Serie Documentos de Trabajo
Superintendencia de Pensiones
Teatinos 317.
Santiago, Chile.

www.spensiones.cl

DOCUMENTO DE TRABAJO N°32

WORK HISTORIES AND PENSION ENTITLEMENTS IN ARGENTINA, CHILE AND URUGUAY

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Septiembre 2009

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September 2009

Abstract: We propose alternative methods to project pension rights and implement them in Chile and Uruguay and partially in Argentina. We use incomplete work histories databases from the social security administrations to project entire lifetime work histories. We first fit linear probability and duration models of the contribution status and dynamic linear models of the income level. We then run Monte Carlo simulations to project work histories and compute pension rights. According to our results, significant swathes of the population would not access to fundamental pension benefits at age 65, if the current eligibility rules were strictly enforced.

JEL: H55, J14, J26

Keyword: density of contributions, pensions, work history

Acknowledgements: We are grateful to David Robalino, Stephen Jenkins, an anonymous referee and participants at the 2008 LACEA meeting and at the regular seminar of the *Instituto de Economía y Finanzas* of the *Universidad de Córdoba*, Argentina, for valuable comments on a previous version. This document is part of a research project financed by the World Bank. The project would have not been possible without the support of David Robalino, who contributed to shape the initial idea and provided valuable comments on earlier versions of this paper. The authors take full responsibility for the content of this document.

Citation: Forteza, A. , I. Apella, E. Fajnzylber, C. Grushka, I. Rossi and G. Sanroman. 2009. "Work Histories and Pension Entitlements in Argentina, Chile and Uruguay." Washington DC: World Bank.

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TABLE OF CONTENTS

List of Tables

List of Figures

I. Introduction

The individual records of contribution are a key component of the reforms that introduced individual savings accounts to the pension schemes in Latin America. Yet, their ability to make the system more transparent and in so doing, to better-protect the working poor in old age is neither widely known nor fully understood.

In Argentina and Uruguay before the reforms, the main social security program administrations had almost no record of individual contributions and could not check whether workers had accumulated the years of contribution legally required to access a pension. Benefits were hence granted on a very informal basis, often appealing to the testimony of witnesses. Not surprisingly, many people managed to receive a contributory pension even when they had not contributed the required number of years. In the Chilean case, even if individual records were available before the 1980 reform, they were distributed among a number of different institutions with heterogeneous contents, formats and integrity.

Gradually systems are being built to populate work history records and increasingly, administrators are gaining the capacity to authenticate individual records of contribution. The outcome has been better enforcement of the conditions required to access a pension. Nevertheless, as the region advances fiscally, a large portion of workers risk not adapting to the new conditions and ending 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 passed in January of 2008 eliminated the condition that workers had to have 20 years of contribution in order to access the minimum guaranteed pension. The Uruguayan parliament has recently passed a law that reduces the required number of years of contribution to access the ordinary pension and the age at which the advanced-age pension is first granted. Colombia is also considering a reform that would reduce the vesting period (or, *Beneficio Economico Periodico*).

In this context, it is important to anticipate the future distribution of pension entitlements both from the social protection perspective and from the fiscal perspective. In the social protection perspective, we want to know to what extent the pension schemes provide effective protection against the risk of poverty in old age. In the fiscal perspective, we want to know what the contingent liabilities of the social security schemes are. In this study we simulate the distribution of pension entitlements in Argentina, Chile and Uruguay using both samples of administrative records of individual contributions to pension schemes and available surveys of socio-economic characteristics. To ensure reliability, alternative estimation methods were implemented and compared, providing a robust perspective of the pensions to be received by workers across Argentina, Chile and Uruguay.

II. Preliminaries: Institutional Organization of Social Security and Aggregate Statistics

2.1 Design of social security

2.1.1 *Argentinean system*

The enactment of Law No. 24241 in 1994 created the *Sistema Integrado de Jubilaciones y Pensiones* (SIJP) (Integrated Pension System). The SIJP includes two regimes: one public, financed on a pay-as-you-go (PAYG) basis and managed by the *Administración Nacional de la Seguridad Social* (ANSES) (National Administration of Social Security), and the other based on individual capitalization managed by private companies (AFJP).[1]

In exchange for managing the workers' mandatory savings, the AFJPs receive a fee proportional to the workers' wage. All workers older than 18 years must contribute to the SIJP, but they have the option of choosing between the traditional PAYG and the individual capitalization regime. In addition, the parameters have gradually been adjusted; for instance, increasing the minimum age of retirement from 55 to 60 years for women and from 60 to 65 years for men, along with contribution requirements increasing from 20 to 30 years.

The newly designed scheme is (partially) financed by an individual contribution equivalent to 11 percent of employee income (or presumed income for the self-employed) and an employer contribution of 16 percent. The public segment is financed by a series of specific resources designated from general tax revenue.

