive to the choice of reference groups, with the possible exception of eastern Germany where the within effect tends to be larger when the 2008 coefficients are selected as the reference category.

6. As a matter of fact, for the whole of Germany and for western Germany the (statistically weak) evidence suggests that the observed changes in establishment characteristics are *per se* favorable

to a higher coverage of firm-level agreements, while for eastern Germany the within effect is again dominant.

7. The equivalent latent variable model is given by $Y_{it}^{*} = X_{it}\beta + u_i + e_{it}$, where Y_{it}^{*} is the latent variable and $e_{it} \sim N(0,1)$, with e_{it} uncorrelated with u_i. Assuming $\Pr(Y_{it} = 1 \mid u_i, X_{it}) = \Pr(Y_{it}^{*} > 0 \mid u_i, X_{it})$, model (1) follows easily.

8. We cannot offer a similar exercise for ρ since it does not depend on the marginal distribution.

9. For the time-varying covariates, we shall ignore possible anticipation and delay effects. We shall also assume that the effect of any continuous variable on the hazard is independent of the level of the variable (i.e. the marginal effect is constant). A model without covariates will be used to obtain the predicted median duration of coverage/'uncoverage' for newly-founded establishments (see Table 7).

10. Formally, the model in equation (3) is PH with time-invariant covariates; the corresponding PH model with time-varying variables is given by $h[t; X(t)] = h_0(t)\exp[X(t)\beta]$ (see Wooldridge, 2002: 693).

11. The year of birth of any establishment in the panel is always known; only the bargaining status in the pre-observation period is unknown.

12. For transitions into collective agreements, this amounts to adding the following sequences to the existing restricted sample of permanents: 0111|11111, 0011|11111, 0001|11111, and 0000|11111. In the case of transitions out of collective agreements, we add the sequences 1111|00000, 1110|00000, 1100|00000, and 1000|00000. The vertical bar in these sequences denotes the 2003 separation point and '1' ('0') signifies coverage ('uncoverage'). The 2004-2008 interval defines the risk period.

13. Note that if one is ready to accept that covered establishments suffer from lower employment, employment growth will be concentrated in non-covered establishments, which fact can only imply an inevitable decline in collective bargaining in the long-run. But in the presence of pro-productive collective voice, the optimal mix of covered establishments in the economy is likely to be non-zero.

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TABLE 1
Variable description and means of the raw sample

Variable		Mean	*n*
Any type of collective agreement	Dummy	0.527	82,137
Sectoral agreement	Dummy	0.458	82,137
Firm-level agreement	Dummy	0.069	82,137
Works council	Dummy	0.340	82,137
Log number of employees	Continuous	3.685	82,137
Use of modern technology	Dummy	0.693	80,146
Proportion of skilled workers	Percent	67.355	82,118
Proportion of female workers	Percent	37.845	82,004
Foreign majority ownership	Dummy	0.072	80,715
Single establishment	Dummy	0.713	81,400
Establishment older than 10 years	Dummy	0.650	81,769
Regional dummies (16)	Dummy		82,137
Industry dummies (37)	Dummy		82,137

Notes: In coding the works council and collective agreement variables, we assumed that if the status in year *t-1*was the same as in year *t+1*, then the status in year *t* was unchanged. This assumption resulted in 0.5 and 3.3 percent of all works council and collective agreement observations, respectively, being recoded.

