BARGAINING ABOUT WAGES:
EVIDENCE FROM SPAIN

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simply becomes:

$$[4.1'] \qquad E_P V(\pi) = E_P\left\{\eta_j y_i - \frac{w_i}{P}\right\}$$

The union seeks the maximization of its expected utility, which is assumed to depend solely on wages and to have the following specific functional form[91]:

$$[4.2] \qquad E_P U = E_P\left\{\frac{1}{m}\cdot\left(\frac{w_i - a}{P}\right)^m\right\}$$

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[92]:

$$[4.3] \qquad L_w = [E_P\{U(w/P)\}]^{\beta}.[E_p\{V(\pi)\}]^{1-\beta}$$

where β is the union bargaining power without a COLA clause. For a given β

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assumption see Ehrenberg et al (1983).

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

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

and assuming the firm is risk neutral[93] the solution to the above problem, the ex-ante wage without COLA, w^*, might be written as follows,

$$[4.4] \qquad w^* = \frac{m\beta\eta_i y_i + (1-\beta)a}{[m\beta + (1-\beta)]P^*}$$

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:

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

Normally, in Spain and other countries, the contingent wage contract takes the following explicit form[94]:

$$[4.5] \qquad w_C = \begin{cases} w_L & \text{if } P \leq PU \\ w_H = w_L + \theta(P-PU) = w_L(1+\theta^*(P-PU)) & \text{if } P > PU \end{cases}$$

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:

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

[94] 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.

$$[4.6] \qquad L_{W_L} = \left\{E_P\{U(\frac{W_L-a}{P}) \;/\; P \le PU\}q + E_P\{U(\frac{W_H-a}{P}) \;/\; P > PU\}(1-q)\right\}^{\beta_C}$$

$$\left\{E_P\{(\eta_j y_i - \frac{W_L}{P}) \;/\; P \le PU\}q + E_P\{(\eta_j y_i - \frac{W_H}{P}) \;/\; P > PU\}(1-q)\right\}^{1-\beta_C}$$

where $q = prob(P \le PU)$ 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[95]. Consequently, the solution to the unrestricted problem, can be written down as,

$$[4.7] \qquad w_L^*(\theta^*, PU) = \frac{m\beta_C \eta_j y_i + (1-\beta_C)a}{[m\beta_C + (1-\beta_C)]P^{**}}$$

where:

$$[4.8] \qquad P^{**}(\theta^*, PU) = q.E_P[\frac{1}{P}/P \le PU] + (1-q).E_P\{\frac{(1+\theta^*(P-PU))}{P}/P > PU\}$$

That is, P^{**} 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 $dw_L^*/d\theta^* < 0$ 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 [4.6]. However, in the case $\beta_C > \beta$, expression [4.6] 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

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[95] 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.

restrictions will make the algebra tedious, we will omit it by means of parameterizing the optimal solution of the restricted version of [4.6]. 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:

$$[4.7r] \qquad w_L^*(\theta^*,PU,k)=\frac{m\beta_C\eta_j y_i+(1-\beta_C)a}{[m\beta_C+(1-\beta_C)]P^{**}}.k \text{ for some } k\leq 1 \text{ if } \beta \text{ is close } \beta_C$$

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[96].

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.

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

[96]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.

possibility of a proportional[97] (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]

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[97] 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[98]. 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[99] 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.

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[98]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).

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

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:

$$[4.9] \qquad E_P\left\{V[\pi(w_L^*(\beta_C).k)]\right\} \geq E_P\left\{V[\pi(w^*(\beta))]\right\}$$

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

$$[4.7'] \qquad w_L^*(k) = \frac{m\beta_C\eta_j y_i + (1-\beta_C)a}{[m\beta_C + (1-\beta_C)]P^{**}}.k \qquad \text{for some } k \leq 1$$

where k $\lesseqgtr$ 1 iff $\beta \lesseqgtr \beta_C$. On the contrary, if [4.9] does not hold, there is no revision clause and the ex-ante wage will be given by:

$$[4.4] \qquad w^* = \frac{m\beta\eta_i y_i + (1-\beta)a}{[m\beta + (1-\beta)]P^*}$$

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, [4.9] 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 [4.9]) and the wage equation is related to [4.4] -for non indexed contracts- and [4.5]-[4.7'] -for indexed contracts. Formally:

$$[4.10] \qquad I = \begin{cases} 1 \text{ if } I^* > 0; \\ O \text{ otherwise} \end{cases}$$

$$[4.10'] \qquad I^* = Z'\varphi + \varepsilon_I$$

$$[4.11] \qquad \Delta w_{NC} = h(X_{NC}, \alpha_{NC}) + \varepsilon_{NC} \quad \text{if } I=0 \text{ (non COLA)}$$

$$[4.12] \qquad \Delta w_C^e = g(X_C, \alpha_C) + \varepsilon_C \quad \text{if } I=1 \text{ (COLA)}$$

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[100]; 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

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

[4.10]-[4.12], which are assumed to be serially uncorrelated jointly normally distributed with covariance matrix:

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

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 [4.10]-[4.12] 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)[101]. 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[102] (i.e., strike activity).

