Duration and Risk of Unemployment in Argentina

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Abstract: After a decade of structural reforms, unemployment rates have tripled in Argentina. This paper is concerned with the measurement of unemployment risk and its distribution. We show the importance of considering re-incidence in the measurement of risk and develop a methodology. Our estimates for Argentina show that, though the typical unemployment spell is short, once re-incidence is taken into account, unemployment risk is high, has risen substantially in the last decade and is shared very unequally in the labor force. This counters the established view that unemployment is a small risk, short-duration phenomenon, which arises when re-incidence is not considered.

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1. Introduction

After a decade of structural reforms in Latin-American economies, there are growing concerns about the social consequences of increased employment volatility and the incidence of unemployment. The sharp rise in unemployment rates in Argentina lends support to this concern. Moreover, there is a perception that unemployment risk is very unequally distributed and that certain groups share an excess burden of the adjustment. This paper is concerned with the measurement of unemployment risk and its distribution. It raises some critical issues concerning the definition of unemployment risk, deals with the corresponding methodological considerations and provides estimates for the Buenos Aires labor market.

A standard method to evaluate unemployment risk is to consider the incidence of unemployment and its duration. Even when incidence may be high, if unemployment spells are short-lived, it is understood that the social cost of a typical unemployment spell is low. This paper shows that this reasoning is misleading when the typical employment spell is also short-lived. In such case, a correct account of unemployment risk must take into consideration the re-incidence of unemployment spells. Our estimates for the Buenos Aires labor market show that, contrary to the view that unemployment spells are short, total expected duration, accounting for repeated spells, is indeed long.

As a point of reference, consider European labor markets, which have experienced high unemployment rates since the mid-seventies. A salient characteristic of the high unemployment era has been the high proportion of long term unemployment. Certainly, this feature made the European unemployment performance particularly problematic: although there have been other periods of high unemployment rates, long-term unemployment seems to be a characteristic of the last decades.[1]

Some developing countries (like Argentina) have also gone through episodes of high unemployment during the 90s. The lack of well-developed social security systems may suggest that most unemployment episodes are of a short-run nature. Indeed, the evidence of high flows in and out of unemployment confirms this hypothesis (see section 2). The Argentine case is particularly striking; the inflow rate is over 2 percent since the beginning of the nineties and grew to 4 percent by the middle of decade (see section 2). As it is well known, *ceteris paribus*, a high inflow rate implies low average unemployment duration. In steady state, the average duration of all episodes of unemployment equals the ratio of the unemployment rate to the inflow rate.

Table 1 presents average unemployment rates as well as short and long-term unemployment rates for OECD countries along with the Argentine figures for the period 1989-1994. This data show that the incidence of long-term unemployment in Argentina is substantially lower than in most European countries and it is similar to that of the US. Thus, Argentina seems to be a country where unemployment is mostly a short-term phenomenon.[2]

[1] That is, controlling for the unemployment rate, long spells of unemployment were less important before the mid-seventies than later (cf. Machin and Manning, 2000). Generous unemployment insurance has been blamed for this long duration. There is ample evidence suggesting there is a positive impact of benefit levels and entitlement duration on the duration of individual unemployment spells (cf. e.g. Narendranathan et al., 1985 and Meyer, 1990). Although some authors argue that these effects are not large (cf. e.g. Atkinson and Micklewright, 1991), there is convincing evidence about the negative impact of the duration of benefits on the outflow from unemployment (cf. e.g. Carling et al., 1996 and Katz and Meyer, 1990).

[2] Indeed, the statistics reported in table 1 refer to the length of the episodes of unemployment in progress. However, due to the extremely high inflow rates observed in Argentina, the average length of all spells is lower than the average length of the episodes in progress (see section 2). Akerlof and Main (1980) present a good discussion of the differences between these two statistics.

