BARGAINING ABOUT WAGES:
EVIDENCE FROM SPAIN

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IV. Economic and econometric framework

Bargaining procedures in Spain are quite different from those in the US or Canada. Normally, employees and the employer have an indefinite contract. Current working and pay conditions are settled in an additional protocol called "convenio" which usually covers a number of years. However, wage increases are negotiated or renegotiated almost yearly. In this sense we will assume that bargaining over wage increases takes place yearly.

The negotiation process starts when the union decides to make a wage increase claim (CLAIM). The institutional setting is such that the firm must respond to the union with an initial counteroffer (OFFER). If this OFFER equals the above CLAIM, there is an immediate agreement. If not, they alternate offers until they reach an agreement[140]. In the meantime the union uses a latent strike threat.

There are several institutional features which condition the strike threat during negotiation. First, it is unusual to strike before CLAIM and OFFER have been announced. Second, the workers must compulsorily announce in advance to the firm the starting date of the strike. Moreover, in some cases they announce the duration of the strike or, in any case, they make an indefinite threat. Third, it is forbidden for the firm to hire temporary replacement workers. And finally fourth, workers must compulsorily guarantee a minimum service level in some key industries (especially in the transportation and the utilities sectors), which severely limits the efectiveness of a strike. Undoubtedly, these institutional features will

[140] See chapter 2 for a description of the main characteristics of the initial CLAIM and OFFER in Spain during the 1985-1990 period.

affect greatly what is observed. For example, the announcement could lower strike activity because the firm could react to a formal threat avoiding a costly work stoppage. Unfortunately, despite the fact that Spanish strike statistics provide information about these institutional features, our sample does not include information about them.

Despite other possibilities, we concentrate here on analyzing the implications of standard asymmetric information theories on wage increase, strike decision and duration joint determination process. The determination of the initial CLAIM and OFFER will be considered separately because they are used to proxy the level of uncertainty about the change in the value of labour for the employer.

Recent development of OSAI models postulate that strike (or holdout) duration is determined by the delay required to credibly establish that the employer's demand price is no higher than the truth. Then, the wage settlement splits the difference between the demand and the supply prices. Card (1990b) and Cramton and Tracy (1992) provide complementary frameworks to signalling models in a non-dynamic context. Whilst the first concentrates in the relationship between strikes and wages, the second considers also the incidence of delaying the agreement. Nonetheless, both papers agree with the general idea that longer strikes should produce lower observed wages. It means that there is a negative sloped concession curve between wage and strike duration. However, whilst in Card's paper there is not any perceptible difference between the wage with and without strike apart from the effect of the work stoppage duration, in Cramton and Tracy both wages may be different and, particularly, the strike wage could be greater.

It should be convenient to formalize a general setup for the subsequent analysis in the spirit of Card's model, but slightly more general than his own econometric specification. As it is implicit in this paper, the relevant wage variable, wage increase (Δw) in our case, strike threat (s^*) and strike duration threat (d^*) are determined in a joint maximization process (without taking into account any sort of selection, for the moment). Consider the solution to that problem is well represented by the following general structure:

$$[5.1] \qquad \Delta w = f(s^*, d^*, z_w, \gamma_w) + e_w$$

$$[5.2] \qquad s^* = g(d^*, \Delta w, z_s, \gamma_s) + e_s$$

$$[5.3] \qquad d^* = h(s^*, \Delta w, z_d, \gamma_d) + e_d$$

$$[5.4] \qquad \begin{cases} d = d^*;\ s = 1 & \text{iff } s^* > 0 \text{ and } d^* > 0 \\ d = 0\ ;\ s = 0 & \text{otherwise} \end{cases}$$

where z_w, z_s and z_d are the relevant set of variables for the wage, strike decision and strike duration equations, respectively. Among other variables, these vectors include proxies for the change in firm's profitability (the change in sales per employee and the level of profits per employee) and also a proxy for the level of bargaining uncertainty (the difference between the initial claim and offer). γ_w, γ_s and γ_d are the relevant parameters and e_w, e_s and e_d are the error terms. Finally [5.4] specifies the most general observability rule of the strike indicator (s) and the strike duration (d).

In this context, three issues will be carefully examined: endogeneity, dynamics and self-selection. The endogeneity of work stoppage variables is a direct consequence of the joint determination. As far as errors in [5.1]-[5.3] are presumably cross-correlated, strike variables cannot be considered exogenous to the wage determination process. Consequently, an IV method is required. As regard dynamics, we could consider several sources for their introduction into the model. First of all, learning or reputation may influence the current outcome of the negotiation process. In particular, reputation may explain why the threat instead of the strike outcome is important in the model. Second, because of the adjustment costs of some variables (for instance, employment or investment). Third, a single negotiation may be embedded in an indefinite bargaining game. Consequently, there is no reason to expect that current negotiation can be isolated from past (or future) negotiation rounds. A third issue to note is that there could be some sort of selection affecting the wage determination process. If so, we need to specify different wage equations for the two regimes.

