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The Wage Effect of an Indexation Clause: Evidence from Spanish Manufacturing Firms*

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Abstract

When looking at wage negotiations, only 50 per cent of the agreements include a revision clause. Assuming that unions are more risk averse than firms, a revision clause may be non–optimal because of high transaction costs and/or the participation of weak unions in the bargaining process. A Switching model of wage increases under COLA and non COLA contracts is estimated for the Spanish manufacturing sector, using collective bargaining data for the 1984–1991 period. The propensity to a COLA clause is higher for nationwide unions than for other unions. Thus, it confirms our initial assessment about the role of weak unions. With respect to the wage increase equations, the possibility of a common wage structure across indexed and non–indexed contracts is strongly rejected by our data. The relevant price variable have been found having a positive, but small impact on wage increases in all the specifications under consideration. For indexed contracts, expected ex–ante total inflation coverage is set in a range of 0.35–0.63 per cent (0.63–0.99 considering the effect of proxy for the inflation threshold). It has also been found that workers pay a significant wage increase premia (wage increase differential) for obtaining a COLA. Such a premia has been estimated in a range of 3.5 to 2.6 per cent.

I. Introduction.

The process of wage determination in a bargaining context is an extremely difficult topic of empirical analysis. There are many reasons which could explain such difficulty. Sometimes the lack of information about it, other times because of the industrial, social or individual relationships involved or the fact that it is linked to decisions such as hours, employment setting or cost of living allowance clauses (COLA) or actions such as the conflicting activity. In our opinion, the correct framework for analyzing wage setting process must consider the above decisions and actions[1].

As mentioned, we consider that there is a potential gain in considering some of these issues which could be playing an important role in negotiations. Our main purpose consists in formulating a model of joint determination of a COLA clause and a wage increase under inflation uncertainty. If we believe that bargainers seek the maximization of a function of their real income, protection against unexpected inflation ought to have an important role to play.

The literature on wage indexing in an uncertainty about inflation context has widely assessed that a firm and a union bargaining over wages could benefit from indexing the contract[2]. Under the assumption that the implicit bargaining costs are relatively low to the potential gain of indexation and the union is more risk averse than the firm[3], Shavell (1976) showed that there is at least one Pareto optimal contract which includes a COLA clause. In fact, the more risk averse the union is, the higher the

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[1] As examples of empirical papers relating wage and employment setting we could mention Alogoskoufis and Manning (1991) and Dorion (1992). Papers relating wage setting and indexation clauses are Card (1986) and Prescott and Wilton (1992). A paper relating the wage increase and the strike activity is Stengos and Swindinsky (1990).

[2] See Shavell (1976), Blanchard (1979), Dazinger (1980, 1983), Ehrenberg et al. (1983), Card (1986) and, more recently, Gottfries (1992).

[3] See Dazinger (1980) for an exposition of the arguments in favour of such maintained assumption. The main argument relies on the fact that there are less opportunities for diversifying human capital against risk than for diversifying a similar amount of other kind of capital.

indexation degree it wants. Clearly, the union ought to be willing to accept a lower expected real wage for getting a COLA clause from the firm.

Note that, in such a context, rejection of the escalator contract must be due to either the firm being relatively more risk averse than the union and/or relatively high transaction cost[4]. It is hard to believe *a priori* that both causes could explain the empirical evidence (in Spain and in other countries) that more than half of the contracts does not include a COLA.

Taking into account that the wage setting is a process of bargaining, there is a very simple explanation for not observing an escalator contract (despite the above two assumptions). It is sufficient for rejecting the clause that the union's wage bargaining power is heterogeneous amongst contracts with and without the clause. In particular, if the union's wage bargaining power without the clause is sufficiently lower (in the manner we will show in the next section) than with it, the firm will prefer not to accept the clause. Thus, rejection of the escalator contract can arise regardless of high transaction costs and the relative higher risk aversion of the firm.

It is not our purpose to investigate on the determinants of bargaining power in negotiation[5], but on the reason of different bargaining power with and without indexation clause. The reason for observing such a heterogeneity in negotiation power is twofold. On the one hand, it maybe the case that some unions (worker councils) are not able to motivate their rank and file to fight against high unexpected inflation values (i.e., far away from inflation target); so, it is difficult to incorporate them into the negotiated wage without a revision clause (on the contrary, it is not very difficult to do so when the union has a escalator contract). In this sense, we will be distinguish between *weak* unions (those with low capacity to motivate their rank and file) and *strong* unions (with high capacity to motivate their rank and file). Note that under this assumption the expected

[4]In fact neither of both assumption is strictly necessary. Recently, Gottfries (1992) showed, in an extension of the Baily (1974) and Azariadis (1975) standard labor contract model, an equilibria in which, provided all other contracts do not are contingent, there is no indexation.

[5]For a recent investigation see the excellent work of Doiron (1992).

wage of workers with a protection clause is, *ceteris paribus*, higher than the expected wage of employees without it (opposite to the intuition).

On the other hand, as far as the underlying bargaining structure is dynamic and there is, at least in Spain, renegotiation of wages (increase) almost yearly, it could be the case that some unions prefer not to ask for a COLA clause but to establish a mechanism for incorporating the deviation of inflation from their expectation into the next wage negotiation round. That is, for some bargaining pairs the inflation coverage is given not through COLA coverage but through what is usually called price catch-up (PCU). For weak unions PCU inflation coverage may be an alternative to avoid costly negotiations. The firm may also prefer PCU setting because in most cases it will imply a new bargaining process, though examining the ex-post conditions, and in many cases, from the point of view of the firm, will imply a better adjustment (for instance, if the past year was a bad year for the firm, it could negotiate relatively lower wages).

Taking the above facts into consideration we will formulate a simple decision model about which kind of contract is binding (COLA contract or not) and which is the ex-ante wage increase. Firms are assumed to be risk neutral (not crucial) and the union maximizes a standard risk averse utility function. It will be assumed, for simplicity, that the bargaining solution is sequential. Agents decide first whether or not there will be an indexation clause in the contract and, second, they set the ex-ante wage increase to hold in each contingency. They take the decision, both having the right to *veto*, about the revision clause by comparing the utility levels they get in both regimes (with and without COLA). Finally, we will also assume that the solution to the negotiation process, in each regime, is represented by a generalized Nash bargaining (GNB) program, in which the weights on the respective utilities in the GNB are the bargaining powers of the two parties.

There are few examples of applied work on this field, mainly, because of the difficulty of getting the data that this kind of model requires (particularly, the lack of information about several key variables, such us the wage-inflation elasticity) and the extreme difficulty of linking any theoretical framework with an econometric setup. Among the recent papers, we highlight Christofides et al. (1980), Hendricks and Kahn (1985), Card (1986), Christofides (1987) and Prescott and Wilton (1992). The empirical part of this study is related, in some sense, to the last paper, a recent work using Canadian data on collective agreements. The structure of Spanish data on collective bargaining is close to Canadian data so that their empirical methodology is easily applicable, although we would like to emphasize several topics not considered there, especially with respect to price variables and the provisions of the COLA contract (i.e, wage-inflation elasticity and inflation threshold -if any-), substantially different in the Spanish case.

Moreover, the consideration of the COLA provisions faces several empirical difficulties, since they are only observable if the contingent event occurs (normally inflation rate above a given ceiling). In this case, the information set only includes the ex-ante and the ex-post wage (if any). Hence the wage-inflation elasticity is only known[6] for the share of COLA clauses that are triggered. Given the fact that the wage-price elasticity might be crucial in the determination of COLA contracts, ignoring this empirical restriction might lead to a serious bias. We will undertake this problem by considering a reduced form model of wage-price elasticity determination. The estimates of such a model will be used to forecast the wage-price elasticity for the whole COLA sample. I will also pay attention to the problem of measuring implicit wage differentials among indexed and non-indexed contracts. In our context, wage increase differentials provide a measure of the cost (or premium) workers must pay to obtain a COLA clause. Apart from their inherent utility, wage increase differentials will be compared with the ex-post contingent compensation to obtain an indirect evaluation of the ex-post wage increase differentials among both kinds of contracts in the sample period.

The basic data source used here is the "Estadística de Convenios Colectivos" (ECC), which has information on *all* the agreements settled in Spain during the 1981-1991 period. From the raw data set we extracted an unbalanced panel data set for the manufacturing sector. We restricted the

[6]And not in a exact form but proxied by using the following formula:
Wage-inflation elasticity= (ex_post wage - ex_ante wage)/(inflation rate).

sample to firms which can be observed for at least four consecutive years. After looking carefully to the data, it must be pointed out that sample selection (by merger, acquisition or missreporting) seems not to be especially important in our data set. To my knowledge, there is only one previous work using this data set[7], although the possibility of following units across time was not used there.

The rest of the paper will have the following structure. In section II we describe briefly the main characteristics of the manufacturing data. In section III a reference model is developed. The econometric setup and methods are described in section IV. The main results of the analysis and an evaluation (at sample means) of implicit ex-ante wage differentials can be found in section V. In section VI we provide a brief summary. After the tables, the wage-elasticity problem is described in the Appendix A and the variables are defined in the appendix B. .

[7] Jimeno (1992), who considers a very simple wage increase determination equation.

III. The data.

From 1981 to 1991 there had been more than forty-two thousand observations on collective agreements which are compiled in the ECC. More than twenty thousand of these agreements correspond to the manufacturing sector. The record has information on wage increases, regular annual hours and cost of living allowance clause to prevent for unexpected inflation, among other variables. Table 1 summarizes the most important results of the bargaining process in those years for the sample. The percentage of COLA contracts ranges from a low of 31.5 to a high of 58.7 per cent in 1990. Notice that there is some evidence for asserting that COLA propensity is related to bargaining unit size (proxied by number of employees). For all the years and both bargaining levels considered, the mean size of bargaining unit under COLA doubles the mean under no COLA.

For all the groups considered the ex-ante wage increase is higher without COLA than with it. Differences are, in general, larger in the 1987-1991 period than in preceding years (1984-1986). However, after revision, ex-post wages under COLA are, except for 1987, higher. This is caused by the high amount of deviation of end of the year inflation with respect to target in those years for Spain (mean 1984-1991: 1.14, with a maximum in 1989 of 3.9 per cent points).

Table 2 shows that COLA contracts lower the mean and the variance of the ex-ante wage increases in all the sampling years. Thus, it is confirming the theoretical guess that risk averse workers exchange wage for a lower variance. In Table 3 we report the distribution of our sample by duration of the contract, percentage of revision clause, percentage of contracts settled after the expiration of the last agreement and mean delay (days from the expiration date of the last agreement). First notice, that 76.7 per cent of contracts are first year observations and close to 20 per cent are second year observations, the rest, around 3 per cent are third year or more contract observations. Consequently, it seems reasonable to assume that wage increase bargaining takes place almost yearly.

As a final point we like to examine persistence of the COLA and the non-COLA decisions in sample. In Table 4 we report the sample percentage of observing a contract with COLA in a given year conditional on having a revision clause in each of the K preceding years. The same concept is reported for the non-COLA contracts. Conditional on a COLA in the previous year, COLA sample percentage (which can be viewed as a sample probability) is 80 per cent. Conditioning to more than one previous COLA contracts increases slowly the conditional sample per cent; to 90 per cent after 6 periods of having a revision clause. The same pattern is observed for non-COLA contracts. Hence, we conclude that although the conditioning probabilities are increasing in time, one period conditioning explains quite well the state dependence in COLA decision.

III. A simplified reference model.

Our main purpose throughout this section is to develop a reference model, under very restrictive assumptions which will allow us to obtain an explicit solution, to justify the special case in which the COLA contract is not accepted by the firm despite the fact that the union is risk averse, the firm is risk neutral and there are not significant negotiation costs. More formal models of several forms of wage indexation can be found in Ehrenberg et al (1983), Card (1986), both dealing with the optimal indexation level and, recently, Gottfries (1992) who worried about indexation to firm's demand in an Insider-Outsider context[8].

Assume we have a competitive firm and a union bargaining over a pay scheme during an annual contract having uncertainty about the level of inflation. The firm, assumed to exhibit constant returns to scale, seeks the maximization of its expected real profits per worker (π):

$$[1] \qquad E_P V(\pi) = E_P\left\{\frac{P_j y_i - w_i}{P}\right\}$$

where P_j is the price firm faces, y_i is the level of output per worker, w_i is the wage and P is an aggregate price level (note that the subindex i will always refer to firm, the subindex j to the industry and the absence of subindex to aggregate variables). To simplify [1], we assume that the firm's price is related to the aggregate price level as follows: $P_j = \eta_j P$. Where η_j is assumed to be constant in a given industry[9]. Therefore, [1] simply becomes:

$$[1'] \qquad E_P V(\pi) = E_P\left\{\eta_j y_i - \frac{w_i}{P}\right\}$$

The union seeks the maximization of its expected utility, which is

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[8] In this work it is argued that current employees (insiders) have few incentives to change to a contingent contract, since nominal demand shocks have small effects on real wages and employment variation is mostly given by variation in the rate of hiring.

[9] We assume this to simplify our exposition. For a model relaxing this strong assumption see Ehrenberg et al (1983).

assumed to depend solely on wages and to have the following specific functional form[10]:

$$[2] \qquad E_P U = E_P\left\{\frac{1}{m}\cdot\left(\frac{w_i - a}{P}\right)^m\right\}$$

where m-1 is the degree of union's relative risk aversion (DRRA) and a is the alternative wage level (quit wage).

