CHANGES IN OCCUPATIONAL MOBILITY, LABOUR REGULATIONS AND RISING PRECARIOUSNESS IN ARGENTINA

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Abstract

Labour turnover of low-tenure workers rose in Argentina during the second half of the nineties, precisely after important changes were introduced in labour regulations.

The analysis of exit rates indicates that the alterations in the labour market institutions apparently had no effect on labour mobility. This fact appears to be associated to the shortage of occupational opportunities which may have increased the incidence of unstable trajectories among those working without coverage. Moreover, the rising unemployment increased the participation of non-registered, highly-mobile workers. Hence, this is a process that by itself led to an increase in overall mobility.

Resumen

Durante la segunda mitad de los noventa se verificó en Argentina un incremento en la movilidad ocupacional que afectó principalmente a los trabajadores de menor antigüedad. Paralelamente, en este período se introdujeron importantes modificaciones en la regulación laboral.

El trabajo evalúa en que medida los cambios en la legislación explican la mayor inestabilidad laboral. Los resultados sugieren que las alteraciones en las instituciones del mercado de trabajo no han tenido influencia sobre la creciente intermitencia laboral, siendo la mayor precariedad y el aumento en las tasas de salida desde un puesto no registrado los factores que explicarían aquel fenómeno.

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Introduction

During the post-war period, Argentina registered moderate levels of open unemployment although the relatively important presence of precarious occupations –informal and non-registered-wage-earners[1] occupations- suggests that certain population groups were subjected to frequent changes in their labour situation. Some evidence coming from previous studies points to an increase in labour turnover during the second half of the nineties with respect to the levels prevailing during previous years.[2] There are, at least, two processes that could have led to such behaviour. On the one hand, changes in labour regulations were carried out by the middle of the decade; these changes involved the reduction of firing costs, the establishment of new types of fixed-term contracts with lower cost and the introduction of the trial period. On the other hand, the employment structure underwent modifications; especially, precarious employment increased its share in total employment and open unemployment rose.

This paper aims at analysing to what extent these two factors –the changes in regulations and in the employment structure– were associated to a rise of labour turnover during the second half of the nineties in Argentina with respect previous years. We study the transitions from occupations estimated with data from the regular household survey (Permanent Household Survey, known as the EPH in Spanish). Even though this type of data faces some limitations for measuring labour transitions, they provide reasonably good evidence of the occupational mobility characteristics. Due to micro–data availability, the analysis will be focused on the 1987-1999 period and on the Greater Buenos Aires area. The paper extends the results of previous researches at least in two dimensions. Firstly, it evaluates not only the effects of the changes in regulations but also the impact of other factors. Secondly, it takes into consideration the movements from jobs to any other state, whereas other studies[3] exclude the movements out of the labour force. In labour markets such as the Argentine, however, the latter are of great importance, mainly for women and young workers.[4]

The paper is divided into six sections. The first one analyses the evolution of the Argentine labour market during the nineties and includes a brief discussion of the main changes in labour regulations. Section 2 examines the methodology as well as the information source to be employed. The selection of the group of workers to be considered in the analysis of mobility is described in Section 3, while Section 4 provides a brief summary of the principal determinants of the probability of leaving a job. Section 5 focuses on the main objective of the paper, i.e. the analysis of changes in exit rates and the influence that the

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[1] In this paper, "Informality" is used to refer to own-account workers as well as those wage earners employed by small -"informal" firms (the ILO approach). Non- registered employees are those wage earners not covered by social security.
[2] See Beccaria and Maurizio (2001)
[3] Hopenhayn (2001) and Galiani and Hopenhayn (2000) also analysed the effects of regulations on instability in Argentina
[4] Similar studies for other Latin American countries can be found in Saavedra and Torero (2000), for Peru, in Kugler (2000) for Colombia, and in Calderón-Madrid (2000) for Mexico.

modifications in regulations and in the employment structure could have had on mobility. The conclusions drawn from the analysis are summarised in the last section.

1 – Labour market and changes in labour regulation in Argentina during the nineties

This section briefly summarises some characteristics of the labour market performance during the period of analysis, including a description of the main changes introduced in labour regulation, specially those potentially affecting occupational instability.

The late eighties were characterised by high macroeconomic instability; the inflation rates were extremely high–with peaks of hyperinflation in 1989 and 1990– and the GDP was stagnant. Real wages were consequently very low while, on the other hand, unemployment only grew slowly and remained around moderate levels (around 6%). Such performance was accompanied by growing hourly-underemployment and informality.

In 1991 a new set of short-term policies and structural reforms was implemented.[5] From this year onwards, important progresses were made towards macroeconomic stability: inflation was rapidly controlled and GDP grew significantly. However, the changes underwent by the productive structure during the initial years –mainly, between 1991 and 1994– in response to the adjustment program impeded the creation of enough new jobs to absorb the growing labour supply. Consequently, the unemployment rate rose, reaching 12% in 1994. On the other hand, as the inflation rate fell during 1991-93, real wages recovered from the very low figures registered during the high-inflation period.

The economic expansion was interrupted as the difficulties for obtaining international financing arose after the Mexican crisis towards the end of 1994. The pre-existing occupational problems were consequently deepened, bringing the unemployment rate close to the 20% mark in Greater Buenos Aires in 1995.[6] As the short-run difficulties in the external sector mitigated, a new period of economic growth started by late 1995, this time associated to a relatively high rate of employment growth. However, some developments in the international capital markets during 1998 once again limited the amount of external financing obtained by the country, which –given the size of the external debt– had an important impact on domestic output and led to a new period of falling GDP.[7]

Most of the additional occupations generated during the decade were precarious, even during the phases of GDP growth. Specifically, registered employees accounted for less than one quarter of the 1991-1999 net increase in total employment in Greater Buenos Aires. That performance led to a further reduction in the already low share of registered wage earners in total employment: from 46% to 43% between those two years.