Table 1: Unemployment rates in OECD and Argentina (%)
1989-1994

Country	Total unemployment rates	Short-tern unemployment rates	Long-term unemployment rates	Long-term unemployment incidence rates
Spain	18.9	9.1	9.7	52
Ireland	14.8	5.4	9.4	64
Denmark	10.8	7.9	3.0	28
Finland	10.5	8.9	1.7	16
France	10.4	6.5	3.9	38
UK	8.9	5.5	3.4	38
Italy	8.2	2.9	5.3	65
Belgium	8.1	2.9	5.1	63
Netherlands	7.0	3.5	3.5	50
Norway	5.5	4.3	1.2	22
Germany (W)	5.4	3.2	2.2	41
Portugal	5.0	3.0	2.0	40
Sweden	4.4	4.0	0.4	9
Switzerland	2.3	1.8	0.5	22
Japan	2.3	1.9	0.4	17
Australia	9.0	6.2	2.7	30
New Zealand	8.9	6.6	2.3	26
Canada	9.8	8.9	0.9	9
US	6.2	5.6	0.6	10
Argentina	7.9	6.7	1.2	15

Notes: These rates are OECD standardized rates with the exception of Denmark and Italy. The data for Argentina refer to the Metropolitan region and follows the ILO definition. Hence, these rates are very similar. Long-term rates refer to those unemployed with duration over 1 year.
Sources: Nickell and Layard (2000) and Authors elaboration based on Nickell and Layard (2000) and the Argentine Household Survey (October).

In this paper we are concerned with the distribution of unemployment risk among different groups of individuals. The type of problem we are concerned with can be illustrated by the following example: Consider the following two situations, both of which result in a 10 percent unemployment rate. In the first case, a given 10 percent of the labor force is unemployed the whole year; in the second, everyone is unemployed once a year for one-tenth of the year. Clearly, the distribution of unemployment differs substantially between the two cases. In the first scenario, the risk of unemployment is completely concentrated among a (relatively) small group of the population, while in the latter it is uniformly distributed among all individuals. Specifically, we deal with the following question: What groups are at risk of being unemployed high proportions of a given period of time?

One could conclude that in countries with high long-term incidence rates, the risk of unemployment is highly concentrated among small groups of workers, while in countries with high turnover and low long-term incidence rates, unemployment risk is more evenly distributed among the population. This paper argues that such simple characterization of labor market behavior would be misleading, at least for Argentina, but most probably for other countries with high turnover rates. We show that even in a country where the inflow rate to unemployment is above 2 percent, the risk of unemployment is relatively concentrated in the population.

As mentioned above, the key in reconciling high inflow rates and concentrated unemployment is the fact that individuals often re-enter unemployment soon after leaving it. It is well known that those individuals with a past record of unemployment are most likely to be currently unemployed, a phenomenon that Heckman and Borjas (1980) have labeled occurrence dependence. Thus, due to multiple spells, unemployment affects some groups repeatedly, which tends to concentrate the risk of becoming unemployed.

In this regard, several authors (cf. e.g. Clark and Summers, 1979; Johnson and Layard, 1986 and Machin and Manning, 2000) have argued that it should be analyzed the distribution of individuals unemployed at a point in time, not according to the duration of the current spell, but according to the amount of time those individuals will be unemployed in a certain period of time. Clark and Summers (1979), for example, estimate that the average person unemployed at a point in time during the period 1965-68 spent one-quarter of those 4 years unemployed.

Especially in countries with high turnover, where the average duration of unemployment is low, a good indicator of long-term unemployment is thus the proportion of time an individual has been unemployed over a certain period of time. In this paper we study the conditional distribution of total unemployment for a two-year period. Using panel data from household surveys for the Buenos Aires area[3] for the period 1989-1998, we estimate a Markov process for transitions from employment to unemployment (and vice versa) that allow for duration dependence. From these estimates we obtain a distribution for the number of incidences and total unemployment time that someone entering unemployment will experience in the following two years.

We find that the median worker entering unemployment in 1998 has a total of 4 unemployment spells in the following two years and total cumulative duration of 7.4 months. A worker with college education experiences half the number of spells and approximately 30 percent less time out of work. In contrast, the median young worker with low schooling exhibits 6 spells of unemployment and cumulative duration of one year out of the two years. Our estimates also show the importance of long-term unemployment: of all workers unemployed at a given point in time, 34 percent spend more than one year of unemployment during the past two years. These figures are much closer to the high numbers found in European economies. Finally, comparing the first and last period of our sample, the median number of spells over the two-year period has doubled while median cumulative unemployment duration increased by 35 percent.

