NBER WORKING PAPER SERIES

THE EFFECT OF UNIONS ON EMPLOYMENT: EVIDENCE FROM AN UNNATURAL EXPERIMENT IN URUGUAY

Adriana Cassoni
Steven G. Allen
Gaston J. Labadie

Working Paper 7501
http://www.nber.org/papers/w7501

NATIONAL BUREAU OF ECONOMIC RESEARCH
1050 Massachusetts Avenue
Cambridge, MA 02138
January 2000

This research was supported by the Interamerican Development Bank. We have received helpful comments and suggestions from Carmen Pages, James Heckman, Daniel Hamermesh, and participants at IDB-sponsored workshops in Washington, DC, and Lima, Peru. The views expressed herein are those of the author and not necessarily those of the National Bureau of Economic Research.

The Effect of Unions on Employment:
Evidence from an Unnatural Experiment in Uruguay
Adriana Cassoni, Steven G. Allen, and Gaston J. Labadie
NBER Working Paper No. 7501
January 2000
JEL No. J5

ABSTRACT

This study examines the impact of unions on wages and employment using data from Uruguay in a period where unions were banned (1973-1984), then legalized with tripartite bargaining (1984-1991) followed by industry-wide or firm-specific bargaining (1992-1997). The relationship between wages and employment shifted significantly across these periods as evidenced by

- Recursive residuals show structural shifts in five of six industries with the shifts coming at the same time as the regime changes.
- Wages are exogenous to employment before 1985, but not afterwards.
- The wage elasticity and the employment-output elasticity fell sharply after 1984.
- Unions significantly raised wages in 1985-1992, but afterwards the change in bargaining structure and increased openness led to concessions.
- Starting in 1985, workers in unionized industries were less likely to be laid off than workers in nonunion industries.

Adriana Cassoni
Department of Economics
Universidad de la Republica
Montevideo, Uruguay
adrianac@decon.edu.uy

Steven G. Allen
College of Management
NC State University
Raleigh, NC 27695-7229
and NBER
steve_allen@ncsu.edu

Gaston J. Labadie
Dean, School of Business
Universidad ORT Uruguay
Montevideo, Uruguay
labadie@athenea.ort.edu.uy

I. Introduction

The subject of how unions affect employment adjustment generates strong opinions. The prevailing view among many economists and policy analysts is that unions prevent labor market forces from operating effectively. Unions take a hard line in bargaining that prevents wages from falling, no matter how high unemployment has gotten. They resist attempts by management to streamline production and introduce new technology. They stand in the way of team-based production by clinging to outdated job descriptions and occupational jurisdictions. They insist on advance notice and severance pay arrangements that make it extremely costly to reduce employment.

Au contraire shout union supporters. Centralized negotiations provide a framework for wage adjustments to take place more rapidly than they would in a world where all bargaining is one-on-one. Unions see the handwriting of technological change on the wall as clearly as management, and also see that management does not think about implementation of new technology in the workplace until installation time. Joint committees provide a framework to make changes more productive by getting full input from employees on how to redesign jobs and processes. Rules on job security admittedly make downsizing more difficult, but other parts of union agreements make labor markets more effective by encouraging long term employment relationships and investments in firm-specific skills.

In Latin America the prevailing wisdom is that the former view is closest to the truth. Even though most markets have been liberalized, the labor market has been what Sebastian Edwards, the former chief economist for Latin America and the Caribbean for at the World Bank, calls "the forgotten sector." Welfare losses come from three main sources: (1) wages set above market clearing levels, (2) lost output and wages from strikes, and (3) rent-seeking activities such as support for protectionism and state ownership of industry. Edwards (1995, p. 286) argues:

> Reforming legislation governing labor-management relations in Latin America is an important unfinished part of the recent structural reforms. In general, a modern and flexible legislation is characterized by incentives to resolve conflicts quickly and fairly. This requires clear rules of the game, modern institutions, an efficient judiciary, and a system in which both parties incur costs if the conflict becomes protracted. In most countries, however, the current situation is far from that: there are asymmetric costs for unions and employers that, in fact, do not penalize delaying agreements.

Given these very strong views, one would think that there would be a massive research literature on how unions affect employment adjustment to changes in wages and output. Think again. Globally this subject has received little attention and in Latin America it has received virtually no attention. In the case of Uruguay, some theoretical work has been developed by Rama (1993a, 1993b, 1994) while there is some recent empirical research (Allen, Cassoni and Labadie, 1994; 1996; Cassoni, Labadie and Allen, 1995). Although there are numerous studies making union-nonunion comparisons for particular countries at particular time periods, they have generally concentrated on wage gains and wage gaps (Blanchflower, 1984; Freeman and Medoff, 1984; Hirsch and Addison, 1986; Lewis, 1986; 1990;) while employment differentials have been somehow neglected[1]. Regarding elasticities of substitution between labor and capital and among different types of labor, research has been even less prolific. In the US, it has been found that they are much lower in union than nonunion establishments (Freeman and Medoff 1982; Allen 1986). Further, Boal and Pencavel (1994) found some evidence suggesting the underlying

[1] An extensive survey can be found in Pencavel, 1991 and Booth, 1995.

production function is different depending on whether the sector is unionized or not. In the UK, Blanchflower, Millward and Oswald (1991) analyzed the impact of unionism on the path of employment growth, finding significant differences, although their result has been criticized for not being robust (Machin and Wadhwani, 1991). Another line of research that has been followed is that related to the influence of unions on the costs of adjusting the level of employment (Burgess, 1988; Burgess, 1989; Burgess and Dolado, 1989; Lockwood and Manning, 1989 are examples). Finally, it has been found that workers in the US have, relative to nonunion workers, lower quit rates (Freeman, 1980), higher layoff rates in the private sector (Medoff, 1979) and lower layoff rates in the public sector (Allen, 1988).

Although all the above papers do illuminate one component or another of the effects of unions on wages and/or employment, they do not address the bottom line questions. For the same establishment or individual with and without union status, does employment adjustment to changes in wages and output vary when the firm is unionized and when it is not? How long does it take to complete the adjustment in these two settings?

This paper examines these issues directly, using evidence from manufacturing industries in Uruguay from 1975 through 1997. Uruguay is well suited for such a study because the economy has gone through a series of regime changes. A military government took over in 1973 and stayed in power through 1984. During and after this regime there were significant changes in labor and trade policy that will allow us to identify the impact of these policies on labor demand parameters.

Collective bargaining was proscribed during the military regime. Labor unions regained the right to bargain collectively with the return of democracy in 1985. As part of its anti-inflation policy, the national government played a significant role in negotiations. Since then, legal regulations of work -which constitute public-order individual rights and therefore cannot be resigned under any circumstance- can be superseded by collective agreements. They can go beyond these restrictions, increasing (but not decreasing) the benefits that workers have in the area of minimum wages, working conditions, job security, and employee benefits. Tripartite negotiations took place at the industry level through Wage Councils, allowing wage adjustment to vary by industry. If an agreement met the government's anti-inflation targets, then it would apply to all firms – even those with nonunion work forces -- in the industry once the agreement was officially endorsed.

The government stopped participating in this system in 1991. Some bargaining is still conducted through industry-wide Wage Councils, but increasingly it is being done at the company level. As a result there are three different bargaining regimes that can be examined in this study: before 1985 when bargaining was banned, 1985-91 when there was tripartite bargaining, and 1992 to the present when the government did not participate in bargaining.

Although the primary focus of this study is on the impact of these regime changes, the role of changes in trade policy cannot be ignored because their timing is correlated with the changes in the structure of collective bargaining. Taxes on traditional exports were substantially cut in the mid-1970s and some initial steps were taken to lowering import barriers, steps which halted in the global recession of 1982. After a temporary increase in 1985, tariffs were gradually decreased starting in 1986. By the end of 1993, the highest tariff was 20 percent. At the same time there were reductions in non-tariff barriers, elimination of some sectoral privileges, and reductions of export subsidies. Accompanying these unilateral policy changes was the creation of the Southern Cone Common Market (MERCOSUR) in 1991. By 1995 a great number of Uruguayan products could be exchanged with MERCOSUR members (Argentina, Brazil, and

Paraguay) without any tariff. As a consequence, there was a sizable increase in the volume of both exports and imports which, as will be shown below, had a significant impact on union behavior.

This study looks at two types of evidence. The primary focus is on estimating labor demand parameters under different bargaining regimes. Using standard techniques, the elasticity of employment to wages and output is estimated and compared across the policy regimes. The model is then extended to examine the dynamics of the adjustment process. A dynamic labor demand model is estimated, letting the length of the lag vary with the extent of openness and with the percentage of workers covered by collective bargaining. The secondary focus is on flows in the labor market. The odds of layoffs and turnover resulting in unemployment are examined to further understand the micro processes by which unions have an effect in the adjustment process.

The paper begins with background on the economy, the labor market, and collective bargaining in Uruguay (Section II), followed by a brief theoretical overview on unions and labor demand (Section III) and a description of the data (Section IV). The labor demand results (Section V) indicate a structural shift in the labor demand function occurred at about the same time as the return of collective bargaining. Wages are weakly exogenous to employment through 1984, but weak exogeneity is rejected afterwards. The elasticity of employment to wages and output fell by more than 50 percent after 1984. There is no change in the amount of time needed for the market to adjust, as indicated by the coefficient of lagged employment. Results from a bargaining model show that union wage demands are highly sensitive to the openness of the economy. These patterns are further analyzed in the mobility and unemployment results (Section VI). The concluding section summarizes and assesses these findings.

II. Background on Uruguay

II.I. Macroeconomic and labor market conditions

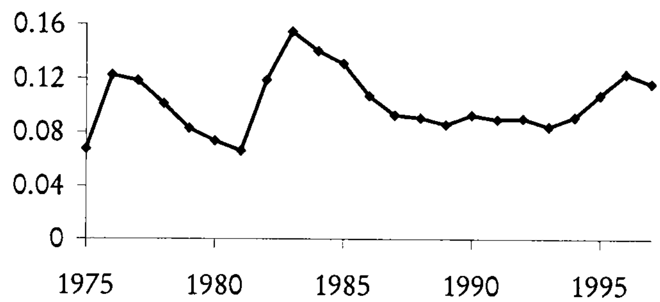
Traditionally, the Uruguayan economy has been subjected to a series of global and regional shocks, particularly those coming from Argentina (Favaro and Sapelli, 1986) and this has continued to be so during the last 25 years. At the beginning of the sample period (1975), the Uruguayan economy was still recovering from the oil shock of 1973 and the ensuing global recession. These conditions were exacerbated by the European Community's decision in 1974 to stop importing beef. Unemployment was above 10 percent in 1976-1978 (see Figure 1).

The economy recovered in the late 1970s in response to a series of steps to liberalize financial markets and promote exports. Growth accelerated when the government adopted a pre-announced schedule of monthly devaluations with the rate of devaluation declining gradually over time. Global economic conditions were not kind to this schedule; by the early 1980s, the net result was a highly overvalued currency which had to be devalued in the global recession of 1982. Unemployment had fallen to 7 percent by 1980-81, but increased to 15 percent in 1983 and stayed above 10 percent through 1986. GDP decreased by 15.9 percent in three years.

By 1988 Uruguay had successfully recovered from this deep recession. The economy grew 8.9 percent in 1986 and 7.9 percent in 1987, supported by an increase in demand from Brazil, which was implementing a stabilization plan (Plan Cruzado). Exports grew and the public deficit decreased to 4.2 percent of GNP in 1987. In 1989, however, the favorable regional environment changed, the public sector deficit grew to 7 percent and a stabilization plan was implemented by the new government elected in 1990. These policies have resulted in a sustained, steady decline in inflation from 129 percent in 1991 to 15 percent in 1997.

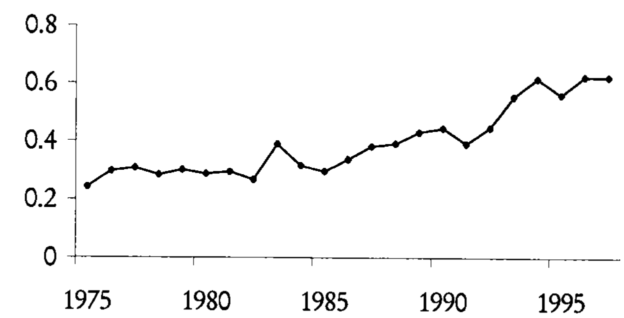
During the 1990s, together with the regional shocks, a domestic stabilization plan and an increase in the openness of the economy had significant effects on macro and industry performance (as shown in Figure 2, where openness is defined as the ratio of imports plus exports to GDP). The Argentinian "Plan de Convertibilidad," imposed in April 1991, improved relative competitiveness for Uruguay, with exports to that country increasing 130 percent in 1991 and 74.3 percent in 1992. Expanded trade with Argentina, no small part of which consisted of tourism, and a deterioration of the real exchange rate meant that growth in the service sector far outstripped growth in goods production. Within the latter, the actual impact depended upon exposure to external competition.

Figure 1: Unemployment rate



Source: National Institute of Statistics (INE).

Figure 2: Openness



Source: Bank of the Republic of Uruguay (BROU); Central Bank of Uruguay (BCU).

Besides the domestic stabilization plan, during the period 1990-1992, a series of trade policy measures consolidated the opening of the economy that had started in the mid-1970s during the military regime, but that was discontinued in the early 1980s. In 1982, the highest tariff was 55 percent and after a temporary increase in 1985, a gradual decrease started in 1986, which ended with a maximum tariff of 40 percent in 1989. The pace of these changes accelerated in 1991-1993. By 1993 the highest tariff was 20 percent. Together with these reductions, many non-tariff barriers and sectoral privileges (like those given to the automotive industry) were removed, and export subsidies were reduced (de Brun and Labadie, 1997).

These unilateral trade policy changes were accompanied by a series of regional tariff reductions as a consequence of the creation of the Southern Cone Common Market (MERCOSUR). By 1995 a great number of Uruguayan products could circulate among its members (Argentina, Brazil, Paraguay and Uruguay) without any tariff. The exceptions to the Common External Tariff are subject to a calendar that was established in December 1994. Economic conditions deteriorated since 1994, largely in response to high unemployment generated by the "tequila effect" in Argentina. Unemployment increased to 11 percent in 1995 and 12 percent in 1996-1997. Unemployment held steady around 9 percent between 1987 and 1994.

The manufacturing sector has been severely affected by all the above. Its share in total output has gone down from 25-27 percent at the beginning of the period to 18 percent in 1997. Employment in manufacturing grew until 1989, but it has decreased significantly since then, to unprecedented levels. This decrease reflects the impact of trade liberalization as some establishments cut back production, whereas others raised productivity to compete

.

II.2. Collective bargaining[2]

When parliament was closed by the military in June 1973, the union confederation CNT launched a general strike. The government reacted by banning union activity and giving employers the right to dismiss anyone who did not return to work. Many union leaders were jailed; the others went into hiding or exile. The union movement began a political comeback in the early 1980s, with a series of demonstrations and general strikes organized by a new confederation, but there was no bargaining until the return of democracy in 1985.