The growth of the unemployment rate over the decade, as well as its short-run movements, were mainly caused by changes in the inflow rate into unemployment. Even though the average duration of unemployment episodes also rose, this factor was of secondary importance, as shown in Table 1. It has to be taken into account that although unemployment insurance was introduced in 1991, its coverage is low –only 5 to 8% of

[5] For a description of the policies set and also of the macroeconomic behaviour of the nineties, see Heymann, 2001.
[6] A detailed analysis of the labour market variables can be found in Damill, Frenkel y Maurizio (2002).
[7] This phase lasted until the end of 2001, when things took a turn for the worst given the financial crisis, the debt default and the devaluation of the peso. Once again, labour indicators deteriorated during this period

unemployed population during most part of the decade.[8] To a large extent, this was precisely due to the relatively reduced share of registered wage-earners in the occupational structure.[9]

Table 1
Unemployment in Greater Buenos Aires: rates, duration and flows
(assuming steady state)

	1991	1999	Change (%)
Average duration of completed spell (months)	2.9	3.1	25.0
Inflow rate (%) [10]	1.8	4.7	188.9
Unemployment rate (%)	5.3	14.7	247.2

Source: Authors' estimation based on data from the PHS

The rise of the inflow-to-unemployment rate is already suggesting changes in the intensity and characteristics of labour mobility. Evidence discussed in previous papers (Beccaria and Maurizio, 2001 and Beccaria, 2001) also points towards the same result: the proportion of people who remain continuously employed during 18 months fell from 61% in the second half of the eighties, to 55% ten years later.

The modifications introduced in labour regulations were an important feature of the nineties. The flexibilisation of labour regulations is a typical component of the structural reform programs adopted in many Latin American countries during the nineties. A new Employment Act was enacted soon after the first new policies were implemented in 1991, and four years later, when unemployment increased sharply, new modifications of greater significance were introduced. The following were the most prominent measures implemented in 1995:

- Employer's social security contributions were reduced; the proportion reached 40% in 1996 and the following years;
- Time flexibilisation agreements could be reached through collective bargaining; this would permit a reduction of overall labour costs;
- Severance payments were reduced for workers with less than two years of tenure: in cases dismissal without just cause, the payments were equivalent to one-month salary for each year of tenure. Until 1995, there had been a minimum severance payment of two months, i.e. those workers with tenure between three[11] and 24 months were eligible for a payment of two months when fired. After that year, this provision was changed and the payment became strictly proportional to the number of months of tenure (i.e. 1/12 of a month salary for each month of tenure);
- New types of fixed-term contracts were authorised in 1991, subject to lower firing costs and employers' social security contributions than those of the typical contracts of

[8] These nation-wide figures have been calculated by comparing the number of unemployed workers receiving unemployment benefits (provided by the Ministry of Labour, see www.anses.gov.ar), and the total number of unemployed workers estimated from the PHS.

[9] Only persons who have been fired from a registered job and who have made social security contributions for a minimum of 12 out of 36 months previous to dismissal are eligible for the unemployment benefit. It excludes lay-offs from sectors such as construction, domestic housekeeping services, the public sector and rural activities, since they have specific regulations. The replacement rate is 50% during the first four months, 42% for the following five months and 35% for the other three. However, the monthly benefit has a floor and a ceiling, which were equivalent to 1 and 1.5 minimum wages, respectively, during most part of the nineties.

[10] As it is usual, entry rate was proxied by the ratio between the number of unemployed with durations of one month or less and the labour force. Average duration of completed spells are computed as the ratio between total unemployed and the number of those unemployed with durations of one month or less.

[11] No severance payment is due before the third month, although one month prior notice is mandatory.

undetermined duration. However, it was not until 1995 that the changes in legislation fostered their use.[12] This new type of contracts were finally eliminated in 1998;

- The trial period was adopted in 1995, which meant the removal of mandatory (one month) prior notice. The maximum duration was three months, although it could be extended to six months insofar as it was made effective by means of collective agreements. Its attractiveness, however, initially derived from the exemption of employer contributions during this period.[13] In 1998 the maximum duration was reduced to thirty days, also extendable to six months. However, from the second month forward, any dismissal was subject to severance payments (equal to half of the regular payments). Since that year, exemptions from social security contributions were restricted to the first month only.

The explicit purpose of "adjusting" labour legislation, especially in 1995, was to facilitate total, or registered, employment growth. Nevertheless, evidence suggests a scarce effect of these measures on employment-product elasticity.[14] As already indicated, the proportion of unregistered salary-earners in total employment rose even after 1995, despite the reduction in direct costs (i.e. employer contribution to social security).

Beyond this apparent scarce impact on the overall employment level, all together the greater use of fixed-term contracts, the trial period and the reduction in severance payments could have risen the degree of instability during the second half of the nineties by lowering exit costs. In particular, the design of the trial period between 1995 and 1998 favoured turnover not only because it eliminated any type of firing costs but mainly due to the exemption of social security contributions during the period. This should have led to an increase in the dismissal of workers near the end of the trial period in order to replace them by new ones.

