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Some Facts and Policy Options[1]

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Abstract

Incomplete and highly fragmented work histories threaten to leave many contributors of the pension schemes in Latin America without the minimum pension guarantee or even without access to the ordinary pension. We propose a methodology to assess this risk, identify vulnerable groups and study potential determinants of the history of contributions using information from the work history records of the social security institutions. We apply this methodology to the largest social security institution of Uruguay, the *Banco de Previsión Social*, and show that the majority of contributors to this institution might not comply with the minimum number of years of contribution that is currently required to access an ordinary pension when they reach the retirement age.

JEL Classification: H55, J14, J26.
Keywords: density of contributions, work history.

Resumen

Historias laborales incompletas y altamente fragmentadas amenazan con dejar a muchos contribuyentes de los regímenes de pensiones de América Latina sin la pensión mínima garantizada, o incluso sin acceso a la pensión (jubilación) común. En el presente estudio, se propone una metodología para evaluar este riesgo, identificar los grupos vulnerables y estudiar los posibles factores determinantes de las historias de contribuciones, utilizando la información de los registros de historia de laboral de las instituciones de seguridad social. Aplicando esta metodología sobre los registros de la principal institución de seguridad social del Uruguay, el Banco de Previsión Social, se obtiene que la mayoría de los contribuyentes de esta institución podría no cumplir con el mínimo de años de cotización requeridos actualmente para acceder a una pensión (jubilación) común al llegar a las edades habituales de retiro.

Clasificación JEL: H55, J14, J26.
Palabras clave: densidad de cotizaciones, historia laboral.

1 INTRODUCTION

The low coverage of the social security systems in Latin America is a motive for concern. Coverage is low in many countries both among the elderly and the working population. The proportion of the elderly who are receiving pensions is a direct indicator of the effectiveness of the systems to provide income security in old age, but the proportion of the economically active population that is contributing is also important as it conditions the access to contributory pensions and the amount of the future benefit. Like in other developing regions, large segments of the economically active population do not contribute to the pension system in Latin America (Gill et al. 2003; Auerbach et al. 2005; Rofman and Lucchetti 2006).

The exact nature, causes and policy implications of the low participation of workers in the contributive pension system are likely to be different depending on whether contribution to social security is a permanent or a temporary status for each worker. Low aggregate coverage might arise because the population is segmented, with some workers contributing and other workers not contributing, or because many workers contribute only part of their working career. In the first scenario, those workers who are contributing to the system are basically protected and are expected to receive a contributory pension. In the second scenario instead, there might be a considerable number of workers who would not be well protected by contributive programs despite of being registered as contributors. Therefore, it is not only the average rate of contributors in the population what matters but also the turnover. A similar issue has recently been raised about labor informality in Latin America. Perry et al. (2007) describe a pretty dynamic picture, where many workers transit between formality and informality. Bosch and Maloney (2006) analyze the transitions between salaried and self-employed workers in Mexico and show that there is indeed a significant turnover.

Furthermore, the turnover of workers between contributing and not contributing is having an increasing impact on the access to contributive pensions in practice as the social security administrations are increasing their capacity to control the fulfillment of the vesting period conditions. Not long ago, the social security administrations had almost no records of the individual contributions they had collected, and could not check whether workers had accumulated the years of contributions legally required to access a pension. Benefits were hence granted on a very informal basis, often appealing to the testimony of witnesses. Albeit difficult to prove, there are many informal tales suggesting that having connections in the social security institutions and in the political system was also helpful to get a pension in some countries. It is no wonder then that many people managed to get a contributory pension even when they had not contributed the required number of years. This state of affairs is gradually changing in several countries in the region as the social security systems started to build work history records as part of the reforms that took place in the eighties and nineties. Social security administrations are increasingly being able to check individual records of contributions, leading to a better enforcement of the conditions to access to pensions. These are mostly good news, but there is also a risk that many workers do not adapt to the new conditions and end up with a very low pension or no pension at all. Aware of this risk, several governments in Latin America are currently considering options to reduce the number of periods of contribution required to access a pension. In the case of Chile, for example, the reform that the government proposed to the parliament in 2007 eliminates the condition that workers had 20 years of contribution to access the minimum pension guarantee. Colombia is also considering a reform that would reduce the vesting period (the so called *Beneficio Economico Periodico* or periodic economic benefit). Options to soften this requirement are also in the agenda of the social security authorities of Uruguay.

Some social security institutions have recently delivered samples of the work history records and some other institutions will probably follow their way in the near future. These rich and large panel datasets can in principle be used to assess the proportion of workers who could reach the required periods of contributions at the retirement age, but the available histories are still partial, usually about 10 to 15 years long. In this paper, we present a methodology to assess the access to contributory pensions building life-time work histories from incomplete histories and present results for the case of Uruguay. The paper aims to answer several questions. How many workers will not comply with the years of contributions required to access a contributory pension at ordinary retirement ages? Which are the most vulnerable groups? What factors condition the probability that a worker accumulates the required years of contribution at the usual retirement ages?

We applied our methodology to a database of the work history records of the largest social security institution of Uruguay, the *Banco de Prevision Social* (BPS). We found that, unless there is a significant change in the patterns of contribution in the future, compared to the period in which the sample was taken (1996-2004), a vast majority of the contributors to this institution will not comply with the years of contribution that are required to access a pension. Our results show that the situation will be particularly severe among low income workers working in the private sector. These gloom projections seem to be at odds with the comparatively high coverage that the old-age pension programs have achieved among the elderly in Uruguay. But it should be noticed that the current levels of coverage were achieved in a period in which the BPS was not really able to enforce the periods-of-contribution condition to access to pensions. This was so because of the lack of work history records. This situation is gradually changing as the institution accumulates records of contribution.

The literature that analyzes access to pensions in Latin America has used a variety of concepts, data sources and methodologies. The most widely used data source is the household survey. Analysts have long been using this data to compute coverage of the labor force as the proportion of the economically active population that contributes to the social security system. Rofman and Lucchetti (2006) provide extensive computations of this and related indicators across several Latin American countries and explore the links between the rate of coverage and several key socio-economic characteristics (like education, sector of activity, size of the firm, etc.). Using the same data source, several researchers have used qualitative dependent variable models to estimate the probability that workers contribute to social security and to analyze what factors impact on this probability (Auerbach et al. 2005; Barr and Packard 2003; Barrientos 1996 and 1998; Bustamante and Paola 2006; Holzmann et al. 2000; Li and Olivera 2005; Packard et al. 2006; Packard 2001; among others). The household surveys are rich data sources but they only provide cross section information at individual level, so these surveys are not suited to the type of questions we aim to answer. Besides, some household surveys provide information about affiliation to social security but not about actual contributions. It has been reported in several countries in Latin America that the number of affiliates is significantly larger than the number of effective contributors at any point in time.

Taking advantage of the recently available work history data bases, several analysts have been looking at the individual density of contributions, i.e. the proportion of lifetime that each worker contributes to social security (Bertranou and Sanchez 2003; Arenas de Mesa et al. 2004; Lagomarsino and Lanzilotta 2004; Bucheli et al. 2005 and 2006; Bravo et al. 2006; Berstein et al. 2005 and 2006). Building on this literature, we use in this paper survival analysis to model the transitions between contributing and not contributing. Survival analysis allows us to adopt more flexible assumptions than those adopted in the literature mentioned above to address this issue. We then use the transition rates to simulate the complete work histories of hypothetical workers using Monte Carlo simulations.

In the next section, we present a brief summary of the Uruguayan pension system. We describe the data in section 3. In section 4 we present the methodologies for the estimation of the hazard rates and the simulation of the work histories. The main results are reported in section 5 and section 6 concludes.

2 THE SOCIAL SECURITY SYSTEM IN URUGUAY

In Uruguay, the first old-age, survival and disability (OASDI) insurance programs were set up in the 19th century. Since then similar programs have spread and became almost universal in the 1950s, incorporating new programs to cover against other risks. In 1967, a public institution called *Banco de Previsión Social* (BPS) was created with the purpose to be in charge of the social security system. Since then, BPS administers four large retirement programs that covered public servants, private workers (with some exceptions), rural workers and domestic workers. In addition, some categories of workers have their own special pension schemes: bank employees, notaries, self-employed university graduates, armed forces personnel, and police force personnel. By 2001, the number of contributors to the BPS represented 89% of the total number of contributors to all social security institutions in the country (Ferreira-Coimbra and Forteza 2004).

Before the social security reform initiated in 1996, the Uruguayan retirement system relied on a pay-as-you-go (PAYG) system financed through payroll contributions paid by employers and employees. Contribution was also mandatory for self-employed workers who were subject to minimum declared earnings.