By restricting or relaxing assumptions, the vast majority of the analyses may be embodied in the above general framework. Herrington (1988), McConnell (1989) and Card (1990b) considered that the error term in [5.1] is independently distributed from the errors in [5.2] and [5.3][141]. Moreover, they substitute the threat variables (s^*,d^*) by realized outcomes (s,d). A linear version of this structure may be expressed as follows:

$$
[A] \qquad \begin{cases} [5.1A] & \Delta w_{it} = \delta^A_{11} s_{it} + \delta^A_{12} d_{it} + z_{wit}\gamma^A_w + e^A_{wit} \\ [5.2] & s^*_{it} = \delta_{21}\Delta w_{it} + \delta_{22} d_{it} + z_{sit}\gamma_s + e_{sit} \\ [5.3] & d^*_{it} = \delta_{31} s_{it} + \delta_{32}\Delta w_{it} + z_{dit}\gamma_d + e_{dit} \end{cases}
$$

where the wage increase settlement is a negatively sloped function of the

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[141]Observationally, this is equivalent to the assumption that strike outcomes are random realizations. This is precisely the assumption imposed in earlier studies for Canada as Coisneau and Lacroix (1977), Auld et al. (1979, 1981) or Ridell (1980).

strike duration, after controlling for other variables which could include some dynamic terms and/or bargaining unit specific effects (z_w). In such a framework, $\Delta w(d=0)$ maybe be understood as the maximum wage increase available for workers. Notice that under Card's formulation the wage without strike is a corner solution of the wage with strike. In such context the probability of observing a strike ($s^*>0$) is fully characterized by the probability that the employer's demand price for labour is lower than a given threshold (i.e, the corner solution for d). Consequently, the observed duration of a work stoppage fully catch the differences between the wage settlement with and without strike. A test against this hypothesis would be a test against Card's view of the joint model. Notice also that even in the case which [A] holds, Card's treatment of the wage equation is only valid if s and d are exogenous to w, which can be checked with a Hausman test. We will return to this key point later.

In such a framework, Herrington (1988) and Card (1990b) found no evidence of any systematic relationship between wage outcomes and strike decision and duration, respectively, whereas McConnell (1989) found some evidence to support such a relationship[142]. In our opinion the assumption about the independence of the errors in the system may cause a wrong perception of findings. If all three outcomes are the result of a unique maximization problem, the errors in [5.1]-[5.3] are probably positively correlated. As a result, a positive bias is expected to arise in the relevant structural coefficients when estimating [5.1] separately from [5.2]-[5.3] or without instrumenting the strike variables.

[142] Two previous studies for Canada, Auld et al (1979) and Lacroix (1986) had found evidence in favour of a positive relationship.

An alternative to the above setup (without considering any kind of sample selection) is to specify the strike variables in terms of threats rather than of outcomes:

$$
[B]\quad \begin{cases}
[5.1B] & \Delta w_{it} = \delta^B_{11} s^*_{it} + \delta^B_{12} d^*_{it} + z_{wit}\gamma^B_w + e^B_{wit} \\
[5.2] & s^*_{it} = \delta_{21}\Delta w_{it} + \delta_{22} d^*_{it} + z_{sit}\gamma_s + e_{sit} \\
[5.3] & d^*_{it} = \delta_{31} s^*_{it} + \delta_{32}\Delta w_{it} + z_{dit}\gamma_d + e_{dit}
\end{cases}
$$

On the other hand, Cramton and Tracy's model with multiple threats suggests that each threat will lead to a sensibly different wage equation, which means that the strike mechanism produces some sort of selection on it. This idea was first considered by Stengos and Swidinsky (1990). The structural system under observed outcomes may be stated as follows:

$$
[C]\quad \begin{cases}
[5.1CA] & \Delta w_{ita} = z_{wit}\gamma^C_{wa} + e^C_{wita} & \text{if } s_{it}=0 \\
[5.1CB] & \Delta w_{itb} = \delta^C_{12} d_{it} + z_{wit}\gamma^C_{wa} + e^C_{wita} & \text{if } s_{it}=1 \\
[5.2] & s^*_{it} = \delta_{21}\Delta w_{it} + \delta_{22} d_{it} + z_{sit}\gamma_s + e_{sit} \\
[5.3] & d^*_{it} = \delta_{31} s_{it} + \delta_{32}\Delta w_{it} + z_{dit}\gamma_d + e_{dit}
\end{cases}
$$

where the selection rule is given by equation [5.4].

This selection mechanism determines whether we observe the wage with strike (w_{ta}) or without it (w_{tb}). This last one should be understood as the outcome under the alternative threat[143]. Notice also that the effect of the strike variable is captured by the constant.

It is far beyond the scope of this chapter to attempt a joint estimation of any of the systems [A]-[C], because their inherent complexity makes the task practically unaffordable. Instead of this, we will consider

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[143] For which we are still assuming that there are no selection problems.

limited information methods, as Card and Stengos and Swidinsky did. However, in the empirical application we will take into account the implications of the joint determination model.