2 - Methodology and sources of information

2.1. Analysis methods

The present paper, as it is traditional in occupational mobility studies, focuses on transitions from jobs. In order to evaluate the changes in the probabilities of leaving a job between the first and second halves of the nineties –which is the objective of the article–, one of the approaches to be considered is the comparison of exit rates. They are defined as the proportion of persons in a job in "t" that are no longer in that job in "t+1", when they may be either employed in other job, unemployed or out of the labour force.

A second approach is the use of duration models; specifically, considering hazard functions. They estimate the conditional probability that a given episode –in this case, the probability that the occupation– ends between "t" and "t+1", given that it has lasted until "t". In this paper, the proportional form proposed by Cox (1972) is used. It allows to evaluate the importance that certain variables may have on the probability of leaving the occupation.

[12] New contracts were introduced that year and the proportion of the firms' work force that could be hired through these types of fixed-term contracts was increased.

[13] Except for the health component.

[14] It must be stressed that no analytical study has been done on the effect of changes in regulations on overall employment evolution. However, there are a few studies that arrive to that conclusion by considering partial approaches (see Beccaria and Galin, 2002; pp. 96-115).

It can be shown that the Cox model consists of two parts:[15] a baseline hazard function that takes into account the individual heterogeneity not included in the explanatory variables, and the proportional component. The former depends on duration while the latter only depends on the explanatory variables. Therefore, the probability of leaving the current state is defined as follows:

$$\lambda(t, x, \beta, \lambda_0) = e^{-\beta' x}\lambda_0(t)$$

where λ_0 is the baseline hazard function, which depends on the duration, x is the vector of covariables –in this paper we will consider gender, occupational category, industry, education and position in the household–, and β is the vector of parameters. In this formulation, the covariables only affect the level of the risk-function basic form, thus producing vertical upward or downward movements. This model allows for functional flexibility and its interpretation is quite straightforward.[16] Cox propose a partial maximum likelihood estimator that makes it possible to estimate β without calculating parameters λ_0.

Even though the Cox model is usually employed –as already said– to evaluate the impact of different covariables on the exit rate, in this paper this function is used to analyse changes between the two periods (1988/94 and 1995/99). This will be done by including a "period" variable in vector x, which takes the value zero for 1988/94 and one for 1995/99.

2.2. The source of information

Data on labour market movements used in this paper come from the regular household survey of Argentina, the Permanent Household Survey (EPH) carried out by the National Statistical Office (INDEC), which covers urban areas and collects information especially on labour market variables. Until 2003, it was carried out twice a year in 31 urban centres, during May and October.[17] The analysis will be restricted to Greater Buenos Aires, given the lack of micro-data for other surveyed areas for the entire period.[18] Although the EPH is neither a longitudinal survey nor does it include retrospective questions, its rotating panel sample allows to draw flow data from it, i.e. a selected household is interviewed in four successive moments or waves. Consequently, by comparing the situation of an individual in a given wave to that of the same person in the following one (i.e. five or six months later), it is possible to assess if he/she has experienced changes in diverse variables, including occupational variables.

Specifically, the data set used in this paper includes data on the occupational situation in wave *t+1* (October of year *j-1* or May of year *j*) of persons employed in wave *t* (May of year *j* or October of year *j*). Consequently, it is possible to assess whether he/she remained employed, became unemployed or left the labour force. Each worker may also be characterised by personal and job attributes (including tenure). In particular, the variable "job tenure" *in t+1* is used in order to identify whether a person employed both in *t* and in *t+1* remained in the same job or moved to another one. When the answer is higher than five months in October (or seven in May), it is considered that the person did not change job. The

[15] For the derivation of these functions and for other specifications of duration models see, for example, Kiefer (1988), Kalbfleisch and Prentice (1980), Lancaster (1990), Heckman and Singer (1984), Cox and Oakes (1985), Hausman and Han (1990), Klein and Moeschberger (1997), Parmar and Machin (1996).

[16] This is a model frequently employed, since it is, approximately, half way between the non-parametric Kaplan-Meier and the parametric models; the latter imply rigid structures.

[17] For a description of the survey's methodology, see INDEC (1996). In 2003, the survey underwent some changes and is now producing quarterly estimates.

[18] Greater Buenos Aires accounts for one third of total population of the country and 55% of total population of the urban centres covered by the EPH.

survey does not investigate the causes associated to job separation, hence, it is not possible to distinguish a dismissal from a voluntary quitting.

In order to have enough observations, and given the aim of the paper, transitions of the entire May 1988-October 1989 to May 1999-October 1999 period were pooled. This set comprises about 23,000 transition pairs. The two additional data sets considered resulted from pooling transitions from May 1988-October 1989 to May 1994-October 1994 and from May 1995-October 1995 to May 1999-October 1999.[19]

Data on movements coming from this source face limitations. Some of these limitations derive from the sampling design itself: 25% of the sampling panel is renewed in each wave, thus allowing to compare only 75% of the sample. Yet, this does not hinder the aim of the paper due to the possibility of pooling the data. Nonetheless, it should be taken into account that the effective proportion of individuals and households that are actually matched using panels from two successive waves is lower than 75% due to attrition. Therefore, even if the number of observations left in the pooled panels is still sufficient, the mentioned phenomenon may introduce biases that have not been researched yet. Another difficulty arises from the fact that not every movement can be captured when matching two successive waves because a transition is identified by comparing two observations in a five or seven-month span. Individuals could have performed two or more symmetrical movements during the inter-wave period –e.g. exiting from unemployment to outside the labour force and then returning to unemployment.