Assume that the negotiation process is sequential. Firstly, both agents decide whether or not the contract will include an indexation clause. Secondly, they set the ex-ante wage increase if there is no revision clause, and the ex-ante wage and the exact form of the contingent part of the contract if there is COLA clause. Naturally, agents decide the first stage by comparing their respective utility levels in the second stage. Consequently, we must solve first the last stage.

Assume that the solution to the implicit optimization without revision clause is well represented by the solution to the following GNB problem[11]:

$$[3] \qquad L_w = [E_P\{U(w/P)\}]^{\beta}.[E_p\{V(\pi)\}]^{1-\beta}$$

where β is the union bargaining power without a COLA clause. For a given β and assuming the firm is risk neutral[12] the solution to the above problem, the ex-ante wage without COLA, w^*, might be written as follows,

$$[4] \qquad w^* = \frac{m\beta\eta_i y_i + (1-\beta)a}{[m\beta + (1-\beta)]P^*}$$

where $P^* = E(1/P)$. The expected utility levels of the firm and the union without indexation clause will be compared with the solution of the contingent wage contract:

$$w_C = w(P) \qquad \text{for some aggregate price index P.}$$

[10]This is a simplification of the utility function specification of Ulph and Ulph (1990).

[11]Firm's status quo position is set, for simplicity, to zero.

[12]This assumption, not crucial for the analysis, will allow us to obtain an explicit solution for the ex-ante wage.

Normally, in Spain and other countries, the contingent wage contract takes the following explicit form[13]:

$$[5] \qquad w_C = \begin{cases} w_L & \text{if } P \leq PU \\ w_H = w_L + \theta(P-PU) = w_L(1+\theta^*(P-PU)) & \text{if } P > PU \end{cases}$$

where w_C is the cola wage, w_L is the ex-ante wage, w_H is the contingent wage level, PU is a prefixed inflation threshold and $\theta^*>0$ is the wage-price elasticity. Notice we have (with respect to the non-COLA contract) two additional parameters θ^* and PU, so the escalator contract is fully characterized by (w_L,θ^*,PU). In what follows we will assume the bargaining procedure restricts the contract to be as above. Under such restriction, the relevant unrestricted objective function for a COLA contract is given by:

$$[6] \qquad L_{w_L} = \left\{E_P\{U(\frac{w_L-a}{P}) \;/\; P\leq PU\}q + E_P\{U(\frac{w_H-a}{P}) \;/\; P>PU\}(1-q)\right\}^{\beta_C}$$
$$\left\{E_P\{(\eta_j y_i - \frac{w_L}{P}) \;/\; P\leq PU\}q + E_P\{(\eta_j y_i - \frac{w_H}{P}) \;/\; P>PU\}(1-q)\right\}^{1-\beta_C}$$

where $q=prob(P\leq PU)$ and β_C is the union bargaining power under a COLA, assumed to be, at least, as great as β. As we are mainly interested in comparing (ex-ante) wage solutions, we shall opt for solving it, for a given pair (θ^*,PU), in terms of the ex-ante wage[14]. Consequently, the solution to the unrestricted problem, can be written down as,

$$[7] \qquad w_L^*(\theta^*,PU) = \frac{m\beta_C\eta_j y_i + (1-\beta_C)a}{[m\beta_C + (1-\beta_C)]P^{**}}$$

where:

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[13]This expression is similar to Card's (1986) basic setup for an escalator clause. See the Appendix A for the most usual formulae in Spain.

[14]It is evident that in our context (a risk neutral firm and no bargaining costs) the optimal solution for θ^* is full indexation and PU = minimum{P}. However, if we assume that they are determined in an additional bargaining stage for sharing the gains of indexation, they may differ from the above solution.

$$[8] \qquad P^{\circ\circ}(\theta^{\circ},PU)=q.E_P[\tfrac{1}{P}/P\leq PU]+(1-q).E_P\{\tfrac{(1+\theta^{\circ}(P-PU))}{P}/P>PU\}$$

That is, $P^{\circ\circ}$ is a weighted (by q) function of the expected inverse low price and the sum of the expected inverse high price and the wage-price elasticity. Notice $dw_L^*/d\theta^\circ<0$ and $dw_L^*/dPU>0$. It is straightforward to show that if $\beta=\beta_C$ then the indexed contract is optimal and the ex-ante wage is the solution to the unrestricted problem [6]. However, in the case $\beta_C>\beta$, expression [6] is no longer the relevant problem. We must consider that both expected firm's profits and the utility level of the union under COLA must be both higher than without it. Given the fact that considering both restrictions will make the algebra tedious, we will omit it by means of parameterizing the optimal solution of the restricted version of [6]. It is obvious, that for β close to β_C the escalator contract is still preferable for both agents. In that case, the ex-ante wage will be given by:

$$[7r] \qquad w_L^*(\theta^*,PU,k)=\frac{m\beta_C\eta_jy_i+(1-\beta_C)a}{[m\beta_C+(1-\beta_C)]P^{**}}.k \text{ for some } k\leq 1 \text{ if } \beta \text{ is close } \beta_C$$

Eventually, for a β sufficiently different from β_C the COLA contract is no longer preferable for both agents. The following proposition will set the condition in which the indexed contract is not chosen despite the fact that the union is risk averse and the firm is risk neutral[15].

Proposition. Assume a risk neutral firm and a risk averse union ($0<m<1$) and assume that price level (P) is distributed as a uniform: $U[P_L,P_H]$. Then the COLA contract, for a given pair (θ^*,PU), will not be chosen by the firm if and only if,

$$\boxed{\beta_C > \beta + r^*}$$

where $r^*>0$ is a function of θ^* and PU.

[15] The result could be easily generalized to the case where the union is just more risk averse than firm. But, in this case, the algebra is tedious and the wage solution does not have an expression as easy as we have in this simple case.

Proof:

-If $\beta_C > \beta$ it follows immediately that $EV[\pi(w^*)] > EV[\pi(w_L^*)]$ for any $\theta^*>0$ and PU in the relevant interval $[P_L,P_H]$. So, in the absence of any additional compensation, it will prefer the non COLA situation.

-Because EU(COLA) > EU(NO_COLA) for the union, we also must consider the possibility of a proportional[16] (or lump sum) compensation, $0<k<1$, from union to firm, making:

[p.1] $$EV[\pi(w^*)] \leq EV[\pi(w_L^*.k)]$$

[p.2] $$EU(w^*) \leq EU(w_L^*.k)$$

for any $\theta^*>0$ and PU $\in [P_L,P_H]$. It is also plain to see that, for a given pair (θ^*,PU), the firm will accept the COLA contract if k is lower than k^M:

[p.3] $$k^M = \frac{qw_L^*E_L-(1-q)w_L^*[E_H+\theta^*(1-PU \cdot E_H)]}{w^*E} = k^M(PU,\theta^*,\beta,\beta_C)$$

where $E = E(1/P)$; $E_L = E(1/P \ / \ P \leq PU)$; $E_H = E(1/P \ / \ P>PU)$. On the other hand, the union will ask for an indexed contract as far as k is higher than k^L:

[p.5] $$k^L = k^L(\theta^*,PU,\beta_C-\beta)$$

which is the solution to: $EU(w^*) = EU(w_L^*.k)$. From [p.1] and [p.2], we know that there will not be a COLA contract iff:

[p.6] $$k^L(\theta^*,PU,\beta_C-\beta) > k^M(\theta^*,PU,\beta_C-\beta)$$

Finally, defining $r(\theta^*,PU) = \beta-\beta_C$, and making use of the facts:

i) k^L is increasing in $r(\theta^*,PU)$
ii) k^M is decreasing in $r(\theta^*,PU)$
iii) $k^L(\theta^*,PU,0) < k^M(\theta^*,PU,0)$ [=1]

[16] We will show our proposition using the proportional assumption because the algebra and the interpretation are easier.

we can state that, a COLA contract is not optimal for the firm if:

$$\beta_C > \beta + r^*(\theta^*,PU)$$

where r^* is the solution of: $k^L(\theta^*,PU,r) = k^M(\theta^*,PU,r)$; $r^* \geq 0$ ■■■

It is straightforward to show that in absence of any COLA cost, a risk averse union will prefer always the COLA contract. Therefore, it will never make use of its right to *veto*. So, in fact, the firm is always taking the COLA decision by means of its *veto* right[17]. It is also clear that if union bargaining power is the same in both situations ($\beta_C \simeq \beta$), the employer will agree (not using its *veto* right) to concede to the union the escalator contract, because $E(w_C^*/P) \leq E(w^*/P)$; that is, the expected real wage under revision clause is lower than the expected real wage without it, thus the COLA contract is preferable. Notice that in the simple case we have pointed out (firm risk neutral, $\beta_C=\beta$ and no bargaining about employment) and as far as a COLA contract represents a lower real wage, it will tend to increase its labour demand. On the contrary, if union bargaining power with COLA (β_C) is sufficiently larger[18] than without it (β), the firm will not agree to it, because $E(w_C^*/P) \geq E(w^*/P)$. Under this alternative, as far as the COLA represents a higher expected real wage, it will imply a relatively lower labour demand.

As a sort of summary, let us write down the basis of our model. Conditional to low negotiation costs, the union will always ask for a COLA and will pressure in negotiation (to push up its β power and k^M) to enforce firm to accept it. The firm will accept it, for a given (θ,PU), iff:

$$[9] \qquad E_P\left\{V[\pi(w_L^*(\beta_C).k)]\right\} \geq E_P\left\{V[\pi(w^*(\beta))]\right\}$$

In such circumstance the ex-ante wage will be given by:

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[17] Notice that the veto assumption is not really important because the union could avoid it by increasing its bargaining power without COLA (β) through adding pressure in the negotiation process (i.e., motivating its workers against the actions of the firm).

[18] To exclude the possibility of a constrained solution.

$$[7'] \qquad w_L^{\circ}(k) = \frac{m\beta_C \eta_j y_i + (1-\beta_C)a}{[m\beta_C + (1-\beta_C)]P^{\circ\circ}}.k \qquad \text{for some } k \leq 1$$

where $k \lesseqgtr 1$ iff $\beta \lesseqgtr \beta_C$. On the contrary, if [9] does not hold, there is no revision clause and the ex-ante wage will be given by:

$$[4] \qquad w^{\circ} = \frac{m\beta \eta_i y_i + (1-\beta)a}{[m\beta + (1-\beta)]P^{\circ}}$$

The strength of the above framework of joint wage and COLA setting is twofold. On the one hand, it does not exclude alternative explanations for the firm rejection of the COLA clause. For instance, if the firm is more risk averse than the union, [9] does not hold, so the COLA contract is rejected by the firm. Hence, our criterion function for COLA decision is not rejecting either of both plausible explanations (we do not consider the implausible case of high transaction costs) for the non-optimality of the COLA contract. On the other hand, it provides a well-defined structural framework for the econometric specification, especially with respect to the COLA criteria function which could be easily extended allowing for more general determinants (i.e. risk aversion of the firm and/or presence of relevant bargaining costs).

IV. The econometric specification.

Following the arguments of the above section it seems adequate to assume that wage increase setting depends on whether or not protection against inflation is negotiated for the relevant period. Our strategy will consist in the formulation of a pseudo-reduced form joint COLA and wage increases determination model. In particular, and assuming linearity in the relevant variables, the COLA decision is related to firm's criteria (given by equation [9]) and the wage equation is related to [4] -for non indexed contracts- and [5]-[7'] -for indexed contracts. Formally:

$$[10] \qquad I = \begin{cases} 1 \text{ if } I^{\circ} > 0; \\ 0 \text{ otherwise} \end{cases}$$

$$[10'] \qquad I^{\circ} = Z'\varphi + \varepsilon_I$$

$$[11] \qquad \Delta w_{NC} = h(X_{NC}, \alpha_{NC}) + \varepsilon_{NC} \quad \text{if } I=0 \text{ (non COLA)}$$

$$[12] \qquad \Delta w_C^e = g(X_C, \alpha_C) + \varepsilon_C \quad \text{if } I=1 \text{ (COLA)}$$

where I is a dummy taking the value one if the contract has an indexation clause; I° is the unobservable latent variable which determines whether the negotiation unit signs a COLA or a non-COLA contract[19]; Z represents the set of exogenous determinants of a revision clause; Δw_{NC} represents the wage increase without COLA; Δw_C^e is the expected wage increase under revision clause; X_C and X_{NC} are, respectively, the (pseudo-reduced form) determinants of wage increase with and without COLA; φ, α_{NC} and α_C are the unknown sets of coefficients; and finally, ε_I, ε_{NC} and ε_C are, respectively, the error in [10]-[12], which are assumed to be serially uncorrelated jointly normally distributed with covariance matrix:

[19]Note that in our theoretical model I^* corresponds to the following expression: $I^* = E_P V[\pi(w_L^*(\beta_C).k)] - E_P V[\pi(w^*(\beta))]$

$$\Sigma = \begin{bmatrix} 1 & \sigma_{INC} & \sigma_{IC} \\ \sigma_{INC} & \sigma_{NC} & \sigma_{NCC} \\ \sigma_{IC} & \sigma_{NCC} & \sigma_{CC} \end{bmatrix}$$

Consistent estimates for both wage increases equations might be obtained by using the two-stage method by Heckman (1976) for selectivity models. Efficient estimates might be also obtained by maximizing the likelihood of the above system. Our preferred alternative will be Heckman's method, because, on the one hand, the restriction on Σ (in fact, $\sigma_{NCC}=0$) cannot be tested without impossing more structure on the form of Σ and, on the other hand, there are some variables potentially endogenous, which makes us to opt for an instrumental variables method. In the following, we will describe, with some detail, the insights of the empirical specification of equations [10]-[12] in the light of the comments of our last two sections.

a. The COLA decision.