[3] This market covers approximately half of the labor force in the country.

2. The facts

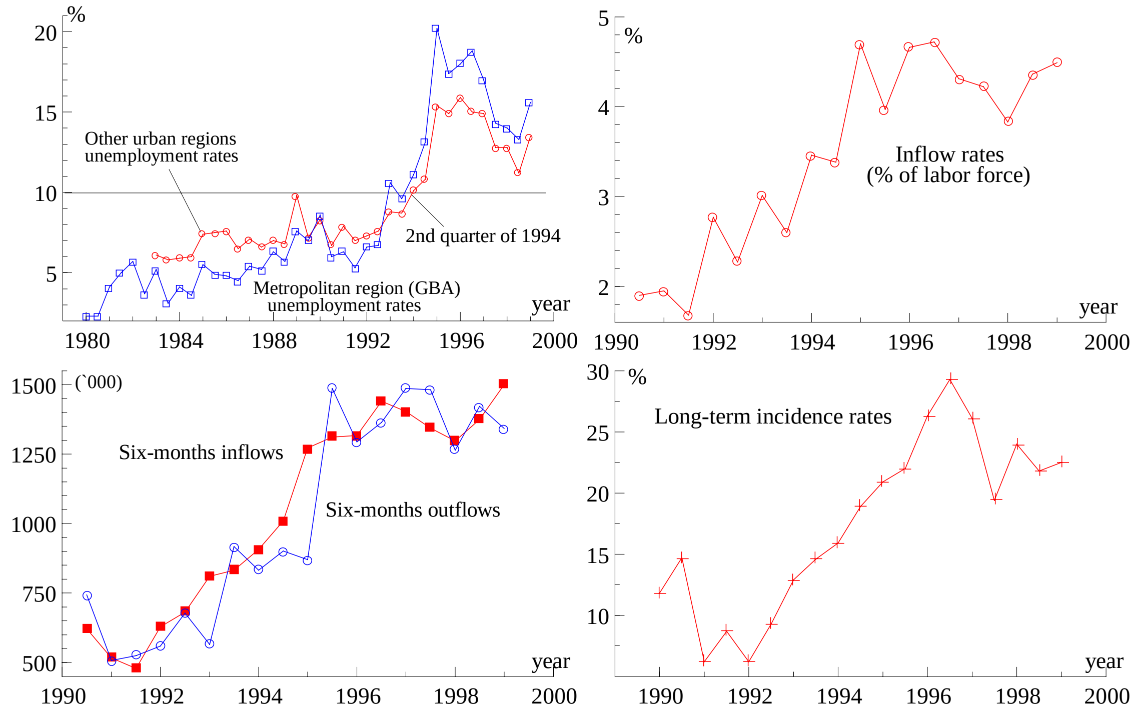
The period that we consider is marked by a sizable increase in the unemployment rates, from less than 10 percent to a peak of 20 percent around the mid-nineties. Although the recovery in the demand for labor in 1997 and 1998 induced a partial reversion of the upward shift; the unemployment figures at the end of the period remained much higher than at the beginning of the period (see Figure 1, panel a).

All major groups in the labor force increased their unemployment rate. The change was the sharpest among high-age individuals, especially females. Although the female participation rate grew since the mid-eighties, that growth accelerated during the nineties (the largest proportional increase in the female participation rates occurred among the oldest groups). However, using transition matrix analysis, we estimate that for the population as a whole, a higher participation numerically explains only a third of the increase in unemployment. Instead, the predominant factor was the increase in job destruction. This result is consistent with the rising trend in the inflow rate observed during the nineties (see Figure 1, panel b).[4,5]

Figure 1 also illustrates an interesting feature of unemployment in Argentina. Contrary to the European experience, where the secular increase in unemployment can be arithmetically accounted for by a rising average duration (a fall in the outflow rate), rather than a rise in the inflow rate, in Argentina both the inflow rate and the long-term

[4] The number of persons unemployed for less than one month is used as a proxy for inflows. This is a useful measure but it does underestimate somewhat the number of inflows, because persons who became unemployed but find a job in less than a month may not be included.
[5] For simplicity, we compute the inflow rate as the ratio of the inflow to the labor force instead to employment. This facilitates steady state computations.