It should be kept in mind that any episode observed in the first wave (in this case, being employed in a given job) may, or may not, conclude before the second wave. Only incomplete durations can be known regarding cases in progress at the moment of the second observation; they are, therefore, censored. Regarding the cases finishing between both observations (e.g. a worker leaving his/her job and becoming unemployed or getting another job), a proxy of the complete duration can be computed. It would be equal to tenure in the first observation plus six months (the period between two waves) minus the duration of the state in the second observation. Notice that the latter situation creates another source of biases. Since the episode's incomplete duration is known in the first case, average estimated durations are lower to the real ones. Heckman and Singer (1984) showed that, given certain assumptions, the complete episode should, on average, double the value of the estimated one. However, given that only employed persons are considered initially, the incomplete episodes are longer than the completed ones. These authors also show that under the same assumption, together with the assumption of independence from state duration, both effects cancel each other out exactly.

A difficulty of a different kind refers to the quality of the "job tenure" measurement in the EPH. After a detailed examination of the data set, some inconsistencies were detected in the declarations. As errors are concentrated in high-duration tenures, these cases were excluded from the analysis (see below). This decision does not affect the objective of the analysis, as it will become clear below.

Despite the limitations just mentioned, the information to be used provides a reasonable picture of labour market dynamics, since it makes possible to identify almost every type of transition experienced by workers.

[19] In 1998 and 1999 the survey was also carried out in August. However, the sampling panel used in these cases was different to the one employed in the waves of May and October; consequently, data sets for August waves were not considered in this paper.

3. The group of workers to be analysed

The analysis will focus on employed men between the ages of 15 and 65, and employed women between the ages of 15 and 60, taking into account the mandatory retirement age in Argentina and seeking to minimise the bias introduced by exits due to retirement.[20] Moreover, the study covers workers who declared up to 60 months of tenure in the first observation, accounting for 60% of total observations. This was necessary in order to reduce measurement errors in the tenure variable that, as already indicated, are concentrated in the high-duration cases. Finally, those initially employed in the construction industry and domestic service were excluded, as there are serious difficulties to identify a "job change" in these sectors.

In order to illustrate the effect of these limitations, Table 2 presents the exit rate**s** for all workers for each of the sectors mentioned in the previous paragraph, for those with tenure higher than five years and, finally, for the groups that are analysed in the rest of the paper.

Table 2
Exit rates to all occupational states / destinations
Workers in active ages *
Entire period (1988-1999)

Tenure in the first observation	Total	Sectors to be considered (excluding construction and domestic service)	Domestic Service	Construction
One year or less	40.32	37.5	48.32	55.72
1 to 5 years	16.98	15.42	23.14	32.56
Total up to 5 years	27.9	25.38	35.87	46.8
More than 5 years	10.93	9.37	17.07	26.98
Total (without restricting tenure)	20.68	18.35	30.22	38.88

* Age 15 -65 for men, and age 15-60 for women.

Stability is lower for the group of workers to be considered here than for all workers. This is mainly due to the fact that the turnover of those with tenure higher than 60 months is low. Notice that the effect of omitting this group is not compensated by the exclusion of Domestic Service and Construction sectors.

As mentioned above, not only exits within the labour force but exits to all destinations will be considered, in contrast to previous papers dealing with this theme in Argentina. In labour markets such as the Argentinean, which lack of an extended system of unemployment insurance, the movements out of the labour force are more common than in other cases.

[20] An early retirement is possible at 55 years old for women and 60 for men, but in these cases, the benefit is lower than in the case of regular retirement. Some estimates were made restricting the sample to individuals within these –lower– age ranges, yet their results were similar to the ones that will be discussed in this paper.

4. Exits from occupation during the whole period

In this brief section we make an analyses of the whole period –1988-1999– in order to show the effects of both the accumulated duration in a given occupation and certain personal and job characteristics on the degree of occupational mobility in Greater Buenos Aires.

Table 3
Exit rates from employment
All workers, entire period (1988-1999)

Tenure	Exit rates (%)
3 months or less	52.08
3-6 months	35.56
6 –12 months	25.35
1 -2 years	18.6
2 - 5 years	13.47
One year or less	37.5
1 – 5 years	15.42
Five years or less	23.38

As expected, exit rates go down as tenure at the initial observation rises (Table 3). More than one half of those workers with three months of job tenure, or less, leave the job within the six-month period between two observations. In contrast, this rate reaches only 15%, on average, for those workers with tenure over 12 months. Similarly to other countries, it can be also shown that the influence of duration on exit rates changes along the survival function, i.e. its marginal contribution diminishes as tenure rises.

Hazard functions using the proportional Cox model[21] also show that gender, education, position within households, age and occupational category have significant effects on exit rates. Regressions were run considering all workers and also for the groups with tenure up to 12 months and up to 24 months (See Annex, Table A.1). The coefficients of the regressions for all workers suggest that the occupational category defines the groups of workers with the greatest differences in instability: registered wage-earners –both private and public– have the lowest exit probabilities, while unregistered wage-earners have the highest. The results are similar for the three tenure groups, although the gaps are narrower for those with tenure lower than12 months.[22]

Unregistered wage earners have low legal firing costs,[23] thus making them attractive for employment in industries with unstable demand and for unstable occupations. An employer may also decide not to register an employee in order to test the worker's ability for longer than the official trial period, or just as an alternative to it. Consequently, the low exit rate of registered wage-earners may be explained by the existence of firing costs as well as

[21] The proportionality test suggests that the effect of tenure is proportional; consequently, the Cox model seems adequate for this case.
[22] Differences in exit rates grow as tenure rises, and reach its minimum in the three-month or less bracket. This result suggests that turnover is high even for unregistered employees and non-wage earners with tenures of more than 12 months. The average exit rate of registered employees is also high for those with low tenures. The latter finding, also identified in other countries, may be indicating the relevance of inadequate matching and low separation costs in low tenure jobs (Beccaria y Maurizio, 2003, Table A.1, Annex)
[23] They would be those derived from fines and payments that the employed would be obliged to pay if the dismissed unregistered employee reported such situation at the Ministry of Labour and/or Justice.

by the fact that they receive more specific training. Moreover, employees who are not covered by social security and also self-employed workers have a greater presence in small scale and informal firms, which are regularly exposed to risks that make them more vulnerable. As they operate with low capital/labour ratios, the decision to interrupt economic activity is easier.