It is well established (see, for instance, Ehrenberg et al. (1983) and Card (1986)) that the probability of observing a revision clause is determined by the welfare gains associated with the COLA contract compared to the non COLA contract. In section III, we stated three different reasons for the non optimality of a COLA clause (given that firm is taking the decision): The firm is more risk averse than the union; the cost of negotiating a COLA is higher than the expected gains (so the net gains are negative); and, finally, the workers' committee is weak (it has lower bargaining power if a COLA clause is not agreed), which implies in our model that the firm's value function under COLA is lower than without it, and it decides not to agree on it.

Therefore, in our empirical specification for the COLA decision, we should consider three different sets of factors. First, those related to firm's risk aversion, like (in the absence of information about firm price level) industry prices (ΔIP_j). In this sense, we expect that the higher (the lower) the elasticity of firm prices with respect to the conditioning variable (inflation) the higher (the lower) the probability of observing a COLA clause (Ehrenberg et al. (1983)). Second, those affecting costs or expected gains, like the amount of unexpected inflation in the previous year ($UNEXPECTED\_INF_{-1}$), price volatility (σ_P) and negotiation length (DEL)[20]. For instance, we expect that the lower the unexpected inflation in past year (or price volatility) the lower the COLA incidence, because the expected gains are relatively smaller. Finally, among the factors explaining whether workers' council is strong or weak, we consider the size of the bargaining unit, measured by the number of employees (EMP), the composition of worker's committee (in terms of weak -non_affiliated and other unions- and tough unions -CCOO and UGT-) and some indirect measures of the strength in negotiation, like negotiation length and specific industry conflicting activity as a proxy of firm's conflicting activity[21] (i.e., strike activity). We expect that the larger the bargaining unit the higher the probability of setting a COLA clause. According to the theory, we expect to show the nationwide unions being stronger than the regional (included in the others group) or the non-affiliated workers.

We also consider some market related factors, like unemployment (U), industry unemployment (U_j), industry productivity ($IPROD_j$), and a measure of specific industry conflicting activity (S_j). Finally, we allow for time dependence in the COLA decision by introducing a dummy that takes 1 if there was a COLA in the past year. In the light of Table 4, we consider that one period time state dependence suffices to explain the dependence of the COLA decision. As a summary we consider the following specification for equation [10']:

$$[10'] \qquad I^{*} = Z^{*\prime}\varphi^{*} + \tau I_{-1} + \varepsilon_I$$

[20] Since we have no information about the spell of negotiation we use to measure it the delay of the signing of the contract with respect to the expiration date of the last contract.

[21] Unfortunately, we have no information about strike activity at firm level in our sample. So, the results in this respect will be merely approximate.

b. The non-COLA wage increase equation.

Our proposal is closely related to most of the previous empirical work on wage determination (see Christofides et al. (1980)). We consider three different sets of variables. The first one includes the industry unemployment rate (u_j) and the inverse of the regional unemployment rate in the quarter preceding the signing of the contract (u_r^{-1}). Both are proxying the excess of demand in their respective labour market. Two variables represent the expected shifts in labour demand and/or supply during the year, the specific industry j productivity during the past year ($IPROD_{j-1}$) and the change in industry prices, ΔP_j, which will be considered as a potential endogenous variable. And, as a key variable, we use the expected (by the time of signing) change in the CPI for the current year[22] (P_m^e), which is a relevant shift variable for the labour supply curve. Following closely the literature on wage increase determination, we consider a price catch-up (PCU) variable (we will describe it more in detail later on in this section) to account for past uncompensated inflation in previous year. This variable might be viewed as an ex-post mechanism for compensating workers against unexpected past inflation (see Christofides et al. (1980) and, recently, Prescott and Wilton (1992)).

Finally, we considered an additional variable, the mean negotiated wage increase in the specific industry in the previous month iw_j^{m-1}, which is a proxy about what other related bargaining units (in the same industry) are doing[23]. It will capture, if any, the "wage spillover". Following the reasoning in McConnell (1989), there are two rationales for including such a variable. First, it will proxy some valuable firm information, sometimes not observable for an econometrician. And second, wage settlements at other firms may enter directly into wage negotiation via reservation wage or via firm profits function, which is usually known as efficiency wage model[24].

[22]See the data appendix for a description on the form of P^e.

[23]See Burton and Addison (1977) for a review of the empirical studies of the correlation between wage settlements.

[24]See Akerlof and Yellen (1986) for a recopilation of earlier efficiency wages models and Layard et al. (1991) for a detailed description of the model and also a summary of findings.

Notice, that under both rationales it is expected to affect positively the negotiated wage.

The second group of variables includes some bargaining specific factors like the size of the bargaining unit measured by the number of employees (EMP), the proportion of extra-hours by regular hours per employee during the past year (XH_{-1}) and the change in the employment level (ΔEMP). The last two are considered in order to account for the firm's potential excess demand of labour, which is expected to add pressure over the negotiated wage increase. We also consider a group of variables related to the bargaining process like the delay during negotiations, which is considered in a quadratic form, and two variables, the productivy clause (C_PROD) and the absenteeism clause (C_ABS), both taking one if they agreed during negotiations, 0 otherwise. It is expected a lower observed settlement if such a clauses are agreed because both are implying a contingent deferred payment. Both will be instrumented to prevent for some endogeneity. The final subset of variables of this group are the proportion of workers representatives that belong to the CCOO and USO unions, to independent groups (INDEP) and to others representatives (OTHERS).

The last variable we consider is an attempt to capture the implicit premium (if any) for renouncing to a COLA clause. Not being easy to identify such a premium, we opt for introducing a dummy which takes the value one if past year agreement included a COLA and zero otherwise (COLA⇒NO_COLA). We expect to observe a positive effect on wage increase, although not very important, because the shift from revision to no revision might be induced by a fall of union bargaining power in a given year and/or a sudden worsening of the firm's performance.

Since we are going to use Heckman's two-stage estimation method, our empirical specification of equation [11] must take into account the potential selectivity bias arising in such an estimation process. More in detail we must consider that:

$$E(\Delta w_{NC} \;/\; I=0) \;=\; h(X_{NC},\alpha_{NC}) \;+\; E(\varepsilon_{NC}/I=0)$$

and:

$$E(\varepsilon_{NC} \;/\; I=0) \;=\; E(\varepsilon_{NC}/I^{*}\leq 0) \;=\; \rho_{INC}{*}\sigma_{NC}{*}{-}\frac{\phi(Z'\gamma)}{1-\Phi(Z'\gamma)} \;=\; \sigma_{INC}{*}\lambda_{NC}$$

where ϕ and Φ are the univariate normal density and distribution functions, respectively and ρ_{INC} the coefficient of correlation between the errors in [11] and [10]. Provided some consistent estimation of λ, the Mill's inverse ratio, say $\hat{\lambda}_{NC}$, the empirical specification is given by:

$$[13] \quad \Delta w_{NCi} = \alpha_o + \mu P^e_m + \alpha_1 PCU + \alpha_2 PROD_{j-1} + \alpha_3 u_j + \alpha_4 u_r^{-1} + \alpha_5 \Delta P_j + \alpha_6 iw_j^{m-1} + \alpha_7 EMP_i + \alpha_8 \Delta EMP_i + \alpha_9 DEL_i + \alpha_{10} DEL_i^2 + \alpha_{11} C\_PROD_i + \alpha_{12} C\_ABS_i + \alpha_{13} XH_{i-1} + \alpha_{14} CCOO_i + \alpha_{15} INDEP_i + \alpha_{16} OTHERS_i + \alpha_{17}[COLA \Rightarrow NO\_COLA] + \sigma_{INC} \hat{\lambda}_{NC} + \varepsilon^o_{NCi}$$

where i subindex is referred for the bargaining unit, j for the sector to which it belongs and m for the month of signing. Notice we have eliminated the subindex for the year of contract to facilitate reading. Finally, the price catch-up variable is defined as follows[25]:

$$[14] \quad PCU = (1-\theta^i_{-1})(P_{-1} - \mu P^e_{m-1})$$

where P_{-1}, is the change in the consumer price index (December to December) during the past year and θ^i_{-1} is the wage-price elasticity agreed in previous year (which is zero if the previous year contract did not include a COLA clause in it). Notice, we are assuming that bargaining unit specific factors, if any, are not relevant in the wage increase specification. Given the above specification, and assuming ε^o_{NCi} is a well-behaved error term, consistent estimates of the set of parameters may be obtained by applying non-linear least squares (NLS) provided that the variables in [13] and the the error term are uncorrelated. If this does not hold, that is, if there is a group of variables correlated with the error term, consistent estimates might be obtained by applying non linear instrumental variables (IV-NLS) to [13].

[25] See Prescott and Wilton (1992) for a motivation.

c. The COLA wage increase equation.

An indexation clause adds some degrees of freedom in the way that monetary compensation takes place. Any full specification of the effect of such a clause should consider the wage-price elasticity (θ^i) and the inflation threshold (PU^i), both assumed to be contract specific[26]. According to the form of the COLA contract pointed in the previous section, the ex-ante (expected) COLA wage equation might be written as follows,

$$[15] \qquad \Delta w^{\varepsilon}_{Ci} = \Delta w_{Ci} + (1-q_i)[\theta^i(E_m(P/P>PU^i)-PU^i)] + \varepsilon_{Ci}$$

where $\Delta w^{\varepsilon}_{Ci}$ is the expected wage increase under a revision clause, Δw_{Ci} is the ex-ante wage increase, $(1-q_i)[\theta^i(E_m(P/P>PU^i)-PU^i)]$ is the contingent inflation compensation (conditional on inflation greater than a given price threshold), $(1-q_i)=p(P>PU^i)$ and ε_{Ci} is a error term. Subtracting the contingent part of the contract in both sides of [15] we have an expression for the ex-ante negotiated wage increase (Δw_{Ci}):

$$[16] \qquad \Delta w_{Ci} = \mu^i P^{**} + g_1(X_{1i}) + \varepsilon_{Ci}$$

where P^{**} is defined as in [8] and μ^i is the (non contingent and contract specific) wage-price elasticity of the ex-ante wage increase (Δw_{Ci}: i.e., without considering the contingent part of the revision contract). Since the coefficient μ^i is potentially different in every contract it will be assumed that it is related to θ^i as follows:

$$[17] \qquad \mu^i + \mu_2(1-q_i)\theta^i = \mu_1$$

This restriction implies that the sum of the non contingent and a linear function of the contingent wage-inflation compensations is constant across contracts and will allow us to deal with [16] without having to consider the contract specific factor, μ^i. We must not confuse either μ^i or θ^i with the total (expected) wage-price elasticity (γ^i) of the expected COLA

[26]The implicit cost of an indexation clause has not been considered.

wage increase (Δw^{e}_{ic}), which is given by:

$$[18] \qquad \gamma^i = \mu^i + (1\text{-}q_i)\theta^i$$

To reach an estimable expression and as we did for the non-COLA equation, we should consider that,

$$E(\Delta w_C \ / \ I=1) = \mu^i P^{**} + g_1(X_{1i}) + E(\varepsilon_C/I=1)$$

and:

$$E(\varepsilon_C/I=1)=E(\varepsilon_C/I^*>0) = \rho_{IC}{}^*\sigma_C{}^*\frac{\phi(Z'\gamma)}{\Phi(Z'\gamma)} = \sigma_{IC}{}^*\lambda_C$$

provided a consistent estimator of λ_C, the Mill's inverse ratio, say $\hat{\lambda}_C$, and taking into consideration the restriction [18], the empirical specification of equation [16] simply becomes:

$$[19] \qquad \Delta w_{Ci} = \mu_1 \ P^{**} - \mu_2(1\text{-}q_i)\theta^i P^{**} + g_1(X_{1i}) + \varepsilon^{*}_{Ci}$$

where:

$$g_1(X_{1i}) = \delta_o + \delta_1 PCU + \delta_2 IPROD_{j-1} + \delta_3 u_j + \delta_4 u_r^{-1} + \delta_4 \Delta P_j$$
$$+ \delta_5 iw_j^{m-1} + \delta_6 EMP_i + \delta_7 \Delta EMP_i + \delta_8 DEL_i + \delta_9 DEL_i^2$$
$$+ \delta_{10} C\_PROD_i + \delta_{11} C\_ABS_i + \delta_{12} XH_{i-1} + \delta_{13} CCOO_i$$
$$+ \delta_{14} INDEP_i + \delta_{15} OTHERS_i + \delta_{16} NOCOLA{\Rightarrow}COLA_i + \sigma_{IC}\hat{\lambda}_{NC}$$

and,

$$PCU = [(1\text{-}\theta^i_{-1})PA_{-1} - \mu^i P^{**}_{-1}] = [(1\text{-}\theta^i_{-1})PA_{-1} - (\mu_1\text{-}\mu_2\theta^i_{-1})P^{**}_{-1}]$$

The description of the variables we are considering in $g_1(.)$ and the description of the PCU variable (considering the restriction [17]) can be found in the previous subsection. As we did for the NON-COLA equation, we opt for introducing a dummy (NO_COLA⇒COLA) that takes the value one if previous year wage increase agreement was not covered by a revision clause to control for the implicit cost of a COLA. We expect to show this variable having a negative effect over the negotiated wage increase. Apart of this, the set of considerations about estimation methods for the non-COLA wage increase equation still apply. Since it is not possible to know for certain $P^{\circ\circ}$, we will consider a range of ad-hoc alternatives for proxying it, which can be found in the Appendix A.