In 1996, the reform modified the BPS retirement program introducing individual savings accounts to complement the PAYG system. Workers with wages below a threshold continue to be served by the PAYG regime administered by BPS unless they explicitly choose to deposit half of their contributions in a personal account. Workers with higher wages must contribute to both pillars. For the amount below the minimum they contribute to the PAYG public system and from there up to a certain maximum, also established by law, they must contribute to individual accounts. There is no mandatory contribution for earnings over the established maximum. Employers' contributions go exclusively to the public PAYG pillar.

The reform introduced other modifications, mainly with the aim of strengthening the link between contributions and pensions, and postponing the retirement age. The minimum retirement age was fixed at 60 for men and women, which meant an increase of 5 years for the latter. Also, the minimum number of years of contribution required to access a full contributive pension was raised from 30 to 35. However, workers can receive a less generous contributory pension with only 15 years of contribution at70 years old. Workers with hazardous occupations and other special categories have special bonus on their count of years of contribution.

The replacement rate was modified, making it more sensitive to both retirement age and years of contributions, in order to induce longer working-lives. Before the reform, the replacement rates ranged from 65% to 80% for women and from 60% to 80% for men. The reform eliminated gender differences and the replacement rate now ranges from 50% to 82.5%, depending on the years of contribution and the retirement age. The average wage used in the benefit formula was modified to include a longer period of contributions. In addition, there is an extra bonus for low-income workers who choose to contribute to individual savings accounts.

Persons who do not satisfy the vesting period requirements to access a contributory pension at 70 years old (15 years of contribution) may be eligible for a means-tested pension program. With the reform, the minimum age required for this benefit was raised from 65 to 70. The means-test assesses the family income, including income of family members who do not live with the beneficiary. The benefit consists on a sum that allows the beneficiary to reach certain monetary income fixed by BPS. This amount is considered as a minimum threshold to cover basic needs. The data indicate that it is lower than the usual contributory pensions.

The reform also established the systematic registration of workers' labor history. Before that, insomuch as these administrative records did not exist, the control of the actual years of contributions was very difficult and done on very informal bases, such as the testimony of witnesses.

3 THE DATA

We used a random sample of the work history records of the BPS, collected in December 2004 by the Labor History Unit of the BPS (ATYR-BPS). Workers in the sample were between 18 and 70 years old and contributed at least one month between April 1996 and December 2004. The sample has 68,997 individuals.

The records are organized in five databases. One file gives personal information on individuals: date of birth, sex and country of birth. Another file reports about the job of each person, particularly the date of initiation of activity and the explicit end of the link between the worker and the firm. A third file reports monthly information about the contributions. In particular, we have information on wages and some characteristics of the job. Also, there is a database reporting information about benefits that allowed us to know the date of retirement. Finally, there is a database that provides information on firms' characteristics (size, industry, etc.), but we do not use it in this paper.

Table 1 shows the number of individuals in the database by cohort, gender and private/public condition. The gender composition is quite similar to the one found by Bucheli (2004) based on the Household Surveys. Specifically, there are 37,822 men (55%) and 31,175 women (45%). We considered as public worker everyone who had worked in the public sector for at least half of the total time he or she had contributed. According to this criterion, we identified 58,617

(85%) private and 10,380 (15%) public workers in the database. While public sector contributors are evenly distributed between sexes (51% are women), private sector contributors are predominantly men (44% are women).

The database contains two variables that report the beginning and end of the job spells. The spells may include some non-working periods –like low season in the case of seasonal workers or workers on sickness or maternity leave–, which are computed for contribution purposes.

A descriptive statistic that has been widely used to describe the contribution patterns is the so-called "density of contributions". This variable measures the proportion of months that a worker contributed over the potential months he could have contributed. The average density of contributions is 60% in our sample. Lagomarsino and Lanzilotta (2004) report a density of contributions significantly higher, of around 75%, between January 1997 and December 2003. They used a sample of the workers registered in the labor history records of the BPS which was collected in the second semester of 1996. This collection choice has problems given that it loses all the workers who entered the database afterwards. In addition, workers with intermittent activities in the BPS and lower densities of contribution are less likely to be included in their sample than workers with higher densities of contributions. As a result, their calculations on the density of contributions are biased-upwards.

The distribution of the density of contributions has two modes and is strongly asymmetric, characteristics also pointed out by Lagomarsino and Lanzilotta (2004). A similar pattern has been reported in Argentina (Farall et al. 2003; and Bertranou and Sánchez 2003). In our sample, 28% of the workers contributed 100% of the period. This was the most frequent density of contributions in the database. Besides, over 40% do not register contributions for at least half of the potential months of contribution (Table 2).

Men present higher densities of contribution than women. Indeed, on average men contributed 61.4% and women contributed 58.0% of the time.

As expected, public sector workers have significantly higher densities of contribution than private sector workers. On average, public sector workers contributed 85.4% and private sector workers contributed 55.3% of the time. While more than two thirds of public employees contributed the whole period, only 21% of private employees contributed that much. Moreover, a significantly smaller proportion of public than of private workers contributed less than half of the time. There is however a considerable number of individuals classified as public employees who present low densities of contribution. This unexpected result responds in part to the classification as public employees of workers who contributed only part of the time as public employees. Also, this estimation does not correct for those activities with special bonuses, i.e. activities that compute more than one year of contribution per every year of actual contribution. This is the case of most teachers.

Additionally, we grouped the individuals in the sample in quintiles of the earnings distribution. In order to avoid the circular reasoning of finding low densities for workers whose low average income is due to few periods of contribution, we calculated the average earnings over periods in which individuals reported strictly positive earnings. Then, we grouped five-year generations workers by sex and we calculated the quintiles of the earnings distribution for each group. We are aware that earnings in uncovered jobs are unknown. If non-contribution is associated with high earnings jobs, the classification of workers by quintile will be non-accurate. Although some future effort may be done to improve the solution of this issue, it is worth to note that studies for Uruguay suggest that non-covered jobs are associated with low earnings (Amarante, 2002; Bucheli and Ceni, 2007). The average density of contribution consistently rises with the quintile of these distributions. Indeed, the average density of contribution is almost 38% for the poorest quintile and more than 80% for the richest quintile.

There are also significant differences according to individuals' age. At 20, the density of contribution is about 30% on average and it continuously increases with age, exceeding 75% when workers are in their fifties. However, there is an important dispersion for individuals of the same age, as it is shown in Table 3.

The business cycle seems to have had a significant impact on the density of contribution. The Uruguayan economy began a recession in 1999 which was followed by the most severe economic crisis in the country's history. The recovery began slowly in 2003, and 2004 was already a year of significant growth. This evolution is reflected in the unemployment and employment rates observed between 1996 and 2004. In this period of significant macroeconomic volatility, the density of contribution mirrored the rate of employment (figure 1).

4 METHODOLOGY

We want to estimate the proportion of workers who would reach the number of periods of contribution required to access a pension at the normal retirement ages and to identify the most vulnerable groups. It is not possible to estimate this proportion directly though, partly because the conditions are changing and workers and firms also change their behavior, but also because we do not have full lifetime work histories in the region. With the incomplete work histories we have, it is not possible at the moment to directly estimate how many periods of contributions workers have accumulated when they reach the retirement age. We propose in this paper a two-stage methodology to estimate the distribution function of the counting of periods of contributions along lifetime. In the first stage, we estimate the rates of transition (or hazard rates) between contributing and not contributing in each period along the lifetime. In the second stage, we simulate the lifetime work histories using the hazard rates estimated in the first stage and compute the distribution functions of the counting of periods of contribution at different ages.

4.1.1 First stage: estimating the hazard rates

Consider a worker who may be in any of two possible states: contributing and not contributing to social security. Depending on circumstances and his choices, the worker will be making transitions between these two states. Let $h_c(t, X_t)$ be the probability that a worker who *does* contribute to social security during month *t* stops contributing in *t+1* and let $h_n(t, X_t)$ be the probability that a worker who *does not* contribute to social security during month *t* starts contributing in $t+1$. These probabilities are the (discrete time) transition rates or hazard rates of the states "contributing" and "not contributing" respectively. The hazard rates may depend on several variables represented by X_t.

In order to identify the hazard functions, different alternative assumptions are usually made in the literature. The most common one is that the hazard rate can be decomposed in two multiplicative terms, one that summarizes the impact of duration in the state, the so called baseline hazard, and a term that summarizes the impact of the covariates X_t. This model has been called *proportional hazard* because the hazard rates of two individuals who differ only in time-invariant covariates maintain a constant ratio, which is in turn proportional to the absolute difference in the covariates.