Schooling is inversely related to the probability of leaving a job, reflecting the highest instability of less educated workers. The effect of this variable rises as tenure increases. On the other hand, industry is a variable that proved to have a low influence on the level of exit rates.[24]

The other variables also show the expected results. Men and household heads face less instability than women[25] and non-household heads, although the gaps are smaller than those detected between categories. The coefficient of the age variable is negative and significant, thus indicating that young workers face the highest exit rates.

When the regressions are computed for workers of a given category (see, for example, Table A.2 for workers with low tenure) the same gaps appear in each of them. This indicates that young persons, women and non-heads are more unstable not only because they work in precarious jobs more frequently but also due to the fact that their exit rates are even higher among private registered jobs (no significant difference can be found among public employees). These results imply, for the case of youths, that their higher labour instability is not fully explained by segregation, a situation that comes from the fact that they are considered less reliable by employers and, therefore, the most unstable jobs are offered to them. Nonetheless, segregation could be underestimated if the high exit rates of young registered wage - earners were indicating that they are employed in the most unstable jobs. Similar considerations could be made for the case of women.

5. Changes in mobility during the nineties

This section deals with the main objective of the paper i.e. the analysis of the changes in both the degree and the characteristics of occupational mobility that occurred between the two periods under consideration.

Table 4
Exit rates from employment in each period
All workers, both periods

Tenure	1988-1994			1995-1999			Significance of "period" in Cox Model
	Exit rates	Intervals 95%		Exit rates	Intervals 95%		
12 months or less	35.8	34.6	36.9	39.5	38.2	40.7	Yes
24 months or less	30.8	29.8	31.7	32.4	31.4	33.4	Yes
Five years or less	25.1	24.4	25.9	25.6	24.8	26.4	Yes

The average exit rate remained fairly constant between both periods but there was arise among low-tenure workers. This increase is statistically significant when considering the confidence intervals for the exit rates (Table 4) in the case of those workers with tenure of 12 months or less, and is practically significant for those with job duration up to 24 months

[24] The regressions not presented here also showed that the size of establishments is not significant either. However, it has to be taken into account that this variable is deficiently measured in household surveys.
[25] Similar results were found by Cerruti (2000a); see also, Rubery, Smith y Fagan (1999).

as well. The dummy variable "period" –introduced in the Cox model in order to estimate changes between periods– shows that the risk of leaving a job was significantly higher during the second half of the nineties compared to previous years for all tenure groups.

Considering that the changes in regulations influencing exit costs –such as the reduction of firing costs and the introduction of the trial period– specially affect low-tenure workers –see Section 1–, it seems reasonable to evaluate to what extent they represented a source of growth for this group's exit rate. The analysis will be concentrated on those workers with tenure up to 12 months, since they registered the largest change in exit rates and, probably, were among the most influenced by de–regulation: the duration of the trial period was no longer than six months, and the impact of the reduction in severance payments on exit rates diminishes as duration approaches 24 months.

In order to evaluate the effect of the changes in firing costs (and other institutions) on labour stability, we will compare the degree of mobility of those directly affected by regulations to the degree of those that were not. The former group consists of private sector wage - earners registered at the social security system; the latter includes the rest of workers –public employees, private sector non-registered wage - earners, and non-wage workers.

Average exit rates remained constant for registered private wage-earners (Table 5), whereas it rose for the rest of workers. Within the latter group, the rate grew for private non-registered workers, whereas no significant changes occurred for independent workers and public employees. The econometric analysis show similar results: the "period" variable in the Cox function is not statistically different from zero in the case of covered wage-earners of the private sector whereas it is positive when the rest of low-tenure workers are considered together. Regarding the latter the regression indicates that the exit probability grew for private workers (non-registered employees and non-wage earners) but not for public workers.

TABLE 5

Changes in exit rates, workers with tenure of 12 months or less

	Exit rates (%)						Significance of the variable "period" in the Cox Model
	1988-1994			1995-1999			
	Average	Intervals 95%		Average	Intervals 95%		
All workers	35.8	34.6	36.9	39.5	38.2	40.7	Yes
Covered by regulations	25.2	23.3	27.1	25.3	23.2	27.4	No
Not covered by regulations	40.5	39.1	42.0	45.1	43.6	46.6	Yes
Public employees	19.1	15.7	22.5	23.0	18.8	27.1	No
Non-registered employees	46.5	44.5	48.5	50.9	49.0	53	Yes
Non - wage earners	38.0	35.3	40.7	41.1	38.0	44.3	Yes
Men	32.5	31.1	34.0	38.6	36.9	40.1	Yes
Covered by regulations	23.2	21	25.4	25.5	22.9	28.1	No
Not covered by regulations	37.8	35.8	39.6	44.6	42.2	46.8	Yes
Public employees	16.4	10.9	21.9	24.7	16.4	33.1	Yes
Non-registered employees	43	40.5	45.4	48.8	46.3	51.2	Yes
Non - wage earners	32.2	28.8	35.5	38.1	34	42.3	Yes
Women	41.2	39.2	43.1	40.8	38.8	42.9	No
Covered by regulations	30.4	26.6	34.2	24.8	21.1	28.5	No
Not covered by regulations	44.5	42.2	46.8	45.7	43.4	48.1	Yes
Public employees	20.5	16.2	24.8	22.3	17.6	27.1	No
Non-registered employees	53.6	50.1	57.1	55.3	51.9	58.7	Yes
Non - wage earners	47	42.6	51.5	45.1	40.2	49.9	No