TABLE 2
Collective bargaining and works council coverage (in percent) by employment and by establishment [establishments with at least 5 employees, cross-section weighted data, 2000 and 2008]

		2000			2008		
		Germany	West	East	Germany	West	East
No collective agreement	Employment	34.3	31.5	48.4	42.8	40.6	54.4
	Establishment	48.9	44.4	67.2	60.8	58.1	71.9
Firm-level agreement	Employment	7.0	6.4	9.9	8.0	7.3	11.3
	Establishment	2.8	2.4	4.5	2.8	2.3	4.5
Sectoral agreement	Employment	58.7	62.1	41.7	49.2	52.1	34.3
	Establishment	48.3	53.2	28.3	36.4	39.5	23.6
Works council	Employment	46.9	48.6	38.7	42.6	44.2	35.0
	Establishment	11.4	11.4	11.4	8.6	8.8	7.9

TABLE 3
Within versus compositional change by type of agreement and by region, 2000 and 2008, weighted data

		Germany		West		East	
	(a) Collective agreements of any type	2000	2008	2000	2008	2000	2008
(1)	Observed coverage rate	51.2	39.2	55.7	41.9	33.1	28.1
(2)	Percentage point change, 2000-2008		-12.0		-13.8		-5.0
(3)	2008 (predicted) coverage based on 2000 coefficients		51.0		55.5		34.1
(4)	2000 (predicted) coverage based on 2008 coefficients	40.0		42.7		28.1	
(5)	Percentage point change due to changes in characteristics based on 2000 coefficients		-0.2 (1.5%)		-0.2 (1.3%)		1.1 (-21.9%)
(6)	Percentage point change due to changes in behavior based on 2000 coefficients		-11.9 (98.5%)		-13.7 (98.7%)		-6.1 (121.9%)
	(b) Sectoral agreements						
(1)	Observed coverage rate (%)	48.5	36.5	53.5	39.5	28.5	23.7
(2)	Percentage point change, 2000-2008		-12.1		-13.9		-4.8
(3)	2008 (predicted) coverage based on 2000 coefficients		47.9		52.8		29.0
(4)	2000 (predicted) coverage based on 2008 coefficients	37.5		40.7		23.8	
(5)	Percentage point change due to changes in characteristics based on 2000 coefficients		-0.7 (5.4%)		-0.7 (5.0%)		0.5 (-9.5%)
(6)	Percentage point change due to changes in behavior based on 2000 coefficients		-11.4 (94.6%)		-13.2 (95.0%)		-5.3 (109.5%)

Notes: The within effect is always statistically significant at the .01 level, other than for panel (c),while the between effect is never statistically significant. The between effect in row (5) is given by row (3) minus row (1) for 2000, and the within effect in row (6) is given by row (2) minus row (5).

TABLE 4
Coverage propensity by type of collective agreement, random-effects probit estimates, weighted data, 2000-2008

	Any collective agreement	Sectoral agreement	Firm-level agreement
Log number of employees	0.977 (0.021)***	0.667(0.021) ***	0.454 (0.022)***
Use of modern technology	0.019 (0.030)	0.033 (0.030)	-0.089 (0.039)**
Proportion of skilled workers	0.004 (0.001)***	0.002 (0.001)	0.004 (0.001)***
Proportion of female workers	-0.001 (0.001)	0.0006 (0.001)	-0.004 (0.001)***
Foreign majority ownership	0.155 (0.085)	0.071 (0.084)	0.062 (0.079)
Single establishment	-0.643 (0.045)***	-0.447 (0.045)***	-0.393 (0.048)***
Establishment older than 10 years	1.176 (0.068)***	1.288 (0.071)***	-0.067 (0.060)
Region dummies	yes	yes	yes
Industry dummies	yes	yes	yes
σ_u	3.714 (0.051)	3.897 (0.053)	2.577 (0.040)
ρ	0.932 (0.002)	0.938 (0.002)	0.869 (0.004)
Wald χ^2	7595.08	7557.03	938.16
Number of observations	80,958	80,958	80,958
Number of establishments	24,018	24,018	24,018

Notes: The model is given by equation (1) in the text. σ_u is the standard deviation of the unobserved effect u_i, and ρ is the latent intra-group (establishment) correlation. The model specification also contains 16 regional dummies, 37 two-digit industry dummies, and 8 year dummies. Standard errors are given in parentheses; ***, **, * denote statistical significance at the 0.01, 0.05, and 0.10 levels, respectively.