In the absence of unions, employers were relatively free to adjust wages and employment. Wage increases were limited to lagged inflation. This policy, along with high unemployment, was accompanied by a 49 percent decrease in real wages from 1973 through 1984. Employment adjustment also became more flexible. Interview evidence compiled by Handelman (1981) indicates that after the ban on unions, many employers used the opportunity to get rid of trade union officials and excess employees. Dismissals of public sector workers also were permitted by law between 1977 and 1984 (Gillespie, 1991). On the supply side, there was a surge in emigration precipitated by political repression and high unemployment. Taking into account all of these factors, it is clear that the Uruguayan labor market was exposed to strong competitive forces during the ban on unions.

Starting in 1985, Uruguay's unique system of Wage Councils was re-instituted. Collective bargaining in the private sector in Uruguay had traditionally operated mainly through a system of trilateral wage

[2] For a general description of labor market institutions in Uruguay, see Cassoni *et al.*, 1995.

councils that set minimum wages by industry and labor category. Wage levels were adjusted three times a year through 1990; since then, accumulated inflation since the last adjustment had to pass a specific threshold for wages to be adjusted. Often the Wage Councils agreed to a formula that will be in effect for 16 to 24 months, allowing adjustment to take place without a formal meeting. If the government delegates gave their consent to the wage agreement, it applied to the entire sector, not just to the firms and unions involved in the bargaining. Government approval usually required keeping wage increases in line with official inflation targets. Direct negotiation between the union and the firm was also practiced, especially in manufacturing.

In 1991 there was a significant change in the structure of negotiations. The government stopped participating in bargaining. The terms of the contract bound only those firms and unions that are actually represented in the negotiation. Wage Councils only meet in a few sectors and the result, to be shown below, has been a sharp drop in union density in the private sector.

Much bargaining now takes place at the company level. Membership is not compulsory and union dues are voluntary in most cases. In 1988, only three years after unions were legal again, the single National Central Union reported a total of 188,000 members and five years later, in 1993, 177,000 members, belonging to 17 federations and 359 unions. In 1996, there were 164,000 in the National Central, but some unions are not members of it. By 1993, 54 percent of the membership belonged to the public sector, which has had the smallest drop in its number of affiliates.

The role that collective agreements play in introducing rigidities could be very significant, varying in degree depending on union density and the specific clauses of the contracts, that include wage adjustments, minimum wages by job categories, length of work day, holidays, job rotation and stability, recognition of union officers, "peace clauses" that preclude strikes under certain circumstances, and other related working conditions. Although there are no explicit clauses regarding severance pay nor restrictions to hiring new workers, unions have generally imposed extra costs to employment adjustment. In some sectors additional compensation beyond that dictated by the labor contract has been a common practice, while in others strikes have worked as a means of getting additional severance pay. Government intervention in collective bargaining is only provided in the case of Wage Councils, and there is no other regulation of the bargaining process, not even in the case of conflict and strikes (for a more detailed description, see Cassoni *et al.*, 1995: 167-70).

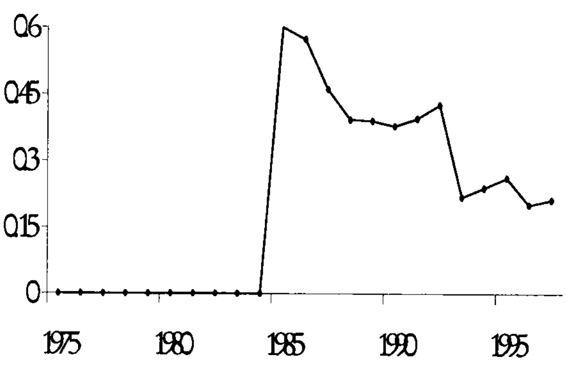
No database up to this date has actually evaluated the impact of the contents of collective agreements. Recently, Ermida *et al.* (1997) and Cedrola, Raso and Perez Tabó (1998) have examined qualitatively the contents of collective agreements for the period 1985-1995. For this study, a data base that covers all collective agreements registered at the Ministry of Labor between 1985 and 1997 has been developed and the contents of its clauses have been quantified to determine the actual nonwage costs resulting from the bargaining process at the industry level.

Using these data it is possible to analyze quantitatively a period in which union behavior was absent (through 1984); a period in which we know the union density of the sector as well as the amount of nonwage costs imposed on all firms in an industry, a consequence of a bargaining structure in place from 1985 through 1991; and a more recent period in which we know union density, but the collective agreements are exclusively binding for those firms and those workers that participated in the negotiation and signed the agreement. The completeness of data for this final period is less clear, since many of these agreements did not have to be registered at the Ministry of Labor (precisely because they did not have to be endorsed by the public authority in order to be binding among the contracting parties).

This study focuses on manufacturing, where there are pronounced changes in union density during the last decade, with no small amount of variation across individual industries. Upon the return of unions in 1985, 60 percent of production workers were covered by collective bargaining agreements (see Figure 3). This initial level probably reflected political support for the role unions played in the return to democracy. Union sustainability hinges on both worker support for collective, as opposed to individual agreements, and on the ability of unionized employers to survive economically. Union density gradually dropped to 40 percent by 1988 and stayed near that level through 1992.

By this point the contracts signed under the old Wage Council system had expired and the impact of trade liberalization was beginning to be felt. The openness ratio jumped from 44 percent in 1992 to 55 percent in 1993 and was above 60 percent for most of 1994-1997. Union density dropped from 42 percent in 1992 to 22 percent in 1993 and has stayed at about that level since. The pattern of union growth and decline has varied considerably across industries, as shown in Table 1.

Figure 3: Percentage union



Sources: National Union Federation (PIT-CNT); National Institute of Statistics (INE).

Union strength remained near 100 percent throughout the sample period in the chemical and oil sectors, which not coincidentally consist largely of state-owned enterprises. In fact union density dropped in all industries after 1992 except in chemicals and oil products. The most dramatic decline took place in metal products and nonmetallic industries, where union coverage in the period dropped to 20% of its original level. At the same time, and particularly in metal products, imports plus exports increased sharply. There also was a considerable drop in union coverage in textiles and apparel and, to a lesser extent, in the paper industry. With the exception of food products, all industries experienced an increase in openness after 1992.

Table 1. Percentage union and openness ratio, by year and industry

Industry	Union 1985	Union 1988	Union 1992	Union 1997	Open 1985	Open 1992	Open 1997
Food, beverages, and tobacco	59	54	55	27	24	24	28
Textiles & apparel	77	54	46	16	49	54	83
Paper products	70	52	44	39	19	19	45
Chemicals and oil products	100	87	100	94	16	44	60
Nonmetallic minerals	48	21	11	10	12	22	36
Metal products	100	43	43	19	76	146	350

Sources: National Union Federation (PIT-CNT); National Institute of Statistics (INE); Central Bank of Uruguay (BCU); Bank of the Republic of Uruguay (BROU).

III. Theoretical framework

This section describes the framework used to analyze possible changes in both elasticities of labor demand and labor dynamics, due to the institutional changes that took place in 1985, that is, the re-appearance of trade unions as "players" in the labor market. In order to do so, the estimable models will be specified so as to measure labor demand elasticities for production workers and the speed of adjustment of labor to its equilibrium level in both regimes.

Through 1984, a competitive model seems suitable to describe the behavior of the labor market. Wage increases were set by the Government from 1968 up to 1979, although from 1977 onwards there were extra shifts in some sectors. In any case, they were exogenous to the firm. Since 1985, it might be possible to approximate the observed employment and wage pairs using the same model, but the institutional framework had changed dramatically. Since that date, the wage level has been the result of a bargaining process that has itself evolved all along the decade. Before 1992, bargaining was a synchronized process, taking place at the industrial sector level through Wage Councils. After that date, it became more heterogeneous as negotiations at the firm level have become quite common, while synchronization has deteriorated.

Given the above institutional changes, the research strategy developed was the following: first a model of labor demand derived from a pure neoclassical static framework was estimated. The wage variable is a cost of labor proxy, including the wage plus nonwage costs - such as health insurance and payroll taxes - as well as other benefits bargained between firms and unions from 1985 onwards.

As will be shown in more detail below, the model was estimated for the whole period and the stability of the parameters was tested. The econometric analysis supported the specification of a different model for the post-1984 period, a model derived from a bargaining framework. A first implication is that wages are not exogenous as in the previous specification, as they are determined jointly by unions and firms through a bargaining process, where firms attempt to maximize profits and unions maximize their members' utility function. Secondly, other variables could enter the model, such as alternative wages or fall-back positions of the parties.

III.1. Labor demand: theoretical framework

We begin with a standard specification for a labor demand equation in a static framework. Assuming a generalized CES production function with three inputs (capital and labor divided in production and non-production workers), maximization of profits would yield a 3-equation system of derived demands for each input. The equation describing the demand for production workers would be:

$$\ln N_t = \alpha_0 + \alpha_1 \ln(w/p)_t + \alpha_2 \ln Q_t \qquad (1.1)$$

where N=employment of production workers, w=wage, p=product price, and Q=output

Hence, the elasticity of substitution between capital and employment (σ) is equal to $-\alpha_1$, while the wage elasticity of labor demand is $-\alpha_1*(1 - s_L)$, with s_L denoting labor's share in value added.

In order to estimate the model, some methodological issues have to be solved. If variables are not stationary, a possible strategy is to estimate the model in differences. A second approach would be to test if the variables involved are cointegrated and if so, the estimation can be carried out in levels. However, as in finite samples the estimators in equation (1.1) are biased, it might be preferable to estimate a dynamic version of the model based on Engle and Granger's representation theorem (Engle and Granger , 1987):

$$\alpha(L)(1-L)\mathbf{Z}_t = -\gamma\beta\mathbf{Z}_{t-1} + d(L)\varepsilon_t \qquad (1.2)$$

where $\alpha(L)$ is a polynomial matrix in the lag operator; **Z** denotes the vector of variables involved (N, w/p, Q); d(L) is a polynomial; and ε_t is a stationary process.

The model can be linearly transformed as an autoregressive-distributed lag model:

$$\alpha_1 (L) y_t = \alpha_2 (L)\, \mathbf{X}_t + \varepsilon_t \qquad (1.3)$$

where $\alpha_1 (L) = 1 - \Sigma^m_{i=1} \alpha_{1i} L^i$; $\alpha_2 (L) = \Sigma^m_{i=0} \alpha_{2i} L^i$ and $(y,\mathbf{X}) = \mathbf{Z}$

The econometric analysis of the model will determine its final dynamic structure. It has been shown that the lag structure of each variable need not be the same (for an extensive discussion of all the above methodological issues, see Banerjee *et al*, 1993).

The fact that variables are non-stationary means that at least some shocks have permanent effects on them. In particular, shocks related to productivity and accumulated knowledge have been generally found to be non-transitory, so that they have long-lasting effects on output and employment (Blanchard and Quah, 1989; Aghion and Saint-Paul, 1993; and references therein). Thus, variables would have a stochastic trend but, if cointegrated, the equilibrium relationship among them would still be stationary and hence stable. The dynamics are the result of agents not being able to adjust instantaneously to equilibrium because of factors such as adjustment costs, price rigidities, etc.

Adjustment costs have been extensively discussed in the literature (Hamermesh, 1993, 1995; Hamermesh and Pfann, 1996) as the source of the observed lags in adjusting employment. They would explain why actual employment (N) differs from its equilibrium level (N^e). If firms maximize expected profits, expectations are static and costs are quadratic, the optimum path of employment would be:

$$N_t = g(N^e - N_t) \qquad (1.4)$$

yielding a demand for labor equation like:

$$N_t = \lambda N_{t-1} + \beta \mathbf{X}_t \qquad (1.5)$$

with $\mathbf{X}_t$ being a vector of variables determining long run labor demand and λ a parameter measuring the speed of adjustment to equilibrium, which is thus assumed to be constant.

III.2. Bargaining models

Since 1985 unions started playing a role in the determination of wages, working conditions and employment. Their role has varied over time, as well as the issues they bargained over. After analyzing all the collective agreements that have been signed since then, it is clear that there have always been negotiations over wages but rarely over employment. Agreements have covered a wide range of other benefits, increasing the annual wage a worker receives; linking the wage to different variables, such as productivity or tenure; and increasing fringe benefits. Working conditions also have been in the bargaining agenda, as well as the length of the working week and year. Although at first sight negotiations looked as if done in stages, this turned out to be false. The procedure followed has generally been one by which at some point unions and firms have bargained over the wage, other benefits and working conditions. Regarding every issue but the wage, agreements have worked as long-term contracts (one year minimum, three years on average). Regarding the wage, however, they were quite short, covering a time period of three or four months, so that most of the contracts were agreements only over the wage.

The above suggests that the most suitable benchmark to analyze the Uruguayan bargaining process is that of a right-to-manage model (for a discussion on this topic, see Pencavel, 1991). The model will be considered as a maintained hypothesis, based on the analysis of all collective agreements. No tests against an efficient contract model will be carried out as it has been extensively proven by now that those tests cannot support one specification against the other (Booth, 1995; Pencavel, 1991)[3]. Thus, the following specification is used:

First Stage: unions and firms bargain over the cost of labor.
$\Gamma(w, w^a, N)$ is the union's utility function, where w is the real wage, w^a is the alternative income, and N is employment. It is assumed that membership status is lost if unemployed; that all members of the union are equally considered by union leaders; and that members care about the real wage surplus over the alternative

[3] For example, the alternative income would enter the employment equation only in the efficient contract model. However, some utility functions can yield a solution to the efficient bargain that excludes the alternative income from the specification. Further, the empirical distinction between both models is not straightforward, as the contract curve may lie on the labor demand curve (Carruth and Oswald, 1987).

income they would earn working elsewhere or being unemployed (de Menil, 1971). A standard specification is then:

$$\Gamma(w, w^a, N, M) = (w - w^a)N^{\phi}$$

where ϕ is a parameter denoting how much weight the union gives to employment in its utility function. If ϕ equals 1, then the model is the rent maximization model (Pencavel, 1991).

$\Pi(p, Q, K, N, p_c, w)$ is the firm's profit function, where K=capital and p_c=price of K. It is assumed that managers maximize revenue minus costs, so that:

$$\Pi(p, Q, K, N, p_c, w) = pQ - wN - p_cK$$

A well known solution to the bargaining problem is given by the maximization over the wage of the generalized Nash bargain, subject to the optimum labor demand that will be set in a second stage:

$$\begin{aligned} &\max_{w} Y = (\Gamma - \Gamma_0)^{\beta} (\Pi - \Pi_0)^{1-\beta} \\ &\text{s. t.} \qquad\qquad\qquad\qquad\qquad\qquad (2.1) \\ &\qquad N = N^{*} \end{aligned}$$

Γ_0 and Π_0 are the fall-back positions of each player. They refer to what the union and the firm would get in the event of no agreement (Binmore, Rubinstein and Wolinski, 1986). If we assume that under this circumstance there will be a strike, then the firm will have zero operating profits and union members will have zero earnings[4].

Second Stage: firms maximize profits.

$$\max_{N, K} \Pi = pQ - wN - p_cK \qquad (2.2)$$

Subject to quite restrictive assumptions, the solution for (2.1) and (2.2) is:

$$\begin{aligned} N^{*} &= N(w/p; Q) \\ & \qquad\qquad\qquad\qquad (2.3) \\ w^{*} &= \eta\, w^a \end{aligned}$$

The first equation is just the result of profit maximization by firms, under a CES production function, for example. However, to get the equation for the wage level, it has to be assumed that when bargaining, firms take capital as given, that is, they have already made decisions on the capital level. Thus, the profit function depends just on employment.