Hence, these results could be indicating that the rise of labour turnover experienced by low-tenure workers between the first and the second halves of the nineties in Argentina was not associated to the reduction in firing costs. This result differs from the conclusions of similar studies carried out in other Latin American countries. For example, by using a similar methodology to the one applied here- difference-in-difference, Kugler (2000) found that covered workers have increased their exit rates more than the control group of those workers that were not affected by the modifications of the firing regulation introduced in Colombia in 1990. Gozaga (2003) used a similar approach for Brazil. He shows that the increase in the strictness of regulations in 1988 and 2001 led to lower turnover.

In a previous paper (Beccaria and Maurizio, 2003) it was found that the overall rise in labour instability between the two periods was mainly explained by the increase in the exit probability of men whereas women did not follow this behaviour (this can be appreciated in Table 5). It could be possible, therefore, that the results just discussed, which are indicating a fairly constant exit rate among all covered employees, were influenced by differences between gender. Women's overall behaviour may have been influenced by the long-run trend towards a rise in labour market stability that usually accompanies the economic participation growth. While it is clear that women participation in economic activity have been expanding during the nineties, the movements in and out of the labour force have diminished more for women than for men between both periods.[26] As part of this trend towards a higher and more stable participation in the labour force, it could be considered that women may have reduced the rate of voluntary quitting as well. Had this occurred, an eventual movement towards higher turnover derived from lower firing costs could have been offset by this other trend.

No attempt will be made here to carry out a detailed analysis of possible variables that affect women's attachment to the labour force. However, it seems worthwhile to perform a similar analysis on the influence of regulation on mobility to the one discussed at the beginning of this section, but controlling for gender (Table 5). Whereas the rise in the exit rate among men between the two periods was not statistically significant for the case of registered private wage-earners, it was for the average of the rest of workers. Within the latter, a significant growth occurred among non–registered employees, whereas no change could be identified for the other two groups. When looking at the "period" variable of the Cox model the same result is reached: the coefficient is not different from zero for registered private wage-earners and it is positive and significant for the control group. This model also indicates a change in the exit probabilities for each of the three groups of workers not affected by the firing regulations.

When the analysis is performed for low-tenure female workers, the differences in exit rates proved to be not significant for the average of all workers as well as for each of the identified groups (Table 5). The Cox model only differs from this result in that it shows a significant rise in mobility among non–registered workers and for the aggregate of the control group.

Therefore, the results coming from both approaches suggest that the regulations didn´t have effect on the increase of exit rates of low-tenure workers between the first and the second halves of the nineties, even controlling for gender. Mobility did not change for covered private employees, whether they are men or women, whereas it grew for non–registered wage-earners and non wage-earners The rise in the overall exit rate, and in the average rate for men, is explained by the growth experienced by these two categories.

[26] Using the same data basis, it can be shown shows that that the rise in the proportion of persons staying in the labour force in *t* and *t+1* (i.e. after five or seven months) between 1988-94 and 1995-98 was larger for women than for men. A differences-in-differences analysis proved that this difference was statistically significant.

Consequently, the gap between the mobility of covered and not covered low-tenure workers (mainly, between registered and non-registered employees) widened during the nineties.

The higher exit rates of non–registered workers (overall and for both genders) during the second period may be associated to the rise of open unemployment. During these years, the lack of occupational opportunities –mainly, in the formal sector– led to frequent movements between short unemployment spells and non-covered jobs, which were of short duration as well; or movements from precarious jobs to other occupations of the same type. Non-registered workers leaving a job have, with respect to registered wage – earners, a lower probability not only of getting another job, but also of obtaining a registered new occupation.[27] Again, it is worthwhile to highlight the low coverage of the unemployment insurance.

The rise in the exit rate of own-account workers suggested by the Cox model –specifically for men– could be associated to the modifications experienced by the structure of the informal sector which, to some extent, could have also been associated to the overall shortage of employment possibilities. Traditionally, a large segment of urban informal activities in Argentina was less precarious than what is stereotypically associated to LDC's informality. This segment consisted of small firms that operated with a certain –albeit low– amount of capital and which produced relatively adequate incomes and provided quite stable occupations. The productive restructuring process of the early nineties negatively affected many of these units, thus provoking a reduction in the size of the informal sector but also inducing changes in its structure. The most precarious activities increased their share, leading to a high average labour instability.

It was already indicated that the rise in the participation of precarious employment in total occupation –mainly non-registered workers– (see Table 6) may be another reason explaining the growth of the overall exit rates of low-tenure workers. Such increase signified a larger presence of highly mobile jobs. This is a process that, *per se,* led to a rise in average exit rates beyond the increase specifically shown by the rate of non-covered workers. The average exit rate changes not only as a result of the shifts in each group's risks but also due to alterations in the labour force structure. In order to evaluate the impact that the shifts in the occupational structure could have had on overall mobility of low-tenure workers, we decomposed the observed change in their average exit rate between both periods into two components: the "structure effect" and the "pure instability effect". The latter effect captures the impact of variations on each group's exit rates.[28] The changes in the employment structure explain 27% of the rise in the average exit rate of all workers, and a similar figure (25%) of the rise in men's exit rate. In both cases, this "structure effect" is mainly explained by the increase in the participation of non-registered employees that could not be offset by a reduction of independent workers. For women, the structure effect produced a rise in the average exit rate but it was compensated by the pure instability effect that came mainly from the reduction in the probability of leaving covered jobs.