TABLE 5
Marginal and joint coverage probabilities and intra-class manifest correlation

	Percentiles				
	0.01	0.25	0.50	0.75	0.99
(a) Any collective agreement					
Marginal probability	0.047	0.340	0.526	0.715	0.965
Joint probability	0.032	0.286	0.467	0.665	0.954
Pearson's r	0.681	0.759	0.764	0.755	0.668
Odds ratio	152.118	59.430	56.248	62.534	182.264
(b) Sectoral agreements					
Marginal probability	0.033	0.260	0.444	0.622	0.921
Joint probability	0.022	0.214	0.387	0.569	0.900
Pearson's r	0.680	0.763	0.774	0.772	0.718
Odds ratio	211.653	71.535	62.705	64.312	124.701
(c) Firm-level agreements					
Marginal probability	0.009	0.032	0.056	0.094	0.280
Joint probability	0.004	0.018	0.033	0.060	0.211
Pearson's r	0.473	0.541	0.572	0.602	0.657
Odds ratio	186.059	82.865	60.012	45.389	28.659

Notes: The reported statistics are obtained using the command *xtrho* in Stata 10, and are described in Rodriguez and Elo (2003). In the case of panel (a), for example, the 95% confidence intervals for the median percentile are (0.527, 0.526), (0.467, 0.468), 0.758, 0.770), and (52.681, 59.098), respectively. See section VI for definitions.

TABLE 6
Beginning- and end-period collective agreement status of newly-founded establishments, 2000-2008, unweighted data

Year of birth	Collective agreement status in year of birth+1		Collective agreement status in year of exit		Year of exit (average)
	Status	*n*	*Anycb*=0	*Anycb*=1	
1999	*Anycb*=0	150	132	18	2002.6
	Anycb=1	124	19	105	2002.6
	Total	274	151	123	2002.6
2000	*Anycb*=0	138	129	9	2003.1
	Anycb=1	118	17	101	2003.4
	Total	256	146	110	2003.2
2001	*Anycb*=0	172	159	13	2004.3
	Anycb=1	112	11	101	2004.1
	Total	284	170	114	2004.2
2002	*Anycb*=0	68	64	4	2005.1
	Anycb=1	38	6	32	2005.9
	Total	106	70	36	2005.4
2003	*Anycb*=0	253	231	22	2006.0
	Anycb=1	198	20	178	2006.0
	Total	451	251	200	2006.0
2004	*Anycb*=0	203	185	18	2006.7
	Anycb=1	195	37	158	2006.7
	Total	398	222	176	2006.7
2005	*Anycb*=0	241	230	11	2007.3
	Anycb=1	178	17	161	2007.3
	Total	419	247	172	2007.3
2006	*Anycb*=0	290	268	22	2007.7
	Anycb=1	201	22	179	2007.7
	Total	491	290	201	2007.7
2007	*Anycb*=0	278			
	Anycb=1	226			
	Total	504			

Notes: A newly-founded establishment in the 2000 (2001, …, 2008) survey is a unit born in 1999 (2000, …, 2007). Consequently, all 2008 births (i.e. establishments born in 2008) are discarded in our subsequent survival analysis. Also note that all establishments born in, say, 2002 but not observed (surveyed) before 2006, for example, are dropped from the sample. In other words, only those establishments that can be followed from the outset (year of birth) are included in the estimation sample. Exit means rotation out of the panel or failure (end of the initial state). *Anycb* is a dummy variable signifying the presence of any type of agreement.