In a discrete time framework, the proportional hazard assumption leads to the following specification (Jenkins 2005, pp 41-2):

$$h(t, X_t) = 1 - \exp[-\exp(\beta' X_t + \gamma_t)]$$

where γ_t summarizes the baseline hazard. This model is known as the *cloglog model*, because a complementary log-log transformation of the hazard is a linear function of the baseline hazard and the covariates:

$$\log[-\log(1 - h(t, X_t))] = \beta' X_t + \gamma_t$$

An alternative identifying assumption that is often used in discrete time models is that the odds ratios are proportional to the absolute difference in the covariates. This assumption leads to the *logistic hazard model*:

$$\log it[h(t, X_t)] = \log\left[\frac{h(t, X_t)}{1 - h(t, X_t)}\right] = \beta' X_t + \alpha_t$$

where α_t summarizes the relative odds of making a transition when $X = 0$.

In order to identify γ_t and α_t in the cloglog and the logistic models, additional assumptions have to be made about the underlying functional forms. Lacking specific theoretical guidance, we adopted the usual practice of using dummies to represent duration and age. Once the empirical duration and age patterns could be identified, we chose some more parsimonious functional forms to facilitate the simulations. We decided to use a polynomial of degree three on age and the log of duration. As the pattern of duration might vary along the life cycle, we included two variables of interaction between duration and age.

The work history data set presents some characteristics that make the estimations of the hazard rates relatively complex. The data contains censoring and truncation, multiple spells and unobserved heterogeneity. We briefly explain how we dealt with each of these complications in what follows.

A spell of contributions is expected to end when the worker transits from contributing to not contributing, but the observed spell can also end because of the end of the observation period. If this happens, we only know that the worker did not transit to the other state before the last period of observation but we do not know whether the worker made or made not a transit afterwards. This right censoring is not a major problem for the estimation of the hazard rates though. It is enough to acknowledge the fact that the only information we have about the last observation is that the individual survived in the state at least until that period.

In our data set, right censoring occurs in two different cases. First, observations are censored at the end of the work history sample. We do not know the contributing state of any worker after December 2004. Second, a worker may die or retire during the period of observation. Death and retirement could be thought of as different states in the context of a competing risks model. In this context, the destination-specific censoring indicators of the states contributing and not contributing take value zero when the individuals die or retire.

There is left censoring when the starting date is not observed. In our data, the spells of not contribution that began before April 1996 are left censored. We did not use these spells in the estimations. There is no left censoring of the spells of contribution in this data set. Even when the spell may have began before the initial observation date we have the information about when these spells began.

There is truncation when some survival times are systematically excluded from the sample. Left truncation occurs when individuals who did not survive enough are excluded and right truncation occurs when individuals who survive too much are excluded. In our data set, there might be left truncation of the spells of contribution. The work history database captures all individuals who contributed at least one month between April 1996 and December 2004. Consider two workers who began to contribute say in January 1990, but one left one year later and never returned and the other one continued at least until April 1996. While the second worker will be registered in the social security database, the first worker with the shorter spell will be excluded.

Unobserved heterogeneity may significantly bias these estimations. Over time, the proportion in the population of individuals with high risk of leaving the state declines. Thus, the average hazard rates of mixed populations tend to decline over time even if the “true” hazard rates of the individuals in the population rise. In order to reduce uncontrolled heterogeneity, we worked separately with several categories of workers whose behavior is likely to differ. We distinguished men and women, public and private workers and quintiles of the labor income distribution. For each category, we run formal tests of unobserved heterogeneity and controlled for the heterogeneity that could still remain modeling it as an individual effect. Assuming that the individual effects are normally distributed with zero mean, the hazard rates for mixed distributions can be estimated using random effects complementary loglog or logit models (Jenkings 2005, pp 84-5). We used the xtcloglog and xtlogit commands in STATA to estimate these models (Stata 2003). The program reports the likelihood ratio test for the null hypothesis that the variance of the individual effects is zero. Frailty is not important if this hypothesis cannot be rejected. Otherwise, frailty matters and the random effects estimation takes it into account.

The strategy to model the probability of contributing used in this paper is more general than previous attempts we could identify in the literature (including our own previous work). The first studies on this subject that we could find assumed that the probability of contributing is independent of previous status, which is a strong assumption as Bucheli et al (2005) explicitly acknowledge (Bertranou and Sanchez 2003; Arenas de Mesa et al. 2004; Lagomarsino and Lanzilotta 2004; Bucheli et al. 2005; Bravo et al 2006). The assumption of independence actually amounts to assuming the following two hypotheses (i) the hazard rates do not depend on duration, and (ii) one minus the hazard rate of the state "contributing" equals the hazard rate of the state "not contributing". These hypotheses not only look strong, but they can also be tested and so there is no need to just assume them as true.

In a recent study, we extended the framework adopted in the papers mentioned above, allowing the probabilities of contributing to depend on the previous period status (Bucheli et al. 2006). We dropped the assumption that one minus the hazard rate of "contributing" equals the hazard rate of "not contributing", but we still maintained the not very appealing assumption that the hazard rates do not depend on duration in the state. This hypothesis was formally tested and systematically rejected in the framework proposed in this paper.

4.1.2 Second stage: simulating the work histories

Our final goal is to build empirical distribution functions of the number of periods of contributions at the usual retirement ages. This can be done analytically, if the probabilities of contributing each period are independent of previous period status, as we have shown in a previous study (Bucheli et al. 2005). But it cannot generally be done analytically when the probabilities of contributing depend on previous periods status and vary along the life cycle. In this case, work histories are determined by a non-homogenous Markov chain. We performed Monte Carlo simulations to overcome this difficulty.

The simulation of the work histories entails building strings of "c" and "n" (for contributing and not contributing, respectively) that adequately replicate the stochastic properties of the observed incomplete histories. The simulated worker contributes in t if either he was contributing in *t-1* and did not make a transition to not contributing or he was not contributing in *t-1* and made a transition to contributing. Let p represent the propensity to make transitions and let us assume that it is drawn from a uniform distribution in the $[0,1]$ interval. The individual contributes in t if $p \geq h_c(t-1, X_{t-1})$ and he was contributing in *t-1*, or if $p \geq 1 - h_n(t-1, X_{t-1})$ and he was not contributing in *t-1*. With this rule, the probability that an individual who contributes in *t-1* also contributes in t is $1 - h_c(t-1, X_{t-1})$, which is the probability of not leaving the state "contributing". The probability that an individual who does not contribute in *t-1* contributes in t is $h_n(t-1, X_{t-1})$, which is the probability of leaving the state "not contributing".[3]

[3] The conditional probability of contributing each month can be thought of as the probability of not leaving the state "contributing", if the individual contributed the previous month, and as the probability of leaving the state "not contributing", if the individual was not contributing the previous month. The former is one minus the (discrete time) hazard rate of the state "contributing" and the latter is the hazard rate of the state "not contributing". In turn, the conditional probability of not contributing can be thought of as the hazard rate of the state "contributing", if the

The above algorithm was applied to the lifetime of each simulated individual. The simulations began at age 18 in the state "not-contributing" and ended at age 70. The number of months of contribution accumulated at any age could then be counted in each simulated work history. Repeating this procedure many times, we got empirical distributions of the counting of the months of contribution at the desired ages.

5 RESULTS

5.1.1 The hazard rates

We estimated the complementary loglog and logit models, getting very similar results. For the sake of brevity, we only present here the results obtained with the complementary loglog model.[4] The results are summarized in tables 4 to 11.

The unemployment rate has a significant positive impact on the hazard rate of contributing and a significant negative impact on the hazard rate of not contributing for most categories of *private* workers. This is to be expected, as in periods of high unemployment workers stop contributing at higher rates because of job loss and transition to informality, and find it more difficult to get formal jobs. However, we do not find this pattern for men in the poorest quintile and women in the poorest two quintiles. Higher hazard rates of contributing and lower hazard rates of not contributing lead to lower densities of contribution during downturns. In the case of Chile, Bertsein et al. (2006) report that the unemployment rate has a negative impact on the density of contributions of men, but positive on the density of contributions of women.

Not surprisingly, the unemployment rate does not impact on the hazard rates of contributing and not-contributing for most categories of *public* workers. However, some middle income public workers show *lower* hazard rates of contributing during recessions. A possible explanation for these contrasting patterns is that while private workers lose jobs during downturns, public workers choose not to quit when it is more difficult to find a job in the private sector.