Finally, we would like to discuss something about the wage-price elasticity (θ^i), for which we face a serious observability problem: it is only observable for triggered clauses. In a previous (and related) work, Prescott and Wilton (1992) opted for setting $\theta^i=0$ for non-triggered clauses. However, this introduces measurement error in [19]. Here, we opt for substituting θ^i for its prediction (in a reduced form model) in an attempt to, firstly, avoid the measurement error problem induced by using $\theta^i=0$ instead of its unknown (not realized) true value, θ^i; and secondly, to avoid the consideration of an additional endogenous factor in the COLA wage equation. A detailed description of the θ forecasting process, which is a basic insight of our modelization, can be found in the appendix A.

V. Empirical results.

a. The COLA decision.

The results, including marginal effects, of our preferred specification are reported in Table 5 for the 1985-1991 period. A set of time and industry dummies is included although their coefficients are not reported[27]. The percentage of correct predictions, 78.3, is comparable to other studies' prediction levels[28]. The dominant effects are the lagged COLA which exhibits a large positive and significant parameter (the marginal effect is set around 0.41) and the size of the bargaining unit measured by the number of employees which affects positively to the probability of observing a COLA clause (marginal contribution: 0.07). Although the result is not reported, dropping these two key variables implies a decrease of about 15 points in the percentage of correct predictions.

All the variables that we have used to proxy the structure of the workers committee have the expected coefficient[29]. Nationwide unions, CCOO and USO have no significant difference with respect to the omitted one, UGT. On the contrary, regional (OTHERS) and non_affiliated (INDEP) union variables have both negative and significant coefficients which implies a marginal contribution to the COLA probability of, -0.04 and -0.12, respectively. The finding is in accordance to the fact that both might be considered weak unions when setting a COLA.

Additionally, multiyear contracts imply a significant increase in indexation probability. The marginal effect is set around 0.07. We also found a significant concave effect of the delay of the negotiations on the probability of observing a COLA. The maximum of probability is obtained around 115 days in column (1) (128 in column (2)) which is slightly lower

[27] Available from the author on request.

[28] For instance, in Prescott and Wilton (1992) the percent of correct predictions ranges between 75 and 87 per cent.

[29] Notice we are omitting the UGT's union proportion, so the coefficients might be understood as a difference with respect to the implicit coefficient of the omitted variable.

than the mean of delay in sample (around 120 days).

Neither of the inflation variables is found to be very significant (at 5% significance level). Price expectation influences little in COLA setting (0.02). However, price volatility has a strong impact on indexation probability. Its marginal contribution is found around 0.06. Consequently, the findings offer support for the idea that workers are risk averse (increases in inflation uncertainty make more attractive the COLA clause).

All the industry variables, we considered, were found, as a rule, to be non-relevant. It is surprising the size (0.26) and the sign of the industry unemployment level (U_j). Our suspicion is that this variable characterizes industry demand better than industry productivity or industry prices do. Finally, the regional unemployment level (in the quarter preceding the signing of the contract) has, as expected, a strong and significant negative effect on COLA propensity (marginal contribution: -0.06).

b. The non-COLA wage equation.

The set of results about the non-COLA wage increase equation is reported in Table 6. The same basic specification is reported under two alternatives for the price catch up variable (PCU_A for (1)-(3) and PCU_B for (4)-(5)) and two estimation methods, NLS and IV-NLS (because the presence of a set of potentially endogenous variables: ΔP_j, C_PROD, C_ABS). The overall fit of the model (measured by the R^2) is not very large (around 0.30 in all the cases). We also report a Hausman specification test under the null hypothesis that the NLS provides consistent estimates of the non-COLA wage equation. The null is not rejected in both columns. The selectivity term, λ, is not relevant in neither column. Thus, it is revealing that the COLA decision and the non-COLA contract are separable. In fact, it suggests a sequential bargaining procedure. Agents, first, decide whether or not the contract will have an indexation clause and, second, they set the ex-ante wage given the decision about the COLA. In such circumstances the COLA should not have any effect on the ex-ante wage without it.

Price expectation has a small (0.05) but significant effect on wages, considerably lower than a previous estimation (0.31) by Prescott and Wilton (1992) with data for Canada. The PCU variable shows very little effect on wage increases and is not significant in any of (1) to (4). Our guess is that such a phenomenon is caused by the low time series variability the PCU variable has in sample. In any case, we must point out that the result is in accordance with previous results for Canadian contract data[30]. The last price variable we considered, industry price changes, has no significant effect on wage settlements. Not considering time dummies (not reported) halves the coefficient of price expectation, changes the sign of industry prices and increases the coefficient (up to 0.30) and also the significance of the PCU coefficient. Thus it confirms our guess about the lack of time series variation of the price variables.

Bargaining size reduces the wage settlement. In our opinion, this a direct consequence of the fact that relatively large BU have more complex pay structures and, consequently, their pressure in wage increase settlements is lower. The change in the employment level has, as expected, a strong and significant effect on wage increases. Among the other BU variables, we did not find any significant effect of bargaining clauses (particularly, the absenteeism clause is negligible), although both have, as a contingent clause, the expected negative coefficient. Concerning union variables, it is found that non-nationwide unions (represented by the proportion of independent workers and other unions) add pressure to non-COLA wage settlement although the effect is only significant at 10% level. The extra-hours per worker (XH_{-1}), as a proxy of firm demand, has a positive coefficient although it is not relevant (at 5 per cent of probability level). Delay in bargaining shows the typical concave effect with a maximum at 300 days. In contrast with COLA determination, the maximum is so far away from the mean. Consequently, the longer the delay the higher the wage increase the union is able to achieve.

Finally, among the rest of industry or regional variables we considered in our basic specification, the industry wage increase mean in the month preceding the signing of the contract ($iw_j^{\tau-1}$) has the greater influence (around 0.19 in all the columns). This variable might be proxying

[30]See Christofides et al. (1980) and Prescott and Wilton (1992).

the effect of the available information to the bargaining unit by the time they decide to set the new contract.

In column (3), we include in our basic specification a dummy for those contracts that have a COLA in past year, which in our opinion is proxying the wage premium for renouncing a COLA clause. As it can be shown, it is not found to be significant. Moreover its introduction changes perceptibly several coefficients but especially that of the selection term. In fact, we have an identification problem with these two variables that cannot be easily solved because of the lack of adequate instruments for the variables.

c. The COLA wage increase equation.

For simplicity we opt for presenting in Table 7 (a -NLS- and b -IV_NLS) a single basic specification, including time and dummies, under a couple of alternatives, AP1 and AP2, for proxying our key price variable (P^{**}) and two alternatives to forecast the wage-COLA elasticity, θ^i. The wage-inflation elasticity, $\hat{\theta}_A$, was obtained by forecasting θ for the whole COLA sample given the estimated model of the observed θ_A (Table A.2(1)), defined as follows: $\theta_A=$ (ex_post wage - ex-ante wage)/(inflation - inflation threshold). Alternatively, $\hat{\theta}_B$, which may be interpreted as the ex-post marginal wage-price elasticity, was obtained by forecasting θ for the whole sample given the model of the observed θ_B (Table A.2(2)), defined as follows: θ_B =(ex_post wage - ex_ante wage)/(inflation) -see Appendix A for technical details. In the same way, the incidence of the price variable (P^{**}) is proxied in two different ways. The first, considering the expected inflation level (P^e), which will be named AP1 and the second by using $\hat{P}^{**}$, which will be named AP2. Additionally, in both cases, we include a proxy for the relevant inflation threshold, (PU-dum). The construction of these variables is detailed in Appendix A.

The coeficient of the selectivity variable is significant across of the columns of both tables, which implies that non-random sampling for indexed contracts is an important feature of our model. The same considerations we made at the start of the non-COLA wage increase section about potentially endogenous variables and testing still apply. A formal test of the hypothesis that COLA and non-COLA contract are driven by the same underlining model was carried out using the specification of Table 6(1) and Table 6(3), excluding the selectivity term. The likelihood ratio statistic, which under the null is distributed as a $\chi^2(45)$, is very large (relative to the critical acceptance value, 33.9, at the 5% significance level) in both cases, 314.6 (specification of Table 6(1)) and 311.5 (specification of Table 6(3)) respectively[31]. Consequently, we have no evidence for accepting a common model for wage increase determination in both COLA regimes. We shall discuss first the set of industry and specific bargaining unit push variables, emphasizing the comparison with the previous section main findings (non-COLA wage increase equation). Thereafter, we'll turn our attention to the price variables and COLA provisions effects.

As we found for the non-indexed contracts, both proxies considered for industry activity level, industry productivity ($IPROD_{j-1}$) and unemployment rate (U_j), are found either insignificant or exhibiting the wrong sign. On the contrary, the industry prices variables (ΔP_j and iw_j^{m-1}) show a sign in accordance with expectations. Industry price change (ΔP_j) influences positively COLA wage increases across all the estimates, though is found relevant in neither NLS nor instrumental variables estimates. Likewise the effect of iw_j^m, the proxy for other bargaining pairs actions, as in the non-COLA case, is always estimated around 0.20, except in the case of the AP2 with is found sensibly higher (0.32). Consequently, we might assess that there is not much difference in the effect of industry variables (overall in price variables) between COLA and non-COLA contracts.

Concerning specific bargaining unit variables the findings are different from those we pointed out in the non-COLA section. The size of the bargaining unit (measured by the number of employees in it, EMP) shows also a negative effect, but significantly smaller. On the contrary, the effect of the change in the employment level is much bigger in size (around 0.50 in all the cases) and significance. Delay in negotiation shows the usual concave function. The estimated coefficients are larger than in the non-COLA

[31] Letting time and industry dummies be different across indexed and non-indexed contracts also gives the same result. In such case the test is distributed as a χ^2_{18}. The statistics are 230.5 (specification in Table 6(1)) and 230.9 (Table 6(3)), respectively.

case, although the maximum effect implies a lower delay (around 190 days).

Both bargaining clauses considered, productivity and absenteeism, have the expected negative coefficient, though they are not found to be significant. Opposite to the non-indexed wage increases case, for indexed contracts it is found that the absenteeism clause has a comparatively greater effect. Thus, there is a divergence in the effect of additional contingent clauses among COLA and NON-COLA contracts. It seems that fixed payments clauses have greater incidence on indexed contracts and variable payments clauses on non-indexed contracts. The effect of last year extra hours (XH_{-1}) also affects positively negotiated COLA wage increase, although the estimated coefficient is perceptibly smaller (roughly a third) than in the non-COLA case. Neither of union's variables has a significant effect (except the proportion of other representatives in Table 7.a(3)). Thus, we can conclude that unions affect the COLA propensity but have no clear effect in the wage increase given by the firm.

In column (2) of Table 7.a and Table 7.b we have included (with respect to column (1) of both tables) a dummy (NOCOLA⇒COLA) trying to capture the implicit cost of a new revision clause. The wage premium workers must pay, in terms of ex-ante wage, for obtaining a COLA clause, i.e., the cost of a new COLA, has been found with the expected negative sign. However, as we observed for non-indexed contracts, the consideration of the above dummy affects dramatically (changing its sign) the selection term. Moreover, when considering instrumental variables estimates (Table 7.b(2)), both are not significant, evidencing some identification problems. Comparing indexed and non-indexed contracts, there is some evidence in favor of the fact that premium is much more relevant for the former. At the end of this section we will turn back to indexation costs through the analysis of implicit wage differentials.

Finally, we turn our attention to discuss findings on price variables. All the price variables considered, under both price variable alternatives (AP1 and AP2) show the expected sign and are significant (except in a very few cases). Notice that our method of considering an overidentified reduced form forecast for the wage-price elasticity (θ) instead of using its observed value avoids the (potential) simultaneity problem between θ and the

COLA wage increase. Consequently, this key variable does not need to be instrumented.

Under the θ_A (the one we consider to be the most appropriate for the Spanish case) both μ_1 and μ_2 (for both price alternatives: AP1 and AP2) are rather small. Only μ_2 is estimated once (for AP2) higher than 0.25. As a consequence the estimated response of ex-ante wage with respect to expected price variable (μ^i) is, in all the cases considered, very small, a low of 0.07 in Table 7.a(1) and a high 0.09 in Table 7.a(3) (see Table 8 for a summary of findings). On the contrary, estimated ex-ante expected COLA wage elasticity, γ^i, is sensibly higher around 0.35 (when using θ_A). In the case where we use θ_B the estimate for γ is perceptibly lower, 0.11. This difference might be surprising, but we must take into account the implicit definition of θ_B, which is clearly a point dependent (of inflation rate) measure. Particularly in the Spanish case, we consider this approach to measuring wage-COLA elasticity, used in other studies, not to be very adequate.

Note that in the case where we compute γ^i under the assumption $\theta_j=1$ ($j=A,B$) the estimate is always around 0.60, not much lower than Prescott and Wilton (1992) sample means reported estimates. This means (in the Spanish case) that ex-ante expected COLA elasticity (γ^i) is given, in a large share, through COLA elasticity (θ^i), that is, contingent compensation, instead of through non-contingent compensation (μ^i). Notice the fact that the null $\mu_1=\mu_2=\mu^\circ<1$ is not rejected. Accepting such a restriction has, at least, two implications. On the one hand, under it, $\mu^i=\mu(1-(1-q_i)\theta^i)$, which implies that "workers purchase a unit of COLA coverage (expressed as a proportion of the expected rate of inflation) by giving up less than an equivalent amount in noncontingent wages increases" (Prescott and Wilton (1992), page 345). On the other hand, notice that in this case γ^i simplifies:

$$[18r] \qquad \gamma^i= \mu^\circ[1-(1-q_i)\theta^i]+ (1-q_i)\theta^i = \mu^\circ+ (1-\mu^\circ)(1-q_i)\theta^i$$

It is straightforward to show that 1-q moves towards one as the threshold moves towards 0 (or towards a minimum relevant price level). Consequently, the lower the PU the higher is γ^i and so, workers are, in the absence of high PU cost, better off choosing a lower PU°, i.e. inflation target.