TABLE 7
Cox proportional hazard model estimates, newly-founded establishments, 2000-2008, unweighted data

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.348 (0.068)***	0.349 (0.092)***
Use of modern technology	-0.500 (0.157)***	0.011 (0.203)
Proportion of skilled workers	-0.004 (0.003)	0.002 (0.004)
Proportion of female workers	-0.007 (0.004)*	0.001 (0.004)
Foreign majority ownership	0.460 (0.273)*	-0.490 (0.449)
Single establishment	0.604 (0.215)***	-0.032 (0.245)
Number of observations	1,787	2,362
Number of establishments	787	1,003
Number of failures	145	117
Wald χ^2	81.47	73.91
Predicted median duration	1.81	2.61

Notes: The hazard function is given by equation (2). The model includes 7 industry dummies and 1 region (western Germany). Clustered standard errors are given in parentheses. The Wald test rejects the null of no joint statistical significance of the model. The (predicted) median duration in the last row of the table is obtained using a PH exponential model without covariates.

TABLE 8
Cox proportional hazard model estimates, restricted sample of permanent establishments, 2004-2008, unweighted data

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.241 (0.074)***	0.050 (0.237)
Use of modern technology	0.150 (0.232)	0.447 (0.447)
Proportion of skilled workers	-0.002 (0.005)	0.007 (0.012)
Proportion of female workers	-0.001 (0.005)	-0.001 (0.009)
Foreign majority ownership	-0.788 (0.598)	0.434 (0.855)
Single establishment	1.002 (0.303)***	-0.431 (0.486)
Establishment age	0.072 (0.280)	-0.694 (0.385)*
Number of observations	3,928	3,051
Number of establishments	821	627
Number of failures	93	35
Wald χ^2	76.89	8,783.72

Note: See notes to Table 7.

TABLE 9
Cox proportional hazard model estimates, restricted sample of permanent establishments, 2004-2008, unweighted data (*counterfactual*)

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.224 (0.074)***	0.055 (0.236)
Use of modern technology	0.175 (0.234)	0.461 (0.445)
Proportion of skilled workers	-0.002 (0.005)	0.008 (0.012)
Proportion of female workers	-0.009 (0.005)*	-0.0001 (0.009)
Foreign majority ownership	-0.810 (0.597)	0.492 (0.853)
Single establishment	0.976 (0.305)***	-0.384 (0.478)
Establishment age	0.123 (0.287)	-0.763 (0.386)**
Number of observations	4,163	3,551
Number of establishments	868	727
Number of failures	93	35
Wald χ^2	75.45	28.22 (0.0133)

Notes: See notes to Table 7.

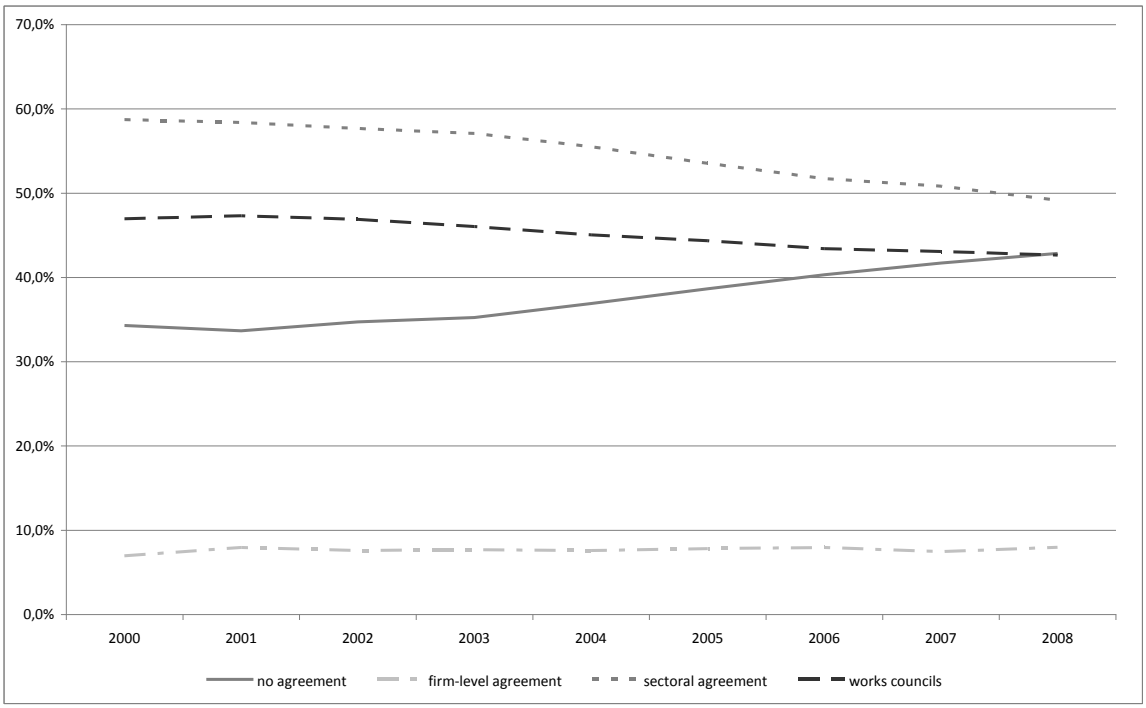
TABLE 10
Cox proportional hazard model estimates, sample of permanent establishments, 2000-2008, unweighted data