The Argentinian pension system has two pillars. The first pillar consists of a flat benefit granted to all who are insured and meet the required age and years of contribution. This benefit, called the *Prestación Básica Universal* (PBU), is administered by ANSES. The second pillar depends on the exercised option (i.e., the regime chosen). The PAYG regime grants an additional benefit, called the *Prestación Adicional por Permanencia* (PAP), equivalent to 0.015 times the average wage of the last 10 years per every year contributed to the public regime after 1994. The capitalization regime grants an annuity, called the *Jubilación Ordinaria* (JO), based on the accumulated balance. Workers who made contributions before 1994 are entitled to the *Prestación Compensatoria* (PC), which is a

[1] Except for one firm, *Nación AFJP*, managed by the *Banco de la Nación Argentina*.

compensatory benefit computed like the PAP. The PC, like the PBU, is independent of the regime chosen by the affiliate.[2] In 2008, the Argentinean government passed a law that reformed the SIJP. The main change was the elimination of the individual accounts pillar.

2.1.2 Chilean System

The current Chilean pension system can be deconstructed into three main pillars: a poverty prevention pillar, a contributory pillar and a voluntary pillar.

The poverty prevention pillar, before the 2008 reform, was based on two components: a means-tested assistance pension, the *Pensión Asistencial* (PASIS) and the minimum pension guarantee (MPG) for individuals who contributed for at least 20 years to the individual capitalization scheme but were not able to finance a minimum amount for retirement. Together, these two programs corresponded to the main government policies aimed at avoiding old-age poverty, and were financed by general revenue. The reform of 2008 replaced these two programs with the New Solidarity Pillar (NSP), providing (i) a basic pension to the 60 percent poorest with no contribution and (ii) a supplement to the 60 percent poorest that self-finance a small pension. For individuals over a certain age, a provision of the reform ensured the option to choose between the MPG and the NSP.[3]

The contributory pillar, previously based on a dispersed number of PAYG schemes, was drastically reformed in 1980 to create a unique national scheme that is based on individual accounts managed by professionals firms, the *Administradoras de Fondos de Pensiones* (AFPs).[4] Funds cannot be withdrawn until retirement, meaning any point after the legal retirement age (65 years for men and 60 for women) or in early retirement if they have accumulated sufficient funds in their account and they receive a minimum replacement rate. Upon retirement, the individual can choose between buying an annuity from an insurance company and receiving a programmed withdrawal stream from the AFP (or different combinations of these two options). In both cases, benefits are actuarially calculated as a function of the individual's lifetime accumulated savings, the potential beneficiaries and (age- and gender-specific) life expectancy.[5]

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[2] Upon initital contribution, the individual is considered an "affiliate".

[3] See Rofman et al. (2008) for a detailed description of the 2008 Chilean reform.

[4] Only the armed forces, military and police remained in their previous PAYG schemes.

[5] A detailed description of the current AFP system can be found in Berstein (2007), available in the English section of www.spensiones.cl. A number of articles have been written about the impact the 1980 Chilean

To complement the compulsory savings made in the contributory scheme, tax incentives are provided for individuals to make additional voluntary contributions in a wide set of financial products.

2.1.3 *Uruguayan System*

The first old-age, survival and disability (OASDI) insurance programs were set up in the 19th century in Uruguay. These programs expanded their coverage and scope in the following decades, becoming almost universal in the 1950s. In 1967, a public institution called *Banco de Previsión Social* (BPS) was created to administer the social security system. Since then, BPS has administered the largest retirement programs in the country, covering public servants, private workers (with some exceptions), rural workers and domestic workers. Some categories of workers have their own special pension schemes: bank employees, notaries, self-employed university graduates, armed forces personnel and police force personnel. Contributors to the BPS represented 89 percent of the total number of contributors to all social security institutions in the country in 2001 (Ferreira-Coimbra and Forteza 2004).

Before the social security reform initiated in 1996, the BPS retirement programs were financed on a PAYG basis through payroll contributions paid by employers and employees and transfers from the central government. Contribution was also mandatory for self-employed workers, who were subject to minimum declared earnings. The 1996 reform modified the main parameters of the BPS programs and introduced an individual savings accounts pillar administered by private firms.

As a general rule, workers whose earnings are below a threshold are exclusively affiliated to the PAYG pillar. However, they can opt to deposit half of their personal contributions into individual savings accounts. Workers whose earnings are over this threshold must contribute to both pillars. For the amount below the threshold, they contribute to the PAYG public system and from there up to a certain ceiling, also established by law, they must contribute to individual accounts. There is no mandatory contribution for earnings over the established maximum. Employer contributions go exclusively to the PAYG pillar.

Furthermore, the minimum retirement age was fixed at 60 for men and women, which meant an increase of five years for the latter. Also, the minimum number of years of contribution required to access an ordinary pension was raised from 30 to 35. Workers can receive a special “advanced-age”

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pension reform may have had on social security coverage, financial development, national savings and economic performance. For instance, see Corbo and Schmidt-Hebbel 2003.

pension served by the solidarity pillar with only 15 years of contribution, but this benefit is smaller than the ordinary pension and eligibility is restricted to individuals who are 70 years or older. Also, since 2001, workers who are 65 years or older can stop contributing to the savings accounts pillar and receive an annuity, regardless of their count of years of contribution. Workers with hazardous occupations and other special categories have a special bonus added to the count of years of contribution.