Finally, the findings about the effect of the proxy for PU° and the price catch-up variable are quite satisfactory. The proxy for PU° has a strong effect, around 0.28 in all the cases. Note this variable, which we were not able to identify in the non-COLA equation (because the time dummies) could be considered as a part of the wage-price elasticity. Adding the effect of the proxy for PU° to the estimated range for the ex-ante wage-price elasticity (γ^i) we obtain a range of 0.63 to 0.90 using θ_A (0.37-0.87 using θ_B), comparable to the Prescott and Wilton previous estimated range. On the other hand, the PCU variable has a small but significant effect on ex-ante wage increase. The estimate ranges from a low 0.06 in column (1) of Table 7.a to a high of 0.12 in column (4) of the same table. In contrast with Prescott and Wilton (1992) we did not found any significant interaction effect between the PCU variable and the proxies for the wage-price elasticity.

d. Ex-ante wage increase differentials.

The set of results we have obtained on ex-ante wage increase setting for indexed and non-indexed contracts will permit us to draw some conclusions about the implicit ex-ante wage increase differentials among both indexation regimes for twenty-two industries. To compute these differentials we apply the methodology that can be found in Stengos and Swidinsky (1990). As they did, we report both differentials, corrected (considering the selection term) and uncorrected[32] (not considering the

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[32]The corrected differential (CD_l, where $l=A,B$ index the prediction for θ) could be expressed as:

$$CD_l = (1/M).\Sigma_{i=1}^{N}\Sigma_{t=1}^{T} \left\{(\Delta\hat{w}_{it}^{\bar{c}l}/I_{it}=0)-(\Delta\hat{w}_{it}^{cl}/I_{it}=1)\right\};\ l=A,B$$

where M is the number of observations, $\Delta\hat{w}_{it}^{cl}$ is the prediction of the COLA model and $\Delta\hat{w}_{it}^{\bar{c}l}$ is the prediction of the non-COLA model, both considering the selection induced by the observed indexation variable (I_{it}). On the other hand, the uncorrected differential is defined as:

$$UD_l = (1/M).\Sigma_{i=1}^{N}\Sigma_{t=1}^{T} \left\{\Delta\hat{w}_{it}^{\bar{c}l}-\Delta\hat{w}_{it}^{cl}\right\};\ l=A,B$$

where both predictions do not consider the selection terms. Consequently, the difference between CD and UD could be expressed as:

selection term) for 22 industries[33]. For comparison purposes, we also report the sample means ex-ante wage differentials. The results of our experiment are reported in Table 9.

The first point to note is that, though there are sensibly differences in the uncorrected case, corrected differentials are estimated rather similar under θ_A and θ_B. Our estimated sample means wage differential is 0.295 wage increase points (3.5% in relative terms) when using θ_A and 0.27 points when using θ_B (2.6% in relative terms). Note the estimated differentials increase with the wage-price elasticity (θ) and are larger in the latest years (1988-1991) of our sample period. The findings are robust to several wage increase equation specifications[34]. To our knowledge there is no previous work estimating ex-ante wage increase differentials for indexed and non-indexed contracts with which to contrast our results.

By big sectors (1-digit SIC classification), our findings also suggest workers pay a positive premium to obtain an indexation clause. However, there are some important differences. Whereas in the Energy sector the implicit wage differentials are practically negligible (0.06 wage increase points when using θ_A and 0.01 when using θ_B), in the Minerals and Chemical they are much bigger (0.43 and 0.41 basis points, respectively).

The most surprising findings are found when looking at industry level (2-digit SIC classification). As it is shown in Table 9, there are several industries for which the premium is negative and, in some cases, extremely large. This is the case of the Mineral Oil Refining (between -12 and -16 percentage points), the Electronic Engineering and the Leather industries. On the contrary, the biggest premia are given in the Non-metallic minerals and the Paper, Printing and Publishing industries.

According with the theory, risk averse workers ought to be willing to accept a lower expected real wage for getting a COLA clause from the firm. It is far beyond the scope of the paper to analyze in deep ex-post wage

$$CD_l - UD_l = \sigma_{CI}\, \hat{\lambda}_{wc} - \sigma_{NCI}\, \hat{\lambda}_{w\bar{c}}$$

[33] Two digits Standard Industrial Code (SIC) classification.

[34] We have make several exercises which confirm our result. For instance, constraining the COLA wage equation to the same specification of the non-COLA wage equation implies a wage differential of 4.2 per cent.

increase differentials but adding the estimated wage differential with the sample mean of the realization of the contingent compensation for COLA contract we are able to obtain an approximate measure of ex-post wage differentials. As it can be seen in Table 9 there are very few industries (2 out of 22) for which the wage increase premium fully compensates the realized contingent compensation (column (6) of Table 9). On the contrary, for most of the industries the premium is much lower than the realized contingent compensation. Our finding for Spain is in contrast with a previous work estimating ex-post wage level differentials by Hendricks and Kahn (1985). They set such the ex-post cost or premium in a range of 1.5 to 2 per cent[35].

Consequently, our empirical evidence (though it has the inconvenience of a very short time series sample period) rejects the standard theory (Shavell (1976)) that COLA contracts imply lower expected ex-post wages. Unfortunately, we are not able to offer strong support for any alternative theory. As we have commented in the COLA decision section there is some evidence supporting our theoretical guess that non-nationwide unions have trouble getting the clause. However, there are other possibilities. The most evident one is that unions worry not only about wages but also about employment level. In such circumstance, the workers could renounce the COLA clause in order to preserve the employment level.

[35]Their estimated range was 1.5 to 22 per cent. However, they consider the larger figure as unlikely.

VI. Concluding remarks.

This work develops and estimates a joint model of wage and revision clause setting in a uncertainty context. Assuming that unions are more risk averse than firms, the non-optimality of the escalator contract may be explained by the presence of weak union in negotiation and also by the existence of an alternative mechanism (ex-post catch-up) to link wages to prices. The set of specific assumptions taken leads to a very simple framework, which leads to a switching model of wage increases under COLA and without it, which was estimated using the well-known Heckman's method for selectivity models. Our estimates suggest that non-random sampling is a salient feature of our model but only for COLA contracts. However, we use a very restrictive set of assumptions about the error structure. As a previous step to the estimation of such a model, we deal with the observability problem of the wage-price elasticity variable by means of using an unconditional forecast for it to avoid simultaneity problems. The empirical part of the study was carried out using Spanish collective bargaining data for the manufacturing sector in the 1984-1991 period.

The estimated probability of observing an escalator contract is found to be higher (5 to 10 %) for nationwide unions than for others unions. Additionally, there is some evidence supporting the fact that non-nationwide unions get worse conditions (lower probability that the clause will be triggered and lower wage-price elasticity) when bargaining the escalator contract. On the other hand, there are no relevant differences in wage setting behaviour (in either COLA or non-COLA equations) between nationwide and other unions. Thus, there is no significant evidence against our initial assessment that regional unions or independent workers' representatives may have distinct bargaining power in wage setting with COLA and without it. Heterogeneity in bargaining power could also be supported by the asymmetry found between the wage premium workers must pay for getting a new COLA and what they receive if they renounce to it. Whilst the first is negligible, the second was found to be significantly positive. However, we shall note that we face some identification problems for this variable.

Both wage increase equations suggest that wage settlements are poorly related to industry performance and market variables (basically unemployment). In practice, much of the predictive power of our model comes from the variables trying to proxy the available information, at the time of signing the agreement, the BU variables and the set of time and industry dummies. Both alternatives to proxy the wage-price elasticity and the price variable performed quite well. However, in our opinion, the most adequate for representing the Spanish case is θ_A.

Ex-ante inflation coverage is estimated in a range of 0.04 to 0.08, depending on the proxy for the wage-price elasticity. However, note that an important share of the ex-ante inflation coverage is captured by the set of time dummies. Our estimate (only identifiable in the COLA equation) suggests this share could be in a range of 0.26-0.28. Hence, a reasonable ex-ante expected inflation coverage for indexed contracts should be in a range of 0.56-0.99. Consequently, inflation coverage in Spain is mostly given on ex-post basis. The catch-up variable has been found non-relevant for non-indexed contracts and having a small but significant coefficient in the complementary subsample. However, this could be explained by the low time series variation this variable has in sample. In this sense, not considering time dummies raises significantly the coefficient (up to 0.30) of the catch-up variable in both subsamples.

Using the set of estimates for the indexed and non-indexed ex-ante wage increase equations we have estimated the implicit wage premium workers must pay for obtaining a COLA clause. Such a premium has been set in a range of 2.6 to 3.5 per cent. By sectors, we found the premium is negligible for the Energy sector and very high for the Minerals and Chemical sector (around 5 per cent). However, we observe that the premium that workers pay to get the indexation clause does not compensate in our sample period (1984-1991) the mean contingent compensation, in contrast with a previous estimate for the US by Hendricks and Kahn (1985).

This observation has several consequences. Firstly, it maybe the case that ex-post wage increases for COLA contracts are greater than for non-COLA contracts. In such a context the employment level should be negatively related to the degree of indexation and, in general, to the COLA clause. Consequently, we cannot offer support to the assumption that workers ought to be able to accept a lower expected wage to obtain the COLA clause. Naturally, we think such a situation is not sustainable for the Spanish economy in the long run. Note that the eighties were transition years for the Spanish economy and, consequently, this apparent contradiction of the theory may be transitory. Our impression is that in the forthcoming years the apparent contradiction may disappear because of the better adjustment of inflation to its expected target. This impression gains some support from our finding that the premium has been increased in the 1988-1991 period.

Table 1. Manufacturing firm level agreements (sample).1983-1991.

year	COLA clause present	#	% COLA	ex ante wage	ex post wage	% triggered COLA	mean hold-out	mean emp
1984	No COLA	352	--	7.5	--	--	0.93	366
	COLA	162	31.5	7.7	7.9	0.22	0.96	748
1985	No COLA	331	--	7.4	--	--	0.90	291
	COLA	294	47.0	7.2	8.0	0.77	0.94	660
1986	No COLA	363	--	8.2	--	--	0.77	340
	COLA	374	50.7	8.1	8.4	0.65	0.70	605
1987	No COLA	489	--	6.9	--	--	0.90	304
	COLA	380	43.7	6.4	6.4	0.01	0.87	654
1988	No COLA	570	--	5.7	--	--	0.75	321
	COLA	368	39.2	5.0	6.5	0.80	0.61	651
1989	No COLA	423	--	6.7	--	--	0.74	315
	COLA	463	52.2	6.3	8.0	0.84	0.75	572
1990	No COLA	330	--	8.3	--	--	0.78	249
	COLA	444	57.1	7.5	8.1	0.50	0.60	580
1991	No COLA	284	--	7.8	--	--	0.77	213
	COLA	371	56.6	7.3	7.5	0.31	0.69	572

SOURCE: "Estadística de Convenios Colectivos". 1984-1991.

Table 2. Mean and variance of ex-ante and ex-post agreements. 1984-1991.

	without COLA ex-ante			non-triggered COLA ex-ante			triggered COLA ex-ante			ex-post			
YEAR	#	Δw	$\sigma_{\Delta w}$	#	Δw	$\sigma_{\Delta w}$	#	Δw	$\sigma_{\Delta w}$	Δw	θ_A	θ_B	PU=T
1984	352	7.53	2.48	126	7.73	1.51	36	7.65	0.61	8.44	0.58	0.09	0.75
1985	331	7.42	2.21	67	7.51	0.99	227	7.15	0.42	8.20	0.96	0.13	0.71
1986	363	8.23	2.48	132	8.48	0.85	242	7.93	0.64	8.40	0.90	0.06	0.49
1987	489	6.90	3.38	376	6.45	1.02	4	5.12	1.06	5.26	0.22	0.03	--
1988	570	5.70	2.86	73	5.98	1.16	295	4.71	1.33	6.66	0.85	0.33	0.65
1989	423	6.73	3.56	73	7.22	2.22	390	6.09	2.72	8.16	0.74	0.33	0.58
1990	330	8.30	2.72	221	8.00	0.88	223	7.01	2.22	8.17	0.83	0.18	0.52
1991	284	7.83	1.81	256	7.65	1.17	115	6.58	1.16	7.12	0.84	0.10	0.79

Keys:

θ_A= (ex_post wage - ex_ante wage)/(inflation -inflation threshold)

θ_B= (ex_post wage - ex_ante wage)/(inflation) -i.e., marganal elasticity-

PU=T ⇒ Inflation threshold equals government's inflation target.

SOURCE: See below Table 1.

Table 3. Agreements by duration, COLA and delay in sample (6884 obs).

Agreement observed in its..	Number	COLA	% delay>0	delay
1 year	3895	0.41	0.96	101
2 years	2364	0.58	0.64	-14
3 years	543	0.56	0.45	-111
4 years	65	0.52	0.37	-272
5+ years	17	0.29	0.29	-177

SOURCE: See below Table 1.

Table 4. Conditional COLA AND NON-COLA sample probabilities after k periods of doing the action.

# of previous years doing the same action	# agre.	p(COLA)	# agre.	p(NON-COLA)
cond. to 1	2303	79.9	2638	76.4
cond. to 2	1392	83.6	1619	81.2
cond. to 3	823	87.4	986	82.5
cond. to 4	475	89.7	530	84.3
cond. to 5	247	89.9	272	85.3
cond. to 6	114	91.2	127	86.6
cond. to 7	34	100.0	45	95.6
unconditional	2856	47.6	3142	52.4

SOURCE: See below Table 1.