[101]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.

[102]Unfortunately, we have no information about strike activity at firm level

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 4.4, 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 [4.10']:

$$[4.10'] \quad I^* = Z^{*\prime}\varphi^* + \tau I_{-1} + \varepsilon_I$$

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 ($IPROD_{j-1}$)

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in our sample. So, the results in this respect will be merely approximate.

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[103] (P^e_m), 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[104]. 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[105]. 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

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

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

[105] 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.

(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 [4.11] 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 [4.11] and [4.10]. Provided some consistent estimation of λ, the Mill's inverse ratio, say $\hat{\lambda}_{NC}$, the empirical specification is given by:

$$[4.13] \quad \Delta w_{NCi} = \alpha_o + \mu P^e_m + \alpha_1 PCU + \alpha_2 PROD_{j-1} + \alpha_3 u_j + \alpha_4 u_r^{-1} + \alpha_5 \Delta P_j + \alpha_6 iw_j^{m-1} + \alpha_7 EMP_i + \alpha_8 \Delta EMP_i + \alpha_9 DEL_i + \alpha_{10} DEL_i^2 + \alpha_{11} C\_PROD_i + \alpha_{12} C\_ABS_i + \alpha_{13} XH_{i-1} + \alpha_{14} CCOO_i + \alpha_{15} INDEP_i + \alpha_{16} OTHERS_i + \alpha_{17}[COLA \Rightarrow NO\_COLA] + \sigma_{INC}{*}\hat{\lambda}_{NC} + \varepsilon^{*}_{NCi}$$

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[106]:

$$[4.14] \qquad PCU = (1-\theta^{i}_{-1})(P_{-1}-\mu P^{e}_{m-1})$$

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

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

the above specification, and assuming ε^{*}_{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 [4.13] 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 [4.13].

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[107]. 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,

$$[4.15] \qquad \Delta w^{e}_{Ci} = \Delta w_{Ci} + (1-q_i)[\theta^i(E_m(P/P>PU^i)-PU^i)] + \varepsilon_{Ci}$$

where Δw^{e}_{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 [4.15] we have an expression for the ex-ante negotiated wage increase (Δw_{Ci}):

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107 The implicit cost of an indexation clause has not been considered.

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

where P^{**} is defined as in [4.8] 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:

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

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 [4.16] 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 wage increase (Δw^e_{ic}), which is given by:

[4.18] $\gamma^i = \mu^i + (1-q_i)\theta^i$

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 [4.18], the empirical specification of equation [4.16] simply becomes:

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

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,

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

The description of the variables we are considering in $g_1(.)$ and the description of the PCU variable (considering the restriction [4.17]) 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^{**}, 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 [4.19]. 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 4.5 for the 1985-1991 period. A set of time and industry dummies is included although their coefficients are not reported[108]. The percentage of correct predictions, 78.3, is comparable to other studies' prediction levels[109]. 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[110]. 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

[108] Available from the author on request.

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

[110] 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.

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 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 increase equation.

The set of results about the non-COLA wage increase equation is reported in Table 4.6. 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[111]. 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

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

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^{m-1}) has the greater influence (around 0.19 in all the columns). This variable might be proxying 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 4.7 (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 4.A.2(1)), defined as follows: θ_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 4.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 4.6(1) and Table 4.6(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 4.6(1)) and 311.5 (specification of Table 4.6(3)) respectively[112]. 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

[112]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 4.6(1)) and 230.9 (Table 4.6(3)), respectively.

relevant in neither NLS nor instrumental variables estimates. Likewise the effect of iw_j^{rp}, 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 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 4.7.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 4.7.a and Table 4.7.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 4.7.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 4.7.a(1) and a high 0.09 in Table 4.7.a(3) (see Table 4.8 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 $\mu_1=\mu_2=\mu^*<1$ 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:

$$[4.18r] \qquad \gamma^i = \mu^*[1-(1-q_i)\theta^i] + (1-q_i)\theta^i = \mu^* + (1-\mu^*)(1-q_i)\theta^i$$

It is straightforward to show that 1-q moves towards one as the