Figure 1: Unemployment in Argentina during the 90s



Notes: Panel b: The monthly inflow is the number of people who, at a point in time, have been unemployed for one month or less. The inflow rate is the monthly inflow divided by the total labor force at that point in time. Panel c: The six month inflow to unemployment is calculated as follows: I(t,t-1) = (21 I(t) + 15 I(t-1))/6; where I(t) measures the inflow to unemployment in period t. Thus, it measures the cumulative inflow to unemployment between t and t-1 as a weighted average of the inflows between those periods. The six month outflow from unemployment is calculated as follows: O(t,t-1) = U(t-1) + I(t,t-1) - U(t); where U(t) is the number of unemployed people in period t. Panel d: the long-term unemployed are those individuals whose current spell is a year or higher. The long-term incidence rate is the proportion of long-term unemployed people in total unemployment at a point in time.
Sources: Panel a: INDEC press reports. Panels b, c and d: Authors elaboration based on the GBA Household Survey.

incidence rate have increased substantially over the nineties (see Figure 1, panels b and d). Notwithstanding, the incidence rates are well below the numbers observed for most countries in continental Europe.

Likewise, the average duration of the current spells has also increased but remained well below a year. As pointed out above, the average (completed) duration of all spells is quite low in Argentina, but has also increased during the nineties.

Table 2: Unemployment duration and flows

Year	Unemployment rate (%) (U/L)	Inflow per month (%) (S/L)	Steady state average completed duration of all spells (months) (U/S)	Average uncompleted duration of current spells (months)
1990	8.6	2.8	3.1	4.4
1991	6.3	2.0	3.1	3.5
1992	6.7	2.8	2.4	3.0
1993	10.6	3.0	3.5	4.7
1994	11.1	3.5	3.2	4.3
1995	20.2	4.7	4.3	6.1
1996	18.0	4.7	3.8	6.9
1997	17.1	4.3	4.0	7.5
1998	14.2	3.8	3.7	7.2
1999	15.7	4.5	3.5	6.6

Source: Authors calculations based on GBA Household Survey, May.

Hence, we observe a secular increase in unemployment, inflows and unemployment duration, which constitute a special configuration that deserves analysis. In particular, as we have seen, even if duration increased, the incidence rate was still low compared to that in Europe. However, it is likely that we observe multiple unemployment spells among those individuals who experience unemployment. This may be inferred from the extremely high number of inflows episodes accumulated in a six-month period (see Figure 1, panel c).[6] In this regard, without re-entry to unemployment, in the course of three years, the whole labor force would have entered unemployment once, implying the lower possible concentration of unemployment risk among the population. However, the incidence of unemployment has been probably much more concentrated in the population due to the existence of multiple spells. We explore this effect in detail in the

[6] Strikingly, the inflow to unemployment remained high even in the recovery after the large macroeconomic shock of 1995, which induced a deep recession.

next section, by estimating the conditional distribution of the length of time an individual was unemployed over a two-year interval.

3. Modeling unemployment risk

This section models the cumulative risk of unemployment. At any point in time, a worker could be in any of two states: Employed (E) or Unemployed (U).[7] A Markov process discussed in detail below determines the transition between these two states. This Markov process allows for duration dependence, i.e. the probability of transition from one state to the other varies with the time spent in the former. The specific parameters of the process also depend on a series of covariates that capture individual characteristics. Consider a worker that enters unemployment. The process described above determines a distribution for the total time spent in the unemployment state in all spells (including the starting one) over the following two years. We focus on this measure of unemployment risk.

To study the conditional distribution of this random variable we estimate the transition probabilities (hazard rates) between the employment and unemployment states. The next subsection details the statistical model estimated.