The parameter η is the mark-up over the alternative income. It can be considered a function of some characteristics of the sector firms operate in, such as the degree of competitiveness and the affiliation rate (Layard, Nickell and Jackman, 1991).

---

[4] There are no legal provisions assuring any income to strikers in Uruguay. They generally ask people for contributions but this cannot be measured.

Finally, the alternative income workers consider as a comparison wage is a weighted average of what they would earn if they got a job in any manufacturing industry; what they would get if they decided to become self-employed; and of what they would receive as unemployment benefits in the event of losing their job. Weights are given by the probability of being in each of the mentioned states, calculated as the annual frequency of each category.

The estimable model proposed is a multivariate model, in which wages are not exogenous but they are set subject to the determination of the level of employment.

III.3. Union impact

In a static framework, unions have an incentive to take whatever steps they can to reduce the wage elasticity of labor demand so that they can bargain for increased wages with less severe consequences for employment. Unions can make product demand less elastic by making fewer options available to consumers through various rent-seeking activities. One way of doing this is to create entry barriers, such as state ownership or regulated entry into markets where establishments are unionized. Tariffs, quotas, and other barriers to free trade also can be used to reduce consumer choice.

The elasticity of substitution between union labor and other inputs can be reduced through collective bargaining. Contracts with unions often spell out the conditions under which work is to be performed, including dictates on minimum crew sizes, limitations on substituting nonunion personnel for work that "belongs" to the union, and limits on technologies that reduce labor hours.

Empirically, it is well known (at least since Marshall) that unions should try to organize the sectors of the economy with the most inelastic demand. In this study we look at the same sectors of manufacturing before and after re-unionization, so this allows us to control for this self-selection into rent-seeking opportunities. This study will be able to establish in a before-and-after framework whether unions are actually able to reduce labor demand elasticities.

The impact of unions on adjustment lags and the elasticity of labor demand to output hinge on a variety of factors. Ignoring adjustment costs for the moment, keep in mind that firms can adjust labor hours to a change in output by changing employment or by changing hours per person. The impact of unions on this tradeoff is not clear *ex ante*. Unions often negotiate for premium rates for overtime that are well above those required by labor legislation, which would by itself lead unionized firms to increase employment more for a given increase in output. However, unions also negotiate for employee benefits that make increasing employment expensive relative to increasing hours. Lower turnover in unionized establishments encourages greater investments in employee training, which in turn increase the cost of hiring an additional person. In a frictionless world, the effect of unions on the employment-hours balance would be an empirical question that would hinge on whether the marginal cost of an extra hour per person is the overtime rate dictated by labor laws or the super-overtime rates from the union contract. If it is the standard overtime rate, then the dominating effect of unions would be through increased costs of hiring an extra person and we would expect a smaller elasticity of employment to output.

A final channel for union influence is the speed at which labor adjustments are made. Unions have numerous methods at their disposal to change the cost of making changes in employment. This can be done with formal contract provisions dictating advance notice or severance pay in case of layoffs or through informal threats of slowdowns or strikes. Another factor leading to slower adjustment of employment to output in unionized establishments is the low rate of voluntary turnover. When attrition is sufficiently high, employment can adjust very quickly through a simple hiring freeze.

The expected duration of layoffs often plays an important role. If the expected duration of a drop in output is expected to be short, unions will not hesitate to opt for layoffs rather than hours reductions so workers can take full advantage of unemployment insurance. The prevailing wisdom is that unions create longer adjustment lags through their impact on advance notice and severance pay.

IV. Data

Before describing the actual definition of variables, some aggregation issues are worth stating. First, the units of observation considered will be manufacturing industries at the two-digit level of aggregation. Six of them can be observed during the period 1975 to 1997: food, beverage & tobacco; textiles and apparel; paper; chemicals and oil products; nonmetallic minerals; and metal products. It is well known that the optimum unit of observation is the establishment as adding up technologies never guarantees that the parameters obtained for the aggregate are what they are sought to be. However, working with industries is not the worst of the alternatives. In a small country like Uruguay, most of the year-to-year variation in industry data is driven by a small number of firms, hence problems related to aggregate data should be fewer than in a large country. Nevertheless, it should be taken into account that this might bias the estimates (Hamermesh, 1993). Second, temporal aggregation does not seem a problem here as quarterly data will be used, so that the lag structure should not understate the true lag structure.

IV.1. Cost of labor: W

The measure to be used in the model has to approximate the total cost of labor for the firm, so that it has to include not only the wage but also nonwage costs. The latter include labor taxes; social security contributions; and bargained costs since 1985. We are omitting, however, all costs related to hiring and firing workers. In order to account for these costs, the labor demand function should be specified contingent on different states of nature, that would imply firing or hiring workers, and a distribution of these states should be also proposed. It can be shown that not specifying a state contingent labor demand might bias downwardly the estimates of the elasticities due to the omission of relevant variables. We will not address this issue empirically as data needed to calculate marginal firing and hiring costs are not available [5].

[5] The law relative to severance pay has not changed in the sample period and the compensation a worker is entitled to is the same for all industries and depends on his/her tenure (none if tenure is less than three months; one wage per year for those working for more than three months and up to a maximum of six). Average tenure for those employed in 1991-1997 (the only years for which the data is available) is between seven and ten years, not varying much between industries. Hence, the expected average severance pay does not change, being between 3.7 and 4.2 wages depending on the industry. As it is not possible (due to the number of observations) to calculate the probability of a worker being laid off for each tenure, this should be calculated as the overall frequency of layoffs and will thus be negatively correlated with employment by definition. Finally, even if we included a tentative measure of average severance pay based on tenure of employees instead of on that of laid off workers, we would be introducing biases which need not be of the same sign along the period. They would depend on the prevailing rules of firing workers and these have been probably different during 1975-1997. The only evidence available was specially collected for the last quarter of 1997. Manufacturing workers that were laid off in the month previous to the survey had an average tenure of 1.5 years, while the average for all unemployed manufacturing workers was 6 years. This suggests that a rule of last in-first out was in

Data on wages are obtained from the Quarterly and Annual Industrial Surveys carried out by the National Institute of Statistics (INE) [6]. Annual data for production workers are available from 1975 up to 1997. Quarterly data, however, are not published (nor processed by the INE) after 1991. Hence, for 1992-1997 the within-year evolution of wages was assumed to follow the same pattern as that stemming from the Wage Survey (INE) for manufacturing workers[7].

Data on nonwage costs were taken from Picardo, Daude and Ferre (1997) and from Cassoni and Ferre (1997). All costs related to health insurance and social security as well as payroll taxes were used to build a factor by which to increase wages for each 2-digit industrial sector. Social security and health insurance contributions are a fixed percentage of wages that has varied over time. On the other side, payroll taxes, first imposed in 1982, have generally varied depending on the level of earnings. Hence, information from the Household Survey (INE) was used to calculate the distribution of workers in the different relevant segments, yearly, for each manufacturing sector. Apart from these factors increasing wages, employers face an annual extra payment of one monthly salary plus 20 days that must be paid before the worker starts his/her annual holidays before the end of the year. Both were also included in the cost of labor.

There are several issues over which unions have bargained since 1985. Among them, supplemental end-of-year bonuses, either related to tenure, productivity, or simply on a general basis; shorter length of the working day; and extra holidays. These negotiations took place at the industrial 2-digit level, so that they vary by industry. Annual premia applying to all workers were directly used to increase the factor built upon the legal rates. Information on extra holidays was used to calculate the percentage increase in costs due to non-working days. If paid vacations were 12 days more per year over the legal standard, the actual monthly wage would be 25/24 times w, instead of w. Where agreements were reached shortening the legal length of the working day or week, the cost of labor was increased by the proportion of legal to bargained hours in the same way as paid vacations.

All the information above stemming from the manufacturing collective agreements signed between 1985 and 1997 was used to build an index increasing the legal cost of labor. This index varied in time and among industries, with an average value for the whole manufacturing sector of 12 percent. Industries with the lowest extraordinary bargaining costs were paper; metallic industries and nonmetallic minerals, for which the increase was around 1 percent on average. Sectors related to food, beverage and tobacco and chemicals have negotiated increases of 12 percent over the legal costs, while those related to textiles have an average percentage premia of 21 percent during the period.

Given all the above, the cost of labor variable was defined as:

Cost of Labor = CL = Wage*(1 + legal nonwage costs + bargained nonwage costs)

---

place. However, during periods of restructuring, as were the late seventies and the early nineties, firms might have got rid of more senior workers, with higher wages and not easily retrainable. Given all these issues, we will omit these costs from the analysis, although they might be reflected in the estimated effect of unions on the labor demand model.

[6] These surveys are carried out using a sample of firms employing 5 workers and more, that stems from the previous Industrial Census. Data collected refer to many variables related to production, employment, and inputs. The Quarterly Survey reports indexes while the Annual Survey publications report values.

[7] The Wage Survey is carried out on a monthly basis to establishments belonging to all economic sectors.

IV.2. Employment: N - Production: Q - Product prices: p

Employment refers to total number of production workers obtained from the Quarterly and Annual Industrial Surveys, at the 2-digit level. An index of production is available on a quarterly basis (INE). The index was then transformed to monetary values using the 1988 Industrial Census and the Annual Industrial Survey (INE). Data on product prices refer to the PPI at the 2-digit level (INE). All data refer to monthly values calculated as an average on a quarterly basis.

IV.3. Some corrections to the official data

In Uruguay, the industrial census is performed every 10 years. Each time a census is done, annual and quarterly surveys update their samples based on the new information. These samples are such that those establishments with more than 50 or 100 workers (depending on their share on the industry value added) are always surveyed (hence, death and births are accounted for), while smaller plants are sampled at the beginning of the period and remained in the sample until the new industrial census is carried out. In 1988, the last national industrial census was performed and its results showed that the samples that were being used in the industrial surveys – stemming from the 1978 census – were severely misrepresenting the different sectors. Annual surveys started including the new information in 1989 while quarterly surveys did so in 1993. However, no correction to the data was done before those dates. The differences in the samples meant that the estimated levels of employment and output for the whole manufacturing sector differed by about 25 percent depending on the sample used. At the 2-digit level there were even broader differences. It was thus decided to correct the official data, discussing and taking advice from those in charge of the surveys at the National Institute of Statistics. Given that the 1982-1983 economic recession had major and different effects depending on the industrial sector, the assumption used to calculate the new data was that the lack of representativeness of the 1978 sample went back to 1984. As other sources showed that the evolution of the variables stemming from the surveys along the post-1984 period was quite correct, the differences in the levels according to both samples were geometrically distributed along those years (1984-1988 for the annual survey; 1984-1993 for the quarterly survey).

IV.4. Degree of openness: OPEN

The index was calculated as total exports plus total imports divided by value added, per manufacturing industry. Data came from the Republic Bank of Uruguay (BROU), that was the authority in charge of registering all foreign exchange activities. Since 1995, the Customs Office has been responsible for collecting the data.

IV.5. Alternative wages: AW

They were calculated using the information on wages in manufacturing as described in IV.1; and that of average income of self-employed individuals according to the Household Survey. The alternative income for a worker in industry "j" was defined as the weighted average of the wage in the rest of the manufacturing industries; the income the worker would receive if he/she becomes unemployed and collects unemployment benefits (50% of his/her current wage); and the average income of self-employed individuals. Weights were defined as the annual frequency of each category as stemming from the Household Survey.

IV.6. Union density: UNION

Union density was calculated using annual number of production workers as reported in the Industrial Surveys and total membership as reported by the National Union Federation after each congress. Congresses took place in 1985, 1987, 1990, 1993 and 1996-97. In intervening years, union membership is estimated through simple interpolation.

Descriptive statistics for the above variables are summarized below in Table 2, differentiating between the pre and post re-unionization subperiods (1975-1984 and 1985-1997). Data for the entire manufacturing sector are reported to indicate overall trends; data for manufacturing industries indicate the heterogeneity of conditions across different markets. Note that with the return of collective bargaining, the market trends are toward greater production, reduced employment, higher wages, and increased openness.

Table 2. Descriptive Statistics

a) *Manufacturing sector*

	1975.1 – 1984.4 Number of observations: 40				1985.1 – 1997.4 Number of observations: 52			
Variable	Mean	S.D.	Max	Min	Mean	S.D.	Max	Min
W	82.02	13.97	103.6	56.81	90.02	21.64	133.3	52.38
LNWC	1.336	0.071	1.426	1.243	1.332	0.031	1.375	1.290
BNWC	1.000	0.000	1.000	1.000	1.123	0.038	1.156	1.000
TLC	109.7	18.6	143.54	72.56	136.2	36.34	203.7	67.79
AW	0.000	0.000	0.000	0.000	42.95	11.67	62.59	24.87
UNION	0.000	0.000	0.000	0.000	0.365	0.129	0.601	0.200
OPEN	0.298	0.036	0.388	0.242	0.468	0.109	0.620	0.295
Q	57.00	6.740	70.00	44.60	60.16	5.097	71.04	49.16
N	108143	14496	129491	86010	104782	19727	129995	71735

b) *Manufacturing industries*

	1975.1 - 1984.4 Number of observations: 240				1985.1 - 1997.4 Number of observations: 312			
Variable	Mean	S.D.	Max	Min	Mean	S.D.	Max	Min
W	86.93	28.84	202.9	41.90	104.8	40.96	246.3	41.25
LNWC	1.337	0.071	1.433	1.238	1.328	0.038	1.383	1.232
BNWC	1.000	0.000	1.000	1.000	1.076	0.096	1.265	1.000
TLC	115.3	35.31	255.6	58.65	151.4	68.23	405.8	53.23
AW	0.000	0.000	0.000	0.000	69.88	21.27	136.7	30.79
UNION	0.000	0.000	0.000	0.000	0.507	0.285	1.000	0.083
OPEN	0.338	0.257	1.149	0.096	0.575	0.657	3.500	0.102
Q	9.431	6.971	27.42	1.598	9.804	6.784	26.69	1.296
N	17661	12763	49715	4167	16543	12292	42150	3897

Notes: W is monthly real wage per production worker in 1988 pesos; LNWC is 1 + percentage increase in wages due to legal nonwage costs; BNWC is 1 + percentage increase in wages due to bargained nonwage costs;TLC are monthly total real labor costs in 1988 pesos; AW is the monthly real alternative wage in 1988 pesos; UNION is percentage union; OPEN is degree of openness; Q is production in 1988 million pesos; and N is number of production workers.

To preview the dynamic patterns of the key variables, the quarterly change in employment for the manufacturing sector is plotted along with the quarterly changes in production and employment (all in logs) in Figures 4 and 5. Employment varied less on a quarterly basis with the return of unionization; the standard deviation of the log change in employment fell from 0.036 in 1975-1984 to 0.028 in 1985-1997. At the same time production became more variable, with the standard deviation of log change rising from 0.063 to 0.095 for the same two periods. This would indicate a strong likelihood that the employment-output elasticity fell after 1985. The story for wages is less clear cut. The standard deviation of the log change in wages decreased from 0.124 to 0.063[8]. Proportionally speaking, the quarterly variation in wages fell by more than the quarterly variation in employment, indicating a possible increase in the wage elasticity of labor demand. However, between 1975 and 1980 there are six nearly consecutive episodes of a sharp (0.1 or larger) increase in wages followed one or two quarters later by an equally sharp decrease in wages. It is doubtful that employers reacted very much to such short term wage shocks.