The replacement rate was also modified in the 1996 reform, making it more sensitive to both the retirement age and years of contribution, in order to induce longer working-lives. Nowadays the replacement rate is the same for both genders and it ranges from 50 to 82.5 percent, 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 contribution. In addition, there is an extra bonus for low-income workers who choose to contribute to individual savings accounts.

It is worth noting that contributors are also entitled to unemployment insurance, a disability pension, sickness and maternity subsidies, family allowances and health care. In all these programs, benefits are conditioned on compliance with eligibility requirements and as a general rule, the self-employed receive fewer benefits.

This system is financed by employer—employee payroll contributions and general taxes. In the 1996 reform, employee contributions were set at 18 percent of monthly earnings, while employer contributions were set at 17.5 percent of total payroll. Over the years, the government introduced exemptions to employer contributions and in 2007, in the context of a tax reform, some of the exemptions were lifted and the general employer contribution rate for social security was reduced to 12.5 percent. Self-employed contributions are based on a minimum declared monthly income.

Not contributing to the social security system facilitates evasion of other labor costs. The most important are the compulsory insurance to cover work injury risk, a progressive tax on wages that until 2006 ranged from 0 to 6 percent, and the income tax introduced in 2007.

Workers who do not contribute may access means-tested benefits: an assistance pension for persons older than 70 years old (which is less generous than the contributory pension), family allowances and free health care in the public system.

It is difficult to assess the costs and benefits of not contributing. The most widely-quoted disincentive for contributing to the BPS has been the lack of an effective control of the number of years of contribution. Insomuch as the administrative records do not exist or are very incomplete, the BPS has had to accept the testimony of witnesses as proof of contributions. Providing incentives for workers to contribute and strengthening the link between contributions and pensions, it should be noted, was one of the aims of the 1996 reform.

2.2 Aggregate statistics

The size of the pension system measured by the expenditure in pensions as a percentage of GDP varies widely in these three countries, ranging from 3 percent in Chile to almost 11 percent in Uruguay. Coverage of the elderly is comparatively large in the three countries, particularly so in the case of Uruguay. Coverage of the labor force is in the order of 60 percent in Chile and Uruguay, and 40 percent in Argentina. The average pension represents roughly half of per capita GDP in Argentina, 44 percent in Uruguay and about one third of per capita GDP in Chile.

Table 1: Aggregate Pension Statistics in Argentina, Chile and Uruguay (2006)

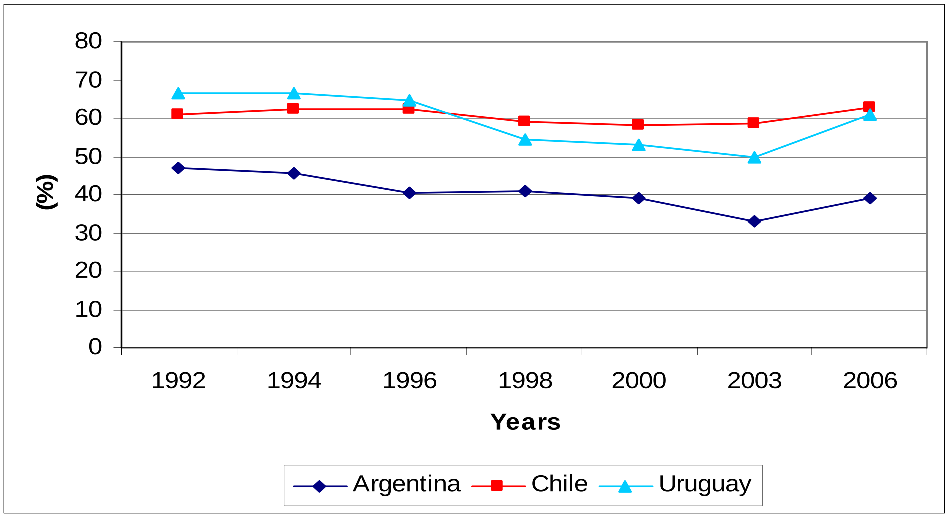
	Argentina	*Chile*	*Uruguay*
Spending on Pensions as % of GDP [A]	6.4	3.1	10.8
Average Pension (% of per Capita GDP)	52.8	33.8	44.0
Coverage of Elderly (as % of Population Aged 65+) [B]	70.5	75.5	85.6
Coverage of Labor Force (%)	39.2	62.7	60.9

Note: [A] Contributory old-age pensions: Uruguay includes old-age, survival and disability; Chilean calculations include contributory pensions (old-age, disability and survivorship) paid from both the PAYG and the AFP systems; [B] Argentinean coverage may have increased after the Pension Inclusion Program implemented by the government in 2007.

Source: Authors' computations based on (i) Instituto Nacional de Estadística, Uruguay; (ii) Rofman, Lucchetti and Ourens (2008); (iii) Ministerio de Economía y Producción de la Argentina; and, (iv) statistics for Chile from www.spensiones.cl, Chilean Central Bank and www.inp.cl.