V. Econometric specification.

The main purpose of the analysis we are going to carry out is to be able to elucidate if it is possible some simplification of the general model considered in the previous sections. In this sense, we are interested in four different (but related) issues. First, we would like to clarify whether or not the dispute outcomes should be considered endogenous to the negotiation process. Secondly, we test the importance of the unobserved heterogeneous effects in the relevant equations to estimate. Third, we evaluate the differences between the strike and non-strike wage equations (under the null of exogeneity of the strike outcome). Finally, we try to detect the presence of some sort of sample selection. By failing to reject some of the maintained hypothesis, the difficulty of the model proposed is reduced and we could estimate it with simple procedures. On the contrary, we need to deal with the whole problem.

a. Strike decision

Consider the reduced form of the equation [5.2], which describes the latent strike decision:

$$[5.5] \qquad s^*_{it} = \alpha_s' X_{sit} + f^s_i + v_{sit} \qquad t = 1, .., T_i;\ i = 1, .., N$$

$$s_{it} = 1(s^*_{it} > 0)$$

being 1(A) an indicator function of the event A, which takes the value one whenever A holds and zero otherwise. X_s is a vector of variables influencing

s^*, α_s the corresponding parameter vector, f_i^s a bargaining unit specific effect and v_s an error term.

Standard qualitative dependent variable models (PROBIT or LOGIT) applied to equation [5.5] provide consistent estimates in the context of cross-sections under some assumptions about the error term. Using panel data, we have to impose in addition to homoskedasticity, that the error terms are serially uncorrelated across individuals and that the unobserved specific effects are negligible (or constant among individuals), for the estimators to be consistent and asymptotically efficient. Consequently, testing will be crucial to ensure the consistency of the results.

If we assume a logistic distribution for the errors in [5.5], we have the usual Logit model which in this case includes unobservable individual effects. We could write the probability for each observation in each time period and form the likelihood function. However, standard maximum likelihood estimation of the relevant parameters will produce inconsistent estimates because of the incidental parameter problem noticed by Neyman and Scott (1948). But, as Chamberlain (1984) pointed out, we can make use of the fact that $\Sigma_{t=1}^{T} s_{it}$ is a sufficient statistic for f_i^s in order to obtain a conditional likelihood which does not depend upon f_i^s.

We have to express the probability of a given sequence of outcomes for each individual $s_i=(s_{i1},s_{i2},..,s_{it})$, conditional on the statistic above, to write the conditional likelihood function and estimate the relevant vector of parameters α_s'. In doing so, we need to account for the fact that we have available an unbalanced panel of bargaining units, although there are no additional difficulties for applying this method to it. After rearranging terms, the corresponding likelihood can be expressed as follows:

$$[5.6] \qquad \log L = \sum_{i=1}^{N} \ln \frac{\exp(\alpha_s \Sigma_{t_i} X_{sit} s_{it})}{\Sigma_{d \in B_i} \exp(\alpha_s \Sigma_{t_i} X_{sit} d_t)}$$

where:

$$B_i = \{d=(d_1,\ldots,d_{T_i}) \;/\; d_t=0 \text{ or } 1 \text{ and } \Sigma_{t=1}^{T_i} d_{it} = \Sigma_{t=1}^{T_i} s_{it}\}.$$

Given the conditional estimates, say α_{sCL}, and the pooled estimates α_{sP}, it is possible to test the null that the fixed effects are negligible, for which we could conduct a Wald test, which is distributed as a χ^2 with k degrees of freedom, k being the number of parameters. Unfortunately, the conditional Logit approach presents several inconveniences. On the one hand, only the subset of units for which there are variation in the dependent variable contribute to the likelihood and, hence, to the estimation of the parameters of interest. This reduces significantly the sample and could cause some identification problems. On the other hand, neither lagged outcome (s_{it-1}) can be included as a regressor in X_{sit}, nor we can account for a random specification in which the individual effects and the explanatory variables are permitted to be correlated (although this is not an important issue given we control for the effects with the conditional procedure). Finally, the estimates are no longer consistent if the error term is not logistic, is not homoskedastic or is serially correlated.

Although we could use the parameter estimates as an intuitive test of the adequacy of the hypothesis, the deficiencies associated to the conditional logit specification make us think in alternative procedures. We also estimate a Linear Probability model (LPM) which has been previously used by Card (1990b) in this same context. Although this method has also several well-known shortcomings, it allows us to consider the lagged outcome as an additional explanatory factor and instrument the equation to obtain consistent. Moreover, we could control for specific effects, whatever fixed or random, by taking first differences in [5.5]:

$$[5.7] \qquad \Delta s^*_{it} = \gamma s_{it-1} + \Delta X_{sit}\alpha_s + \Delta v_{dit}$$