[27] For the whole 1988 – 1999 period, 44% of those non-registered wage earners in "t" that leaves a job are working in "t+1"; this proportion rises to 53% for registered employees. Among the former group, only 19% of those getting a new job enter a registered occupation while the figure rises to 44% for the latter group.

[28] The decomposition formula is:

$$(\Delta PS)/PS = \underbrace{\left[\sum_i \left(\Delta(PS_i)x\frac{E_i}{E}\right)\right]}_{EFECTOINESTAB}/PS + \underbrace{\left[\sum_i \left(\Delta\left(\frac{E_i}{E}\right)xPS_i\right)\right]}_{EFECTOESTRUCTURA}/PS$$

where PS is the average exit rate; PS_i is the exit rate of group "i"; E is total employment; and E_i is the employment of group "i".

Table 6
Employment distribution according to category and gender

	May 1988 - Oct 1994			May 1995 - Oct 1999		
	Men	Women	Total	Men	Women	Total
Registered wage- earners	54.5	56.3	55.1	51.5	53.6	52.3
Unregistered wage- earners	17.9	18.2	18.0	23.0	22.7	22.9
Non-wage- earners	27.6	25.4	26.9	25.5	23.7	24.8
Total	100.0	100.0	100.0	100.0	100.0	100.0

5. Conclusions

The evidence analysed in this paper suggests that occupational instability rose between the first and second halves of the last decade of 20th century in Greater Buenos Aires. The increase in occupational turnover was concentrated among men, while women's exit rate from occupations remained constant. The latter result may have been affected by the persistence of the long-term trend towards a higher and more stable female participation in the labour force. During these years, women increased their participation rate and extended their time within the economically active population, by engaging in an active search and/or remaining unemployed after leaving a job. It could be suggested that had it been possible to discount the effects of that trend, a process of increasing occupational instability would have emerged among women as well.

Labour regulations appear as a potential source of the rising mobility, since they experienced various modifications by the middle of the decade. The most relevant were the reduction of severance payments for low-tenure workers, the introduction of the trial period and the relaxation of restrictions on the usage of fixed-term contracts. The analyses of exit rates changes and the Cox model indicate, however, that those alterations in the labour market institutional framework had apparently no effect on turnover. The turnover actually rose among workers not affected by firing regulations, especially non–registered employees but also –according to one of the approaches used– male non–wage earners.

The exit rate growth of employees not covered by labour regulations (overall and for each gender) appears to be associated to the particular labour market behaviour which, during the second half of the nineties, was characterised by a shortage of occupational opportunities. Even if the labour market performance regularly shows a high level of mobility due to the large share of non-covered workers (and the consequent low coverage of the formal unemployment insurance), that circumstance should have increased the incidence of unstable trajectories among those usually working without coverage. These unstable trajectories entail movements from short unemployment spells and non–covered jobs, which are of short duration as well; or from precarious jobs to other occupations of the same type.

Moreover, the rising underemployment also induced an increase in the participation of non-registered, highly mobile, workers in the employment structure. This is a process that by itself led to an increase in overall mobility.

The worsening of the labour situation could have also affected the informal activities that, traditionally, have been less precarious in Argentina than in other LDCs countries. One of the consequences of that behaviour could have been the rise of the share of more precarious jobs among own-account workers, thus leading to a higher labour turnover of this group -on average.

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Annex[29]

TABLE A.1
COX MODEL. WHOLE PERIOD

	1-12 months				1-24 months				1-60 months			
Variable	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|
Education												
Complete sec., incomplete tert.	**-0.2680**	0.000	**-0.1907**	0.000	**-0.3390**	0.000	**-0.2412**	0.000	**-0.3896**	0.000	**-0.2533**	0.000
Tertiary completed	**-0.6999**	0.000	**-0.4237**	0.000	**-0.8524**	0.000	**-0.5238**	0.000	**-1.0604**	0.000	**-0.6909**	0.000
Industry												
Trade	**0.0343**	0.344	**-0.0810**	0.033	**0.1148**	0.000	**-0.0518**	0.131	**0.1975**	0.000	**-0.0391**	0.210
Transp. + Finan.	**0.0110**	0.799	**-0.1537**	0.000	**0.0690**	0.076	**-0.1573**	0.000	**0.2116**	0.000	**-0.0967**	0.007
Other services	**0.0017**	0.973	**-0.1480**	0.004	**0.0468**	0.301	**-0.1637**	0.000	**0.1439**	0.000	**-0.1432**	0.001
Gender	**0.1741**	0.000	**0.2276**	0.000	**0.1846**	0.000	**0.2502**	0.000	**0.2264**	0.000	**0.3218**	0.000
Head of Household	**-0.1934**	0.000	**-0.1984**	0.000	**-0.2225**	0.000	**-0.2219**	0.000	**-0.3191**	0.000	**-0.2934**	0.000
Age	**-0.0114**	0.000	**-0.0084**	0.000	**-0.0148**	0.000	**-0.0111**	0.000	**-0.0205**	0.000	**-0.0155**	0.000
Non covered by regulations	**0.5465**	0.000			**0.6476**	0.000			**0.7020**	0.000		
Non- Regist Priv			**0.7716**	0.000			**0.9609**	0.000			**1.1294**	0.000
Public empoyees			**-0.3597**	0.000			**-0.4333**	0.000			**-0.5318**	0.000
Non- Wage earners			**0.4101**	0.000			**0.5026**	0.000			**0.5746**	0.000
Period	**0.1922**	0.000	**0.1704**	0.000	**0.1244**	0.000	**0.0916**	0.001	**0.0757**	0.002	**0.0538**	0.028
Observ.	12079		12079		17659		17659		26788		26788	