Since the early nineties, no noticeable improvement in the coverage of the labor force has occurred in any of the three countries (Figure 1). Furthermore, coverage experienced a decline in Argentina and Uruguay between 1992 and 2002-3 and only in recent years has there been some recovery, partially explained by the business cycle. In Chile, coverage of the labor force has remained virtually stagnant.

Figure 1: Coverage of the Labor Force in Argentina, Chile and Uruguay



Source: Authors' computations based on Rofman, Lucchetti and Ourens (2008).

Many workers have highly incomplete histories of contribution, which negatively impact on their pension rights. Analysts have been looking at the density of contribution – the proportion of the potential months of contribution in which a contribution is effectively recognized – to assess the risk that workers run of not accessing pensions or the risk of receiving an inadequately low pension even if eligible. The average density of contribution was 55 percent in Argentina (1994-2001), 51 percent in Chile (1981-2004) and 58 percent in Uruguay (1996-2004).[6] The gender gap is more pronounced in Chile than in the other two cases (Table 2, Table 5 and Table 8). We provide more detailed information and analysis of these statistics by country in the following section.

[6] For these computations, we considered individuals aged between 18 and 70 years. It is important to mention that some of these individuals could be, during some periods, out of the labor force. It is difficult to determine, from administrative data, the work status (employed in the informal sector, unenmployed or inactive) during periods when people are not contributing.

III. Work History Records

3.1 Argentinean data

Our sample comes from the Labor Histories database that was developed by the National Direction of Social Security Policies (DNPSS) of the Ministry of Labor and Social Security. In the past, the database was used only partially for labor and social security research (Bertranou and Sánchez 2003; De Biase and Grushka 2003; DNPSS 2003).

The SIJP information is based on (i) the sworn statements (*Declaración Jurada*) of employers that are required before depositing the social security contributions of employees and (ii) payments made by the self-employed (independent and *monotributistas* workers).[7]

As in other countries, the work history records in Argentina were designed taking into account only the immediate operational needs of the social security administration. Because of this, the records are relatively poor in terms of socio-demographic information (DNPSS 2003). Nevertheless, the database provides some basic information about all those who contributed to the SIJP.

The original database includes sworn statements covering a 90-month period between July 1994 and December 2001, following more than 8 million workers (or affiliates) who made at least one contribution to the SIJP since 1994. Of this population, the Social Security Secretariat randomly selected a sample of 10,000 registered workers who were still alive and active in December 2001.[8] Pension rights, it is important to note, are based on employer statements, not actual contributions.

To organize the data, the following matrix was established: a Unique Code of Tax or Labor Identification (CUIT/L) associated the record to a worker; one row was designated to one worker, equaling 10,000 rows; and, one column represented one month, thus 90 columns. To designate a contribution made within a given month, a "1" populated the respective cell; if no contribution was made, a "0" was entered. Hence, the sum across a row determines the worker's accumulated contributions in the window of observation. To this was added a set of variables from the worker registry: age, type of worker (classified into three categories: employee, self-employed and mixed),

[7] The *monotributo* is a simplified regime. The contributions in this regime go entirely to the public program.

[8] Unfortunately, this sampling procedure may introduce some biases linked to attrition: workers who are not captured in the sample because of death or inactivity are likely to differ from the captured workers.

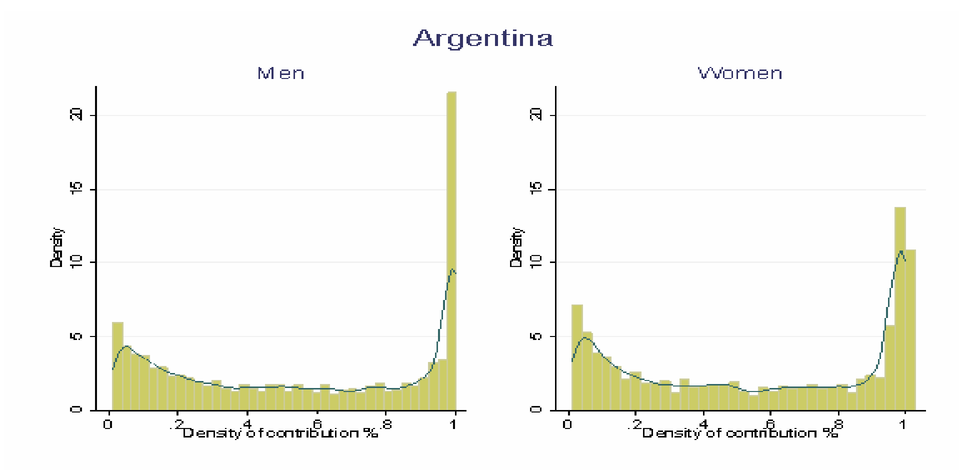
taxable income, pensions regime (PAYG or capitalization) and jurisdiction of residence. This data is available only in December 2001, so the database is cross-section regarding these variables.

The density of contribution to the social security system has a bimodal distribution in the population, with the two modes close to zero and one (Figure 2). This pattern has been reported in previous studies (among others, De Biase and Grushka 2003). As expected, young individuals tend to have comparatively low densities of contribution (Table 2).