	Leaving any type of collective agreement	Joining any type of collective agreement
Log number of employees	-0.367 (0.054)***	0.224 (0.122)**
Use of modern technology	0.245 (0.165)	0.193 (0.272)
Proportion of skilled workers	-0.006 (0.003)**	0.0004 (0.005)
Proportion of female workers	-0.005 (0.003)	-0.011 (0.006)**
Foreign majority ownership	-0.630 (0.422)	0.623 (0.462)
Single establishment	0.648 (0.198)***	-0.337 (0.298)
Establishment age	-0.212 (0.169)	0.074 (0.247)
Number of observations	7,486	5,697
Number of establishments	922	675
Number of failures	193	82
Wald χ^2	147.56	31.45

Note: See notes to Table 7.

FIGURE 1
Collective bargaining and works council coverage, 2000-2008
(establishments with at least 5 employees; cross-section weighted data)

(a) Coverage by employment



(b) Coverage by establishment

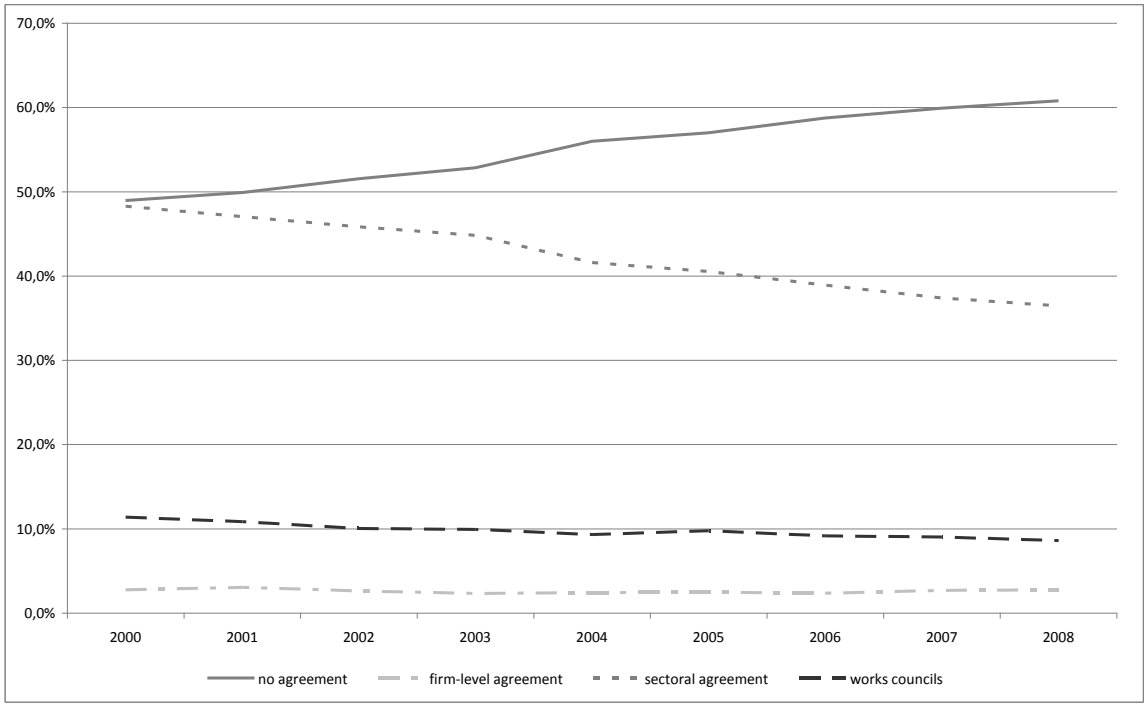


FIGURE 2
Collective bargaining and works council coverage by employment, 2000-2008, cross-section weighted data

(a) With less than 250 employees (b) With at least 250 employees

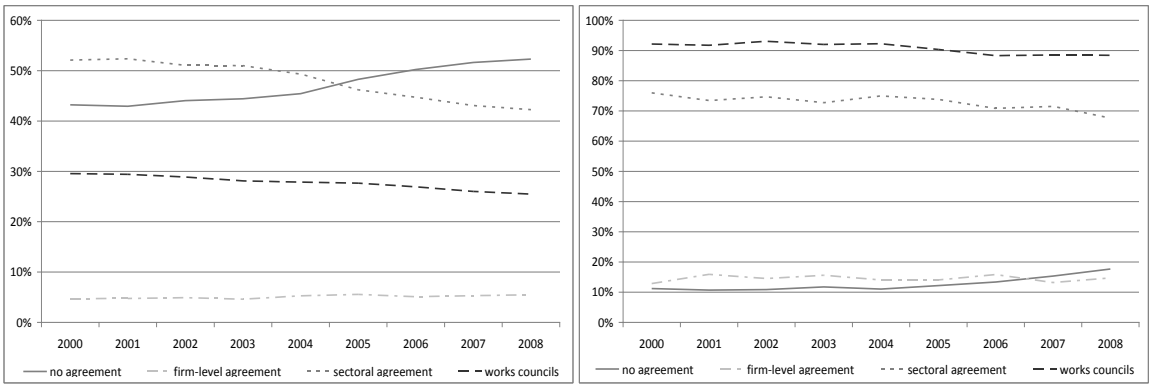


FIGURE 3
Collective bargaining and works council coverage by establishment, 2000-2008, cross-section weighted data

(a) With less than 250 employees (b) With at least 250 employees

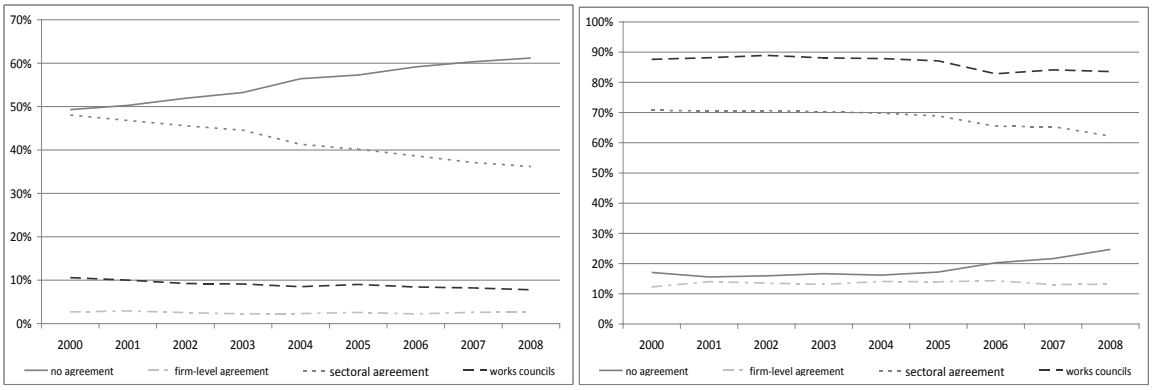


FIGURE 4
Collective bargaining and works council coverage by employment, 2000-2008, cross-section weighted data