3.1. Estimating the hazard functions.

Generally, the duration of unemployment is modeled from the specification of the conditional probability of leaving unemployment (the hazard function). Such hazard function models have been extensively used in the economic literature over the last two decades (cf. e.g. Lancaster, 1990 and Heckman and Singer, 1984).

[7] We do not model transition in and out of the labor force.

A useful feature of the analysis of both employment and unemployment duration is that the standard job search and job match theory provides a framework for the specification and interpretation of these models in an economic meaningful manner. Within this framework, hazard models of the type considered in the statistical survival analysis literature are directly interpretable in terms of well-established economic models. For example, the conditional probability that an unemployed leaves unemployment in a small interval can be viewed as the product of two probabilities: the probability of receiving a job offer and the probability that the job offer is acceptable.

However, the economic theory is not informative about the shape of these conditional exit probabilities. Hence, we take as a starting point Cox´s (1972) proportional hazard form for them. We construct a piecewise constant baseline hazard function where the hazard functions are assumed constant within duration intervals and varying between them. Consider a grid of duration periods $\{0 = t_0 < t_1 < ... < t_J\}$, and for $j = 1,...,J$ let $\Delta_j = t_j - t_{j-1}$ denote the length of each of the J intervals. The hazard rates are constant within each of these duration intervals.

Let $J(t) = \max \{j \mid t_j < t\}$, so that $t_{J(t)} \leq t < t_{J(t)+1}$. Given a vector of covariates $\mathbf{x} = (\mathbf{x}_1, \mathbf{x}_2)$ and parameters $\beta = (\beta_0, \{\beta_j\}_{j=1,...,J})$, the hazard rates are given by:

$$h(t; \mathbf{x}, \beta) = g(\mathbf{x}_1, \beta_0)\, h_{J(t)}(\mathbf{x}_2, \beta_{J(t)}) \qquad (1)$$

where the specification adopted for the hazard functions is log-linear: $h(t; \mathbf{x}, \beta) = \exp(\beta_0' \mathbf{x}_1) \exp(\beta_{J(t)}' \mathbf{x}_2)$.

Given the above specification, the survival function S(t) satisfies:

$$S(t, \mathbf{x}, \beta) = \exp\left(- g(\mathbf{x}_1 \ \beta_o) \left[\underset{1 \le j \le J(t)-1}{} h_j(\mathbf{x}_2, \beta_j)\Delta_j + h_{J(t)}(\mathbf{x}_2, \beta_{J(t)})(t - t_{J(t)}) \right] \right) \quad (2)$$

Our data consist of spells that may have been completed or continued between two consecutive interviews. For both types of spells, we have information on elapsed duration at the time of the first interview, which we denote by t_0 months. In case of incomplete spells, elapsed duration in the second interval is given by $t_1 = t_0 + 6$, since the survey takes place every six months. In case of complete spells, the information is limited due to interval censoring. Letting δ denote the duration of the new spell, all we know is that $t_1 \in [t_o, t_o + 6 - \delta]$.

The conditional probability of a continuing spell, is given by $S(t_o + 6)/S(t_o)$ and the conditional probability of a completed spell is given by $[S(t_o) - S(t_o + 6 - \delta)]/S(t_o)$. Letting I_o denote the set of individuals with continuing spells and I_1 those with completed spells, the likelihood function is given by:

$$\begin{aligned} \ln L(\beta; \mathbf{x}) = & \underset{i \in I_o}{} \left[\ln S(t_i + 6; \mathbf{x}_i, \beta) - \ln S(t_i; \mathbf{x}_i, \beta) \right] + \\ & \underset{i \in I_1}{} \left[\ln \{ S(t_i; \mathbf{x}_i, \beta) - S(t_i + 6 - \delta; \mathbf{x}_i, \beta) \} - \ln S(t_i; \mathbf{x}_i, \beta) \right] \end{aligned} \quad (3)$$

Finally, note that by restricting our estimates to conditional probabilities, we circumvent the problems associated to length bias sampling and non-stationary of flows. This model has the property that some covariates affect the hazard rate proportionally while others affect the same differently depending on the duration interval where the hazard rate is evaluated. The specification adopted also allows us to circumvent the problem of interval censoring of the duration data. Additionally, it is needed in order to capture the extremely high transitions rates observed at the period of initial duration of both employment and unemployment.