The density of contribution seems to be positively correlated with the income level (Table 2). In order to avoid the circular reasoning of finding low densities for workers whose low average income is due to a few periods of contribution, we calculated the average earnings over periods in which individuals reported strictly positive earnings. Then, we grouped workers in clusters of five years and by sex, and calculated the quintiles of the earnings distribution for each group. We found consistently higher densities of contribution in higher income quintiles.

There are no obvious differences between the two pillars of the Argentinean mixed system, as the affiliated to the individual accounts pillar have basically the same density of contributions as the affiliated to the PAYG pillar. Self-employed show a smaller density of contribution as other workers.

Figure 2: Argentina—Distribution of the Density of Contributions (July 1994-December 2001)



Source: Authors' Computations.

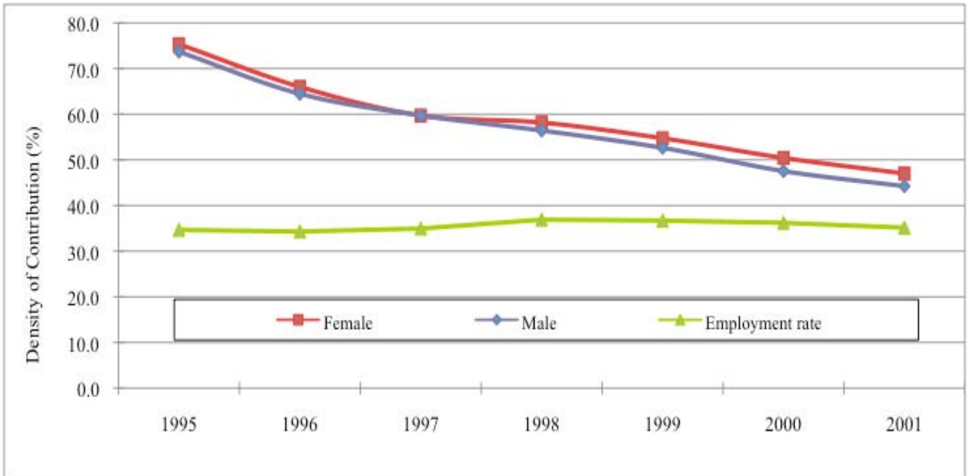
Table 2: Argentina—Contribution Densities of SIJP Contributors (July 1994-December 2001)

Characteristics	*Mean*	*Median*	*Percentage of Contributors with Densities (d) …*				
			d<25%	*25% ≤ d < 50%*	*50% ≤ d < 75%*	*75% ≤ d < 100%*	*d=100 %*
Total	55.0	56.7	30.7	15.2	13.8	28.5	11.8
Sex							
Men	55.0	55.6	30.5	15.5	14.4	27.6	12.0
Women	56.9	61.2	29.7	14.7	12.9	31.4	11.3
Regime							
PAYG	57.2	60.0	28.1	16.8	13.6	31.2	10.4
Capitalization	54.8	55.6	31.2	15.0	14.0	28.1	11.8
Type of Worker							
Employed	68.3	78.9	14.7	15.4	17.0	38.0	14.9
Self-Employed	61.5	65.6	19.7	18.6	18.6	33.5	9.6
Income Bracket							
Poorest Quintile	44.1	36.9	39.5	19.9	15.7	17.5	7.5
2nd Quintile	51.4	48.0	32.6	18.0	15.7	24.1	9.7
3rd Quintile	54.5	54.5	30.9	15.4	14.8	28.4	10.5
4th Quintile	58.6	65.6	27.9	13.7	12.9	33.3	12.2
Richest Quintile	67.7	88.9	22.3	8.8	10.2	39.5	19.3
Age							
20	34.7	20.0	41.7	16.7	4.2	8.3	29.2
35	69.2	83.3	15.8	10.5	15.8	47.4	10.5
50	68.0	85.0	21.4	7.1	7.1	42.9	21.4

Note: The sample window is 90 months. Each income bracket includes the minimum of the interval.
Source: Authors' computations.

There are no clear signs that the business cycle has impacted on the densities of contribution in Argentina (Figure 3). Rather, the data shows a steady decline of the density across the period for which we have data, 1994-2001.

Figure 3: Argentina—Density of SIJP Contributions and Rate of Employment (July 1994-December 2001)



Source: Authors' Computations.

Table 3 and Table 4 summarize statistics on the duration of the spells of contributing and not contributing on to the Argentinean SIJP. The spells last on average only 12 months. There are no large differences between sexes, regime (PAYG/capitalization), or type of worker (employed, self-employed). In turn, the duration of the spells of contribution (lack of contribution) consistently increase (decrease) with income level. The distribution of the spells duration is strongly skewed to the right: the median is much smaller than the mean duration. For example, the median duration of the spells of contributing in the whole population was only six months and the median duration of the spells of not contributing was only seven months. About 40 percent of the spells of both contributing and not contributing lasted less than six months. In summary, most contributors to the Argentinean SIJP show frequent interruptions in their contribution status.