Figure 4: Log changes in employment and wages

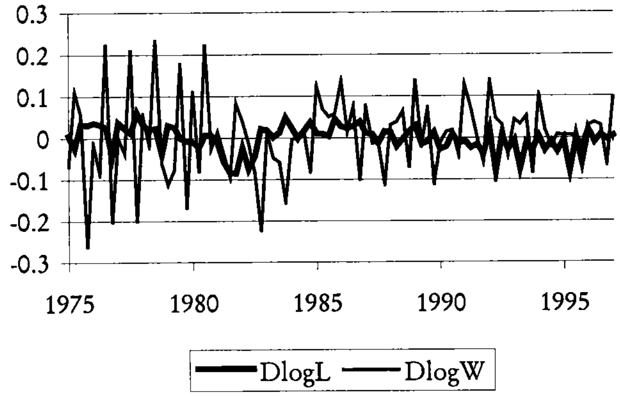
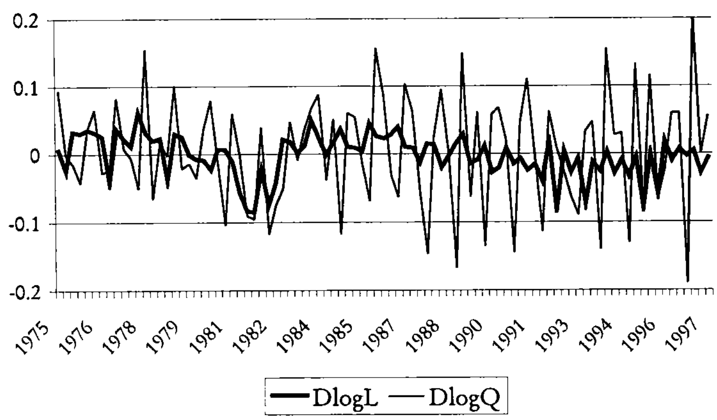


Figure 5: Log changes in employment and production



[8] The null hypothesis of no change in the variance is rejected at the 5 percent level for all three variables and at the 1 percent level for wages and output, with F(38,51)=3.85 for log wages, F(38,51)=2.27 for log output and F(38,51)=1.65 for log employment.

V. Labor demand: empirical results[9]

V. 1. Specifying a model for the whole period

To determine whether and how much elasticities and adjustment lags of labor demand in the manufacturing sector changed after the return of collective bargaining, we must first establish the appropriate specification of the empirical model. The quarterly data on the six manufacturing industries described in previous sections was used. To estimate equation (1.1) as it stands, the stationarity of the variables must first be established, which we did by estimating the order of integration of employment, labor costs, and output for each manufacturing industry in the 1975-1997 period. All variables are nonstationary but their first differences are stationary, so that they are integrated of first order -I(1). The unit root tests used to perform the analyses were those proposed by Fuller (1976), known as Augmented Dickey-Fuller tests (ADF). The models over which the tests were performed include only a constant and lags of the dependent variable (except in one case where seasonals and a time trend were also included; for details, see Table 1 in the appendix, available from the authors upon request). These results are somewhat expected. Regarding employment, output and real wages, accumulated knowledge and productivity shocks have been found to generate stochastic trends in these variables as was mentioned in section III.

Two other key variables used in the analysis – the degree of openness and union density – also are nonstationary (results available from authors upon request). The nonstationarity of the degree of openness could be interpreted in similar terms, with external shocks and trade policies in the root of the result. Finally, the most likely explanation for the stochastic trend found in the union density variable should be linked to membership dynamics and insider-outsider arguments (Blanchard and Summers, 1986). Given the statistical properties of the data, one possible strategy is to estimate the model in differences.

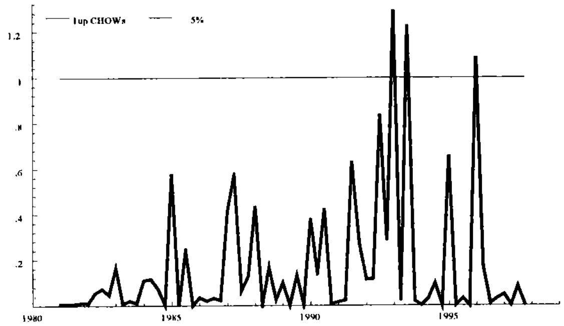
The institutional framework depicted in previous sections suggests, as a second step, the analysis of the stability of the parameters in time. The model in differences was thus estimated industry by industry, using recursive least squares (RLS) and assuming wages and output are exogenous. The results, depicted in Figure 6, show there are structural breaks in the labor demand equation in all industries except nonmetallic minerals. The timing of the breaks is not identical in each industry, but breaks can be identified at some point in the early 1980s as well as at another point around 1991-1993. These dates can be clearly related to the major economic crisis in 1982-84; the end of the military regime in 1985; and the end of government participation in the Wage Councils.

A third stage of the analysis involved using the pooled cross section-time series data set. Given the nonstationarity of the variables and the instability of the parameters, the model was specified in differences with the parameters shifting in various combinations of 1983, 1985 and 1993 and estimated by ordinary least squares (OLS). Elasticities were imposed to be the same for all six industries while wages and output were taken as exogenous variables. These results are reported in Table 3.

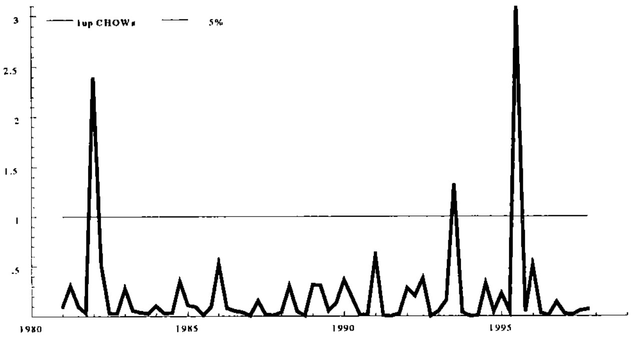
The first three columns test for a single break in 1983, 1985, and 1993. The null hypothesis of no shifts cannot be rejected for 1983 and 1993, but is rejected for 1985. The output coefficient falls from 0.141 in 1975-1984 to 0.073 in 1985-1997. The wage coefficient becomes smaller in absolute value terms,

[9] We are grateful to John E. Driffill for his useful comments as well as to Fernando Lorenzo for his econometric advice.

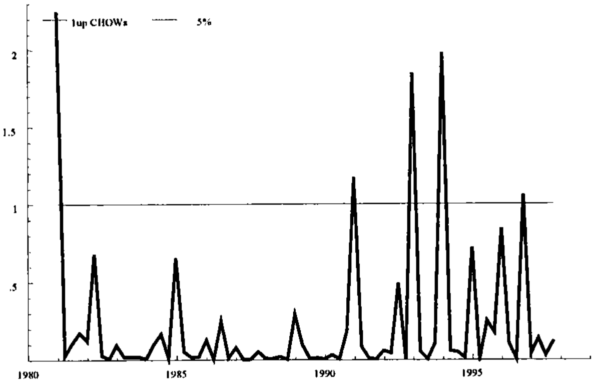
Figure 6: Recursive residuals, by industry



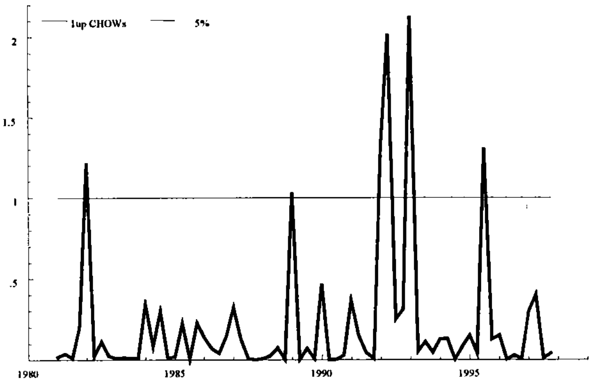
Food, beverage & tobacco: breaks in 1992-93



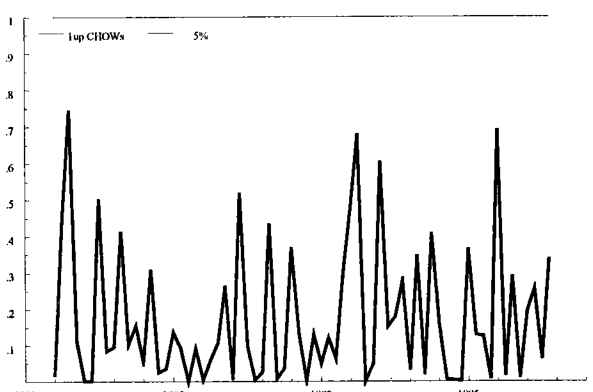
Textiles & apparel: breaks in 1982, 1995



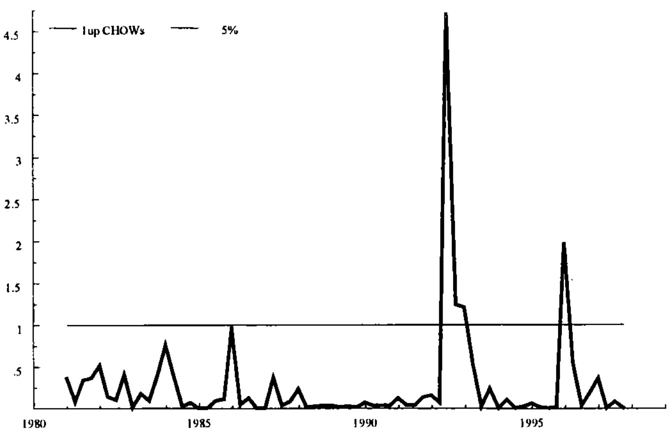
Paper products: breaks in 1991- 92



Chemicals: breaks in 1982, 1993



Nonmetallic minerals: no breaks



Metal products: break in 1992

Table 3: Estimation in differences, manufacturing industries, 1975-1997

Dependent variable: $\Delta N_t = N_t - N_{t-1}$
Sample: 1975 – 1997
Number of Observations: 534

Structural Breaks	1983	1985	1993	1983&1993	1985&1993
Variables					
ΔN_{t-1}	0.01256	0.04312	0.02399	0.01316	0.03833
	(0.0604)	(0.0564)	(0.0468)	(0.0593)	(0.0557)
ΔN_{t-2}	0.16010	0.15278	0.09722	0.16832	0.15082
	(0.0605)	(0.0564)	(0.0469)	(0.0594)	(0.0559)
ΔQ_t	0.15244	0.14078	0.12545	0.15180	0.14259
	(0.0306)	(0.0263)	(0.0203)	(0.0300)	(0.0260)
ΔW_t	-0.08480	-0.10309	-0.09007	-0.08364	-0.10675
	(0.0296)	(0.0234)	(0.0224)	(0.0291)	(0.0264)
$\Delta NdY1_{t-1}$	-0.01144	-0.07197	-0.08663	0.01955	-0.06430
	(0.0833)	(0.0835)	(0.1043)	(0.0943)	(0.0995)
$\Delta NdY1_{t-2}$	-0.11730	-0.10185	0.03228	-0.21630	-0.20982
	(0.0837)	(0.0839)	(0.1051)	(0.0950)	(0.1002)
$\Delta QdY1_t$	-0.05929	-0.0679	-0.05994	-0.04599	-0.04991
	(0.0374)	(0.0349)	(0.0388)	(0.0410)	(0.0430)
$\Delta WdY1_t$	0.01234	0.05570	0.09638	-0.01371	0.04344
	(0.0424)	(0.0244)	(0.0704)	(0.0449)	(0.0507)
$\Delta NdY2_{t-1}$				-0.24582	-0.18643
				(0.1233)	(0.1293)
$\Delta NdY2_{t-2}$				0.03721	0.04908
				(0.1213)	(0.1273)
$\Delta QdY2_t$				-0.03598	-0.02352
				(0.0430)	(0.0473)
$\Delta WdY2_t$				0.13014	0.09706
				(0.0747)	(0.0793)
R^2	0.0930	0.1028	0.0922	0.1398	0.1344

Note: $\Delta X = X_t - X_{t-1}$. N is number of production workers; W is the real labor cost of a production worker; Q is production. $\Delta XdY1$ is ΔX multiplied by a dummy variable that is equal to 1 in the subperiod starting in Y1 (Y1= 1983; 1985 or 1993 according to the column). $\Delta XdY2$ is ΔX multiplied by a dummy variable that is equal to 1 in the subperiod starting in 1993. Standard errors are in parenthesis below each estimated coefficient.

going from –0.103 to –0.047. The sum of the two lagged employment coefficients falls from 0.196 to 0.022. The models in the last two columns test for multiple break points. Having established a shift in the early eighties, these results examine whether there was an additional shift in 1993. In the fourth column breaks in 1983 and 1993 are included while in the fifth the shifts take place in 1985 and 1993. The joint null of no breaks is rejected in both cases.

Finally, cointegration techniques were also applied . When variables are nonstationary the estimation of the model in levels has been proven to be misleading, unless the variables are jointly stationary, that is, they are cointegrated. Hence, cointegration (CI) tests were then done to see if an equilibrium relationship could be sustained for the whole period. Both Engle and Granger (EGM) and Johansen (JM) methods (Johansen, 1988) were used, specifying various models that differ in the number of lags included, as well as in the inclusion of seasonal dummies or a constant. Cointegration between employment, production and labor costs was rejected for all industries according to at least some of the tests performed (Table 2 in the appendix). In those cases in which CI cannot be rejected, the graph of the CI relation shows it is not stationary, so that it is probably spurious, as it is the existence of a structural break in the relation that makes the statistics significant (see Graphs 1 to 4 in the appendix).

In summary, all the above analyses suggest 1985 stands out on both institutional and statistical grounds as the date at which a structural change in labor demand behavior took place. There is also some evidence of a further shift in the nineties. These break points will be used in the remainder of the paper.

V.2. Specifying a model for each subperiod

Once we determined 1985 as the breaking point, we first repeated the analysis of order of integration and cointegration of variables for each subsample and each industry. For 1975-1984 and 1985-97, every variable is I(1) within each subperiod. Details are reported in Table 3 in the appendix. Second, for 1975-1984, the tests using EGM and/or JM report a CI relation for at least one model (see Table 4 in the appendix)[10]. For 1985-1997 no CI among employment, labor costs and production can be found in any industry, for any model using EGM. However, CI is not rejected in any industry once variables that would reflect a bargaining framework -- alternative wages, bargained costs, degree of openness and union density -- are included. The existence of an equilibrium relation between the variables – according to the non rejection of CI- would state that shocks, having a long lasting effect on each of the individual variables, alter equlibrium only in a transitory way. In the first subperiod, the result is consistent with a standard neoclassical labor demand framework. In 1985-1997, however, the need to include other variables to achieve CI suggests that the framework in which labor demand has been determined actually changed. One possibility is to link the existence of a stochastic trend in the residuals to not having modelled technical change. One might argue that this is partially captured when adding the degree of openness: increases in openness would force the different industries to invest in new technology once they face greater competitive pressures; or it would force firms with older technologies to close, so that on average technical progress would be observed. However, as not only openness but variables accounting for bargaining are included in the CI relation, there is also evidence supporting that a bargaining framework is in place to analyze the labor demand schedule in 1985-1997.