3.2. Hazard rate estimates

Our sample consists of the matched rotating panels from May 1989 to October 1998. There are a total of approximately 64.000 individuals in the sample, evenly distributed throughout the sample period, of which over 44.000 have multiple observations. We further restrict the sample to those individuals with ages between 21 and 65 years old. Additionally, the estimate of the hazard rate from employment is conditional only on those individuals who are salaried employed with initial tenure under 5 years and that are still in the labor force the following period.[8]

The proportional hazard function from unemployment is a function of a set of personal characteristics while the piecewise baseline hazard function is a function of a set of dummy variables measuring duration dependence periods.[9]

The proportional hazard function from employment is a function of a set of personal characteristics and a dummy variable capturing the employment size of the firm where the individual works. It is 1 if the firm size is large (i.e. more than 50 employees). The piecewise baseline hazard function also varies by duration segment. The link

[8] Therefore, we exclude from the sample self-employed, owner-managers and unpaid workers

function of these baseline hazard is modeled as a linear function of a dummy variable for the period 1995-1998, a dummy variable that measures whether or not the worker is entitled to any of the employee benefits or contributes to the social security system associated to regular employment contracts in the formal sector (Dformal), the interaction of these two dummy variables and an intercept term.

The differential effect on employment stability postulated for the period 95-98 is due to the changes in the labor market legislation of 1995. This reform introduced a trial period for all employment contracts and a wide set of fix-term contracts. There is evidence that this type of reforms increase employment volatility. Cabrales and Hopenhayn (1997) present evidence for Spain that shows a significant increase in the hazard rate from employment after the rules for temporary employment were substantially relaxed. Additionally, there are well established theoretically arguments that shows that lower job matches termination costs implies higher turnover rates (cf. e.g. Bertola and Rogerson, 1997 and Hopenhayn and Rogerson, 1993).

In our empirical models, the age of the individual, sex (a dummy that equals one if the individual is male) and the level of education capture the individual characteristics. The schooling information is categorical. There is a set of dummy variables that measure the maximum level of the educational system attended by an individual and whether or not it has been completed. The educational categories are incomplete primary school, primary school, high school drop-outs, high school, incomplete tertiary degree and tertiary degree (Schooling i, i = 1,...,6). The base category in the likelihood functions is the incomplete primary school (Schooling 1).

[9] It is worth to note that in Argentina, the proportion of insured unemployed is extremely low (cf. Galiani and Nickell, 1999).

Table 3: Modeling the probability of leaving unemployment

Maximum likelihood estimates: single-risk model

Variable	Coefficient	P-value	Risk ratio
Age	-0.0154 *** (0.0022)	0.0001	
Sex	0.5232 *** (0.0536)	0.0001	1.687
Schooling 2	-0.1416 ** (0.0835)	0.0450	0.868
Schooling 3	-0.4348 *** (0.0912)	0.0001	0.647
Schooling 4	-0.3294 *** (0.0980)	0.0004	0.719
Schooling 5	-0.4023 *** (0.1112)	0.0001	0.669
Schooling 6	-0.2990 *** (0.1285)	0.0100	0.742
0 – 3 months	0.1053 (0.1773)	0.2763	
3 – 6 months	-0.4654 *** (0.1788)	0.0046	
6 – 12 months	-1.9962 *** (0.1838)	0.0001	
12 – 24 months	-1.9444 *** (0.1890)	0.0001	
Period effects	yes		
Mean log-likelihood	-0.695		
Number of cases	3073		

Notes: *** if the variable is statistically significant at the 1 percent level. ** if the variable is statistically significant at the 5 percent level.

Tables 3 and 4 respectively present the estimate of the probability of leaving unemployment and the estimate of the probability of leaving employment. For each model we report the coefficients, their standard errors, the probability value and the risk ratio. Naturally, the latter statistic is only reported for dummy variables.