Age has an impact on the hazard rates of contributing and not-contributing as well. Young workers tend to have higher hazard rates of both contributing and not contributing than middle-aged workers, implying that the turnover rates are particularly high at the beginning of the working career. Senior workers tend to show higher hazard rates of contributing and lower hazard rates of not contributing than middle-aged workers. This is to be expected, as senior workers are more likely to stop contributing and less likely to restart contributing after they stopped than workers who are in the middle of their working careers and have not yet arrived to the retirement age.

In both states, the probabilities of making a transition reduce as the workers spend time in the state. Indeed, duration has a highly significant negative impact on the hazard rates of both contributing and not-contributing. In most regressions the coefficient for duration was found to

individual was doing so in the previous month, and as one minus the hazard rate of the state "not contributing", if the individual was not contributing in the previous month.

[4] The results obtained with the logit model are available from the authors upon request.

be negative and different from zero at 0.1% significance level. The hazard rates drop initially fast and tend to stabilize for long durations. The estimated coefficient varies across categories, but it is worth noting that even in the public sector duration has a highly significant impact on the hazard rates.

Young workers tend to have not only higher but also more resilient hazard rates than workers aged 30 to 59 years. Like their more mature fellows, young workers show decreasing probabilities of making a transition out of any of the two states as they spend time in the state, but the rate at which these probabilities decline is lower in the case of workers below 30 than in the case of workers aged 30 to 59. In other terms, duration has a higher negative impact on the hazard rates of contributing and not contributing when workers are between 30 and 59 years old than earlier in their lifetime. The coefficient of the variable that captures the interaction between duration and the dummy variable for age 30 to 59 is negative in all regressions and significantly different from zero at 0.1% significance level in most regressions. The pattern is less clear for senior workers (aged 60 and above): some categories present negative, some positive and some non significant coefficients multiplying the corresponding duration-age interaction variable.

Table 12 provides some simulated hazard rates of contributing in a typical year, i.e. with the unemployment rate at the mean of the period. For instance, the average 40 years old man working in the private sector and belonging to the richest quintile has an expected hazard rate of about 3.8% in the first month and only 0.7% in the 36^{th} month of contribution. The same individual at 25 would have an expected hazard rate of 4.9% in the first month and 1.0% in the 36^{th} month of contribution.

Table 13 provides some simulated hazard rates of not contributing in a typical year. The average 40 years old men working in the private sector and belonging to the richest quintile has an expected hazard rate of about 6.8% in the first month and only 2.0% in the 36^{th} month he spends in the state "not contributing". The same individual at 25 would have an expected hazard rate of 7.8% in the first month and 4.0% in the 36^{th} month in the state not contributing.

The simulated hazard rates also indicate that lower income workers tend to have higher risk of leaving the state "contributing" and lower risk of leaving the state "not contributing". Therefore, low income workers are less likely to contribute than high income workers because they have higher probabilities of leaving the state contributing and lower probabilities of entering this state.

Workers in the public sector present higher hazard rates of not contributing than private workers. They also present lower hazard rates of contributing. Thus, public workers have higher probabilities of contributing than private workers because they have lower probabilities of leaving the state contributing and higher probabilities of entering into this state.

Considering these results altogether, some distinctive patterns emerge. First, the hazard rate of both states decline with duration. As workers spend time in any of the two states, the chances of staying in the state rise. History and luck during the first periods in the state seem thus to be crucial for the fate of the working career. Second, young workers have higher risk of quitting the state contributing but also of quitting the state not contributing, i.e. they are more mobile.

Third, economic downturns raise the risk that private workers stop contributing and reduce the chances that they start contributing.

5.1.2 The simulations

According to our estimations, there is a serious risk that a sizeable proportion of the workers registered in the work history records of the BPS will not be able to accumulate the 35 years of contribution required to be eligible for a pension when they reach the usual retirement ages. If the frequencies of monthly contributions observed between 1996 and 2004 remain unchanged, only about 21% of contributors will have made the required 35 years of contributions by the age of 60 and about 29% by the age of 65 (Table 14). The 35-years-of-contributions condition will be binding for a vast majority of contributors.

As expected, the problem is more serious for women than men. While only 27% of women satisfy the requirements at 65, 34% of men manage to do so at the same age. These figures fall to 20% and 25%, respectively, at 60.

Public employees are more likely to meet the requirements than private workers. Only 20% of men and 16% of women working in the private sector satisfy the access condition when they turn 65 years old. In contrast, 74% of men and 72% of women working in the public sector satisfy the condition at the same age.

We have also found big differences between workers in different income brackets. Low-income workers are much less likely to meet the accessibility requirements than higher-income workers at any given age. In fact, private workers in the poorest quintile are practically out of the system. In contrast, 95% of the men and 86% of the women in the richest quintile working in the public sector would access a pension when they turn 65 years old.

We also estimated the probability of accumulating 30 rather than 35 years of contribution at the retirement ages (Table 15). The difference between these two estimations provides a rough approximation to the *direct* impact on pension coverage of changing the minimum number of years of contribution required to access a pension. This measure does not take into account the *indirect* effects of such changes, though. It is possible, for example, that a reduction in the number of periods of contribution required to access a pension induce some workers to contribute less periods. It is also possible that an increase in the number of periods of contribution required to access a pension, like the one approved in 1995, provides incentives for some workers to contribute more periods. If this is the case, the *total* impact on coverage of changing this condition to access a pension will be lower than the *direct* impact estimated here. Therefore, the estimated difference between the probabilities of contributing 30 and 35 years at the retirement provide an upper bound to the total impact of this change.

Given certain probabilities of transition between contributing and not contributing, the proportion of individuals who accumulate 30 years of contributions is necessarily higher than the proportion of individuals who accumulate 35 years of contributions at any given age. If the density of contributions observed between 1996 and 2004 remain unchanged almost 35% and 41% of the population registered in the labor history records will accumulate 30 years of contributions when they turn 60 and 65 years old, respectively (Table 15). The difference between the proportion of workers who would contribute 30 and 35 years is around 13 percentage points for both ages considered. The difference is similar for men and women. Changes of the number of years of contribution required to access a pension seem to impact more strongly on individuals with high and medium densities of contribution than on individuals with low densities. Indeed, workers with very low densities of contributions are not likely to contribute either 30 or 35 years.

According to these results, the increase in the vesting period passed in 1995 will likely cause a sizable decrease in the proportion of contributors who will fulfill this requirement. This tightening of the conditions to access a pension is therefore likely to cause a significant reduction in the proportion of individuals that can retire at any given age, and an increase in the average retirement age. Obviously, for these phenomena to take place, the social security administration has to actually check individual records of contributions and enforce the pension access conditions. This was not warranted in the past, but the ability of the administration to do it is growing gradually as information based on labor history records is being accumulated.

The simulations are sensitive to the business cycle. We showed in the previous section that the unemployment rate impact on the hazard rates and hence the simulated work histories depend on the assumptions made about the unemployment rate. The simulations presented so far are based on the average unemployment rate in the period covered by the work history database, i.e. between April 1996 and December 2004, which was 13.45%. In order to analyze the sensitivity of the results to this variable, we estimated the proportion of workers who would accumulate 30 and 35 years of contribution with the unemployment rate at the 1981-2006 average, which was 11.20% (tables 16 and 17).

As expected, we find a higher proportion of workers accumulating the required years of contribution when we run the simulations with the lower unemployment rate. For example, the proportion of men in the private sector who would have accumulated 35 years of contribution at the age of 65 grows from 20 to 30% (tables 14 and 16). In the case of women, this figure would rise from 16 to 23%. However, the poorest workers are practically unaffected by the change in the unemployment rate. Most of them are basically out of the system with any of the two unemployment rates used in the simulations. The business cycle has a more significant impact on the work history of workers in higher quintiles of the distribution.

The work histories of public employees do not seem to be very sensitive to the business cycle. The proportion of workers who accumulate 30 and 35 years of contribution at the retirement ages is very similar when the simulations are run with the two different unemployment rates.

Overall, these results suggest that even though the densities of contribution would be higher in periods with “normal” unemployment rates than in the period in which the work history sample was taken, the proportion of workers who would comply with the required years of contribution at the usual retirement ages would still be very low. Running the simulations with the average unemployment rate of the last 25 years, we still get that only 30% of men and 23% of women would have accumulated 35 years of contribution at the minimum retirement age, i.e. at 60 (table 16).