Table A.2
Cox model, for each category. Whole period

(a) Covered by regulations (private registered employees)

	Total		Men		Women	
Variable	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|
Education						
Complete sec., incomplete tert.	**-0.2769**	0.000	**-0.3019**	0.001	**-0.2327**	0.061
Tertiary completed	**-0.5929**	0.000	**-0.6237**	0.003	**-0.4883**	0.039
Industry						
Trade	**0.0033**	0.968	**-0.0448**	0.675	**0.0734**	0.593
Transp. + Finan.	**-0.0154**	0.858	**0.0212**	0.833	**-0.1112**	0.506
Other services	**-0.2381**	0.061	**-0.1950**	0.210	**-0.3309**	0.132
Women	**0.1412**	0.069				
Head of Household	**-0.2006**	0.026	**-0.2198**	0.040	**-0.2322**	0.250
Age	**-0.0097**	0.018	**-0.0069**	0.184	**-0.0161**	0.022
Period	**0.0363**	0.590	**0.1484**	0.071	**-0.1933**	0.101
Observ.	3616		2519		1097	

(b) Non covered by regulations

	Total		Men		Women	
Variable	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|
Education						
Complete sec., incomplete tert.	**-0.2723**	0.000	**-0.2608**	0.000	**-0.2939**	0.000
Tertiary completed	**-0.8154**	0.000	**-0.8386**	0.000	**-0.7844**	0.000
Industry						
Trade	**0.1267**	0.003	**0.0250**	0.666	**0.2372**	0.000
Transp. + Finan.	**0.1121**	0.030	**0.0161**	0.800	**0.2257**	0.013
Other services	**0.1245**	0.026	**0.0337**	0.634	**0.2353**	0.010
Women	**0.1592**	0.000				
Head of Household	**-0.1769**	0.000	**-0.2312**	0.000	**-0.1472**	0.129
Age	**-0.0117**	0.000	**-0.0082**	0.001	**-0.0150**	0.000
Period	**0.2386**	0.000	**0.2950**	0.000	**0.1733**	0.001
Observ.	7878		4736		3142	

[29] The base category for Education is "Not completed secondary level or less"; it is "Registered wage earners of the private sector" for Occupational category , and "Manufacturing" for Industry.

(b.1) Private non-registered employees

Variable	Total		Men		Women	
	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|
Education						
Complete sec., incomplete tert.	**-0.1175**	0.016	**-0.1133**	0.082	**-0.1095**	0.144
Tertiary completed	**-0.3221**	0.019	**-0.6440**	0.007	**-0.0923**	0.592
Industry						
Trade	**-0.0142**	0.784	**-0.0717**	0.293	**0.0525**	0.512
Transp. + Finan.	**-0.1013**	0.089	**-0.0899**	0.214	**-0.1803**	0.098
Other services	**-0.0998**	0.132	**-0.1023**	0.210	**-0.1026**	0.375
Women	**0.2122**	0.000				
Head of Household	**-0.1149**	0.060	**-0.1122**	0.133	**-0.1180**	0.339
Age	**-0.0079**	0.000	**-0.0079**	0.008	**-0.0079**	0.021
Period	**0.2170**	0.000	**0.2525**	0.000	**0.1595**	0.021
Observ.	4776		3173		1603	

(b.2) Non-wage earners workers

Variable	Total		Men		Women	
	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|
Education						
Complete sec., incomplete tert.	**-0.3529**	0.000	**-0.4022**	0.001	**-0.3218**	0.009
Tertiary completed	**-0.7473**	0.000	**-0.5517**	0.022	**-1.0010**	0.000
Industry						
Trade	**-0.0334**	0.728	**0.0602**	0.662	**-0.1239**	0.354
Transp. + Finan.	**-0.1410**	0.268	**-0.1443**	0.368	**-0.0225**	0.922
Other services	**0.0526**	0.661	**0.0993**	0.547	**-0.0046**	0.979
Women	**0.1704**	0.052				
Head of Household	**-0.3602**	0.000	**-0.3478**	0.006	**-0.5001**	0.010
Age	**-0.0085**	0.010	**-0.0066**	0.191	**-0.0104**	0.019
Period	**0.1512**	0.029	**0.2517**	0.009	**0.0502**	0.614
Observ.	2174		1281		893	

(b.3) Public employees

Variable	Total		Men		Women	
	Coef.	P>|z|	Coef.	P>|z|	Coef.	P>|z|
Education						
Complete sec., incomplete tert.	**-0.0565**	0.753	**-0.1284**	0.672	**0.0182**	0.937
Tertiary completed	**-0.1823**	0.353	**-0.4062**	0.341	**-0.0804**	0.734
Women	**0.1887**	0.314				
Head of Household	**-0.0595**	0.789	**-0.5030**	0.153	**0.1700**	0.575
Age	**-0.0134**	0.094	**0.0081**	0.615	**-0.0193**	0.043
Period	**0.2590**	0.075	**0.5640**	0.037	**0.1534**	0.371
Observ.	928		292		646	