In both cases the demographic covariates are highly significant. The hazard rate from employment decreases monotonically in age and the level of education. For example, the hazard rate from employment decreases 12 percent with 10 additional years to the mean sample age and it is 60 percent lower for someone with a tertiary degree. The hazard rate from employment is 22 percent higher for males. Finally, it is also lower if

Table 4: Modeling the probability of leaving employment

Maximum likelihood estimates: single-risk model

Variable	Coefficient	P-value	Risk ratio
Age	-0.0121 *** (0.0019)	0.0001	
Sex	0.2043 *** (0.0450)	0.0001	1.227
Schooling 2	-0.1290 ** (0.0686)	0.0299	0.879
Schooling 3	-0.2005 *** (0.0734)	0.0031	0.818
Schooling 4	-0.3497 *** (0.0784)	0.0001	0.705
Schooling 5	-0.4558 *** (0.0861)	0.0001	0.634
Schooling 6	-0.7320 *** (0.0973)	0.0001	0.481
Size	-0.1227 *** (0.0425)	0.0019	0.885
0 – 3 months			
Constant	-0.9560 *** (0.3861)	0.0066	
D95-98	0.3269 *** (0.1150)	0.0022	1.387
Dformal	-0.9129 *** (0.2174)	0.0001	0.401
Dformal * D95-98	0.2021 (0.2782)	0.2337	1.224
3 – 6 months			
Constant	-1.6869 *** (0.3814)	0.0001	
D95-98	0.1902 ** (0.1073)	0.0381	1.209
Dformal	-0.4193 *** (0.1325)	0.0008	0.658
Dformal * D95-98	-0.0978 (0.1543)	0.2629	0.907
6 – 12 months			
Constant	-3.7338 *** (0.4006)	0.0001	
D95-98	0.2089 * (0.1582)	0.0934	1.232
Dformal	-0.4404 *** (0.1660)	0.0040	0.644
Dformal * D95-98	0.3027 * (0.2288)	0.0929	1.354
12 – 24 months			
Constant	-3.6111 *** (0.3853)	0.0001	
D95-98	0.2851 (0.1131)	0.0059	1.330
Dformal	-0.3663 *** (0.1168)	0.0009	0.693
Dformal * D95-98	-0.5408 *** (0.1771)	0.0010	0.582
24 – 60 months			
Constant	-4.0442 *** (0.3843)	0.0001	
D95-98	0.3581 *** (0.1200)	0.0014	1.431
Dformal	-0.2689 *** (0.0075)	0.0075	0.764
Dformal * D95-98	-0.4695 *** (0.1593)	0.0016	0.625
Mean log-likelihood	-0.471		
Number of cases	14854		

Notes: *** if the variable is statistically significant at the 1 percent level. ** if the variable is statistically significant at the 5 percent level. * if the variable is statistically significant at the 10 percent level.

the worker is in a large firm and has a formal job. The hazard rate diminishes 12 and 40 percent respectively.

Regarding the hazard rate from unemployment, the schooling effect is not monotone and it is somewhat ambiguous. It appears that it makes a difference to finish primary school, but apart from that, there are little risk differences. The hazard rate from unemployment decreases 15 percent with 10 additional years to the mean sample age and it is considerable higher for males (70 percent).

Both hazard rates present strong negative duration dependence. The exit rate from employment increased about 40 percent after more flexible contracts where introduced in 1995.

In table 5 we present the mean survival rate both in employment and unemployment. The unemployment survival rates confirm that the unemployment duration of a spell is extremely low in Argentina. The employment survival rates are also extremely low and explain why we observe the remarkably high levels of turnover in the labor market documented in section 2. Clearly, an individual that is unemployed at least once in a period of two years is most likely to face multiple spells during that period.

Table 5: Mean survival rates (%): 1989-1998

Duration	Employment	Unemployment
3 months	48.2	21.9
6 months	30.0	10.3
1 year	26.6	7.6
2 years	21.0	4
5 years	23.2	0

Notes: Hazard rates are monthly and constant in the interval defined by two adjacent rows.