The simulations presented so far were based on computations of the periods of service for regular jobs. There are some occupations that compute more than one period of service per period of effective contribution though. This is the case of most teachers and individuals working in risky activities. For example, workers handling radioactive substances register 3 periods of service every 2 periods of effective contribution, and hence they are required to contribute a significantly smaller number of periods to be eligible for a pension. The special regimes are relatively frequent among some groups of public employees, but they are much less frequent among private employees. Hence we do not expect these regimes to have a significant impact on the estimations of the proportion of private employees who would qualify for a pension. Women in the richest quintile register the largest incidence of special regimes among private workers, and yet only around 3% of them have special bonuses for at least half of the time they have contributed.

In contrast, over 30% of every income group of women working in the public sector has a special bonus at least half of the total time they have contributed. This figure exceeds 46% when it comes to women in the richest quintile. Workers that present the lowest incidence of these regimes in the public sector are men in the fourth highest quintile, and almost 12% of them work in activities with bonus for at least half of the time.

We computed a proxy for the number of periods of service of public workers in activities with special bonuses. A worker who contributed one period in any of these activities will be granted more than one period of service (the exact number depending on the bonus legally attached to the activity). This correction of the original estimation increases the estimated proportion of public workers who would satisfy accessibility requirements for every subgroup of public employees. As expected, the difference between the estimations that take into account special bonus and those which do not is much bigger among women than men. The proportion of public workers who accumulate 35 years of contribution at 60 years old rises by 3.4 and 15.6 percentage points, men and women respectively. In the case of women in the fourth highest quintile, the difference between the two estimations is as high as 25 percentage points at the age of 60 and 12 percentage points at the age of 65. These results suggest that special bonuses may have a great impact on the proportion of public workers who satisfy the accessibility conditions, in particular in the case of women.

6 DISCUSSION AND POLICY OPTIONS

The low percentage of contributors to the BPS who would fulfill the access conditions according to our estimations is at odds with the current high coverage of the system in old age. The inability of the institution to enforce the legal conditions to access to contributory pensions explains the puzzle. Because of the lack of hard data on contributions made before the reform, the BPS cannot verify that those who claim a pension are actually entitled to it. Furthermore, the BPS had to issue several norms to facilitate the process of recognition of pension rights. By one of these norms, the testimony of witnesses is accepted as a proof of activity and contributions. There is also a law that validates all declarations of activity made before June 1978. More recently, a resolution of the directory of the BPS established that the ordinary pension can be granted to anyone who can prove at least half of the number of years legally required. The claimant only has to declare that the years he cannot prove were completed before the inception of the work history records, i.e. before April 1996. These norms represent a radical loosening of the conditions to access to a pension.

The very short working careers identified in this paper poses a challenge to the sustainability of the system and/or the adequacy of the pensions. Indeed, the ability of the BPS to provide adequate pensions is compromised by these short histories of contributions.

The diversity of work histories identified in this paper also represents a challenge for the design of the benefit formula. Old-age pension schemes are designed to provide insurance basically against uncertain longevity, but the diversity of contribution trajectories we identified suggests that Uruguayan workers might also face considerable uncertainty about their working careers. In principle, the pension benefit could be designed to address this source of uncertainty: the pension should be basically flat relative to the periods of contribution. Then someone who was luck enough to have a long working career would contribute more than someone who could not have that career and both would receive the same pension. But unlike death, the event of not having a formal job is not only -and maybe not mainly- a matter of chance, it is also a matter of choice. If this is so, a benefit formula that provides the same pension regardless of the number of periods of contribution could significantly distort incentives, inducing individuals to choose shorter contribution trajectories. This is a clear case of moral hazard. Because of this, most pension schemes provide only partial protection against the risk of having short working careers.[5]

The BPS will likely have to loosen the years-of-contribution condition as the institution monitoring capacity increases. On paper, the BPS has comparatively tough access conditions. This is much less so in practice, because of the already mentioned inability of the institution to enforce the legal requirements. The tightening of the access conditions that took place after 1996 has not caused a dramatic drop in the number of pensioners simply because it is not fully enforced, but as the administration is increasing its capacity to monitor contributions, this requirement will likely become increasingly binding. Recognizing this situation, the authorities of the institution are currently evaluating a proposal to loosen the years-of-contribution condition.

Other things equal, a loosening of an eligibility condition should have a negative impact on the budget of social security, since more workers would access to the benefit. The government would then have to increase taxes and this could raise equity and efficiency issues.[6] However, a *de jure* loosening must not lead to a *de facto* loosening, if enforcement is improved. In the Uruguayan case, the motivation for the change in the norm is that the condition is expected to be better enforced as the administration accumulates work history records plus the observation that the existing condition is too tough to be enforced in practice. Therefore, the loosening of the *de jure* access condition is not meant to cause a *de facto* loosening.

The gap between *de jure* and *de facto* pension policies creates political risks in pensions. The policy becomes discretionary and not transparent. The lack of enforcement of the *de jure*

[5] The issue has been analyzed by Diamond and Mirrlees (1986). Gruber and Wise (2002) provide empirical evidence of the distortions in the retirement decision introduced by pensions. Valdés (2002) provides a survey of the literature.
[6] See Valdés (2002, chapter 1) for a survey of the literature on efficiency and equity in social security.

eligibility conditions opens the window for patronage and clientelism. The unwritten procedures to grant pensions in this policy regime are likely to raise uncertainty. This political risk is likely to be reduced as the administration increases its capacity to enforce the formal rules of the pension system. However, there is still another potential source of political risk: the vesting condition might not be credible. Workers might not believe that the condition will remain unchanged if many workers do not comply with it. The workers incentives to comply with the vesting condition would be eroded if they thought that the condition will be loosened. In turn, the government would be forced to loosen the condition if many workers did not contribute the required minimum number of years.[7] Setting a realistic eligibility condition and a sensible benefit formula that provides the right incentives is likely to ameliorate the credibility problem.

On the other hand, to acknowledge the moral hazard problem, the replacement rates to be set for less than 35 years of contribution must be sufficiently sensitive to the number of periods of contribution.

A loosening of the years-of-contribution requirement to access a pension would raise the number of individuals who receive the minimum pension. There are two reasons for this. First, this policy would probably increase the proportion of low income workers who would fulfill the years-of-contribution condition. They are particularly likely to benefit from this policy because, as we have shown, they typically have low densities of contributions. With less stringent conditions, many of these workers might become eligible for an ordinary pension, but their low contributions would not allow them to receive more than the minimum. Second, allowing contributors to receive a pension with fewer contributions would raise the number of pensioners receiving the minimum pension if the replacement rates were set to match the reduced contributions.

Individuals receiving the minimum pension have basically no incentives to contribute above the minimum number of years effectively required. Therefore, a loosening of the years-of-contribution requirement would inevitably reduce the already low incentives that low income workers have to contribute. These distortions could be ameliorated redesigning the benefit formula such that pensions are always an increasing function of contributions. The pension scheme could provide a subsidy to low income workers to help them reach a desired pension, but the level of this target pension should be an increasing function of the contributions made. The subsidy would in turn be a decreasing function of the contributions. Unlike the current flat minimum pension, this subsidized pension would reward workers who contributed more with a larger pension. In the case of Suiza, the target pension increases with the number of years of contribution (Valdés, 2002). The government of Chile introduced a related solution in the social security reform passed in Parliament in January 2008 (CAPRP 2006, Valdés 2007). In this Chilean case, the minimum pension does not depend on the number of years of contribution, but it depends on the self-financed pension which in turn depends on the accumulated contributions in the savings accounts.

[7] This is an example of the Samaritan's dilemma first analyzed in the economic literature by Buchanan (1975). For a formal presentation of the dilemma see Forteza (1999 and 2001).

7 CONCLUDING REMARKS

Large segments of Latin American population are not covered by the old-age pension programs. Some workers stay out of the system all their lives, never contributing or receiving pensions. Other workers contribute, but many of them do not reach the minimum number of periods of contribution required to qualify for a pension or for a minimum pension guarantee. There is increasing evidence of the existence of a significant number of workers making frequent transitions between formality and informality. Highly fragmented and incomplete histories of contribution to social security risk leaving these workers with no pension rights. This risk has probably been rising recently as the social security administrations have been increasing their ability to enforce the fulfillment of the qualifying conditions.

We propose in this paper a methodology to estimate the proportion of workers who would accumulate the number of periods of contribution required to access a pension, and to assess the impact of different explanatory variables (like the business cycle) on the fulfillment of this condition. The estimation represents a significant challenge because the information that is currently available to estimate the number of periods of contribution of each worker is incomplete. We propose a two-steps methodology to overcome this difficulty. In the first step, we estimate the probabilities of transition between contributing and not contributing at each age, using survival analysis. In the second step, we perform Monte Carlo simulations of the histories of contribution using the probabilities of transition estimated in the first step.