We consistently estimate [5.7] using a GMM-IV estimation method proposed by Arellano and Bond (1991), provided N is large and T fixed. The LPM specification also permits us to test for the importance of the firm specific effects. We use a Sargan-difference test proposed by Holtz-Eakin (1988) for autoregressive models and extended by Arellano (1993) allowing for the presence of exogenous regressors. The test is specially useful when the vector X_{sit} includes the lagged outcome (s_{it-1}) as above. Notice that, by construction, at least lagged outcomes are correlated with the effects and, consequently, the model in levels does not allow us to obtain consistent estimates of the relevant vector of parameters. However, under the null that the specific effects are irrelevant, both the model in levels [5.4] and the differenced equation [5.7] provide consistent estimates. This test accounts for the lack of orthogonality between the errors in levels and the lagged outcomes by means of a comparison of the Sargan orthogonality test under the null that both the model in levels and in differences are consistent, $\hat{S}_o$, and under the alternative that only the model in first differences provides consistent estimates, $\hat{S}_a$. This difference is distributed as a χ^2 with r degrees of freedom, where r is the number of additional orthogonality restrictions implied by the model in levels.

b. Strike duration

Consider the reduced form of the equation [5.3], which describes the latent strike duration and the corresponding observability rule:

$$[5.8] \qquad d^*_{it} = X_{dit}\alpha_d + f^d_i + v_{dit} \qquad t = 1, .., T_i;\ i = 1, .., N$$

$$d_{it} = d^*_{it}1(d^*_{it}>0)$$

where 1(.) is defined as in [5.5]. X_d is a vector of variables influencing d^*, α_d is the corresponding parameter vector, f^d_i a negotiation unit specific effect and, finally, v_d is the reduced form error term.

As the work stoppage duration and the strike decision could be closely related, we pose two joint determination models: the well-known Tobit and the selectivity model. The Tobit model assumes that the same set of parameters characterize the decision and the length of a strike and makes use of the whole sample (i.e. including zeros). Alternatively, the selectivity model assumes that there is always a length threat (d^*_{it}) which only is observed if a strike is produced ($s^*_{it}>0$). In both cases, we are assuming the firm specific effects are irrelevant and the errors in [5.8] are normally distributed. The Tobit model will be estimated by maximum likelihood. In order to estimate the selectivity model we apply a two stage method due to Heckman (1976) in the subsample of positive observations.

For comparative purposes we will present, an estimate of an heterogeneous Weibull hazard for positive durations. To characterize the distribution of the duration we start with the conditional survival function $S(d_{it}/v)$:

$$S(d_{it}/v) = \exp(-\lambda_{it} * d_{it})^{1/\sigma} * v$$

where $\lambda_{it} = \exp(-X_{dit}\alpha_d)$, σ is the standard deviation of the duration and the random variable v is the heterogeneity effect, distributed as a gamma with parameter $1/\theta$. Integrating $S(d_{it}/v)$ with respect to v we obtain:

$$S(d_{it}) = [1 + \theta(\lambda_{it} * d_{it})^{1/\sigma}]^{-(1/\theta)}$$

finally, the estimable hazard function $H(d_{it})$ may be expressed as:

$$[5.9] \qquad H(d_{it}) = S(d_{it})^{\theta} . \frac{\lambda_{it}}{\sigma} . (\lambda_{it} d_{it})^{(1-\sigma)/\sigma}$$

Note that this approach will produce consistent estimates if strikes are randomly observed and their duration is exponentially distributed.

c. Initial claim and offer

The Spanish bargaining framework is such that bargaining starts at the time the union makes its initial *claim* (Δw^c). After it, the firm decides either to accept such a claim or counteroffering, which will be formally called initial *offer* (Δw^o). In recent signalling models (Cramton and Tracy (1991, 1992), the claim is assumed to be a function of what union expected a unit of labour is worth for the firm ($\Delta \hat{q}_{it}$). Consequently, a linear equation for the initial claim can be written as:

$$[5.10] \qquad \Delta w^c_{it} = \alpha \Delta \hat{q}_{it} + \beta_1 X^c_{it} + f^c_i + u^c_{it}$$

$$t = 1, .., T_i; \; i = 1, .., N$$

where X^c_i considers variables expected to affect the union claim, as the union's strength, the alternative wage or the wage structure (in terms of the base wage weight and bonuses), f_i^c is a BU specific component and u_i^c is a serially uncorrelated error term. The observed initial firm's wage increase offer (Δw_{it}^o) is made taking into account the above claim[144], knowing the true value of Δq_{it}:

$$[5.11] \qquad \Delta w_{it}^o = \gamma_1 \Delta q_{it} + \gamma_2 \Delta w_{it}^c + \beta_2 X_{it}^o + f_i^o + u_{it}^o$$
$$t = 1,..,T_i; \; i=1,..,N$$

where X_i^o takes into account variables expected to affect the employer offer, f_i^o is a firm specific component and u_i^o is the corresponding serially uncorrelated error term. It is important to note that the institutional setting is such that firm must counteroffer immediately. This will have strong consequences on the amount of information the offer reveals. Consequently, in the light of a recent model by Cramton and Tracy (1991,1992) it will be surprising to show that the initial firm's offer is a Rubinstein offer.