Table 3: Argentina—Duration of Spells of Contribution in Sample

Characteristics	Mean	Median	*Percentage of Spells in Sample with Durations...*				
			< 6 Months	*6-12 Months*	*12-24 Months*	*24-36 Months*	*>36 Months*
Total	11.8	6	43.8	23.6	16.4	6.8	9.5
Sex							
Men	11.6	6	44.9	23.3	16.0	6.5	9.5
Women	12.3	7	40.8	24.4	17.5	7.8	9.5
Regime							
PAYG	11.3	6	45.8	23.3	15.9	6.3	8.8
Capitalization	13.8	8	35.1	24.9	18.7	9.1	12.3
Type of Worker							
Employed	13.1	7	39.0	25.0	17.1	7.3	11.6
Self-Employed	11.9	7	42.5	21.9	19.1	8.2	8.3
Income Bracket							
Poorest Quintile	8.4	5	55.2	22.6	13.5	5.0	3.8
2nd Quintile	10.9	6	45.0	23.6	18.0	7.0	6.4
3rd Quintile	13.0	8	37.4	24.6	19.6	8.3	10.2
4th Quintile	14.9	8	32.5	26.0	18.4	8.8	14.4
Richest Quintile	18.8	12	22.8	26.2	18.7	9.1	23.2

Note: The sample window is 90 months. Only spells within the observation window were computed. Each income bracket includes the minimum of the interval.

Source: Authors' computations.

Table 4: Argentina—Duration of Spells of No Contribution in Sample

Characteristics	Mean	Median	*Percentage of Spells in Sample with Durations...*				
			< 6 Months	*6-12 Months*	*12-24 Months*	*24-36 Months*	*>36 Months*
Total	12.4	7	39.5	22.5	20.6	9.4	8.1
Sex							
Men	12.2	7	40.0	22.7	20.7	9.2	7.6
Women	13.1	8	38.7	21.8	20.2	10.0	9.3
Regime							
PAYG	12.3	7	39.2	22.7	20.9	9.4	7.8
Capitalization	13.1	7	41.1	21.1	18.9	9.2	9.8
Type of Worker							
Employed	9.8	6	46.6	22.7	19.4	7.1	4.1
Self-Employed	9.9	6	45.2	23.2	20.1	7.8	3.7
Income Bracket							
Poorest Quintile	11.3	7	39.5	24.0	22.5	8.7	5.4
2nd Quintile	10.2	6	43.3	24.4	20.4	8.2	3.8
3rd Quintile	10.1	6	45.0	23.1	19.8	8.0	4.1
4th Quintile	8.9	5	51.1	22.0	17.8	5.8	3.4
Richest Quintile	7.2	4	60.5	18.8	14.3	4.1	2.4

Note: The sample window is 90 months. Only spells within the observation window were computed. Each income bracket includes the minimum of the interval.

Source: Authors' computations.

3.2 Chilean data

We have access to the *Base de Historias Previsionales de Afiliados Activos, Pensionados y Fallecidos* (Affiliates Pension Histories database, HPA), populated with individual contribution records, for a representative sample of participants in the pension system from 1981 to 2006.[9] The HPA includes the complete contribution history (in the AFP system) for a sample of approximately 24,000 individuals, representative of the stock of affiliates of the system in July 2002.[10] In addition, the dataset also includes information on the recognition bonds held by the sampled individuals.[11]

Similar to the Argentinean case, the density has a bimodal distribution among Chilean women, with the two modes close to zero and one (Figure 4). In contrast, a monotonically increasing distribution is found among Chilean men, with a large mode at 100 percent. As Table 5 shows, the average Chilean density is 51.4 percent, but with significant differences between men (57.9 percent) and women (43 percent).[12] The average density increases with age (from 29.9 percent at age 20 up to 63.6 percent at age 50) and with position in the income distribution (from 34.9 percent in the poorest quintile up to 64.9 percent in the richest one). [13] Low densities of contribution are usually associated with inactivity, self-employment and informal work. This is particularly strong in the first quintile of the income distribution, where almost half of the workers present densities of contributions below 25 percent.

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[9] The sample was originally drawn as the basis for the Social Protection Survey, a panel instrument that was taken in 2002, 2004 and 2006 for a large fraction of the individuals in the sample. The Social Protection Survey was inspired by the American Health and Retirement Survey. For more information on the Chilean version, see www.proteccionsocial.cl.