Table 5. The COLA clause decision. 1985-1991.

Variable	PROBIT (1) coef.	t-stat	Marginal contribution
CONSTANT	-1.59	(2.99)	-0.441
$COLA_{-1}$	1.48	(34.6)	0.412
DELY	-0.54	(0.64)	-0.015
CCOO	-0.06	(0.72)	-0.015
USO	0.02	(0.11)	0.005
OTHERS	-0.15	(2.01)	-0.042
INDEP	-0.43	(4.15)	-0.119
EMP	0.14	(8.46)	0.071
IDEL	0.35	(4.22)	0.097
IDEL*IDEL	-0.12	(5.16)	-0.032
MULTIYEAR	0.24	(4.96)	0.067
U_j	0.26	(1.36)	0.071
u_r	-0.22	(2.46)	-0.061
P_j	-0.02	(0.04)	-0.005
$IPROD_j$	-0.31	(1.04)	-0.085
$IPROD_{j-1}$	0.29	(0.86)	0.085
$STRIKE_j$	-0.01	(0.40)	-0.003
P^e	0.02	(1.72)	0.007
UN_INF_{-1}	.003	(0.31)	0.001
σ_P	0.22	(1.81)	0.060
Time_dummies	Yes		--
Industry_dummies	Yes		--
Time_Span	1985-1991		--
Observations	4941		--
Cola>0	2461		--
Log_L	-2441.6		--
%Correct Prediction	78.3		--

Table 6. The non-COLA wage increase equation. 1985-1991.

Est. method	(1) NLS coef. t-st.	(2) 2S-NLS (IV) coef. t-st.	(3) NLS coef. t-st.	(4) NLS coef. t-st.	(5) 2S-NLS (IV) coef. t-st.
Constant	3.10 (2.17)	3.11 (2.18)	3.15 (2.20)	3.14 (1.87)	3.19 (1.89)
P_e (μ)	0.05 (1.93)	0.04 (1.90)	0.05 (2.02)	0.05 (1.95)	0.05 (1.92)
PCU (θ_A)	0.02 (0.91)	0.02 (0.93)	0.02 (0.97)	----	----
PCU (θ_B)	----	----	--	0.02 (0.22)	0.03 (0.19)
$IPROD_{j-1}$	0.51 (1.21)	0.56 (1.25)	0.54 (1.27)	0.52 (1.24)	0.58 (1.30)
U_j	-0.13 (0.42)	-0.11 (0.34)	-0.10 (0.31)	-0.11 (0.37)	-0.09 (0.27)
IP_j-$IP^{\ddagger}_{j-1}$	-0.18 (0.17)	-0.78 (0.33)	-0.22 (0.20)	-0.19 (0.17)	-1.00 (0.42)
iw_j^{m-1}	0.25 (2.26)	0.25 (2.27)	0.24 (2.19)	0.24 (2.24)	0.25 (2.25)
u_r^{-1}	0.01 (0.43)	0.01 (0.52)	0.02 (0.66)	0.01 (0.47)	0.02 (0.56)
EMP	-0.09 (2.56)	-0.08 (2.37)	-0.06 (1.40)	-0.09 (2.46)	-0.08 (2.25)
ΔEMP	0.34 (2.42)	0.33 (2.37)	0.34 (2.42)	0.34 (2.41)	0.33 (2.36)
DEL	0.33 (4.04)	0.32 (4.00)	0.35 (4.11)	0.33 (4.07)	0.33 (4.02)
DEL∘DEL	-0.05 (2.52)	-0.05 (2.47)	-0.06 (2.62)	-0.05 (2.58)	-0.05 (2.53)
C_PROD‡	-0.09 (1.15)	-0.13 (1.36)	-0.08 (1.07)	-0.09 (1.12)	-0.13 (1.34)
C_ABS‡	0.01 (0.16)	-0.05 (0.44)	0.01 (0.15)	0.02 (0.18)	-0.05 (0.43)
XH_{-1}	0.05 (1.92)	0.05 (1.92)	0.05 (1.89)	0.05 (1.89)	0.05 (1.89)
CCOO	0.10 (0.84)	0.10 (0.82)	0.09 (0.77)	0.11 (0.86)	0.10 (0.84)
INDEP	0.30 (1.95)	0.28 (1.84)	0.23 (1.37)	0.29 (1.91)	0.27 (1.80)
OTHERS	0.20 (1.64)	0.19 (1.59)	0.17 (1.36)	0.20 (1.64)	0.19 (1.58)
COLA⇒NOCOLA	--	--	0.35 (0.88)	--	--
$\hat{\lambda}_C$	0.07 (0.56)	0.06 (0.53)	0.42 (1.01)	0.12 (1.08)	0.12 (1.09)
Time dum.(6)	Yes	Yes	Yes	Yes	Yes
Ind dum.(22)	Yes	Yes	Yes	Yes	Yes
Obs	2119	2119	2119	2119	2119
Log_L	-3917.6	----	-3917.2	-3918.0	----
R^2	0.305	0.304	0.305	0.305	0.304
σ	1.55	1.55	1.55	1.55	1.55
Haussman(DF)	----	3.01 (38)	--	----	2.63 (39)

‡: Instrumented variables (by using lags) in both, columns (2) and (4).

Note: t-statistics have been obtained from sample covariance of θ.

Table 7.a. The COLA wage increase equation for a sample of Spanish' manufacturing firms. NLS estimates. 1985-1991.

	(1)	(2)	(3)	(4)
θ pred.from⇒	T_A.2.(1):$\hat{\theta}_A$	T_A.2.(1):$\hat{\theta}_A$	T_A.2.(1):$\hat{\theta}_A$	T_A.2.(2):$\hat{\theta}_B$
Price alt:	**AP1**	**AP1**	**AP2**	**AP1**
	coef.t-st.	coef. t-st.	coef.t-st.	coef.t-st.
Constant	1.80 (1.66)	1.54 (1.41)	1.31 (1.21)	1.69 (1.33)
AP1:P^e (μ_1)	0.11 (4.64)	0.10 (4.11)	---	0.11 (5.87)
AP2:P_1^{**} (μ_1)	---	---	0.23 (4.96)	--
AP1:$\hat{\theta}P^e$ (μ_2)	0.16 (4.49)	0.13 (3.29)	---	1.20 (9.94)
AP2:$\hat{\theta}P_1^{**}$(μ_2)	---	---	0.53 (4.96)	--
AP1 :PU-dum	0.28 (9.08)	0.28 (9.00)	---	0.26 (8.34)
PCU	0.06 (4.94)	0.06 (5.18)	0.07 (4.22)	0.11 (1.49)
$IPROD_{j-1}$	-0.01 (0.05)	0.01 (0.04)	-0.05 (0.17)	0.06 (0.22)
U_j	-0.09 (0.38)	0.01 (0.05)	-0.06 (0.29)	-0.07 (0.32)
IP_j-IP_{J-1}	0.82 (1.22)	1.02 (1.08)	0.98 (1.45)	0.76 (1.15)
iw_j^{m-1}	0.21 (2.25)	0.20 (2.10)	0.27 (3.11)	0.16 (1.73)
u_r^{-1}	0.01 (0.55)	0.02 (1.08)	0.01 (0.71)	.003 (0.19)
EMP	-0.04 (1.66)	0.01 (0.14)	-0.05 (2.39)	-0.02 (1.12)
ΔEMP	0.50 (3.87)	0.51 (3.97)	0.50 (3.93)	0.48 (3.80)
DEL	0.67 (8.24)	0.78 (8.13)	0.80 (10.9)	0.50 (6.33)
DEL•DEL	-0.17 (6.24)	-0.21 (6.38)	-0.20 (7.70)	-0.14 (5.36)
C_PROD	-0.07 (1.22)	-0.06 (1.13)	-0.06 (1.09)	-0.04 (0.69)
C_ABS	-0.09 (1.45)	-0.09 (1.46)	-0.11 (1.95)	-0.09 (1.47)
XH_{-1}	0.02 (1.88)	0.02 (1.80)	0.02 (1.69)	0.02 (1.27)
CCOO	0.13 (1.46)	0.11 (1.24)	0.16 (1.69)	0.17 (1.85)
INDEP	-0.07 (0.52)	-0.22 (1.38)	-0.02 (0.01)	-0.06 (0.41)
OTHERS	0.08 (0.91)	0.05 (0.49)	0.15 (1.65)	0.05 (0.51)
NOCOLA⇒COLA	--	-0.61 (2.06)	--	--
$\hat{\lambda}_C$	-0.28 (3.57)	0.40 (1.19)	-0.22 (2.96)	-0.13 (1.90)
Time dum(6)	Yes	Yes	Yes	Yes
Ind.dum(22)	Yes	Yes	Yes	Yes
Obs.	2182	2182	2182	2182
Log_L	-3323.0	-3319.8	-3322.0	-3288.4
R^2	0.486	0.487	0.486	0.502
σ	1.12	1.12	1.12	1.10
$\mu_1=\mu_2$ (χ_1^2)	3.54	0.97	24.6	--
$\mu_2=1$ (χ_1^2)	--	--	--	2.815

Table 7.b. The COLA wage increase equation for a sample of Spanish' manufacturing firms. Instrumental variables NLS estimates. 1985-1991.

	(1)	(2)	(3)	(4)
θ pred.from⇒	T_A.2.(1):$\hat{\theta}_A$	T_A.2.(1):$\hat{\theta}_A$	T_A.2.(1):$\hat{\theta}_A$	T_A.2.(2):$\hat{\theta}_B$
Price alter:	AP1	AP1	AP2	AP1
	coef. t-st.	coef. t-st.	coef. t-st.	coef. t-st.
Constant	1.83 (1.66)	0.95 (0.63)	1.27 (1.16)	1.66 (1.29)
AP1:P^e (μ_1)	0.11 (4.56)	0.07 (1.56)	---	0.11 (5.82)
AP2:P_1^{ee} (μ_1)	---	---	0.23 (4.94)	--
AP1:$\hat{\theta}P^e$ (μ_2)	0.16 (4.41)	0.05 (0.38)	---	1.19 (9.83)
AP2:$\hat{\theta}P_1^{ee}$ (μ_2)	---	---	0.53 (4.94)	--
AP1 :PU-dum	0.28 (9.06)	0.27 (8.28)	---	0.26 (8.32)
PCU	0.06 (4.94)	0.06 (5.44)	0.07 (4.20)	0.12 (1.54)
$IPROD_{j-1}$	-0.01 (0.02)	0.05 (0.19)	-0.08 (0.29)	0.05 (0.18)
U_j	-0.08 (0.35)	0.21 (0.50)	-0.10 (0.41)	-0.08 (0.35)
IP_j-IP_{J-1}†	0.79 (0.55)	0.87 (0.60)	1.54 (1.08)	0.93 (0.66)
iw_j^{m-1}	0.21 (2.23)	0.16 (1.52)	0.27 (3.07)	0.16 (1.72)
u_r^{-1}	0.01 (0.54)	0.05 (1.00)	0.01 (0.72)	.003 (0.18)
EMP	-0.04 (1.63)	0.10 (0.62)	-0.05 (2.30)	-0.02 (1.07)
ΔEMP	0.50 (3.87)	0.54 (3.88)	0.51 (3.95)	0.48 (3.81)
DEL	0.68 (8.22)	1.01 (2.51)	0.80 (10.9)	0.50 (6.34)
DEL∘DEL	-0.17 (6.23)	-0.28 (2.04)	-0.20 (7.68)	-0.14 (5.37)
C_PROD†	-0.07 (0.97)	-0.05 (0.74)	-0.05 (0.77)	-0.04 (0.64)
C_ABS†	-0.12 (1.53)	-0.12 (1.61)	-0.15 (1.99)	-0.11 (1.49)
XH_{-1}	0.02 (1.90)	0.02 (1.54)	0.02 (1.97)	0.02 (1.29)
CCOO	0.13 (1.43)	0.07 (0.57)	0.15 (1.67)	0.17 (1.83)
INDEP	-0.07 (0.52)	-0.54 (0.95)	-0.01 (0.02)	-0.06 (0.41)
OTHERS	0.09 (0.92)	-0.04 (0.23)	0.15 (1.67)	0.05 (0.53)
NOCOLA⇒COLA†	--	-2.01 (0.85)	--	--
$\hat{\lambda}_C$	-0.28 (3.58)	1.97 (0.75)	-0.23 (2.98)	-0.08 (1.95)
Time dum.(6)	Yes	Yes	Yes	Yes
Ind.dum.(22)	Yes	Yes	Yes	Yes
Obs	2182	2182	2182	2182
R^2	0.486	0.482	0.486	0.502
σ	1.12	1.13	1.12	1.10
Hausman(df)	0.75(40)	1.21 (41)	1.27(44)	22.11(10)
$\mu_1=\mu_2$ (χ_1^2)	3.45	0.79	26.4	---
$\mu_2=1$ (χ_1^2)	--	--	--	2.597

†: Intrumented variables (by using lags of the variables).

Table 8. Findings about ex-ante price elasticity.

		price elasticity			
price variable		$\hat{\theta}_A$ NLS	$\hat{\theta}_A$ IV-NLS	$\hat{\theta}_B$ NLS	$\hat{\theta}_B$ IV-NLS
NON-INDEXED	μ^i	0.05	0.04	0.05	0.05
INDEXED: AP1	μ^i	0.07	0.07	0.04	0.04
	γ^i	0.35	0.35	0.11	0.11
	$\gamma^i(\theta_j=1)$	0.62	0.62	0.59	0.59
INDEXED: AP2	μ^i	0.08	0.08	--	--
	γ^i	0.36	0.36	--	--
	$\gamma^i(\theta_j=1)$	0.63	0.63	--	--

Table 9. Estimated wage increase differentials by industry.