Estimation of the joint model [5.10]-[5.11] depends crucially on the assumptions about the structure of the error processes. Least squares on any of both equations will produce inconsistent estimates as long as there are

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[144]The initial offer is linked to an underline initial offer (Δw_{it}^{o*}) by the following selection rule:

$$\Delta w_{it}^o = \Delta w_{it}^{o*} \quad \text{if } \Delta w_{it}^{o*} \leq \Delta w_{it}^c$$
$$\Delta w_{it}^o = \Delta w_{it}^c \quad \text{if } \Delta w_{it}^{o*} \geq \Delta w_{it}^c$$

Consequently, Δw_{it}^o is truncated from above by the union's initial claim. As far as in the sample this kind of truncation is rather small (less than 4% of the contracts) we shall proceed like if there is no truncation. Then, we will assume that always $\Delta w_{it}^o = \Delta w_{it}^{o*}$. Relaxing such hypothesis leads to a complicated simultaneous equation alternative framework. The offer equation must be estimated using symmetrically trimmed least squares, for instance, (Powell (1986)). In our opinion, the set of assumptions to be made in order to implement such a method does not compensate the implicit gain given by its consistency.

variables potentially correlated with the error term or the specific effect. The errors may also be cross correlated not because of a joint maximization problem but because of their relationship to a common unexpected firm specific demand shock (ξ_{it}). We solve these problems by using an IV estimator over the first differences equations of the system:

$$[5.12] \qquad \Delta w^{c}_{it} = \alpha \Delta \hat{q}_{it} + \beta_1 \Delta X^{c}_{it} + \Delta u^{c}_{it}$$

$$[5.13] \qquad \Delta w^{o*}_{it} = \gamma_1 \Delta q_{it} + \gamma_2 \Delta w^{c}_{it} + \beta_2 \Delta X^{o}_{it} + \Delta u^{o}_{it}$$

Under the assumption that the errors in [5.11]-[5.12] are serially uncorrelated, all the variables dated t-2 and earlier are valid instruments to obtain consistent estimates of the parameters of the system. The first differences form also controls for the existence of unobserved effects which are potentially correlated with the explanatory variables. Note that the equation by equation estimates are consistent but not efficient although to get efficient estimates we consider the joint error structure in a panel data context (see chapter 3 for a brief comment).

However, given the fact that we have a wage increases model in levels we must consider the possibility that the specific effects are not significantly different. In such circumstance, all variables dated t-1 and earlier are valid instruments. As in the case of the LPM model, under the null that specific effects are irrelevant, both the system in levels [5.10]-[5.11] and in first differences [5.12]-[5.13] provide consistent estimates. A Sargan differences test is applicable to discriminate between them. This diagnostic is distributed as a χ^2 with r degrees of freedom -r is the number of additional orthogonality restrictions implied by the model in levels.

d. The wage increase equation.

We have described in the previous section several alternatives for analyzing the wage increase equation. As far as the treatment will be rather similar we present the estimation methods using the structural equation [5.1] in framework [A], which is in the spirit of our simplest model:

$$[5.14A] \quad \Delta w_{it} = \delta^{A}_{11} s_{it} + \delta^{A}_{12} d_{it} + z_{wit}\gamma^{A}_{w} + f^{w}_{i} + u^{A}_{wit}$$

$$t = 1,..,T_i;\ i = 1,..,N$$

We decompose the error term in a firm specific effect, f^{w}_{i} and a mixed component u^{A}_{wit}. Note that the vector of explanatory variables, z_{w}, could include some variables potentially correlated with the error term and/or the lagged outcome. Consequently, as far as the structure of equation [5.14A] is similar to the structure of equation [5.10] or [5.11], the estimation procedure is exactly the same. Therefore, consistent estimates may be attached by applying GMMIV to the first differenced of equation [5.14].

Under the assumption that error term in [5.14] is serially uncorrelated, all variables dated t-2 and earlier should be considered valid instruments. Note that the Sargan differenced test is also valid for assessing the necessity of taking first differences in the system. Before going to the testing procedure let us make some remarks about the form of dealing with the wage increases equation in framework [C]. First, note that:

$$[5.15A] \quad E(\Delta w_{ita}/s_{it}=0) = E(\Delta w_{ita}/s^{*}_{it}\leq 0) = z_{wit}\gamma^{C}_{wa} + E(e^{C}_{wita}/s^{*}_{it}\leq 0)$$

$$[5.15B] \quad E(\Delta w_{itb}/s_{it}=1) = E(\Delta w_{itb}/s^{*}_{it}\geq 0) = z_{wit}\gamma^{C}_{wb} + E(e^{C}_{witb}/s^{*}_{it}\geq 0)$$

where in general neither $E(e^{C}_{wita}/s^{*}_{it}\leq 0)$ nor $E(e^{C}_{witb}/s^{*}_{it}\geq 0)$ are expected to be zero. Under the assumption that the errors in the wage increase equations and the error in the underline strike decision equation are jointly normal, these expected values are given by the following expressions:

$$[5.16A] \quad E(e^{C}_{wita}/s^{*}_{it}\leq 0) = -\sigma_{w_a s} * \frac{\phi(\mu^{s}_{it})}{\Phi(\mu^{s}_{it})} = -\sigma_{w_a s}\lambda_{w\bar{s}}$$

$$[5.16B] \quad E(e^{C}_{witb}/s^{*}_{it}\geq 0) = \sigma_{w_b s} * \frac{\phi(\mu^{s}_{it})}{1-\Phi(\mu^{s}_{it})} = \sigma_{w_b s}\lambda_{ws}$$

where $\sigma_{w_a s}$ and $\sigma_{w_b s}$ are the covariances of the error in the strike equation [5.2] and, respectively, the wage without and with a strike, μ^{s}_{it} is the baseline of the strike equation and, finally, $\lambda_{w\bar{s}}$ and λ_{ws} are the well-known Mills' inverse ratio of, respectively, the non-strike and the strike regimes. Consequently, the equivalent to equation [5.14A] in framework [C], say [5.14Ca] and [5.14Cb] does not have well-behaved errors. However, this problem can be easily solved by adding to these equations a consistent estimate of the $\lambda_{\bar{s}}$ and λ_{s}, say $\hat{\lambda}_{\bar{s}}$ and $\hat{\lambda}_{s}$. However, note that no longer the standard errors resulting after estimation are valid. Hence, they should be considered with some caution.

Our main purpose with the testing procedure is to confront the simplest framework ([A]-[C]) for analyzing wage increase setting. We would like to answer the questions mentioned in the previous section which constitute a criticism on most of the work in this field, because it is normally assumed that strike outcome is the relevant work stoppage variable and it is also exogenous. In an initial stage, we analyze the endogeneity issue by means of a comparison of the estimates of the wage equation for the whole sample with and without instrumenting the dispute variables. In a second stage, we worry about the relevance of the work stoppage threat variable (s^*) in the wage equation. In both stages we use a Sargan differenced test as proposed by Arellano and Bond (1991). In a third stage, we compare by means of a Wald test the estimates of the wage equation (excluding the intercept and strike variables) in the full sample to the estimates using the non-strike sample to account for the possibility of different structural parameters. Under the null of exogeneity of the strike outcome, this Wald test provides invaluable information about the significance of such differences.

The above approaches provide consistent estimates in absence of selection problems in the strike process. Under the presence of selection, we need first to estimate the decision to strike and then correct the wage equation with the corresponding inverse Mill's ratio. Before doing so, we perform a variable addition test for selection bias appropriate to panel data, in a dynamic context, following the approach by Wooldridge (1994). The procedure may be stated as follows. First, we estimate T decision equations (in fact T-1) using standard discrete choice models. Then, we compute the inverse Mill's ratio for each observation in each time period. In a second stage, we estimate by IV the equation [5.14Ca] adding the selection term for those observations for which $s_{it} = 1$. Finally, we test the null that the effect of the selection term is zero. We use a first differences method because some of the explanatory variables are predetermined or endogenous as opposed to the method used by Wooldridge (1994). However, given we are interested in the wage equation for both strike regimes and the small sample size the strike bargaining units, we conduct this test using not only the subsample where $s^*_{it} = 1$ but that in which $s^*_{it} = 0$.

VI. Empirical results.

a. Data and variables

As mentioned in section I, the basic data source we use in this chapter is the NGCE, an annual survey carried out by the Spanish Ministry of Economy, recording information about collective bargaining of firms with more than two hundreds workers as well as other questions about firm performance and pay structure. The available sample covers a time span of 6 years, from 1985 to 1990. From the original data set of four thousand records we select those observations which contain information about claim, offer, agreement and the starting and the ending date of the negotiation process. It is necessary to select these data in order to be able to use some controls concerning price expectations and wage signal, which we consider as key variables in our framework. The Data Appendix provides a detailed explanation of the main characteristics of the data set, the definition of the variables and some descriptive statistics.

We are going to summarize the variables we use in the empirical application and their expected effects. As Card (1990b) shows those variables with a positive effect on wages via their effect on the profitability of the firm should decrease the probability and conditional duration of strikes. On the contrary, variables which have a positive effect on wages via their effect on alternative wage opportunities available to workers and that increase the dispersion of the unobservable components of profitability should increase both the probability and duration of work stoppages. There is no complete agreement about which are the variables that pick up all these effects and, as a result, our reduced form specification will consider several alternatives to capture them.

We begin with those which affect the reduced forms of strike decision and duration. We proxy the change in productivity by means of the lagged change in real sales per employee ({ΔSALES-ΔP}(-1)). As proxies for the level of profits of the firm, we use the lagged value of real profits per employee (B/P(-1)) as well as the lagged proportion of hiring to employment (HIRING(-1)) and a dummy taking the value of one if past profits are positive (DB(-1)). All these controls are expected to affect negatively dispute rates. We are using lagged variables not only to avoid simultaneity problems but because we think that the position of the firm more than the expected results is important to the union (at least in the context of asymmetric information). In addition, we include the percentage of sales in the local market (LSALES) as an indirect measure of competitive pressure and the share of capital in hands of foreign agents (CAPEXT), of the public sector (CAPPUB) and of national agents (omitted) in order to account for differences in the firms bargaining power.