(a) Service sector (b) Manufacturing

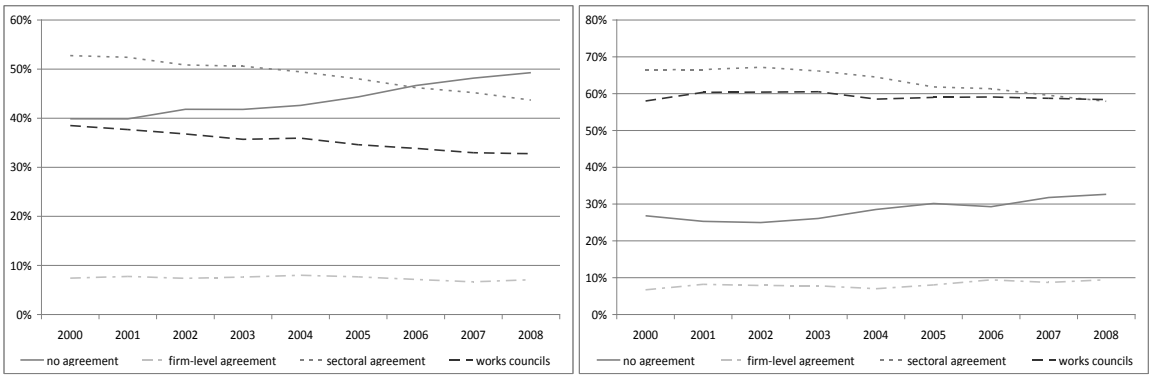


FIGURE 5
Collective bargaining and works council coverage by establishment, 2000-2008, cross-section weighted data