Using this methodology on a database of the main social security institution of Uruguay, the BPS, we found that the majority of contributors to this institution will not reach the required number of years of contribution to access a pension at the normal retirement age. Not at least if the current probabilities of transition between contributing and not contributing remain in the future the same as in the period of observation (1996-2004). The required number of periods of contribution to access a pension in Uruguay is 35 years, which is unusually high for a developing country, but even when we estimate the proportion of contributors who would accumulate 30 years of contribution we get disappointingly low figures.

As expected, the business cycle has a sizeable impact on the probabilities that a private worker transits between contributing and not contributing. This consideration might be particularly relevant in our case, because Uruguay went through one of the most severe recessions of its history during the period in which the work history database was built. Nevertheless, when we repeated the simulations using the average unemployment rate of a longer period (1981-2006) we still got that the majority of private workers would not accumulate the required 35 years of contribution at the normal retirement age (60 years).

Incomplete and highly fragmented histories of contribution are much more pervasive among low than among high income workers. This is hardly surprising, but the size of the differences is really striking. Consider for example the case of men working in the private sector. They have almost a 64% chance of reaching the required 35 years of contribution at 65 years if they belong to the richest quintile, but only 1% chance if they belong to the poorest quintile. The same figures for women are 56% and 4% for the richest and poorest quintiles, respectively.

These results suggest that this pension scheme requires some reform. In particular, the condition of having accumulated 35 years of contribution to access to a pension should be revised.

Several social security systems in Latin America seem to be facing similar challenges as the BPS in Uruguay and have now similar databases as the one used in this paper. The densities of contribution estimated in previous studies for countries like Argentina and Chile are similar to the ones we found for Uruguay. But the densities of contribution alone do not tell us what the probability is that a worker qualifies for a pension when he reaches the retirement age. The methodology proposed in this paper to estimate these probabilities and the factors that impact on them might thus be useful to assess the situation in other countries of Latin America as well.

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TABLES AND FIGURES

Table 1: Number of individuals in the database

Generation	Public Sector			Private Sector			Total		
	Men	Women	Total	Men	Women	Total	Men	Women	Total
1925-1930	109	35	144	339	158	497	448	193	641
1931-1935	291	91	382	801	371	1,172	1,092	462	1,554
1936-1940	534	236	770	1,517	855	2,372	2,051	1,091	3,142
1941-1945	532	435	967	1,830	1,398	3,228	2,362	1,833	4,195
1946-1950	577	611	1,188	2,235	1,786	4,021	2,812	2,397	5,209
1951-1955	607	713	1,320	2,718	2,137	4,855	3,325	2,850	6,175
1956-1960	767	842	1,609	3,066	2,626	5,692	3,833	3,468	7,301
1961-1965	663	766	1,429	3,417	2,862	6,279	4,080	3,628	7,708
1966-1970	414	577	991	3,684	3,146	6,830	4,098	3,723	7,821
1971-1975	305	523	828	4,427	3,951	8,378	4,732	4,474	9,206
1976-1978	146	281	427	3,154	2,642	5,796	3,300	2,923	6,223
1979-1987	110	215	325	5,579	3,918	9,497	5,689	4,133	9,822
Total	5,055	5,325	10,380	32,767	25,850	58,617	37,822	31,175	68,997
Source: Authors' computations using a sample of the work history database of the BPS									

Table 2. Distribution of the population in the database according to the proportion of time in which the individuals contributed. Several groups of workers.

	Less than 50%	Between 50 and 75%	Between 75 and 87.5%	Between 87.5 and 100%	The entire period (100%)	Total	Average
Total	41	14	7	10	28	100.0	59.8
Sex							
Men	40	14	7	10	29	100.0	61.4
Women	43	14	7	9	27	100.0	58.0
Sector							
Public	13	8	5	7	67	100.0	85.4
Private	46	15	8	10	21	100.0	55.3
Income bracket							
Poorest quintile	67	11	5	6	11	100.0	37.9
2nd quintile	52	15	7	8	18	100.0	50.7
3rd quintile	41	17	9	10	23	100.0	59.4
4th quintile	28	15	9	12	36	100.0	71.1
Richest quintile	17	12	8	11	52	100.0	80.4
Note: Each bracket includes the minimum of the interval.							
Source: Authors' computations using a sample of the work history database of the BPS							

Table 3. Distribution of the population in the database according to the percentage of time in which the individuals contributed. Select ages.

Age	Less than 50%	Between 50 and 75%	Between 75 and 87.5%	Between 87.5 and 100%	The entire period (100%)	Total	Average density of contribution
20	72.6	10.8	4.7	3.5	8.4	100.0	30.2
25	54.0	14.9	7.0	6.1	18.0	100.0	47.5
30	40.2	12.5	7.0	7.9	32.4	100.0	59.9
35	33.7	10.1	6.1	6.6	43.5	100.0	66.2
40	30.0	8.6	5.0	6.3	50.1	100.0	70.1
45	27.3	8.5	4.5	6.1	53.6	100.0	72.6
50	25.2	7.5	4.5	6.1	56.7	100.0	74.7
55	23.2	7.9	4.7	6.1	58.1	100.0	76.5
Note: Each bracket includes the minimum of the interval.							
Source: Authors' computations using a sample of the work history database of the BPS							

Table 4. Hazard rates of contributing. Men, private sector.

	Poorest Quintile	Second quintile	Third quintile	Fourth quintile	Richest Quintile
Log of duration	-0.2894427	-0.1733976	-0.2329976	-0.3040770	-0.4348213
	***	***	***	***	***
Age	0.0937066	-0.2099143	-0.1087928	-0.1651087	-0.2539522
	**	***	***	***	***
Age^2	0.0015901	0.0041604	0.0019068	0.0033169	0.0056954
	*	***	*	***	***
Age^3	0.0000095	-0.0000278	-0.0000111	-0.0000225	-0.0000414
		***		*	***
Log of duration * age 30 to 59	0.1175832	-0.1420757	-0.1039667	-0.1082590	-0.0567912
	***	***	***	***	***
Log of duration * age 60+	0.0768316	-0.0595662	-0.0587710	0.0789813	0.2642472
	**	*	*	**	***
Unemployment Rate	0.8109195	2.7337720	3.1980562	5.2060900	7.7134725
		***	***	***	***
Constant	0.3849802	0.1205960	-1.2783098	-0.9505544	-0.5971543
			***	*	
Legend: * p<0.05; ** p<0.01; *** p<0.001					
Note: The table summarizes the results of running a random effects complementary loglog model for the state "contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.					
Source: Authors' computations using a sample of the work history database of the BPS					

Table 5. Hazard rates of contributing. Women, private sector.

	Poorest Quintile	Second quintile	Third quintile	Fourth quintile	Richest quintile
Log of duration	-0.3288185	-0.1423755	-0.2527221	-0.3045152	-0.4635652
	***	***	***	***	***
Age	-0.2867760	-0.1992915	-0.1476158	-0.1825108	-0.1954506
	***	***	***	***	***
Age^2	0.0060549	0.0040793	0.0025086	0.0033899	0.0028882
	***	***	*	**	
Age^3	-0.0000425	-0.0000286	-0.0000140	-0.0000195	-0.0000085
	***	**			
Log of duration * age 30 to 59	-0.0813947	-0.1429265	-0.0985824	-0.1034000	-0.0198979
	***	***	***	***	
Log of duration * age 60+	0.1067427	-0.0652421	-0.0523352	-0.0094198	0.1603543
	**				**
Unemployment Rate	-1.0632649	-0.1813454	2.6109855	4.2608904	6.7140508
			***	***	***
Constant	1.0930710	0.2825117	-0.5671934	-0.5644223	-0.6543332
	***				
Legend: * p<0.05; ** p<0.01; *** p<0.001					
Note: The table summarizes the results of running a random effects complementary loglog model for the state "contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.					
Source: Authors' computations using a sample of the work history database of the BPS					

Table 6. Hazard rates of contributing. Men, public sector.