The availability of information about initial bargaining positions allows us to construct a proxy for profit uncertainty. In this sense, we use the difference between initial claim and offer (DCO), which may also be interpreted as an estimation of the expected change in productivity. This difference should increase the incidence and the duration of disputes. We consider that this variable is exogenous to current strike decisions because initial positions are announced at the beginning of the negotiation process.

We accounted for potential differences between union power by including the percentage of the workers within the council which belong to CCOO

(CCOO), regional (REG) and other unions (OTHER). We also consider a dummy which takes into account the presence of a single union in the workers council (SINGLEUN). The reason for this last variable is the fact that a single union workers council has no coordination problems and, as a result, could have a greater negotiation power. Consequently, its effect, as regards dispute rates, is expected to be negative.

In order to capture the effect of the negotiations timing we consider the length of the negotiation (LNEG) and a dummy for observed delays (DELY). Both are expected to affect positively strike activity. Bargaining unit *status quo* may be well represented by the lagged relative wage ($\{W-W_j\}(-1)$), the lagged quotient of effective plus overtime hours by regular hours (HOURS(-1)), and the lagged amount of tenure payments divided by the wage bill (TENP(-1)). The first variable should decrease the dispute activity because it increases the cost of a work stoppage. The second one may be interpreted as a measure of effort and has no clear effect a priori. However, if it acts as a signal, for instance, the expected sign should be negative. Finally, a higher level of tenure payment indicates higher seniority of workers. Workers with a high level of seniority are expected to show lower levels of dispute activity because they show a higher wage level and a higher bargaining power. The bargaining unit size is controlled by the employment level in the last year (N(-1)). We expect a positive influence on dispute rates. We also include the concession of a cost of living allowance clause (COLA) whose expected effect is a decrease in dispute activity because of the reduction in uncertainty about real wage increase.

The incidence of the market conditions has also been considered carefully. The number of days lost by strike per employee in the industry

(S_j) acts as a proxy to the aggregate bargaining pressure. It is expected to push up strike activity. An increase in the regional unemployment rate (u_r) should decrease dispute activity levels because a drop in the alternative wages. The lagged industry employment level (Δe_j) should increase both the incidence and duration of disputes. Additionally, we proxy prices by the expected price (EXPECT) level and its variance (STIPC). According to Cramton and Tracy (1991), as far as the current wage erodes by inflation, the strike threat becomes much more attractive. The role of inflation uncertainty is less clear.

Although there is no agreement about the inclusion of dynamics within the specifications when using discrete choice and/or duration models, we have tried to include it in order to account for the learning process of firms and unions. In fact, the occurrence of a strike in previous years could change the probability of observing a current strike. Moreover, the consideration of the lagged outcome improves the predictive power of the model. Finally, all the specifications contain year, quarterly and industry dummies, though the latter are not identifiable when applying first differences estimation methods[145].

Regarding the wage equation, the first group of variables we consider are proxying the change in firm's demand price for labour. On the other hand, the change in the level of sales per employee and an index of hours in the past year are intended to proxy firm's demand and excess of demand levels, respectively. We include the lagged level of profits per employee in order to control not only the change but the level. All three are expected

[145]This is the case in the conditional Logit and the first differenced LPM.

to add upwards pressure over the negotiated wage. In addition, we include the percentage of sales in the local market as an indirect measure of competitive pressure. Although we cannot construct a direct measure of profits or productivity volatility we use the difference between the union initial claim and the firm initial offer as a proxy of it which as above also control for the agents uncertainty. We use the share of capital in hands of foreign agents, of the public sector and of national agents in order to account for differences in the firms bargaining power. The differences among workers council bargaining power as well as the characteristics of the bargaining unit, its pay structure, the timing of negotiations (we add a dummy -RETARD- if the negotiation process starts after the expiration of the last agreement about wage increases) and the incidence of the market conditions have also been considered as in the above equations.

The higher the expected price level[146] the higher is expected to be the negotiated pay increase. Moreover, the mean negotiated wage increase in the same industry in the previous month represents an information that agents could use about other bargaining units actions and it could capture the wage spillover (see Chapter 4). It should contribute to the improvement of our specification in at least two directions. First, it offers some demand information not directly observable to the econometrician at industry and regional levels. Second, other firms wage settlements may enter directly into wage negotiations through the reservation wage and/or the profit function. The two variables are dated at the time of signing in the

[146]In order to proxy the expected price level we use an ARIMA forecast as explained in Chapter 4.

agreement equation and at the starting date of the negotiation process in the claim and offer equations and are expected to put upwards pressure over the outcome. Finally, we also consider carefully the possibility of dynamics in the wage equation as well as a full set of time bargaining dummies (year and seasonal) and industry dummies (one digit CNAE classification).