3.3. Unemployment risk

In this subsection we consider the risk of unemployment for an individual that enters unemployment. The objective is to evaluate how this risk is distributed among the labor force. Although the risk associated to a single spell is low, the extremely low employment retention rates induce multiple spells that may spawn a high level of unemployment risk.

Table 6 presents some location moments of the distribution of the time an individual that enters unemployment will spend unemployed over two years. Additionally, in the last column of the table we add the median of the distribution of unemployment incidences (repeated spells) over two years. In the first row we present these moments for the average individual that enters unemployment in 1998. Over that period, the probability of staying out of work more than 7 months is higher than 0.5. The expected mean time out of work is 8 months. If we compare this statistics with those corresponding to the average individual that entered unemployment in 1989, we observe that the entire distribution shifted to the right. The comparison of these two rows gives us a quantification of the increase in unemployment risk. For example, the median time an individual that enters unemployment will spend unemployed over two years has increased 35 percent. The expected median number of spells over two years has doubled, from 2 to 4. Actually, someone who enters unemployment expect to experience 4 spells of unemployment over a period of two years.

Table 6: Unemployment risk:
Moments of the distribution of the time an individual that enters unemployment will spend unemployed over two years

	Total time in two years				
	First quartile	Median	Third quartile	Mean	Number of incidences Median
Average individual (1998)	3.9	7.4	10.9	8.0	4
Average individual (1989)	2.4	4.8	8.5	6.1	2
Average incomplete primary school (1998)	5.9	8.8	11.1	8.6	4
Average tertiary degree (1998)	2.2	4.2	7.7	5.8	2
Average incomplete primary school 18 years old (1998)	7.2	9.4	11.0	8.9	6
Average females (1998)	3.9	8.2	13.4	9.3	3

Table 6 also presents these statistics for several demographic groups. As can be seen, a worker with tertiary degree experiences half the number of spells and approximately 30 percent less time out of work than an average unemployed. Females stay out of work longer even though they expect to experience fewer spells of unemployment. This is due to their lower hazard rate for exit from unemployment. Finally, the young unskilled face extremely high risk of unemployment: the median youth unskilled worker has 6 spells of unemployment and remains jobless over a year out of two.

What is the importance of long-term unemployment? As indicated above, standard measures of long-term unemployment underestimate the importance of total incidence through multiple spells. Accordingly, a new definition is called for. We will say that an unemployed worker is *long-term unemployed* if he has been in that state for more than one year during the last two years.

We construct a theoretical sample of unemployed workers by performing a Monte Carlo simulation of the estimated model. All explanatory variables where set to their sample mean values except for the year dummy set to 1998. A total of 10,000 sample paths were generated of 264 periods (months) each.[10] Our sample comprises all those paths that concluded in unemployment. For each path in this sample, we calculate the total time spent in unemployment during the last 24 periods. The mean value is 10.2 months, the median 8.6 and the percentile 0.75 is 14.5 months. Of all unemployed, 34 percent had been in that state for more than one year during the 2-year window: the long-term unemployed. This is more than twice the figure obtained without taking into account re-incidence and is close to the long-term unemployment figures for France.

[10] By taking a long sample path, we avoid the initial conditions problem.

4. Conclusions

This paper shows the importance of considering unemployment re-incidence in the analysis of unemployment risk. We show that although the duration of a typical unemployment spell in Argentina is very short, the average individual that entered unemployment in 1998 had a probability higher than 0.5 of experiencing a total of 4 or more unemployment spells over two years and cumulative unemployment of over a third of this two-year period. The number of incidences and cumulative unemployment is 50 percent larger for the high-risk group comprising young workers with low schooling. Finally, the risk of unemployment has increased considerably throughout the decade, doubling the median number of incidences and increasing cumulative duration by 35 percent. This is explained mostly by a declining survival time in employment.

Our estimates also indicate that, accounting for re-incidence, the fraction of long-term unemployed is close to the high numbers encountered in European economies. This obviously counters the view that unemployment is a small risk, short-duration phenomenon, which arises when re-incidence is not considered. Unemployment risk is high, has risen substantially in the last decade and is shared very unequally in the labor force.

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