	First to third quintiles	Fourth quintile	Richest quintile
Log of duration	-0.3727189	-0.5196524	-0.4197703
	***	***	***
Age	0.1439406	-0.1937280	-0.7369508
			***
Age^2	-0.0050416	0.0011224	0.0126937
			***
Age^3	0.0000464	0.0000182	-0.0000535
	*		*
Log of duration * age 30 to 59	-0.1876139	-0.1255769	-0.1251013
	***	*	*
Log of duration * age 60+	0.1048698	0.1198424	0.0690889
Unemployment Rate	-1.9631319	-7.5710245	1.2571905
		***	
Constant	-3.3626559	1.7447998	6.5513771
	*		***
Legend: * p<0.05; ** p<0.01; *** p<0.001			
Note: The table summarizes the results of running a random effects complementary loglog model for the state "contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.			
Source: Authors' computations using a sample of the work history database of the BPS			

Table 7. Hazard rates of contributing. Women, public sector.

	First to third quintiles	Fourth quintile	Richest quintile
Log of duration	-0.4151730	-0.5243146	-0.3705551
	***	***	***
Age	-0.0083652	-0.1476185	-0.6206072
			***
Age^2	-0.0009809	0.0007865	0.0109162
			***
Age^3	0.0000139	0.0000166	-0.0000441
Log of duration * age 30 to 59	-0.1508591	-0.0584505	-0.1852682
	**		***
Log of duration * age 60+	-0.0825813	0.0370954	-0.1605181
			*
Unemployment Rate	-2.7404130	-5.4138448	-4.9070827
		***	**
Constant	-1.3561026	0.7599235	5.7893112
			***
Legend: * p<0.05; ** p<0.01; *** p<0.001			
Note: The table summarizes the results of running a random effects complementary loglog model for the state "contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.			
Source: Authors' computations using a sample of the work history database of the BPS			

Table 8. Hazard rates of not contributing. Men, private sector.

	Poorest Quintile	Second quintile	Third quintile	Fourth quintile	Richest Quintile
Log of duration	-0.1041627	-0.1529350	-0.1929391	-0.1797715	-0.1896528
	***	***	***	***	***
Age	-0.0944476	-0.0353684	-0.0810192	-0.1428829	-0.2010818
	**		*	**	**
Age^2	0.0031665	0.0012634	0.0029490	0.0042937	0.0054311
	***		**	***	**
Age^3	-0.0000331	-0.0000150	-0.0000324	-0.0000434	-0.0000501
	***	*	***	***	**
Log of duration * age 30 to 59	-0.0493634	-0.0736875	-0.1153040	-0.1129735	-0.1591143
	***	***	***	***	***
Log of duration * age 60+	0.0058412	-0.1939569	-0.2200634	-0.4019849	-0.5466212
		***	***	***	**
Unemployment Rate	-4.2159658	-6.9818616	-8.3620510	-8.5833914	-7.3123943
	***	***	***	***	***
Constant	-1.8366212	-1.4785787	-0.6600086	0.2700982	0.8917057
	***	***			
Legend: * p<0.05; ** p<0.01; *** p<0.001					
Note: The table summarizes the results of running a random effects complementary loglog model for the state "not contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.					
Source: Authors' computations using a sample of the work history database of the BPS					

Table 9. Hazard rates of not contributing. Women, private sector.

	Poorest Quintile	Second quintile	Third quintile	Fourth quintile	Richest Quintile
Log of duration	-0.0359524	-0.1297845	-0.1447861	-0.1558657	-0.1745559
	*	***	***	***	***
Age	-0.1981453	-0.1005464	-0.1011123	-0.3057562	-0.0926945
	***	*	*	***	
Age^2	0.0059132	0.0027917	0.0029081	0.0089867	0.0021093
	***	*	*	***	
Age^3	-0.0000556	-0.0000261	-0.0000291	-0.0000864	-0.0000206
	***	**	**	***	
Log of duration * age 30 to 59	-0.0373614	-0.0662170	-0.0887558	-0.0808894	-0.1499035
	*	***	***	***	***
Log of duration * age 60+	0.0221732	-0.0774699	-0.0455249	0.0404465	-0.1036621
Unemployment Rate	-4.5108393	-7.8090439	-9.1639798	-8.3978575	-6.5399603
	***	***	***	***	***
Constant	-1.1470445	-0.9862775	-0.5565680	1.6332923	-0.4557131
	*	*		**	
Legend: * p<0.05; ** p<0.01; *** p<0.001					
Note: The table summarizes the results of running a random effects complementary loglog model for the state "not contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.					
Source: Authors' computations using a sample of the work history database of the BPS					

Table 10. Hazard rates of not contributing. Men, public sector.

	First to third quintiles	Fourth quintile	Richest quintile
Log of duration	-0.1992291	-0.2915872	-0.2252352
	***	***	**
Age	0.1074041	0.2929000	0.1454418
Age^2	-0.0007330	-0.0064759	-0.0045434
Age^3	-0.0091990	0.0000297	0.0000304
Log of duration * age 30 to 59	-0.2233859	-0.0350227	-0.3123191
	***		**
Log of duration * age 60+	-0.4234219	0.0451607	-0.5037006
	*		
Unemployment Rate	-1.1692456	-4.7774109	-4.5383666
Constant	-4.8210151	-5.5054553	-2.8916851
	**	*	
Legend: * p<0.05; ** p<0.01; *** p<0.001			
Note: The table summarizes the results of running a random effects complementary loglog model for the state "not contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.			
Source: Authors' computations using a sample of the work history database of the BPS			

Table 11. Hazard rates of not contributing. Women, public sector.

	First to third quintiles	Fourth quintile	Richest quintile
Log of duration	-0.3168436	-0.2706937	-0.2462103
	***	***	***
Age	0.2215887	0.8067520	1.0567885
		***	***
Age^2	-0.0030227	-0.0191354	-0.0278665
		***	***
Age^3	0.0031970	0.0001387	0.0002164
		***	***
Log of duration * age 30 to 59	-0.1607203	-0.2582388	-0.2046164
	**	***	*
Log of duration * age 60+	-0.2455673	-0.6326308	-1.0714887
		*	
Unemployment Rate	1.5727253	3.9966634	-1.9364036
		*	
Constant	-6.3802454	-13.1087930	-14.0971390
	***	***	***
Legend: * p<0.05; ** p<0.01; *** p<0.001			
Note: The table summarizes the results of running a random effects complementary loglog model for the state "not contributing". The null hypothesis of zero variance of the individual effects was tested using a likelihood ratio test built in the xtcloglog STATA command. The hypothesis was rejected at the standard levels of significance in all the regressions.			
Source: Authors' computations using a sample of the work history database of the BPS			

Table 12. Simulated monthly hazard rates of contributing (%)

	Age in years and Duration in months			
Category	Age = 25 Duration = 1	Age = 25 Duration=36	Age = 40 Duration = 1	Age = 40 Duration=36
Men, private sector, poorest quintile	12.8	4.7	9.5	2.3
Men, private sector, 2nd quintile	7.2	3.9	4.7	1.6
Men, private sector, 3rd quintile	7.5	3.3	5.6	1.7
Men, private sector, 4th quintile	6.8	2.3	4.9	1.1
Men, private sector, richest quintile	4.9	1.0	3.8	0.7
Women, private sector, poorest quintile	9.9	3.2	6.4	1.5
Women, private sector, 2nd quintile	7.0	4.3	4.8	1.8
Women, private sector, 3rd quintile	7.5	3.1	4.9	1.4
Women, private sector, 4th quintile	6.3	2.1	4.3	1.0
Women, private sector, richest quintile	5.0	1.0	3.0	0.5
Men, public sector, 1st – 3rd quintiles	8.2	2.2	5.0	0.7
Men, public sector, 4th quintile	4.3	0.7	1.7	0.2
Men, public sector, richest quintile	1.0	0.2	0.3	0.04
Women, public sector, 1st – 3rd quintiles	9.3	2.2	6.3	0.8
Women, public sector, 4th quintile	5.3	0.8	2.8	0.4
Women, public sector, richest quintile	1.4	0.4	0.6	0.09
Note: These simulations have been done setting the unemployment rate at the April-1996-December-2004 period average, which was 13.45%.				
Source: Authors' computations using a sample of the work history database of the BPS				

Table 13. Simulated monthly hazard rates of *not* contributing (%)

	Age in years and Duration in months			
Category	Age = 25 Duration = 1	Age = 25 Duration=36	Age = 40 Duration = 1	Age = 40 Duration=36
Men, private sector, poorest quintile	3.6	2.5	3.9	2.2
Men, private sector, 2nd quintile	6.2	3.6	6.1	2.7
Men, private sector, 3rd quintile	8.1	4.1	8.9	3.0
Men, private sector, 4th quintile	8.3	4.4	7.8	2.8
Men, private sector, richest quintile	7.8	4.0	6.8	2.0
Women, private sector, poorest quintile	2.0	1.8	2.3	1.7
Women, private sector, 2nd quintile	3.9	2.5	3.8	1.9
Women, private sector, 3rd quintile	5.1	3.1	4.6	2.0
Women, private sector, 4th quintile	5.5	3.2	5.5	2.4
Women, private sector, richest quintile	6.8	3.7	4.9	1.6
Men, public sector, 1st – 3rd quintiles	5.4	2.7	8.3	1.9
Men, public sector, 4th quintile	8.6		5.4	1.7
Men, public sector, richest quintile	10.2	4.7	4.8	0.7
Women, public sector, 1st – 3rd quintiles	8.1	2.7	13.4	2.6
Women, public sector, 4th quintile	10.5	4.1	12.2	1.9
Women, public sector, richest quintile	13.0	5.6	5.8	1.2
Note: These simulations have been done setting the unemployment rate at the April-1996-December-2004 period average, which was 13.45%.				
Source: Authors' computations using a sample of the work history database of the BPS				

Table 14. Percentage of workers who contribute *35* or more years at the ages of 60 and 65 in the simulations. Estimation using the average unemployment rate observed between 1996 and 2004 (13.45%).