Regarding the claim and offer equations, there is no special reason for considering them extremely different as the wage specification, though we think there are some important changes (current strike variables are not entering the claim and offer equations, for instance). As already pointed out the wage increase setting and initial positions are expected to be a balance of firm's profitability and aggregate conditions.

b. Strike decision and duration of strikes

Table 5.4 reports the most relevant estimates for the reduced form strike decision equation [5.5]. The specification has been estimated using Probit (column (1)), conditional Logit (2) and LMP models in levels (3) and first differences (4). Whilst the Probit has been estimated using the initial sample of 2207 observations, the rest have been estimated using an unbalanced subsample of 1712.

We begin by explaining the sequence of tests. The pooled Probit and pooled Logit models are rejected in favour of a year by year Probit and a pooled Logit, respectively. Note that the last diagnostic is poorly determined because of the reduced sample size. Finally, after controlling for unobserved heterogeneity it is not rejected the negligibility of bargaining unit specific effects in the LPM specification. These diagnostics seem to suggest the presence of heterogeneity more than specific effects problem.

The main findings about the predictions of signalling models are quite satisfactory. The higher the initial disagreement the higher the probability of observing a strike. If we assume that this variable is proxying the unobservable component of profitability, the finding is in accordance with the theory. On the other hand, an increase in sales per employee in the past year affects negatively to work stoppage probability as expected, though the coefficient is not significant.

Regarding the union variables, the presence of a single union in the workers committee seems to reduce the probability of a strike. The omitted union proportion which corresponds to UGT presents a lower propensity to make a strike. The apparent positive effect that the negotiation length have in the first three columns vanishes when considering the first differenced LPM. On the contrary, it is clear that the delay in reaching an agreement increases the probability of observing a conflict.

Both the size of the bargaining unit and the past relative wage have significant effects on the probability of strikes. Whilst the former contributes to an increase the latter reduces it, which coincides with previous findings (Vroman (1989) and Card (1990b)). On the other hand, the lagged hours and the pay structure variables do not show any significant effect. Concerning to other firm variables, the results are not clear. The level of real profits per employee do not show any significant impact on the decision. However, the profits dummy has the correct negative sign although it is in general not significant. The effect of the share of sales in the domestic market is negative (except in the conditional LOGIT estimation) and both, the proportion of foreign and public ownership have a positive effect.

As far as there is not much time series variation in our sample, the results about upper firm level variables cannot surprise. According to the theory, strike activity is procyclical. Consequently, the unemployment variable should have a negative coefficient and the change in industry employment a positive one. Our findings are contradictory to the expected signs, the unemployment variable is found to affect positively and the change in the industry employment negatively. Similarly to Card (1990b), industry strike activity level increases the likelihood of firm strike, although the estimated coefficient is not significant in the differenced models (Conditional Logit -column (2)- and LPM -(4)). With respect to prices, we found that price expectations do not affect significantly contract strikes and price volatility reduces the likelihood of a strike. Finally, the quarterly dummies indicates that a strike is more likely to be observed in the second and third quarters which is normally when negotiation takes place.

Table 5.5 presents alternative models to explain the determinants of, work stoppages length threats (columns (1) and (2)) and observed duration (columns (3) and (4)). As noted in the previous section the panel treatment is clearly unaffordable because of the small sample size corresponding to positive durations[147]. Consequently, both results and diagnostics must be taken with caution. Although the power of our testing is limited, column (1) indicates that there is significant non-random sampling[148]. In such

[147]334 observations of 237 different bargaining units. Consequently, there are only 96 observations contributing to the estimation of the vector of relevant parameters, which severely limits the accuracy of the estimates.

[148]The coefficient of the inverse of Mills' ratio is significant in column (1)

circumstances, neither conditional LS (column (3)) nor heterogeneous Weibull model (column (4)) produce consistent estimates. In the absence of additional evidence, the comments apply to the unconditional estimates. Apart from this, columns (2) and (4) indicate that there is strong heterogeneity in the sample, as expected.

Looking at the results, we first concentrate in the evidence about predictions of signalling models. The difference between claim and offer, which is a proxy for the dispersion of the unobservable component of profitability, affects positively strike duration. The estimated coefficients imply that a 1% increase in such difference increases duration in 1.1% unconditionally and increases 0.1% the expected length. On the contrary, there is no significant negative effect on duration of either the mean level of productivity or the level of profits.

Union controls do not seem to show relevant effects on duration. The negotiation length and delay seems to increase overall the length of a strike. The size of the bargaining unit, contributes to an increment in the duration of a work stoppage. The higher the past relative wage the lower the length, as expected. This last result is in accordance with the theoretical predictions of Cramton and Tracy (1992), because a higher past wage increases the opportunity cost of a long conflict.

The proportion of sales in the domestic market has a negative effect on duration. This means that exporter firms tend to suffer longer conflicts. Likewise, the coefficient of the proportion of foreign ownership suggests that foreign firms also suffer longer strikes. Regional and industry factors

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and there is strong correlation between strike decision and duration errors.