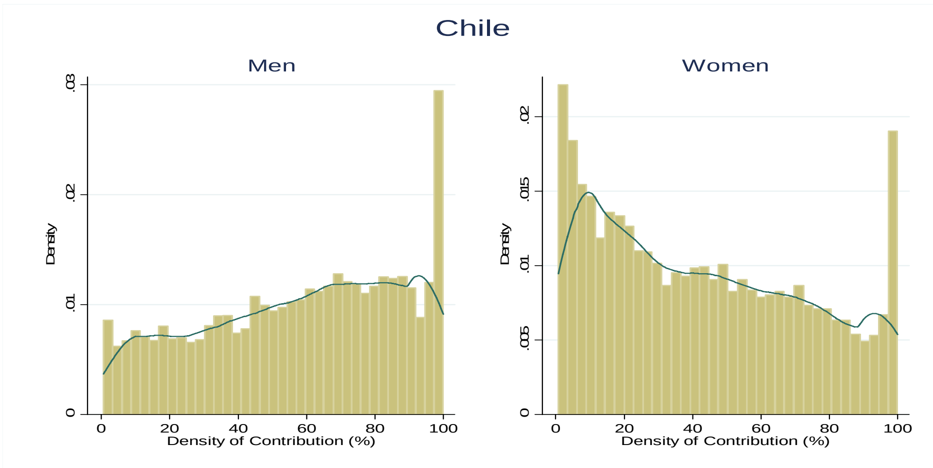
[10] Upon initial contribution, the individual is considered an “affiliate”.

[11] The recognition bonds are obligations assumed by the State to those individuals who contributed to the earlier PAYG system but then switched to the AFP scheme. They are calculated as a function of the contributions made to their previous regime (number, timing and covered wages) and they earn a 4 per cent annual real interest rate until the person reaches the minimum retirement age (60 for women, 65 for men). At that point the State deposits the value of the bond in the individual account.

[12] This is consistent with the level of labor force participation among Chilean women, considered relatively low for the region, given the economic development level of the country.

[13] In the cases of Chile and Uruguay, we used the estimated individual effects of labor income equations, rather than labor income itself, to compute income quintiles (see equation (1)). These effects are meant to capture some permanent characteristics of individuals, mainly education and skills. The quintiles were then constructed within groups defined by cohort and gender. Cohorts in the Uruguayan case were constructed by 5-years intervals according to the year of birth. In the Chilean case, they correspond to the year of birth.

Figure 4: Chile—Distribution of the Density of Contribution (1981- 2004)



Source: Authors' Computations.

Table 5: Chile—Contribution Densities of Contributors in Dataset Several Groups of Workers

Characteristics	*Mean*	*Median*	*Percentage of Contributors with Densities (d) …*				
			d<25%	*25% ≤ d < 50%*	*50% ≤ d < 75%*	*75%≤ d < 100%*	*d=100%*
Total	51.4	52.2	25.5	22.4	24.9	26.0	1.3
Sex							
Men	57.9	61.3	17.4	21.0	27.9	32.1	1.6
Women	43.0	39.2	36.0	24.2	20.9	18.0	0.9
Income Bracket							
Poorest Quintile	34.9	27.4	47.1	25.0	15.6	11.9	0.5
2nd Quintile	46.3	44.6	28.8	26.9	24.8	18.7	0.8
3rd Quintile	52.2	52.8	21.5	25.1	28.4	24.0	1.0
4th Quintile	59.8	63.9	15.9	19.5	28.7	34.1	1.8
Richest Quintile	64.9	70.6	13.0	15.0	27.2	42.3	2.5
Age							
20	29.9	0.0	61.9	9.8	7.5	5.2	15.6
35	55.3	72.7	38.1	6.4	7.1	4.3	44.1
50	63.6	100.0	30.9	4.5	5.8	7.0	51.9

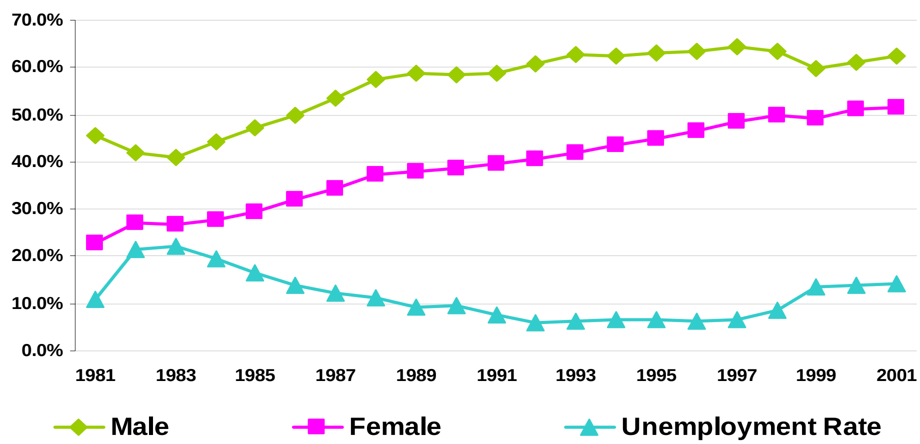
Note: The sample window is 288 months. Each income bracket includes the minimum of the interval. Density was measured from the 18 years of age, except for the individuals who switched to the new system, in which case, density was measured from the moment they switched.

Source: Authors' computations.

Yearly measures of average contribution densities represent a precise idea of the degree of formal employment in the economy. In the Chilean case, the average contribution density of men presents a clear reflection of the unemployment rate in the economy (Figure 5); contribution density decreases in times of high unemployment and increases as the unemployment rate recovers. Female contribution density, on the contrary, tends to exhibit a monotonically increasing path during the

1981-2001 period, coinciding with the slow increase of labor force participation among Chilean female workers.