Sector	CD_A	UD_A	CD_B	UD_B	mean $\Delta w_{\bar{c}}-\Delta w_c$	mean $\Delta w_c^{ep}-\Delta w_c$
ALL (Sample means)	0.29	0.11	0.28	0.27	0.278	0.756
ALL (1988-1991)	0.45	0.28	0.36	0.35	0.350	0.950
ALL(Low θ)†	0.23	--	0.01	--	0.278	0.756
ALL(High θ)‡	0.38	--	0.44	--	0.278	0.756
.. Energy	0.06	-0.07	0.00	0.03	0.044	0.962
1.Extractives	0.24	0.03	0.18	0.15	-0.045	0.839
2.Mineral Oil Refining	-0.62	-0.71	-0.78	-0.72	-0.769	1.151
3.Utilities	0.22	0.16	0.16	0.25	0.361	0.976
.. Minerals and Chemical	0.43	0.23	0.41	0.38	0.428	0.852
4.Extraction of Metalic Ores	0.32	0.13	0.35	0.34	0.634	0.779
5.Iron and Steel	0.34	0.16	0.29	0.29	0.820	1.221
6.Ext. Non-metalic Minerals	0.68	0.48	0.63	0.60	0.619	0.866
7.Chemical industry	0.25	0.10	0.23	0.25	0.110	0.810
.. Metal Processing	0.27	0.09	0.26	0.25	0.214	0.724
8.Manuf. of Metal Products	0.39	0.19	0.43	0.42	0.291	0.611
9.Machinery and mech en'ring	0.11	-0.02	0.09	0.12	0.159	0.784
10.Electrical engineering	0.38	0.23	0.35	0.37	0.198	0.777
11.Electronic engineering	-0.29	-0.39	-0.42	-0.36	-0.425	0.889
12.Motor vehicles	0.22	0.11	0.22	0.27	0.128	0.659
13.Other Transport Equipment	0.33	0.21	0.16	0.20	0.444	1.004
14.Instrument Engineering	-0.17	-0.49	-0.22	-0.32	-0.620	0.617
.. Other Manufacturing ind.	0.30	0.04	0.30	0.23	0.157	0.611
15.Food, Drink and Tobacco	0.16	-0.10	0.17	0.11	0.041	0.562
16.Textile industry	-0.04	-0.35	0.00	-0.10	-0.149	0.644
17.Leather industry	-0.25	-0.61	-0.26	-0.39	-0.926	0.344
18.Footwear and Clothing	0.30	0.01	0.42	0.33	0.054	0.364
19.Timber Cork & Wooden Fur.	0.38	0.10	0.44	0.37	0.062	0.408
20.Paper, Printing an Pub.	0.60	0.42	0.55	0.54	0.607	0.779
21.Rubber and Plastic .	0.20	0.04	0.15	0.16	-0.010	0.637
22.Other Manufacturing sec	0.23	-0.03	0.39	0.32	0.103	0.491

†: We use quartile 1 ⇒ $\theta_A=0.369$, $\theta_B=0.030$;

‡: We use quartile 3 ⇒ $\theta_A=0.693$, $\theta_B=0.153$;

KEYS:

CD_j: Corrected differential under θ_j, j=A,B.

UD_j: Uncorrected differential under θ_j, j=A,B.

mean $\Delta w_{\bar{c}}-\Delta w_c$: Sample means difference between ex-ante non-COLA and COLA wage increases (in percentage points).

mean $\Delta w_c^{ep}-\Delta w_c$: Sample means difference between ex-ante COLA and ex-post COLA wage increase (in percentage points).

Appendix A. Proxying the price expectation (P^{**}) and the wage-price elasticity (θ^{l}) forecast.

The most important inconvenience of a COLA contract is induced by its intrinsic nature. As far as it is a contingent contract it is difficult to know its provisions. In fact neither in Spain nor in other countries there is exact information about the provisions of the contract unless it becomes triggered, in which case ex-post wage increase also becomes observable. Having information about ex-post wage increase it is possible to infer an approximate measure of one of the contract provisions, the wage-price elasticity. However, the problem persists since there is a share of contracts for which we know nothing about the implicit wage-price elasticity (for those that the clause is not triggered). We will attempt to avoid this lack of information by means of a simple modelization of the COLA provisions. Additionally, we will take advantage of this modelization to construct a proxy for the driven price variable of the contingent wage increase, P^{**}, which is given in equation [8]. Finally, at the end of the appendix we present a brief description, for illustrative purposes, of the most common contingent clauses in Spain in recent years.

a. The wage-price elasticity forecasting process.

Assume there is an underline linear reduced form model for the log of the wage-price elasticity ($\ln\theta^{*}$) and inflation threshold (PU^{*}) determination.

[a.1] $\ln\theta^{*} = X_{\theta}\delta_{\theta} + \varepsilon_{\theta}$

[a.2] $PU^{*} = -X_{PU}\delta_{PU} + \varepsilon_{PU}$

where X_{θ} and X_{PU} are the (assumed) exogenous vectors of variables affecting, respectively, $\ln\theta^{*}$ and PU^{*}; δ_{θ} and δ_{PU} are vectors of parameters and, finally, ε_{θ} and ε_{PU} are error terms normally distributed with covariance:

$$cov(\varepsilon_\theta,\varepsilon_{PU}) = \begin{bmatrix} \sigma^2_\theta & \\ \sigma_{\theta,PU} & \sigma^2_{PU} \end{bmatrix}$$

On the other hand, the inflation threshold PU° is not observable. We observe θ° if the threshold (PU°) is lower than inflation rate (P, a random variable). Notice that our model is similar to the wage and participation or hours model (see García (1991) for a description), so the estimation technique will be exactly the same. First we will estimate a probit model for the probability that PU°≤P. Rewriting [a.2] as:

$$[a.2'] \qquad P\text{-}PU^\circ = X_{PU}\delta_{PU} + P\text{-}\varepsilon_{PU}$$

the probability of observing θ° (p(Y=1)) is given by:

$$[a.3] \qquad Prob(TR=1) = prob(P\text{-}PU^\circ > 0) = \Phi(\frac{X'_{PU}\delta_{PU}}{\sigma_{P\text{-}\varepsilon_{PU}}})$$

where TR takes one if the clause is triggered (zero otherwise) and Φ is the distribution function of the standard normal. Estimates of equation [a.1] are reported in Table A.1. Given the estimates of the first stage Probit model, consistent estimates of the unknown parameters of equation [a.1] might be found by applying, in a second stage, LS to the following extended equation (for taking into account the selectivity bias expecting to arise in such a model) in the subsample for which we observe lnθ° (i.e., P>PU°):

$$[a.4] \qquad \ln\theta^\circ = f_\theta(X_\theta,\delta_\theta) + \sigma_{1\theta}\lambda(\frac{X'_{PU}\delta_{PU}}{\sigma_{(\varepsilon_{PU}\text{-}\varepsilon_P)}}) + \varepsilon_\theta$$

where $\lambda = \frac{\phi(.)}{\Phi(.)}$ is the well-known inverse of the Mill's ratio. Once we have estimates for the parameters of equation [a.1] we turn our attention to state the correct method for forecasting correctly the wage-price elasticity, θ°, for the whole sample. Given the fact that the COLA wage increase (Δw_{ic}), the optimal wage-price elasticity and the inflation threshold are jointly determined in the same maximization process, we will follow, an instrumental approach. The purpose of this approach is to obtain a "good" instrument for the wage-price elasticity in the wage increase equation by means of computing an unconditional forecast for all the COLA contracts (without taking into account $\hat{\lambda}$, which may be endogenous). The potential usefulness of such a method is evident, because we are solving simultaneously the unobservability of the wage-price elasticity and we are controlling for its potential endogeneity in the wage equation (see García (1991) for details[36]). Consequently, the forecast method is given by:

[a.5] $\qquad \widehat{\ln\theta} = X_\theta \hat{\delta}_\theta \qquad$ for all the COLA contracts

and consequently:

$$\hat{\theta} = \exp(\widehat{\ln\theta})$$

Unfortunately, even in the set of triggered contracts it is not possible to know for certain the implicit wage-price elasticity of the contract. To proxy it we tried two different alternatives, θ_A and θ_B, defined as follows,

θ_A= (ex_post Δw - ex_ante Δw)/(inflation rate - inflation threshold)

θ_B=(ex_post Δw - ex_ante Δw)/(inflation rate)

From the definition, it is straightforward to show that $\theta_A \geq \theta_B$. In previous work, mostly done for Canadian contract data, the usual proxy was θ_B (see Christofides et al (1980) and Prescott and Wilton (1992)). But in Spain, indexation clauses often include an inflation threshold, hence, we think that θ_A might be a better proxy. Unfortunately, for building θ_A we need to know the implicit inflation threshold, usually unknown (except for the AES-like clauses). We solve this additional problem by using a search method for the inflation threshold in the sample of 1375 triggered clauses. Given the fact we have two different proxies, θ_A and θ_B, for the wage-price

[36]This work is an application of the Nelson and Olson (1978) general method to a two equation model of hours and wages.

elasticity, we estimate and forecast equation [a.4] under these two alternatives which are also considered in the main model for comparative purposes.

b. Proxying $P^{\circ\circ}$.

As we stated in section III the relevant price variable for indexed contracts ($P^{\circ\circ}$) can be defined as follows:

$$P^{\circ\circ}=q.E_P[\tfrac{1}{P}/P\leq PU]+(1-q).E_P\{\frac{(1+\theta^{\circ}(P-PU))}{P}/P>PU\}$$

That is, $P^{\circ\circ}$ is a weighted (by q, the probability that inflation rate is lower than a given threshold) highly nonlinear unknown function of the expected inverse low price and the sum of the expected inverse high price, the wage-inflation elasticity and the inflation threshold. Given the fact that is it rather impossible to know for certain such a variable we opt for by breaking it into pieces and making use of our set of estimates for equations [a.1] and [a.2]:

AP1: variable 1: P^e

variable 2: PU-dum=(1-TR)$\circ$Target+TR$\circ$P

AP2: variable 1: $\hat{P}^{\circ\circ}=\hat{q}\frac{1}{P^e-\sigma_p\hat{\lambda}}+(1-\hat{q})\frac{(1+\hat{\theta}^{\circ}(P-Target))}{P^e+\sigma_p\hat{\lambda}}$

variable 2: PU-dum=(1-TR)$\circ$Target+TR$\circ$P

where $\hat{q}$ = 1 - prob(TR=1)

P^e being the expectation of the inflation rate at the signing of the contract, σ_p is the inflation rate standard deviation in the 5 years period preceding the signing of the contract; Target is government's beginning of the year inflation rate target; the rest of variables are defined as above.

c. The commonly used indexation clauses in Spain.

The most typical clauses for setting ex-post wage increases (Δwr) are:

1. $\Delta wr - \Delta w = \theta_1(P-PU)$ iff $P > PU$

where Δw, is the ex-ante wage; θ_1 is the wage-inflation elasticity; P is the inflation rate (normally December to December) and, finally, PU is a given inflation rate threshold.

2. (AES:85-86) $\frac{\Delta wr}{\Delta w} = \frac{P}{PU}$ iff $P > PU$

which is equivalent to:

$$\Delta wr - \Delta w = \theta_2(P-PU) \text{ ; and } \theta_2 = \frac{\Delta w}{PU} \qquad \text{iff } P > PU$$

obs: if $\Delta w \lesseqgtr PU \Rightarrow \theta_2 \lesseqgtr 1$

Notice this clause might imply a wage-price elasticity greater than unity, so, in case of generalization, could be extremely inflacionist. In fact, this kind of clause induced inflation pressure on the Spanish economy during those years.

3.(1989 onwards):

$$\Delta wr = P + K \qquad \text{without any ceiling}$$

which is equivalent to:

$$\Delta wr - \Delta w = \theta_3(P-PU)$$

where $\theta_3 = 1$; $K = iwr - P$; $P - PU = iw - K$

The typical COLA clause in Spain is type 1 (about 37 % of the clause were of this type in 1990-91), although there is also a substantial share of clauses of type 3 (currently about 20 %), although in 1985-1986 the most common was the AES-like clause (in 1985-1986 above 75 % of all the cola contract included it), B type. Since then the share of this special clause has been reduced to less than 5 %.

Table A.1. The triggered COLA clause probit. 1984-1991.

Variable	(1) coef	t-stat
Constant	-1.31	(1.16)
DELY	-0.11	(0.74)
EMP	0.08	(3.27)
CCOO	-0.05	(0.43)
USO	-0.38	(1.51)
OTHERS	-0.36	(3.02)
INDEP	-0.48	(2.67)
$COLA_{-1}$	0.15	(2.06)
RENEGOTIATION	0.51	(3.92)
MULTIYEAR	0.21	(2.65)
DEL*DELY	-0.05	(0.36)
DEL^2*DELY	0.02	(0.39)
$IPROD_j$	-0.35	(0.80)
$IPROD_{j-1}$	0.68	(1.34)
ΔIP_j	-0.01	(0.67)
S_j	-0.04	(0.96)
u_j	-0.12	(0.45)
u_r	-0.11	(0.78)
P^e	-0.12	(0.58)
σ_P	0.41	(2.41)
$UNEXP\_INF_{-1}$*DELY	0.01	(0.91)
Time_dummies	Yes	
Industry_dummies	Yes	
Obs	2461	
Cola>0	1380	
Log_L	-1098.4	
%_Correct Pred.	0.792	

Table A.2. The COLA elasticity model. 1984-1991.