To further establish whether the return of collective bargaining was a likely cause of the observed change in parameters, we then performed exogeneity tests on wages. In the competitive model wages are assumed to be exogenous to the firm and industry, while in the bargaining model they are not. In the latter case they would be set either simultaneously or subject to the determination of employment. Using a Hausman test (1978) in which the OLS estimate of the wage parameter is compared to a Seemingly Unrelated Regressions estimate (SUR), weak exogeneity of wages cannot be rejected in the

[10] EGM was preferred due to the number of observations available. JM was used for paper and chemicals to check if a CI relation could be found.

first subperiod while it is rejected in the second[11]. The SUR estimator is calculated using lags of the wage as instruments in both subperiods. For 1985-1997, however, the test was also performed including bargaining variables (degree of openness and union density). Further, given the evidence on the existence of instability in the nineties, the statistics were also calculated including a dummy variable in the equations, which takes the value 0 before 1993 and of 1 after that date. The values of the statistics for the different models are reported in Table 4. The results provide further support for estimating a standard neoclassical labor demand model for 1975-1984 and a bargaining model for 1985-1997.

Table 4: Weak exogeneity tests for the wage 1975-1984 and 1985-1997

	1975-1984		1985-1997
Model 1	3.02		5.9
Model 2	-----		90.4
Model 3	-----		294.2
Model 4	-----		226.6
Hausman Statistic			
95% confidence		3.84	

Note: Each model contains 5 industry dummies and a constant. In models 1 and 2 labor demand is specified as a function of wages and output, using 4 lags of every variable. In models 3 and 4 a dummy variable for 1993 is also included in both the labor demand and the wage equations. In models 1 and 3, instruments used for the wage are just its lags while in models 2 and 4 instruments for the wage include bargaining variables.

Given all the above results, the estimated models are as follows:

1975-1984: $\ln N_t = \alpha_0 + \alpha_1(L)\ln(w/p)_t + \alpha_2(L)\ln Q_t + \alpha_3(L)\ln N_{t-1}$

1985-1997: $\ln N_t = \beta_0 + \beta_1(L)\ln(w/p)_t + \beta_2(L)\ln Q_t + \beta_3(L)\ln N_{t-1}$

$\ln(w/p)_t = \gamma_0 + \gamma_1(L)union + \gamma_2(L)open + \gamma_3(L)\ln(w^a) + \gamma_4(L)\ln(w/p)_{t-1}$

where N refers to number of production workers; w/p are real labor costs (which after 1985 include bargained costs); Q is production; union is union density; open is degree of openness; and w^a is the alternative wage. The order of the polynomials in the lag operator will be tested empirically, starting with polynomials of order 4. The bargaining model is a recursive, two-equation model, so gains in efficiency can be achieved through simultaneous estimation. To avoid possible endogeneity bias with respect to output, lagged values of Q (up to two lags), seasonals and industry dummies were used as instruments for this variable in the estimation for both subperiods. Hence, estimation was done using Instrumental

[11] The Hausman statistic is: $T(b_{OLS} - b_{SUR})^2 Var(b_{OLS} - b_{SUR})^{-1}$ where b is the estimator, by OLS or SUR, and T is the number of observations. It is distributed as a $\chi 2$ with 1 degree of freedom.

Variables (IVE) in the first subsample and three stages least squares (3SLS) in the second, using PCGive and PCFiml 9.0 software (1996). The dataset is the pooled cross section – time series one described above. Fixed effects per industry are always allowed for. Elasticities are imposed to be equal for all industries, so that the estimates reflect the average elasticities for the whole manufacturing sector.

V.3. Main results

For both subperiods Table 5 reports three simple versions of the labor demand model. Starting with a model including up to four lags for every variable, sequential reductions were performed. Further, the different coefficients were allowed to vary in 1993 in order to check for possible shifts. We only report the post-1993 shifts that were significant. The model in column (1) includes a two-quarter lag of employment, whereas the model in column (2) includes a one-quarter lag. In column (3) a measure of openness is added to the model in column (2) to test if increased openness was affecting the estimates.

The wage equation for the bargaining model allows the wage bargain to vary by industry after 1993. This was done to test whether the change in the bargaining structure has had an overall impact on wage demands and whether the effect varies by industry. Residuals are not autocorrelated but they are heteroscedastic. Thus, standard errors were calculated according to White (1980). Although normality is rejected, hypothesis testing results should be robust to non-normality given the sample size (Spanos, 1986; Ch.21.2).

As can be seen by comparing columns 1 and 2 of the labor demand results within each subperiod, employment from one quarter ago has an effect on employment in the current quarter but employment from two quarters ago has no impact. Further, the degree of openness is not only statistically non significant but does not alter the estimates of the relevant elasticities. Accordingly, our attention will focus on the results for column 2. These show three major results:

1. The output coefficient falls from 0.093 in 1975-84 to 0.040 in 1985-97.
2. The wage coefficient falls (in absolute value) from –0.102 in 1975-84 to –0.039 in 1985-97.
3. There is no significant change in the impact of lagged employment between these two periods.

The wage equation results show that the effect of union density on wages decreased significantly after 1992, although the extent of this change varies per industry. A key finding in the wage equation results is that bargained wages fall with increased openness. The effect is rather small, however, a 50 point change in openness being associated with a 1.5 percent change in the bargained wage.

Because of the different approaches taken to estimating the IVE labor demand and the 3SLS bargaining model, one might wonder if these findings are sensitive to the choice of estimation method or to the inclusion/exclusion of variables in the model. To put the two subperiods on an equal footing, both models were nested in a 2-equation system and estimated by 3SLS. In order to do so, each variable was multiplied by two binary variables - one for 1975-1984, another for 1985-97 - so that X75 equals X in 1975-1984 and 0 after that date and X85 is equal to X in 1985-1997 and 0 before that date. Tests of significance of coefficients and tests of coefficients being equal before and after 1985 were performed and they all re-inforce the previous results (see Table 5 in the appendix).

Table 5. Estimates of labor demand and wage equations - manufacturing industries

Labor demand equation: dependent variable: N_t

	Sample: 1975 – 1984			Sample: 1985 – 1997		
Model	(1)	(2)	(3)	(1)	(2)	(3)
Variables						
Constant	1.4969	1.3840	1.5638	1.3403	1.3630	1.3526
	(0.2980)	(0.3012)	(0.3338)	(0.2333)	(0.2187)	(0.2186)
N_{t-1}	0.90382	0.88844	0.87473	0.79468	0.86921	0.87186
	(0.1299)	(0.0315)	(0.0330)	(0.0625)	(0.0218)	(0.0202)
N_{t-2}	-0.01477	-----	-----	0.07809	-----	-----
	(0.1181)	-----	-----	(0.0588)	-----	-----
Q_t	0.09074	0.09304	0.09092	0.03912	0.04024	0.03309
	(0.0261)	(0.0244)	(0.0239)	(0.0244)	(0.0245)	(0.0173)
W_t	-0.10000	-0.10180	-0.09865	-0.04098	-0.03886	-0.03882
	(0.0227)	(0.0182)	(0.0174)	(0.0178)	(0.0184)	(0.0172)
DUMMY93	-----	-----	-----	-0.03957	-0.04019	-0.0393
	-----	-----	-----	(0.0123)	(0.0126)	(0.0122)
IND.31	-0.04217	-0.04499	-0.07533	0.08076	0.08336	0.08755
	(0.0285)	(0.0271)	(0.0357)	(0.0287)	(0.0287)	(0.0250)
IND.32	0.03857	0.03757	0.02439	0.08019	0.08335	0.08357
	(0.0247)	(0.0267)	(0.0296)	(0.0206)	(0.0206)	(0.0202)
IND.34	0.02271	0.02498	-0.03521	-0.05909	-0.06096	-0.06533
	(0.0276)	(0.0273)	(0.0442)	(0.0209)	(0.0214)	(0.0238)
IND.35	-0.10358	-0.10557	-0.15528	-0.04310	-0.04563	-0.04006
	(0.0242)	(0.0221)	(0.0409)	(0.0246)	(0.0249)	(0.0201)
IND36	-0.04382	-0.04285	-0.10538	-0.07504	-0.07684	-0.08307
	(0.0243)	(0.0233)	(0.0460)	(0.0279)	(0.0283)	(0.0279)
Qr.1	-0.01536	-0.01524	-0.01451	0.00098	-0.00019	-0.00111
	(0.0127)	(0.0127)	(0.0127)	(0.0081)	(0.0080)	(0.0080)
Qr.2	0.00815	0.00783	0.00846	0.01122	0.01031	0.00996
	(0.0079)	(0.0082)	(0.0082)	(0.0059)	(0.0058)	(0.0053)
Qr.3	-0.01340	-0.01323	-0.01286	-0.01589	-0.01778	-0.01793
	(0.0069)	(0.0067)	(0.0068)	(0.0072)	(0.0069)	(0.0067)
OPEN	-----	-----	-0.07185	-----	-----	-0.00090
	-----	-----	(0.0532)	-----	-----	(0.0092)
Number of Observations:	228	228	228	300	300	300
R^2	0.9946	0.9947	0.9947	0.9967	0.9967	0.9967
AR 1-4	3.3058	3.5757	3.9374	1.2294	1.7403	1.7430
	[0.5080]	[0.4665]	[0.4145]	[0.8732]	[0.7834]	[0.7829]
Normality	143.0	138.0	131.7	60.4	56.6	56.7
	[0.0000]**	[0.0000]**	[0.0000]**	[0.0000]**	[0.0000]**	[0.0000]**
Xi^2	2.9151	2.272	2.309	1.5052	1.7656	1.5585
	[0.0002]**	[0.0067]**	[0.0039]**	[0.0353]*	[0.0074]**	[0.0247]**

Wage equation: dependent variable: W_t

Sample: 1985 – 1997 Model	(1)	(2)	(3)
Variables			
Constant	-0.29674	-0.27408	-0.27471
	(0.1068)	(0.1041)	(0.1041)
W_{t-1}	0.36874	0.43003	0.43033
	(0.0563)	(0.0401)	(0.0402)
W_{t-2}	0.07493	-----	-----
	(0.0433)	-----	-----
Aw_t	0.71198	0.72145	0.72126
	(0.0523)	(0.0540)	(0.0540)
$OPEN_t$	-0.02471	-0.02426	-0.02424
	(0.0107)	(0.0107)	(0.0107)
$UNION_t$	0.15515	0.15477	0.15470
	(0.0227)	(0.0229)	(0.0229)
$UNION93_t$	-0.23953	-0.23432	-0.23437
	(0.0693)	(0.0703)	(0.0703)
$UNION93_t$ *Ind.31	0.05711	0.06146	0.06161
	(0.0846)	(0.0862)	(0.0862)
$UNION93_t$ *Ind.32	-0.14993	-0.14841	-0.14815
	(0.0784)	(0.0809)	(0.0809)
$UNION93_t$ *Ind.34	-0.04242	-0.03842	-0.03838
	(0.0745)	(0.0763)	(0.0762)
$UNION93_t$ *Ind.35	0.17082	0.17512	0.17504
	(0.0616)	(0.0627)	(0.0626)
$UNION93_t$ *Ind.36	-0.89888	-0.89890	-0.89809
	(0.2909)	(0.2934)	(0.2935)
DUMMY93	0.10029	0.10001	0.09997
	(0.0332)	(0.0331)	(0.0331)
Qr.1	-0.04555	-0.04357	-0.04358
	(0.0107)	(0.0109)	(0.0109)
Qr.2	0.01220	0.02054	0.02056
	(0.0091)	(0.0086)	(0.0086)
Qr.3	0.01208	0.00984	0.00985
	(0.0085)	(0.0083)	(0.0083)
Number of Observations:	300	300	300
R^2	0.9780	0.9782	0.9782
AR 1-4	1.9425	1.6430	1.6429
	[0.7530]	[0.7928]	[0.7927]
Normality	7.74	7.85	7.85
	[0.0209]*	[0.0198]*	[0.0198]**
Xi^2	1.9445	2.0968	1.9892
	[0.0014]**	[0.0006]**	[0.0010]**

Notes: N is number of production workers; W is the real labor cost of a production worker; Q is production; Aw is the alternative wage; UNION is union density; OPEN is the degree of openness; Qr"j" is a dummy variable for quarter "j"; Ind."i" is a dummy variable for industry "i"; DUMMY93 is a dummy variable equal to 1 in 1993-1997; UNION93 is UNION multiplied by DUMMY93. Industries reported are: food, beverage & tobacco (31); textiles and apparel (32); paper (34); chemicals (36); nonmetallic minerals (36); metal products (38). Models (1) and (2) differ in that the former includes 2 lags of the dependent variable, while the latter only includes 1. Model 3 includes the variable OPEN in the labor demand equation. Variables are in logs, except for UNION; OPEN and binary variables. Corrected (according to White, 1980) standard errors are in parenthesis below each estimated coefficient. AR 1-4 is a test of autocorrelation of order 4 in the residuals; Normality is Jarque-Bera's test; Xi^2 is a test for hereoscedasticity of the residuals, using all variables and their squared value in the model for the variance.

In Tables 6 and 7, labor demand elasticities and results for other relevant parameters are summarized, using model (2) of the previous table. Confidence intervals are also reported. These results show that the wage elasticity of labor demand dropped from -0.69 in 1975-84 to -0.22 in 1985-97. The employment-output elasticity fell by more than 50 percent, from 0.83 to 0.31. The estimated speed of adjustment is the same in both periods, about 5 quarters, so that there is no evidence that the return of bargaining lengthened the amount of time needed for employment to adjust, which is contrary to what one might generally expect[12].

Table 6: Labor demand parameters- manufacturing industries 1975 - 1997

Short run estimates

	1975-1984		1985-1997	
Variable	Estimate	Confidence Interval	Estimate	Confidence Interval
Production	0.09304	(0.045, 0.141)	0.040243	(0.007, 0.087)
Labor Costs	-0.10180	(-0.137, -0.066)	-0.03886	(-0.075, -0.003)
Lagged empl.	0.88844	(0.827, 0.950)	0.86921	(0.826, 0.912)

Long run estimates

	1975-1984		1985-1997	
Variable	Estimate	Confidence Interval	Estimate	Confidence Interval
Production	0.8339	(0.525, 1.143)	0.3077	(0.080, 0.536)
Labor Costs	-0.9125	(-1.368, -0.457)	-0.2971	(-0.534, -0.060)
Labor share (s_L)	0.248		0.257	
Wage elasticity of labor demand	-0.69		-0.22	

Note: s_L is equal to the wage bill (all wage and nonwage costs included) divided by value added. The wage elasticity of labor demand is equal to $-(1-s_L)*\sigma$, where σ is the elasticity of substitution between capital and labor and is given by the estimated coefficient of the wage in the labor demand equation.

Although the estimates might be downwardly biased due to the omission of hiring and firing costs, the evidence of a decline between both subperiods is quite robust. The smaller responses of employment to changes in output and wages are consistent with collective bargaining restricting the options available to employers. Once unions reappeared and started playing a role in wage setting, the rules of the game changed. Costs of hiring and firing workers were expected to increase because of union resistance. Employment would not adjust to changing output demand as before because of increased uncertainty on the reaction of unions. Hence, there would have been more labor hoarding during slowdowns and increased use of overtime work during upswings than when unions were not active.