(a) Service sector (b) Manufacturing

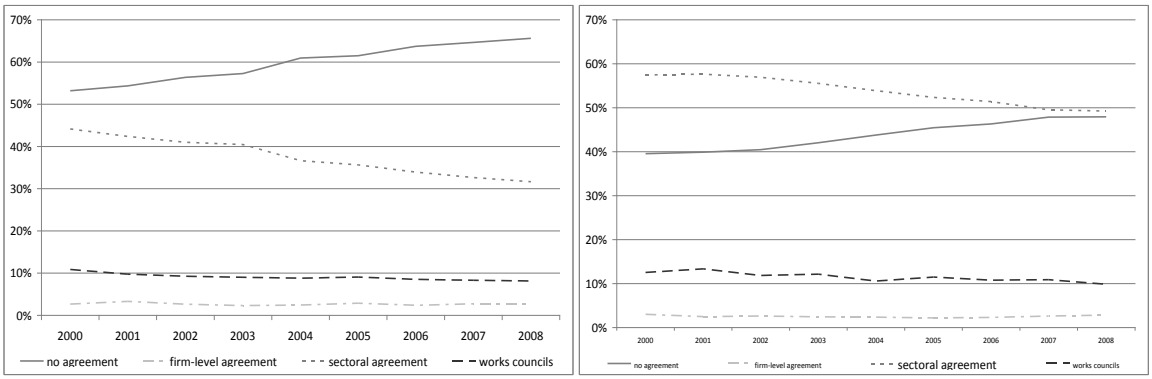
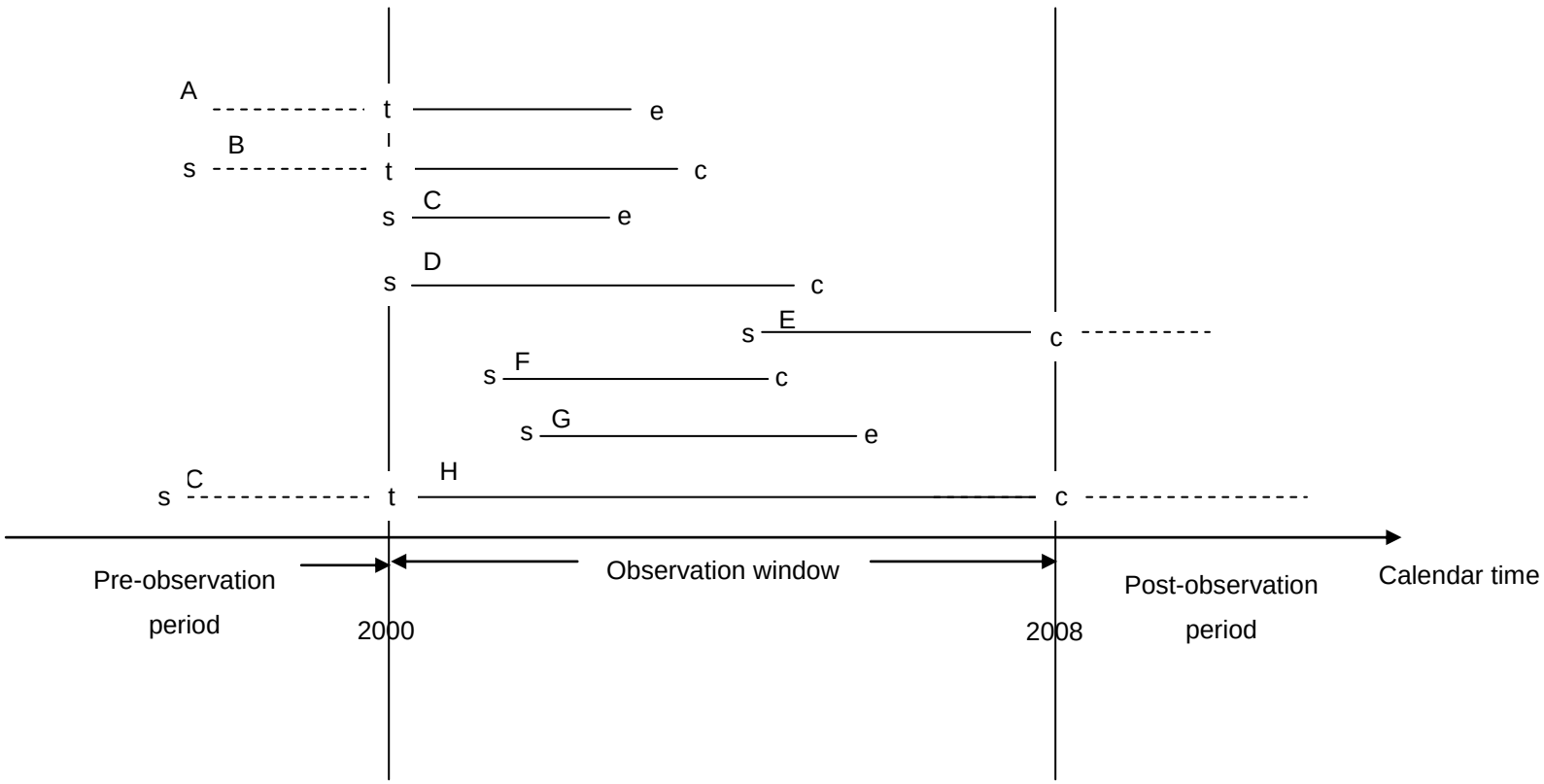


FIGURE 6

Schematic of the observation window and censoring



Legend: t – left-truncation point
c – right-censoring point
s – starting time of the event (or entry to a state)
e – ending time of the event (or exit from a state)