	60 years old	65 years old
Total	21.4	28.6
Men	25.3	34.3
Women	20.2	27.4
Men, private sector	11.6	20.0
Poorest quintile	0.1	1.0
Second quintile	3.8	12.4
Third quintile	6.2	16.3
Fourth quintile	19.6	35.1
Richest quintile	52.8	63.7
Women, private sector	9.3	15.8
Poorest quintile	1.0	3.5
Second quintile	1.0	3.3
Third quintile	4.2	11.8
Fourth quintile	15.4	27.8
Richest quintile	45.8	56.4
Men public sector	67.2	74.3
First to third quintile	17.1	35.3
Fourth quintile	63.6	79.2
Richest quintile	92.7	95.1
Women public sector	58.1	72.1
First to third quintile	16.9	39.8
Fourth quintile	50.0	68.9
Richest quintile	79.2	86.0
Source: Authors' computations using a sample of the work history database of the BPS		

Table 15. Percentage of workers who contribute *30* or more years at the ages of 60 and 65 in the simulations. Estimation using the average unemployment rate observed between 1996 and 2004 (13.45%).

	60 years old	65 years old
Total	34.7	41.0
Men	40.4	48.8
Women	33.9	39.9
Men, private sector	26.9	36.4
Poorest quintile	2.0	5.2
Second quintile	19.1	32.6
Third quintile	25.4	39.1
Fourth quintile	48.7	59.8
Richest quintile	73.6	78.8
Women, private sector	20.6	28.5
Poorest quintile	6.3	10.6
Second quintile	5.9	12.9
Third quintile	16.6	26.5
Fourth quintile	38.4	48.9
Richest quintile	66.5	72.0
Men public sector	81.5	83.3
First to third quintile	43.0	55.2
Fourth quintile	84.3	88.4
Richest quintile	96.7	97.0
Women public sector	80.0	84.3
First to third quintile	45.4	60.1
Fourth quintile	79.7	83.6
Richest quintile	92.9	93.3
Source: Authors' computations using a sample of the work history database of the BPS		

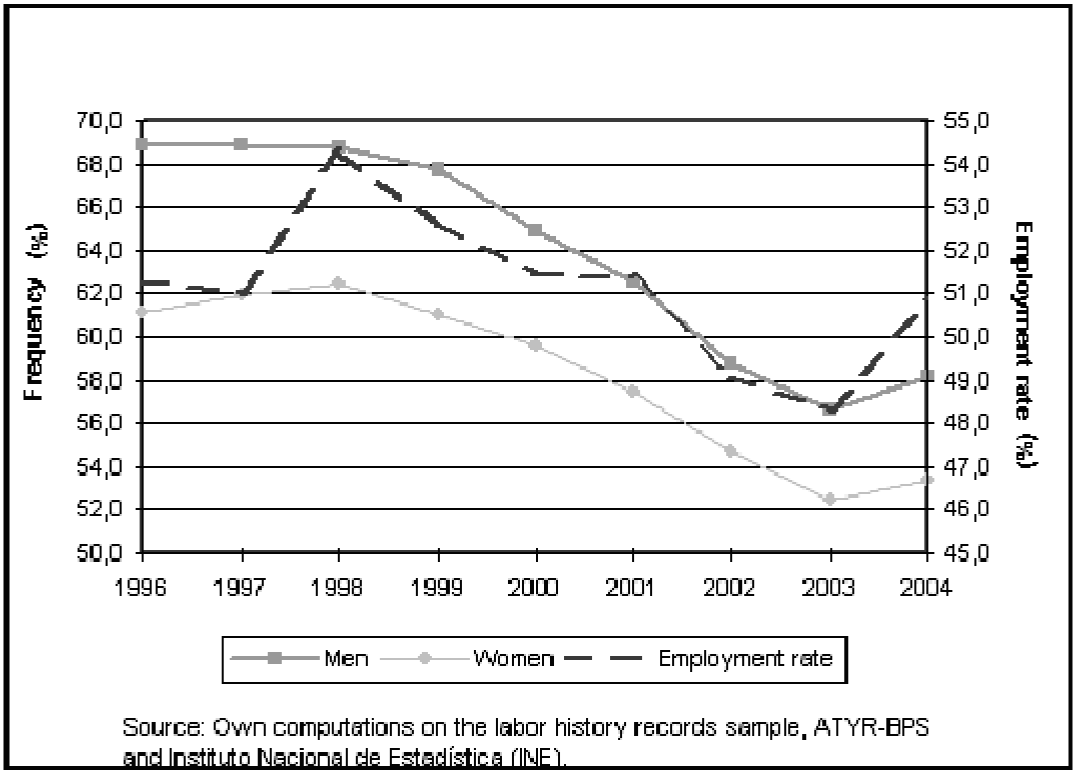
Table 16. Percentage of workers who contribute 35 or more years at the ages of 60 and 65 in the simulations. Estimation using the average unemployment rate observed between 1981 and 2006 (11.20%).

	60 years old	65 years old
Total	25.2	34.7
Men	30.3	41.8
Women	23.1	32.8
Men, private sector	18.0	30.0
Poorest quintile	0.2	1.1
Second quintile	7.1	21.5
Third quintile	11.8	29.3
Fourth quintile	36.3	55.8
Richest quintile	67.9	79.3
Women, private sector	13.8	23.4
Poorest quintile	1.2	4.0
Second quintile	1.0	5.1
Third quintile	8.2	21.6
Fourth quintile	26.6	44.5
Richest quintile	61.0	73.4
Men public sector	68.2	74.6
First to third quintile	16.5	34.9
Fourth quintile	64.8	79.7
Richest quintile	94.2	95.9
Women public sector	55.2	69.4
First to third quintile	14.3	35.6
Fourth quintile	43.3	63.5
Richest quintile	79.2	85.9
Source: Authors' computations using a sample of the work history database of the BPS		

Table 17. Percentage of workers who contribute 30 or more years at the ages of 60 and 65 in the simulations. Estimation using the average unemployment rate observed between 1981 and 2006 (11.20%).

	60 years old	65 years old
Total	41.3	48.9
Men	48.8	58.2
Women	39.3	47.5
Men, private sector	37.8	49.1
Poorest quintile	2.1	5.6
Second quintile	30.4	48.6
Third quintile	41.6	59.1
Fourth quintile	69.2	79.6
Richest quintile	86.3	90.8
Women, private sector	28.5	38.6
Poorest quintile	7.1	11.9
Second quintile	8.7	18.9
Third quintile	29.1	44.1
Fourth quintile	54.9	67.0
Richest quintile	80.1	84.6
Men public sector	82.1	83.8
First to third quintile	42.1	54.9
Fourth quintile	85.7	89.3
Richest quintile	97.3	97.6
Women public sector	77.1	82.1
First to third quintile	40.7	55.8
Fourth quintile	74.0	79.3
Richest quintile	92.8	93.4
Source: Authors' computations using a sample of the work history database of the BPS		

Figure 1. Density of contribution and rate of employment



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Incomplete and highly fragmented work histories threaten to leave many contributors of the pension schemes in Latin America without the minimum pension guarantee or even without access to the ordinary pension. We propose a methodology to assess this risk, identify vulnerable groups and study potential determinants of the history of contributions using information from the work history records of the social security institutions. We apply this methodology to the largest social security institution of Uruguay, the Banco de Previsión Social, and show that the majority of contributors to this institution might not comply with the minimum number of years of contribution that is currently required to access an ordinary pension when they reach the retirement age.

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