[12] An exception is the paper by Lockwood and Manning (1989), in which the opposite result is found.

After 1992 the structure of bargaining changed, so that firm level negotiations became quite common in some industries. The effect of this institutional change is captured in both the labor demand and the wage equations, but in different ways. In 1993 the labor demand equation has shifted in, while the other estimated coefficients are stable. Regarding the wage, the estimated effect is an overall increase in wages but along with a reduction of the impact of union power on the mark-up that is different per industry. Industries that have experienced a greater reduction of this positive effect are those in which firm level negotiations have become more common. Hence, while no significant change is detected in chemicals (35) - a concentrated industry in which public firms are present – in nonmetallic minerals (36) union power has become less effective in increasing the mark-up over alternative income. The estimated long run effect of unions is to increase wages by 1.5 percent per each 10 percent increase in coverage in 1985-92. Given the changes that took place in the nineties, the average effect is almost null for the whole period[13]. The indirect effect of unions over employment *via* wages is such that an increase in coverage of 10 percentage points is associated with a 0.8 percent decline in labor demand before 1993.

Table 7: Impact of key variables on real labor costs - manufacturing industries 1985 - 1997

	Short run		Long run	
Variable	Estimate	Confidence Interval	Estimate	Confidence Interval
Openness	-0.02426	(-0.045, -0.003)	-0.04256	(-0.075, -0.010)
Alt.Wage	0.72145	(0.616, 0.827)	1.26580	(1.175, 1.356)
Lagged Wage	0.43003	(0.351, 0.509)		
Union 1985/92	0.15477	(0.110, 0.200)	0.27154	(0.215, 0.328)
Union 1993/97				
Ind.31	-0.01809	(-0.176, 0.140)	-0.03174	(-0.328, 0.265)
Ind.32	-0.22796	(-0.384, -0.072)	-0.39995	(-0.722, -0.078)
Ind.34	-0.11797	(-0.246, 0.010)	-0.20698	(-0.451, 0.037)
Ind.35	0.09557	(0.031, 0.159)	0.16767	(0.062, 0.273)
Ind.36	-0.97846	(-1.585, -0.372)	-1.71670	(-2.756, -0.677)
Ind.38	-0.07955	(-0.215, 0.056)	-0.13957	(-0.387, 0.108)

Note: Industries are food, beverage & tobacco (31); textiles and apparel (32); paper (34); chemicals (36); nonmetallic minerals (36); metal products (38).

[13] These effects are calculated at the mean value of UNION

As almost every parameter changed, a simulation was done using both models in order to capture all possible effects. First, the wage was calculated for the period 1985-1997 using an ARIMA(4,1,0) model estimated using data for 1975-1984. Comparing the average value of the estimated wage with the actual average value, the result is that wages were 46 percent higher than what they would have been had no institutional changes occurred. For 1975-1984, the same exercise shows that actual wages in the period were 18 percent lower than what they would have been if there had been bargaining over wages and a union density equal to its average value in 1985-1997 (see Figures 7 and 8).

Figure 7
Labor costs 1975-1984 assuming the existence of unions

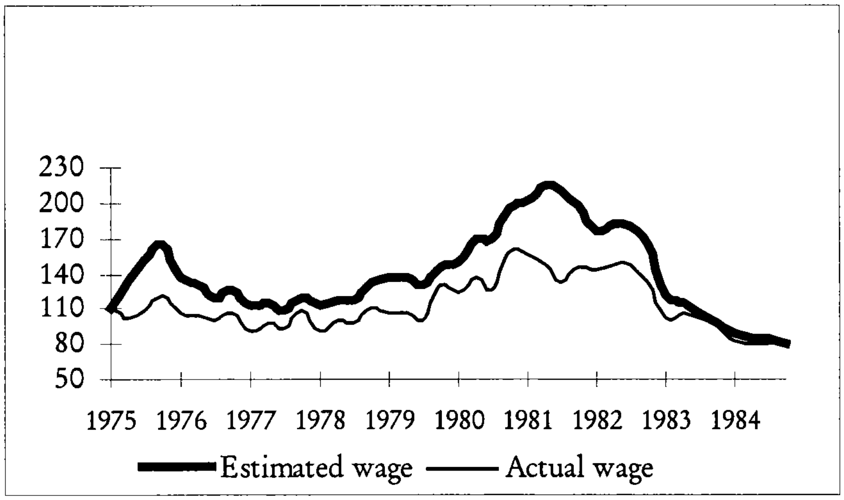
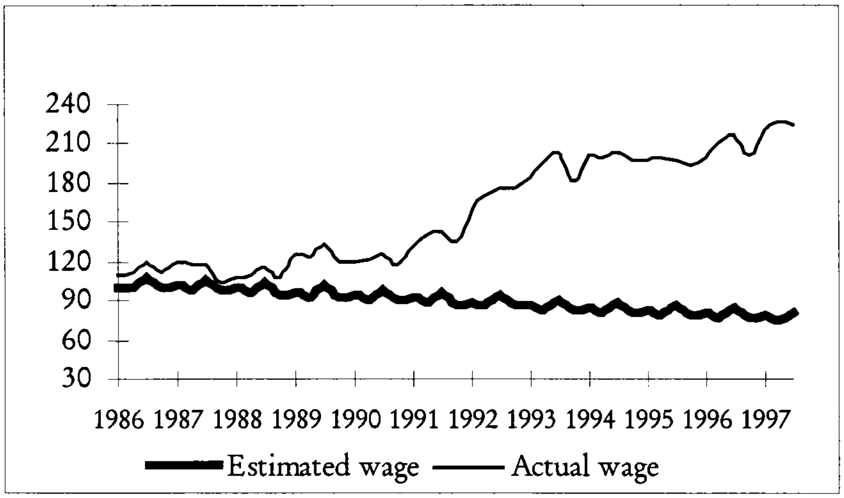


Figure 8
Labor costs 1986-1997 assuming there were no unions



Second, using actual wages and the two specifications of the labor demand equation, the estimated effect of the different regimes on labor demand is that the employment level in 1985-1997 was 9 percent higher than what it would have been according to the 1975-1984 model. This is the combined effect of the decrease in the output and wage parameters. Accordingly, in 1975-1984, employment would have been 5 percent higher than its observed level if elasticities had been those stemming from the bargaining model (see Figures 9 and 10).

Figure 9
Employment 1975-1984 assuming the existence of unions but using actual wages

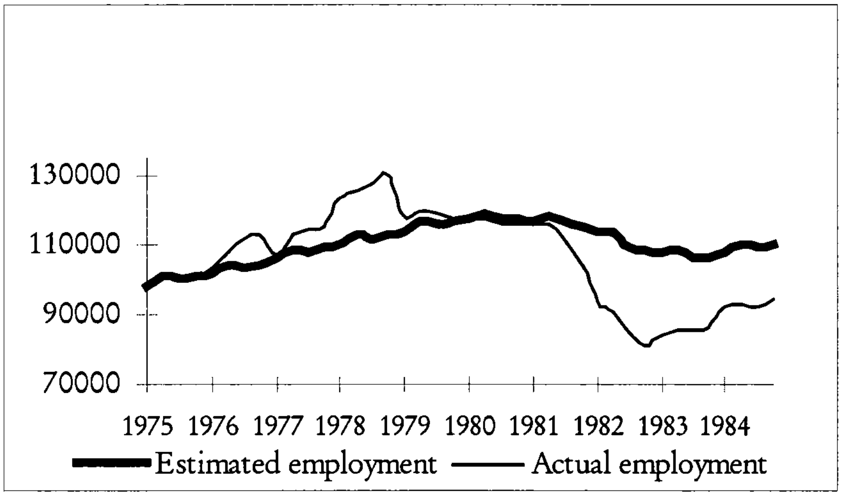
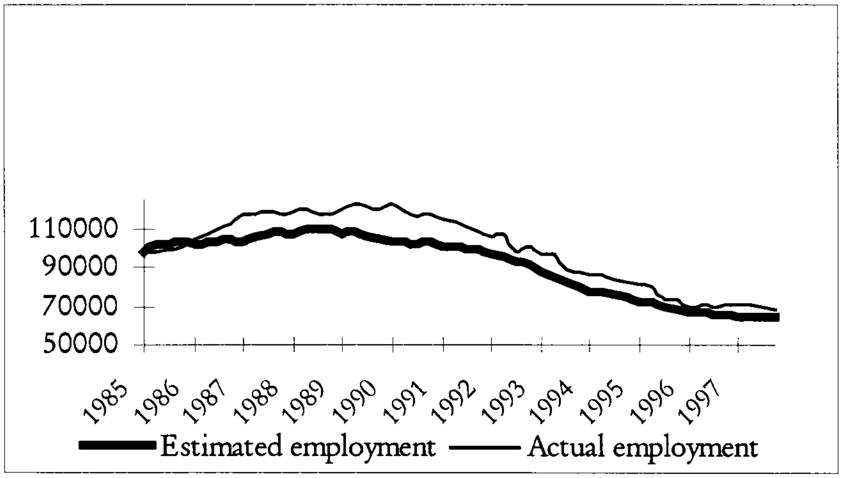


Figure 10
Employment 1986-1997 assuming there were no unions but using actual wages



Finally, considering both the estimated wage level and the change in elasticities, the employment level in 1985-97 was 24 percent lower than what it would have been if wages had followed the 1975-84 ARIMA(4,1,0) model and elasticities had been those according to the 1975-84 labor demand equation. In 1975-1984, on the contrary, if wages had been those predicted by the bargaining model and elasticities had had the values estimated with this same model, then the employment level would have been 1 percent lower than what it actually was (see Figures 11 and 12). In summary, unions could have prevented wages to fall as much as they did before 1985 at the cost of a 1 percent employment loss; while if unions had not been reinstated, employment would have been 24% higher but at the cost of a much lower level of earnings.

Figure 11
Employment 1975-1984 assuming the existence of unions

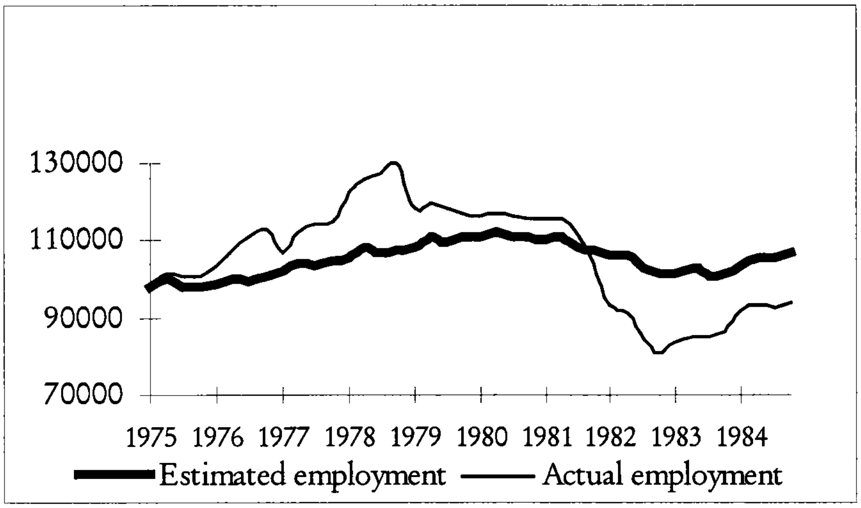
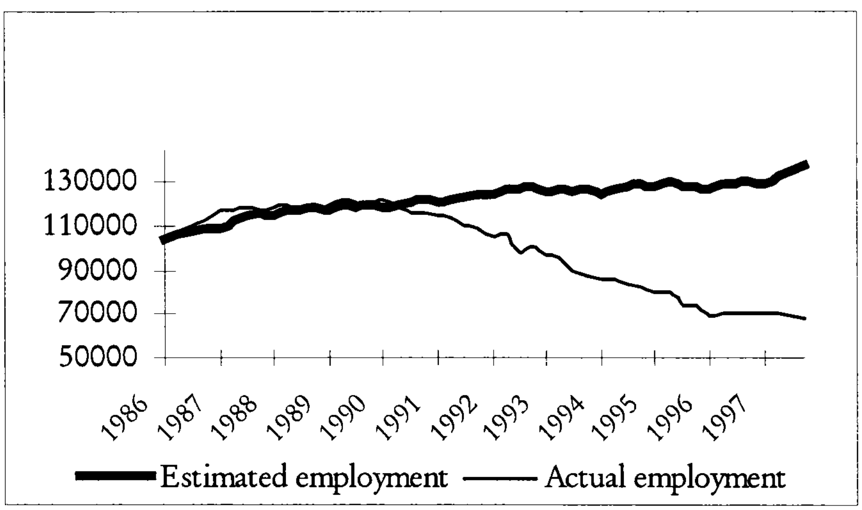


Figure 12
Employment 1986-1997 assuming there were no unions



The above figures should be considered only as rough indicators. The ARIMA(4,1,0) model for wages, as well as the 1975-1984 model for employment, do not take into account the estimated shifts in these variables that took place in 1993. The employment shift might be linked to structural reforms and productivity gains brought by increased openness, whereas the wage shift seems to reflect a rational response by unions to a changed economic environment. In any case the simulations do not take these factors into account and thus are likely to overstate union impact on wages and employment for 1993-1997.

All the results discussed above stem from a model for the whole manufacturing sector using industry data, and in which output and wage elasticities of labor demand were assumed to be the same for all industries. A natural question is if this last assumption holds and, if not, if it is biasing the results significantly. To address the issue, all the coefficients were allowed to vary per industry in both subperiods and the restriction imposed was tested. In 1975-84, the hypothesis of common elasticities and speed of adjustment was not rejected. For 1985-97 the wage elasticity and the lagged employment coefficient were statistically equal among industries while a unique output elasticity was rejected. Paper and chemicals have a significantly smaller elasticity. However, the average elasticity for the manufacturing sector estimated using this model is only slightly higher while the wage elasticity and the parameter accounting for the speed of adjustment do not show important biases[14]. Although such similarity between industries is not expected to hold *a priori*, the statistical result supports the estimation procedure followed using the pooled cross section – time series data set. Further, the decline in the elasticities holds, no matter the amount in which they decreased might be overstated.

Even though no direct bargaining over employment has been observed, all our findings suggest that unions have had an effect on labor demand. This has taken place through two mechanisms. First, re-unionization changed the way wages were set. Bargaining over wage levels has been done taking into account the likely effects on the labor demand schedule and outside opportunities for those that would eventually be unemployed. Industries that have been most exposed to competition have registered lower mark-ups than the rest. Union membership, which has declined systematically all along the period, raised the mark-up during the eighties. At the beginning of the nineties, and probably as a consequence of the progressive decentralization of bargaining and non-enforcement of contracts, this effect has vanished in some industries while in others it has even become negative. Increased openness also has tempered wage demands by unions.

Second, unions have effectively altered the labor demand choice set for employers. Output and wage elasticities have gone down and union resistance is one of the probable causes. As unions forced wages up and more limits were posed to pass that increase onto prices, firms have been forced to adjust employment to cyclical variations of demand less than before. Further, expected union resistance has been probably in the root of a smaller adjustment of employment to wage increases. As a result of all these changes, wages are higher and employment is lower today than what they would have been if no institutional changes had taken place.

[14] Results are available upon request.