Dependent	(1) log θ_A coef.	t-stat	(2) log θ_B coef.	t-stat
CONSTANT	0.50	(0.13)	-2.01	(0.40)
DELY	-0.07	(0.77)	-0.11	(1.39)
EMP	.003	(0.33)	.003	(0.21)
EMP-EMP$_{-1}$	-0.12	(1.70)	-0.06	(0.63)
CCOO	0.02	(0.42)	0.05	(0.71)
USO	0.05	(0.43)	-0.14	(1.00)
OTHERS	-0.09	(1.74)	-0.20	(2.79)
INDEP	-0.02	(0.30)	-0.05	(0.47)
RENEGOTIATION	0.29	(5.12)	0.45	(6.02)
XH$_{-1}$	0.01	(1.69)	.002	(0.14)
MULTIYEAR	0.17	(4.70)	0.19	(3.90)
C_ABS	0.02	(0.69)	0.04	(1.02)
C_PROD	0.02	(0.64)	0.05	(1.43)
RH$_{-1}$	-0.13	(0.26)	-0.16	(0.24)
DEL*DELY	0.01	(0.33)	-0.07	(0.91)
DEL^2*DELY	.005	(0.34)	0.01	(0.58)
IPROD$_j$	-0.12	(0.77)	0.02	(0.10)
IPROD$_{j-1}$	-0.08	(0.42)	0.08	(0.31)
ΔP_j	-0.36	(1.04)	-0.01	(1.69)
S_j	-0.03	(1.64)	-0.03	(1.03)
P^e	.003	(0.35)	-0.02	(1.37)
σ_P	0.36	(4.86)	0.49	(5.03)
UNEXPECTED INF$_{-1}$*DELY	-.001	(0.17)	0.01	(1.25)
$\hat{\lambda}$ (Table a.1)	0.51	(3.79)	0.62	(3.49)
Time_dummies	Yes		Yes	
Industry_dummies(22)	Yes		Yes	
Obs	1375		1375	
Estimation Method	OLSQ		OLSQ	
R^2	0.19		0.60	
σ	0.44		0.591	

Appendix B. Data and Variables.

All the collective agreements in Spain have to be registered in order to be enforceable. As there are some information requirements we know a small set of basic variables for each bargaining unit. The number of collective agreements in the raw manufacturing firm level dataset (the ECC) is very large (14777). From it we have obtained an unbalanced panel of negotiation units in the 1981-1991 period. In Table B.1 we describe the resulting sample. There we distinguish between the sample resulting from considering consecutive observations (CS), which correspond to the sample used in this chapter [37], and the general sample (GS), which is a 32 per cent larger than the first. The former represents the 76 per cent of the later. Note that any large firm is more likely to be followed across time than any small one. The (corrected) probability of exiting the CS sample[38] ranges from a high of 29 to a low of 17 per cent.

The sample used in estimation result from constraining the CS sample to more than four consecutive observations (because we need some lagged information). The resulting sample is an unbalanced panel of 1290 (6884 observations) manufacturing firms running from 1981 to 1991. The 1981-1983 period is not used in estimation because it has some shortcomings in information. As a result, the sample has 4941 observations. For the sake of simplicity we only present the definition of the variables used here, although the data questionnaire is available on request. In Table B.2 it can be found some basic statistics for the key variables in each subsample.

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[37] Constrained to four or more observations.

[38] $p(\text{exit after } i \text{ obs}) = \frac{\#\text{BU with } i \text{ obs}}{\sum_{j=i}^{11} \#\text{ BU with } j \text{ obs}} * (1-q)$. Where q is the probability the BU was in the last sample period.

Definition of the variables.

Bargaining unit variables: (source: ECC)

ΔW_i: Ex-ante wage increase settlement.

ΔWR_i: Ex-post wage increase (only for effectively revised contract).

JP_i: Annual working hours settlement.

EMP_i: Membership, i.e. number of employees at the settlement date.

XH_{-1}: Number of overtime hours during last year.

RH: Yearly number of regular hours.

PUB_i: 1 if the bargaining unit belongs to the public sector. 0 otherwise.

$MULTIYEAR_i$: 1 if the agreement will last for more than a calendar year.

BY_i: 1 if bargaining finishes before expiratory date of last one.

DEL_i: Mean delay (in days) from the expiratory date of the last agreement until the settlement.

$CCOO_i$: Percentage of workers council that represents the CCOO union.

UGT_i: Percentage of workers council that represents the UGT union. (omitted)

USO_i: Percentage of workers council that represents the USO union.

$INDEP_i$: Percentage of workers council that does no represent a union.

$OTHERS_i$: Percentage of workers council that represents other unions.

$COLA_i$: 1 if agreed any cost of living allowance clause.

C_PROD_i: 1 if agreed any productivity clause.

C_ABS_i: 1 if agreed any union presence clause.

θ_A= (ex_post wage - ex_ante wage)/ (Inflation rate-PU[39])

θ_B= (ex_post wage - ex_ante wage)/ Inflation rate

$DELY_i$: 1 if current contract was signed after the date of effectiveness.

$RENEGOTIATION_i$: 1 if agreement is a renegotiation of a multiyear contract.

$NOCOLA{\Rightarrow}COLA_i$: 1 if in the last year there was no COLA and currently there is.

$COLA{\Rightarrow}NOCOLA_i$: 1 if in the last year there was COLA and currently there is not.

TR_i: 1 if the cola clause is triggered.

[39]PU is not reported on the basic data set. But given the characteristics of most of the Spanish COLA contracts it is possible to obtain it by looking at the wage increase data, inflation rate and target.

Industry and regional variables:

u_r: Regional unemployment rate in the quarter preceding the signing of the contract (17 regions). (source: EPA)

u_j: Industry unemployment rate (annual). (source: EPA)

S_j: Number of days lost by strike per employee in the industry. (source: EH)

$\Delta\bar{w}_j^m$: Monthly mean wage increase at the j industry. (source: ECC)

IO_j: Industry Output Index (1972=100, for 22 industries). (source: BE)

IE_j: Employment level in the j industry. (source: EPA)

IP_j: Industry Price Index (1976=100, 22 industries). (source: BE)

$IProd_j$: IO_j/IE_j

Inflation variables:

Target: Inflation rate target at the beginning of the year. (source: BE).

P: Inflation rate (December to December) of the year. (source: BE).

σ_P: Inflation rate standard deviation. It was proxied by using 5 years monthly data standard deviation. (source: BE)

P^e: Weak inflation rate expectation in the month preeceding the signing of the contract. Forecast was made by means an ARIMA model with ten years monthly data. Thus, our forecast is based on data available to agents at the time of signing the contract.

UNEXPECTED_INF: $P-P^e$

PCU: Price catch-up.

Data sources:

-Banco de España: *Boletín Estadístico* (BE).

-Ministerio de Trabajo:

Estadística de Convenios Colectivos (ECC). Recording Tape. 1981-1990.

Estadítica de Huelgas y Cierres Patronales (EH). Recording Tape. 1986-1990.

-Instituto Nacional de Estadística: *Encuesta de Población Activa* (EPA). Several issues.

Table B.1. Resulting sample after matching the ECC.

	GS sample # BU	mean emp.	CS sample # BU	mean emp.	Exit prob. (corrected)
Unmatched	2779	313.8	--	--	--
2 obs.	1128	339.1	745	284.4	29.1
3 obs.	707	460.1	504	324.8	28.7
4 obs.	314	390.3	249	381.6	20.7
5 obs.	291	417.7	238	344.5	25.2
6 obs.	155	499.9	134	448.9	19.6
7 obs.	115	531.8	105	492.3	19.3
8 obs.	79	202.3	77	203.6	17.5
9 obs.	98	696.3	98	695.4	25.7
10 obs.	55	713.7	55	713.7	17.8
11obs.	111	588.3	101	588.2	--

Table B.2. Some descriptive statistics.

	all the sample		cola sample		non-cola sample	
OBS:	4941		2182		2119	
	mean	stdev	mean	stdev	mean	stdev
<u>BARGAINING UNIT VARIABLES</u>						
Δw	6.9385	1.7191	6.8585	1.5473	7.1360	1.8442
COLA	0.4980	0.5000	0.0000	0.0000	0.0000	0.0000
θ_A	0.2355	0.4053	0.4622	0.4638	0.0000	0.0000
$\hat{\theta}_A$	0.2596	0.3144	0.5165	0.2412	0.0000	0.0000
θ_B	0.0598	0.1317	0.1145	0.1600	0.0000	0.0000
$\hat{\theta}_B$	0.0549	0.0945	0.1071	0.1042	0.0000	0.0000
q	0.0000	0.0000	0.4471	0.3097	0.0000	0.0000
TR	0.2793	0.4487	0.5527	0.4973	0.0000	0.0000
λ_C	0.0000	0.6939	0.5428	0.4133	-0.5597	0.4192
emp(logs)	5.0792	1.3313	5.3969	1.3450	4.7891	1.2348
(DEL/100)•DELY	0.9018	0.8718	0.8685	0.8139	0.9765	0.9265
$(DEL/100)^2$•DELY	1.5732	2.8947	1.4165	2.2703	1.8117	3.4673
DELY	0.7464	0.4351	0.7218	0.4482	0.7942	0.4043
NOCOLA⇒COLA	0.0000	0.0000	0.2442	0.4297	0.0000	0.0000
COLA⇒NOCOLA	0.0000	0.0000	0.0000	0.0000	0.1887	0.3914
C_PROD	0.3734	0.4837	0.4028	0.4905	0.3407	0.4740
C_ABS	0.2679	0.4429	0.3056	0.4608	0.2373	0.4255
XH1	0.5048	1.8170	0.6909	2.2317	0.3513	1.2987
RH(-1)	7.4908	0.0318	7.4890	0.0315	7.4919	0.0315
CCOO	0.3300	0.3350	0.3505	0.3160	0.3163	0.3526
INDEP	0.1095	0.2445	0.0874	0.1983	0.1321	0.2807
OTHERS	0.2224	0.3492	0.1975	0.3254	0.2431	0.3676
RENEGOTIATION	0.2608	0.4391	0.2882	0.4530	0.2085	0.4064
MULTIYEAR	0.0350	0.1838	0.0453	0.2081	0.0292	0.1685
Extractives	0.0305	0.1721	0.0265	0.1608	0.0339	0.1812
Mineral Oil Refining	0.0085	0.0918	0.0100	0.0999	0.0051	0.0718
Utilities	0.0686	0.2528	0.0976	0.2968	0.0391	0.1940
Metallic Ores	0.0348	0.1833	0.0348	0.1833	0.0344	0.1824
Iron and Steel	0.0170	0.1292	0.0164	0.1274	0.0165	0.1274
Non-Metallic Minerals	0.0813	0.2734	0.0756	0.2644	0.0887	0.2844
Chemical industry	0.0933	0.2908	0.1063	0.3083	0.0877	0.2830
Manuf. of Metals	0.1396	0.3466	0.1356	0.3425	0.1477	0.3549
Machinery and Mech.	0.0378	0.1908	0.0430	0.2030	0.0311	0.1737
Electrical Eng'ring	0.0487	0.2154	0.0513	0.2207	0.0467	0.2110
Electronic Eng'ring	0.0208	0.1428	0.0256	0.1581	0.0151	0.1219
Motor Vehicles	0.0580	0.2339	0.0705	0.2561	0.0467	0.2110
Other Transport Eq.	0.0315	0.1748	0.0375	0.1902	0.0268	0.1618
Instrument Eng'ring	0.0046	0.0680	0.0032	0.0565	0.0066	0.0810

Table B.2. (CONT)

	all the sample		cola sample		non-cola sample	
OBS:	4941		2182		2119	
	mean	stdev	mean	stdev	mean	stdev
Food, Drink and Tob.	0.1428	0.3499	0.1040	0.3053	0.1680	0.3739
Textile	0.0072	0.0850	0.0041	0.0641	0.0103	0.1013
Leather	0.0080	0.0896	0.0041	0.0641	0.0117	0.1080
Footwear and Cloth.	0.0064	0.0802	0.0041	0.0641	0.0094	0.0967
Timber Cork, Wooden	0.0261	0.1594	0.0187	0.1358	0.0349	0.1836
Paper, Printing & Pu	0.0649	0.2464	0.0632	0.2434	0.0703	0.2557
Rubber and Plastic	0.0538	0.2257	0.0563	0.2306	0.0500	0.2180
Other Manufacturing	0.0147	0.1206	0.0105	0.1021	0.0184	0.1344
INDUSTRY AND REGIONAL VARIABLES						
Δw_j^{m-1}	7.1527	0.9012	7.1659	0.9007	7.1511	0.9026
P_j	6.1073	0.1980	6.1245	0.2088	6.0996	0.1878
P_{j-1}	6.0741	0.1920	6.0916	0.1996	6.0674	0.1840
$IPROD_j$	0.0733	0.7518	0.1398	0.7316	0.0056	0.7566
$IPROD_{j-1}$	0.0596	0.7517	0.1281	0.7311	-0.008	0.7560
u_j	-2.072	0.5650	-2.161	0.5674	-2.010	0.5506
u_r	-1.488	0.6272	-1.456	0.6993	-1.472	0.6224
S_j	0.5379	1.2447	0.5230	1.1916	0.5724	1.3366
PRICE VARIABLES						
P^e	5.3147	1.9970	5.2389	1.9141	5.2890	1.9626
UNEXPECTED INF(-1)	2.1641	2.9670	2.0127	2.7123	1.7652	2.8753
σ_P	2.1825	0.4205	2.1320	0.4269	2.1824	0.4338

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