VI. The effect of unions on employment adjustment: other evidence

Two lines of research were followed in order to further analyze the impact of unions on the labor market, focusing on how they affect employment adjustment. In the first approach, a dynamic labor demand model was estimated to determine how much of an impact unions have on the employment path to equilibrium. Second, the effect of unions on layoff rates was examined, performing the analysis for the whole economy and not just for the manufacturing sector.

VI.1. Dynamic labor demand

Although the conventional wisdom is that unions lengthen the time needed for labor demand adjustment, the estimates in Table 5 show no change between 1975-84 and 1985-97. It is possible that a simple before-after comparison fails to yield an accurate assessment of union impact because of other changes that were taking place in 1985-97. In particular, the opening of the Uruguayan economy could have shortened adjustment lags at the same time that unionization was lengthening them.

A useful approach to test for this is to consider that the speed of adjustment is itself a function of some variables related to changing the level of employment. This would imply that in equation (1.5) $\lambda = \lambda(\mathbf{ac}_t)$, with $\mathbf{ac}$ a vector of variables accounting for adjustment costs, that would hence interact with lagged employment (Burgess, 1988; Burgess, 1989; Burgess and Dolado, 1989). Variables that will be included are union density and the openness ratio. Union density should reflect implicit costs of firing workers as well as potential restrictions on the hiring of new workers. On the other hand, when there is more openness, it is quite likely that employers are subject to greater competitive pressure to compete both abroad and domestically, leading to shorter adjustment lags.

Three models are considered in Table 8: (1) percentage union interacts with lagged employment, (2) openness interacts with lagged employment, and (3) complete interactions among percentage union, openness, and lagged employment. Linear effects are included (and interacted in the third model); otherwise, the interaction coefficients could be interpreted as proxies for linear effects omitted from the model. As shown in columns 1 and 2, the simple interactions with percentage union and openness are positive and negative respectively, as expected, but estimated with little precision. In the model with complete interactions in column (3), the coefficient for lagged employment decreases with openness but also with percentage union. In any case, neither coefficient is estimated with acceptable degree of precision.

Openness also was included in the wage bargaining equation and consistently had a negative effect on the negotiated wage. An increase in openness from 30 to 60 percent is associated with a 1.7 percent reduction in wages.

Table 8: Estimates of dynamic labor demand equations, 1985-1997

	(1)	(2)	(3)
N_{t-1}	0.876	0.889	0.902
	(0.028)	(0.025)	(0.038)
Q_t	0.056	0.055	0.073
	(0.016)	(0.017)	(0.019)
W_t	-0.034	-0.032	-0.060
	(0.015)	(0.016)	(0.018)
N_{t-1}*UNION	0.001		-0.065
	(0.028)		(0.059)
UNION	0.005		0.522
	(0.266)		(0.532)
N_{t-1}*OPEN		-0.027	-0.102
		(0.024)	(0.052)
OPEN		0.248	0.908
		(0.222)	(0.474)
OPEN*UNION			-0.882
			(1.226)
N_{t-1}*OPEN*UNION			0.108
			(0.134)
SEE	0.044	0.044	0.043

Note: each equation also contains three seasonal dummies, five industry dummies, a shift parameter for the 1993-97 period, and a constant. Each equation is estimated by 3SLS along with a wage equation that includes the same variables as in column 2 of Table 5.

These results are not as robust as those related to the static labor demand. However, as a first approach to employment dynamics, they do suggest that low openness in 1985-1992 led to longer lags for employment adjustment, compared to 1975-1984. However, with the increase in openness in the 1990s, the lags quickly shortened and by 1997 were shorter than they were in 1984. The overall effect for the entire 1985-1997 period is an average lag similar to that prevailing for 1975-1984, which is what we reported in Table 5. Regarding the union effect, no conclusions can be drawn at this stage. A more complex specification for the λ parameter might be needed, possibly accounting for different effects per industry from 1993 onwards, but this is beyond the scope of this paper.

VI.2. Labor mobility

The results in Section V showed that labor demand became less responsive to changes in output and wages after the return of collective bargaining in 1985. What we do not know is if unions affect employment adjustment, given that the results in Section VI.1 were not robust. This is an important question because the success of any policy designed to make the Uruguayan labor market more flexible will hinge on the precise mechanisms through which unions influence employment.

Unionization could be correlated with slower employment adjustment because higher wages for union members virtually eliminate all voluntary turnover. In this case policy would have to be directed toward the bargaining power of unions. Another possibility is that union threats of wildcat strikes are sufficient to prevent management from ever conducting layoffs. In this case policy would need to be targeted at strike behavior.

We do not have data on flows in and out of the firm that would allow us to decompose employment changes into changes in hiring, layoffs, and quits (including retirement) and get more meaningful insights into union influence. As a second-best strategy, this section examines household survey data on unemployment, quits, and layoffs. If unionization has resulted in smaller responses of employment to changes in wages or output, this effect should be echoed by some combination of a lower hiring rate, a lower quit rate, and a lower layoff rate. The emphasis here will be on the effect of unions on layoff rates[15]. The analysis will be done using data for all economic sectors, not just manufacturing. Because of possible ambiguity in the responses to questions about events initiating unemployment, we report results not just on layoffs, but also for all unemployment and quits resulting in unemployment. In addition to gaining insights into how unions influence employment adjustment, these results will help determine whether the conclusions for manufacturing industries are generalizable for the entire economy.

Further, there is a sizable literature on how unions affect various forms of turnover, which provides some insights into models of union behavior. It has been well established since Freeman (1980) that unionization is associated with much lower quit rates. The reasons for this correlation remain unsettled. Freeman argues that if unions provide employees with a constructive channel for settling disagreements at the workplace, they will be less likely to quit. Pay compression within unionized workplaces further reduces the quits of employees who receive the largest relative benefits - the least skilled and the youngest. Others have argued that the correlation simply reflects the fact that the wage level is an inadequate measure for the rents received by union members; with a better measure of such rents, the impact of unions would vanish.

The impact of unions on layoff rates is more difficult to predict. In a framework where the welfare of each union member gets equal weight in the aggregate union preference function, one would expect the union to push for job security for all members. Today, however, most models presuppose the existence of politically dominant coalitions of workers within the union, usually the most senior members. In such a setting, only members of the dominant group are sheltered from layoffs. A further complicating factor is the availability and level of unemployment benefits. In the US unions have traditionally used layoffs by seniority and unemployment benefits to buffer their most senior members from economic fluctuations. Medoff (1979) found that layoff rates are actually higher for union members than nonunion workers. However, in the public sector in the US, Allen (1988) found that union members were less likely to be laid off than nonunion workers.

The unemployment insurance system in Uruguay provides a much lower replacement rate of income (50%-75% of the previous monthly wage, capped up to seven times the minimum) than comparable systems in the US and EC, so one might expect Uruguayan unions to place a higher priority on avoiding layoffs at all costs. Unions can significantly increase the transactions costs associated with

[15] Data on new hires are only available since 1991 and quits are only observed if they result in unemployment.

layoffs. The obligations of a nonunion employer are limited to severance pay. Unions can create additional costs, including work stoppages and slowdowns.[16]

Due to the changes that have taken place in the structure of bargaining, it is important to allow the effect of unions on layoffs to vary within the 1985-97 period. This is done by estimating a separate model for each year.

Models and data

To estimate the impact of collective bargaining on labor mobility, ideally one would estimate hazard rates for employment, both overall and separately for quits and layoffs. There are no panel surveys of households in Uruguay and no repeated cross sections with data on completed spells of employment (or unemployment for that matter). The monthly household survey can be used to identify experienced workers who are unemployed, the sector and industry of their last job, and the reason for leaving their last job. Workers are defined as laid off if they say they lost their last job because they were "fired" or "plant or company closed." Workers who say they were suspended from work or who are receiving unemployment insurance are excluded from this definition because of the ambiguity of whether a separation has occurred. All other separations leading to unemployment (including those who gave a reason for leaving their last job that was coded as "other") are defined as quits.

A new questionnaire was adopted for the household survey in 1991. The survey items used to identify layoffs were not changed, but the wording and number of options for quits were substantially altered. This makes it impossible to estimate models of specific types of quit behavior, such as quits for low pay or quits to return to school. As long as quits are defined broadly, there does not seem to be any significant break in the series between 1990 and 1991 (see Table 9).

The quit-layoff distinction can be problematic empirically. Interviews with employers about the reasons for a separation would no doubt yield different answers than the household survey. In addition the timing of the decision and the stated rationale can conceal as well as reveal, e.g., workers may quit to avoid any stigma associated with dismissal. Accordingly, this study also examines the odds that an experienced worker is unemployed, ignoring the stated reason. This latter measure includes individuals who have been suspended or who are receiving unemployment benefits.

A clear limitation of this approach is that we do not observe cases where persons leave their job without an intervening spell of unemployment. This is unlikely to be too much of a problem with the analysis of layoffs – even with advance notice (which is not required in Uruguay), very few job losers are able to find a new position before their old one ends. Difficulties are more likely in interpreting the results on quits. Simple errors in measurement make this study less likely to reject any null hypothesis. Nonetheless, we must emphasize that the results here deal with only one dimension of quit behavior – quits followed by unemployment – and may not be generalizable to all quits.

Probit equations for unemployment, quits, and layoffs are estimated over all experienced wage and salary workers in the household survey for 1981, 1982, 1984, 1986, 1990, and 1991-1996. Data for 1981 and 1982 are available only for the second half of the year. Union membership is not measured in

[16] Initially we planned to include union contract data on severance payments beyond those required by national regulations as part of this exercise. Upon studying the contracts, however, we learned that such payments are negotiated on an as-needed basis, rather than being part of an explicit contractual arrangement. These arranged payments can be quite substantial, as shown in our case study of the banking sector in Cassoni *et al.* (1995).

the household survey. It was calculated as stated in section IV for each two-digit industry in manufacturing and for one-digit industries outside manufacturing.

Other variables included in the analysis include employment in the public sector, age, age squared, sex, years of schooling, percentage informal in the last industry of employment (as indicated by lack of health coverage[17]), marital status (one variable indicating married, spouse present), industry (two variables flagging manufacturing and construction, the industries with the most layoffs and unemployment), and quarter (three variables).

Means and trends

Descriptive statistics for the main dependent and explanatory variables are reported in Table 9.

Table 9 Means and standard deviations of variables used in labor mobility analysis

Year	1981	1982	1984	1986	1990	1991	1992	1993	1994	1995	1996
Unemployed	0.068	0.124	0.118	0.090	0.081	0.077	0.082	0.074	0.080	0.104	0.121
(yes=1)	(0.251)	(0.329)	(0.322)	(0.287)	(0.273)	(0.267)	(0.275)	(0.262)	(0.272)	(0.305)	(0.326)
Laid off	0.009	0.026	0.028	0.017	0.015	0.021	0.023	0.019	0.024	0.036	0.049
(yes=1)	(0.097)	(0.158)	(0.166)	(0.128)	(0.121)	(0.143)	(0.150)	(0.136)	(0.154)	(0.185)	(0.216)
Quit &	0.045	0.066	0.077	0.066	0.054	0.054	0.058	0.054	0.055	0.068	0.071
Unemployed	(0.208)	(0.248)	(0.267)	(0.248)	(0.227)	(0.227)	(0.235)	(0.226)	(0.227)	(0.252)	(0.256)
(yes=1)											
Percentage	0.340	0.326	0.324	0.299	0.261	0.248	0.235	0.222	0.216	0.208	0.168
union	(0.224)	(0.218)	(0.211)	(0.176)	(0.158)	(0.157)	(0.151)	(0.146)	(0.154)	(0.158)	(0.142)
N	5385	5329	10196	10088	9233	10794	10668	10822	10944	11354	11475

Unemployment ranges between 7 and 12 percent over the sample period. These figures are lower than reported unemployment rates, mainly because first time job seekers are not included in our sample (also we exclude self-employed and unpaid workers). The peak periods of unemployment are the global recession of 1982-84 and 1995-96, when Argentina was experiencing very high unemployment. Quit unemployment tracks the overall unemployment rate fairly closely.

The percentage of the labor force on layoff is much higher in the 1990s than in the 1980s, reaching a maximum of 5 percent in 1996. This is almost twice as large as the layoff percentage in 1982, even though the overall unemployment rates for experienced workers in the two years are both around 12 percent. Layoff percentages in 1992 and 1994 averaged 2.4 percent, despite low overall unemployment near 8 percent. In comparison, the layoff percentage in 1982 was 2.6 percent, although the unemployment rate was 12.4 percent.

[17] There is legal mandatory health coverage by social security for those who work in the private sector. The Household Survey poses questions that refer specifically to health. Hence, those that report not having the mandatory coverage are defined as informal workers for the purposes of this study.

Union density varies significantly through the sample period. Keep in mind that there was no collective bargaining through 1984; the values reported in Table 9 for 1981-1986 are based on 1987 data. They are used to control for unobserved industry effects that are correlated with union density. (They vary from year to year because of changing industry composition of employment.) The mean value of percentage unionized dropped from 30 percent in 1986 to 17 percent in 1996. The decline is fairly gradual, except for a sharp four-point drop between 1995 and 1996. Union density declined by 10 or more percentage points in food and beverages, textiles and clothing, transportation and communication, and financial services in the 1990s.

Probit results

The impact of unions varies markedly over the time period (Table 10). It is no surprise that in 1981 and 1982 the percentage union variable (which is acting in those years as a proxy for union sentiment or conditions making a sector conducive to union organization) is uncorrelated with layoff, quit, or unemployment odds. By 1984 unions had become a powerful political force, organizing strikes and demonstrations in an effort to pressure the military government to step down. Percentage union is associated with lower odds of unemployment in 1984, all of which results from fewer quits into unemployment in unionized industries. This behavior probably reflects anticipation of a return to democracy and the restoration of collective bargaining. There is also some indication that layoff rates in sectors of the economy that were to become unionized had already become slightly lower than layoff rates in sectors that were to stay nonunion. The union coefficient in the layoff probit increased from –0.005 in 1982 to –0.016 in 1984 (although the latter effect is not estimated with a high degree of precision).

The picture changes further by 1986, when percentage union (based on 1987 data, as in 1981-82 and 1984) is now strongly associated with lower layoff odds. Unionization is associated with a 1 to 2 percentage point reduction in layoff odds in most years between 1986 and 1994. This may seem modest in absolute terms, but keep in mind that the mean layoff rate varied between 1.5 and 2.4 percent over this period. Assuming a mean unionization of 25 percent, a mean layoff rate of 2 percent, and a union-nonunion difference in layoff rates of 1.5 percent, this implies that the odds of layoff for a union worker are 0.9 percent versus 2.4 percent for a nonunion worker.

Even though union density was declining and unionized companies were more exposed to global competition, the estimated effect of unions on layoffs actually stayed quite strong in 1995 and 1996. The aggregate layoff rate increased to 3.6 percent in 1995 and 4.9 percent in 1996. The impact of unions increased to 3.3 percent in 1995. Based on this coefficient and the means of the key variables, this result implies that the layoff odds for a union worker were 1.0 percent versus 4.3 percent for a nonunion worker. Compared to previous years, all of the increase in layoff risk was borne by nonunion workers. The results for 1996 are too strong; the union coefficient in the layoff and unemployment probits is larger than the mean layoff and unemployment rate.

The increase in the union-nonunion gap in layoff rates might seem puzzling in light of the decline in union density. One might argue that union members are self-selecting into firms with lower turnover, but the model controls for the odds that a worker is in the informal sector, where mobility would be greatest. Another possibility is that the remaining union workers have higher tenure (relative to nonunion workers) in 1995 and 1996 than in previous years, but the data show that this difference (1.2 years) is the same in 1995 and 1996 as in 1991-1994. The most logical possibility is that the unions that did survive until 1995-96 were the most powerful ones. Layoff rates in those firms stayed at 1 percent while those in the nonunion sector increased dramatically.

Table 10. Transformed coefficients and standard errors of union coefficient in probit estimates

Year	Unemployed (yes=1)	Laid off (yes=1)	Quit & unemployed (yes=1)
1981	0.008	-0.006	0.011
	(0.024)	(0.006)	(0.020)
1982	0.059	-0.005	0.029
	(0.032)	(0.013)	(0.024)
1984	-0.109	-0.016	-0.066
	(0.026)	(0.010)	(0.021)
1986	-0.065	-0.027	-0.025
	(0.025)	(0.009)	(0.021)
1990	-0.065	-0.011	-0.018
	(0.021)	(0.006)	(0.016)
1991	-0.032	-0.008	-0.023
	(0.019)	(0.008)	(0.016)
1992	-0.053	-0.016	-0.034
	(0.021)	(0.009)	(0.018)
1993	-0.054	-0.014	-0.035
	(0.019)	(0.007)	(0.016)
1994	-0.040	-0.021	-0.004
	(0.019)	(0.008)	(0.016)
1995	-0.088	-0.033	-0.047
	(0.026)	(0.013)	(0.021)
1996	-0.135	-0.084	-0.041
	(0.027)	(0.016)	(0.020)

Note: Coefficients indicate change in probability resulting from a change in fraction unionized; standard errors appear in parentheses. Control variables include fraction employed in informal sector (as proxied by health insurance coverage), employment in public sector (yes=1), age, age squared, male, years of schooling, married, industry (dummies for manufacturing and construction), and quarter.
Source: Household Survey, INE, 1981-1997

The overall implication of these results is that the layoff odds of a worker in a unionized sector were less than one percent from 1986 through 1996. Any fluctuation in aggregate layoff rates reflected adjustments by nonunion employers.

Even with near-zero layoff rates, unionized employers still have some freedom to make changes in employment if the quit rate is sufficiently high. This is not the case in Uruguay. Instead, employer flexibility in unionized enterprises is especially hampered by very low quit rates. The size of the effect varies from year to year between 1984 and 1994, but in most years it is close to a 3 percentage point difference in the odds of quitting and becoming unemployed between workers in a fully unionized and a fully nonunion industry. The impact of unions on quits resulting in unemployment rises in 1995 and 1996.

In summary, the change in labor demand elasticities observed at the industry level of aggregation is no optical illusion. What we see at the micro, individual worker level matches what we see at the industry level – employment now adjusts less than it would have in the absence of unions. What makes these results all the more convincing is that the same measure of union density used for 1986-1996 has no effect on unemployment, quits, or layoffs when applied retrospectively to 1981-1982. Effects on quits (and a weak effect on layoffs) begin to appear in 1984, as unions became more active, but the full effect on layoffs is not present until bargaining had officially resumed.

VII. Conclusions

This study has examined a unique situation in Uruguay where before-after comparisons about the impact of collective bargaining can be made. During the period under study there were three distinct regimes: (1) 1975-1984 when bargaining was banned, (2) 1985-1991 when there was tripartite bargaining, and (3) 1992-1997 when there was bargaining without government involvement. During the third regime the economy became much more open, which would presumably also have an effect on bargaining results.

We have reported strong evidence of a change in economic behavior after 1985. Recursive residuals show structural shifts in five of six industries with the shifts coming at about the same time as the regime changes. These breaks are also significant in a model specified in differences. Cointegration of employment, output, and labor costs is rejected for the whole period. Wages are exogenous to employment before 1985, but not afterwards.

Based on this evidence we estimated a standard IVE labor demand model for 1975-1984 and a right-to-manage bargaining model for 1985-1997. The results showed that the long run wage elasticity of labor demand and the employment-output elasticity fell sharply. The bargaining model results indicated that unions significantly raised wages in 1985-1992. Afterwards the change in bargaining structure and increased openness had a pronounced effect on bargaining outcomes. Labor demand shifted to the left from 1993 onwards. The union wage differential vanished in 1993 in four industries where there were sharp increases in openness and sharp declines in percentage union. Wages in the chemical and oil industry were not affected very much. Although that industry became more open, it has remained heavily unionized, which is no doubt a consequence of state ownership.

What would have happened to wages and employment had the ban on unions been maintained? To build a counterfactual, we estimated an ARIMA(4,1,0) model of wages for 1975-1984, which was used to project a wage path through 1997. Actual wages have been significantly higher than the simulated "nonunion" wage, based on average values for 1985-1997. Taking into account the higher wage level and the reduced elasticities, employment in 1985-1997 was lower than it would have been if unions had not returned.

Because of possible skepticism regarding the use of industry rather than establishment data, we also examined the effect of unionization on turnover, using household survey data. These results showed that workers in unionized industries were much less likely to be laid off starting in 1985 than workers in nonunion industries. Before 1985, no such pattern is present, indicating that unionization is not acting as a proxy for other industry characteristics associated with high layoff odds.

The following picture emerges from these results. Unions returned on the scene as a political and economic force in 1985 and for two years more than half of Uruguay's workers were union members. Union density settled down to about 40 percent in 1987-1992 and unions were able to successfully negotiate higher wages and were able to protect against job loss by reducing employment elasticities. It would be useful to know the precise mechanisms through which unions reduced employment adjustment. It is doubtful that unions had much effect on consumer choices, since no steps were made to expand state ownership or de-liberalize trade when unions returned. The most likely channels through which unions had an impact were restrictive work practices and the threat of strikes or slowdowns in situations where layoffs were thought possible.

In the 1990s the end of tripartite bargaining, trade liberalization, and the recession in Argentina forced unions to make compromises at the bargaining table. Faced with an adverse shift in labor demand, unions reduced their wage demands to preserve jobs. Percentage union declined to 20 percent as many unionized establishments were no longer economically competitive and others were forced to increase productivity to survive. When a few more years of data become available, it would be fruitful to determine if elasticities had returned to their 1975-1984 values.

This study has focused on wage and employment effects of unions. To get a more complete view of the overall impact of unions on the Uruguayan economy, further study of strikes would be necessary to get a lower bound estimate of work hours lost to strikes. These would include not just strikes against employers in the context of bargaining over wages, but also strikes – including employer-specific, sector-specific and general strikes – that take place when a bargaining agreement is in effect.

Finally, this study has not discussed the benefits that result from successful union-management cooperation. Future work should carefully examine this matter. Not only because of a need to focus as carefully as possible on labor demand and bargaining, but because the structure of the system of labor relations has become increasingly decentralized in Uruguay, and unions are apparently changing their utility function when they bargain at the firm level under competitive pressures.

BIBLIOGRAPHY

Aghion, P. and G. Saint-Paul, 1993. "Uncovering Some Causal Relationships between Productivity Growth and the Structure of Economic Fluctuations: a Tentative Survey". *Working Paper* 4603, NBER.

Allen, S., 1986. "Union Work Rules and Efficiency in the Building Trades," *Journal of Labor Economics*, Vol.4/2: 212-242.

Allen, S., 1988. "Unions and Job Security in the Public Sector," in *When Public Sector Workers Unionize* (pp. 271-296), Richard B. Freeman and Casey Ichniowski (Eds.) Chicago: University of Chicago.

Allen, S., A. Cassoni, and G.J. Labadie, 1994. "Labor Market Flexibility and Unemployment in Chile and Uruguay". *Estudios de Economía*, Vol. 21, special number: 129-146, Santiago de Chile.

Allen, S., A. Cassoni and G.J. Labadie, 1996. "Wages and Employment after Reunionization in Uruguay". *Cuadernos de Economía*, Año 33, No. 99: 277-239, Santiago de Chile.

Banerjee, A., J. Dolado, J.W. Galbraith and D. Hendry, 1993. *Co-integration, Error-Correction, and the Econometric Analysis of Non-stationary Data.* Oxford University Press.

Binmore, K., A. Rubinstein and A. Wolinsky, 1986. "The Nash Bargaining Solution in Economic Modelling". *Rand Journal of Economics*, 17: 176-88.

Blanchard, O.J. and D. Quah, 1989. "The Dynamic Effects of Aggregate Demand and Supply Disturbances". *The American Economic Review*, Vol.79/4: 655-673.

Blanchard, O.J. and L.H. Summers, 1986. "Hysteresis and the European Unemployment Problem". In S. Fischer (ed.), *NBER Macroeconomic Annual*, Vol.1, Cambridge Mass.: MIT Press.

Blanchflower, D.G., 1984. "Union Relative Wage Effects: a Cross-section Analysis using Establishment Data". *British Journal of Industrial Relations*, 22: 311-332.

Blanchflower, D.G., N. Millward and A.J. Oswald, 1991. "Unionism and Employment Behaviour". *The Economic Journal*, 101: 815-834.

Boal, W.M. and J.H. Pencavel, 1994. "The Effects of Labor Unions on Employment, Wages and Days of Operation: Coal Mining in West Virginia". *Quarterly Journal of Economics*, 109: 267-298.

Booth, A.L., 1995. *The Economics of the Trade Unions.* Cambridge University Press.

Burgess, S.M., 1988. "Employment Adjustment in the UK Manufacturing". *The Economic Journal*, 98: 81-93.

Burgess, S.M., 1989. "Employment and Turnover in UK Manufacturing Industries, 1963-82". Oxford *Bulletin of Economics and Statistics*, 51/2: 163-192.

Burgess, S.M. and J. Dolado, 1989. "Intertemporal Rules with Variable Speed of Adjustment: An Application to UK Manufacturing Employment". *The Economic Journal*, 99: 347-365.

Carruth, A. and A. Oswald, 1987. "On Union Preferences and Labour Market Models: Insiders and Outsiders". *The Economic Journal*, 97: 431-45.

Cassoni, A., G.J. Labadie and S. Allen, 1995. *Reforming the Labor Market in a Liberalized Economy*. Chapter Four: Uruguay. Ed. by Gustavo Márquez, Centers for Research in Applied Economics-Interamerican Development Bank, Washington, D.C.

Cassoni, A. and Z. Ferre, 1997. "Costos No Salariales en el Mercado de Trabajo del Uruguay". *Documento de Trabajo*, 8/97, Departamento de Economia, Facultad de Ciencias Sociales, Universidad de la República.

Cedrola, G., J. Raso and F. Perez Tabó, 1998. *La Negociación Colectiva en el Uruguay, 1996-1997*. I.L.O., Santiago (in press).

de Brun, J. and G.J. Labadie, 1997. "Mercado Laboral, Apertura y Recesión: La Experiencia Uruguaya de los Noventa". In: *Mercados laborales en los 90: cinco ejemplos de América Latina.* CIEDLA and Konrad Adenauer Foundation, Buenos Aires.

de Menil, G., 1971. *Bargaining: Monopoly Power versus Union Power*. Cambridge Mass.: MIT Press.

Edwards, S., 1995. *Crisis and Reform in Latin America.* Oxford University Press.

Engle, R.F. and C.W.J. Granger, 1987. "Cointegration and Error Correction: Representation, Estimation and Testing". *Econometrica*, 55: 251-276.

Ermida, O., G. Cedrola, J. Raso y F. Perez Tabó, 1997. "Estructura y Contenido de la Negociación Colectiva en Uruguay". *Document Series*, I.L.O., Lima.

Favaro, E. and C. Sapelli, 1986. *Shocks Externos, Grado de Apertura y Política Doméstica.* Banco Central del Uruguay.

Freeman, R., 1980. "The Exit-Voice Tradeoff in the Labor Market: Unionism, Job Tenure, Quits, and Separations," *Quarterly Journal of Economics*, Vol.94/4: 643-674.

Freeman, R. and J.L. Medoff, 1982. "Substitution Between Production Labor And Other Inputs In Unionized And Nonunionized Manufacturing," *Review of Economics and Statistics*, Vol.64/2: 220-233.

Freeman, R. and J.L. Medoff, 1984. *What Do Unions Do?* New York: Basic Books.

Fuller, W.A., 1976. *Introduction to Statistical Time Series.* New York: John Wiley.

Gillespie, C.G., 1991. *Negotiating Democracy*, Cambridge: Cambridge University Press.

Hamermesh, D.S., 1993. *Labor Demand.* New Jersey: Princeton University Press.

Hamermesh, D.S., 1995. "Labour Demand and the Source of Adjustment Costs". *The Economic Journal*, 105: 620-34.

Hamermesh, D.S. and G.A. Pfann, 1996. "Adjustment Costs in Factor Demand". *Journal of Economic Literature*, Vol.34/3: 1264-1292.

Handelman, H., 1981. "Labor-Industrial Conflict and the Collapse of Uruguayan Democracy", *Journal of Interamerican Studies and World Affairs*, Vol.28/4: 371-394.

Hausman, J.A., 1978. "Specification Tests in Econometrics". *Econometrica*, 46: 1251-1271.

Hirsch, B.T. and J.T. Addison, 1986. *The Economic Analysis of Unions: New Approaches and Evidence.* Boston: Allen & Unwin.

Johansen, S., 1988. "Statistical Analysis of Cointegration Vectors". *Journal of Economic Dynamics and Control*, 12: 231-254.

Layard, R., S. Nickell, and R. Jackman, 1991. *Unemployment. Macroeconomic Performance and Labor Market.* New York: Oxford University Press.

Lewis, H.G., 1986. *Union Relative Wage Effects: a Survey.* University of Chicago Press.

Lewis, H.G., 1990. "Union/Nonunion Wage Gaps in the Public Sector". *Journal of Labor Economics*, 8(1, part 2): S260-S328.

Lockwood, B. and A. Manning, 1989. "Dynamic Wage-Employment Bargaining with Employment Adjustment Costs". *The Economic Journal*, 99: 1143-1158.

Machin S. and S. Wadhwani, 1991. "The Effects of Unions on Organisational Change and Employment". *The Economic Journal*, 101: 835-854.

Medoff, J.L., 1979. "Layoffs and Alternatives under Trade Unions in U.S. Manufacturing". *The American Economic Review*, Vol.69/3: 380-395

Pencavel J., 1991. *Labor Markets under Trade Unionism: Employment, Wages and Hours.* Basil Blackwell Ltd.

Picardo, S., C. Daude and Z. Ferre, 1997. "Indice del Costo de la Mano de Obra: 1982-1995". *Documento de Trabajo* 1/97, Departamento de Economia, Facultad de Ciencias Sociales, Universidad de la República.

Rama, M., 1993a. "Organized Labor and the Political Economy of Product Market Distortions", *mimeo*, The World Bank - CINVE.

Rama, M., 1993b. "Institucionalidad Laboral y Crecimiento Económico en el Uruguay", *mimeo*, Academia Nacional de Economía, Montevideo.

Rama, M., 1994. "Bargaining Structure and Economic Performance in the Open Economy", *European Economic Review*, 38: 403-415.

Spanos, A., 1986. *Statistical Foundations of Econometric Modelling*. Cambridge Univ. Press.

White, H., 1980. "A Heteroskedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heteroskedasticity". *Econometrica*, 48: 817-838.