Figure 5: Chile—Density of Contribution and Rate of Unemployment



Source: Authors' Computations.

In the Chilean sample, the spells of contribution last on average 28.3 months while the spells of no contribution last on average 22.7 months (Table 6 and Table 7).[14] The average duration of the spells of contribution is longer in the Chilean than in the Argentinean and Uruguayan databases. We do not want to push the comparative perspective at this stage though, as the window of observation is longer in the Chilean database, so longer durations are to some extent expected.

Nevertheless, the duration of spells of contributions (and no contributions) has notable variety. Spells of contribution among men and women are very similar in duration, yet women show longer spells of no contribution. This is, once again, probably associated with Chilean women withdrawing from the formal labor force to care for children. It is important to note that the distribution of contribution spells is highly skewed to the right, with the median being only seven months. This large difference can only be explained by a small number of long spells (21 percent exceed 36 months). Furthermore, the average duration consistently increases with income level (ranging from 16.7 months in the first quintile to 54.1 months in the upper quintile). Unlike in Argentina, and rather surprisingly, the average and median spells of no contribution remain relatively constant between the second and the richest quintile. About 43 percent of the spells of contribution and 37 percent of the spells of no

[14] To construct these measures, care was given to avoid false interruptions in contribution spells, which usually originated in delayed payment on the part of employers. Contribution gaps smaller than two months were in most cases corrected in the data.

contribution lasted less than six months. This brief description of the Chilean dataset shows that, like in Argentina, the histories of contribution to the Chilean AFPs show very frequent interruptions.

Table 6: Chile—Duration of Spells of Contribution in Sample

Characteristics	*Mean*	*Median*	*Percentage of Spells in Sample with Durations...*				
			< 6 Months	*6-12 Months*	*12-24 Months*	*24-36 Months*	*>36 Months*
Total	28.3	7.0	42.6	17.3	12.8	6.3	21.0
Sex							
Men	28.4	8.0	42.0	17.4	13.4	6.4	20.8
Women	28.0	7.0	43.7	17.2	11.7	6.1	21.3
Income Bracket							
Poorest Quintile	16.7	5.0	54.0	18.3	11.4	4.8	11.4
2nd Quintile	21.6	7.0	43.3	19.3	14.0	6.8	16.6
3rd Quintile	25.7	8.0	40.7	18.3	14.0	7.0	20.0
4th Quintile	35.7	10.0	36.2	16.0	13.3	6.9	27.6
Richest Quintile	54.1	15.0	33.5	12.2	10.4	6.2	37.6

Note: The sample window is 288 months. Only spells within the observation window were computed. Each income bracket includes the minimum of the interval.

Source: Authors' computations.

Table 7: Chile—Duration of Spells of No Contribution in Sample

Characteristics	*Mean*	*Median*	*Percentage of Spells in Sample with Durations...*				
			<6 Months	*6-12 Months*	*12-24 Months*	*24-36 Months*	*>36 Months*
Total	22.7	8.0	36.8	23.3	14.3	7.4	18.1
Sex							
Men	17.8	7.0	40.8	24.2	14.6	7.1	13.3
Women	30.6	11.0	30.4	22.0	13.9	7.9	25.9
Income Bracket							
Poorest Quintile	27.1	10.0	30.6	24.2	15.7	8.0	21.5
2nd Quintile	22.5	8.0	36.5	24.0	14.0	7.8	17.8
3rd Quintile	20.7	8.0	38.4	24.1	14.4	7.0	16.1
4th Quintile	20.4	8.0	40.2	22.3	14.2	7.2	16.1
Richest Quintile	21.7	8.0	41.0	20.9	12.9	6.6	18.7

Note: The sample window is 288 months. Only spells within the observation window were computed. Each income bracket includes the minimum of the interval.

Source: Authors' computation.

3.3 Uruguayan data[15]

We used a random sample of the work history records collected in December 2004 by the *Unidad de Historia Laboral* (Labor History Unit) of the *Banco de Previsión Social* (ATYR-BPS). Workers

[15] This section draws heavily on Bucheli, Forteza and Rossi (2007).

in the sample contributed at least one month between April 1996 and December 2004. The sample has 68,997 individuals.

The records are organized in five databases. The first database provides personal information: date of birth, sex and country of birth. Job profiles comprise the second, particularly the date of initiation of activity and the explicit end of the link between the worker and the firm. A third reports monthly information about the contributions, wages and some characteristics of the job. A separate database contains information about benefits, including the date of retirement. Finally, the fifth database contains information about contributions sent to administrators of pension funds (AFAP).

Like the Chilean records of work histories, these databases provide detailed information about monthly contributions to social security, gender, age, and sector of activity. Unfortunately, we do not yet have a survey of socio-economic characteristics of contributors to social security for Uruguay. Hence, we lack some important socio-economic characteristics like education and characteristics of the families. However, it is worth noticing that the availability of a longitudinal sample with many time periods allows us to estimate the effect of time invariant characteristics in a parsimonious way.

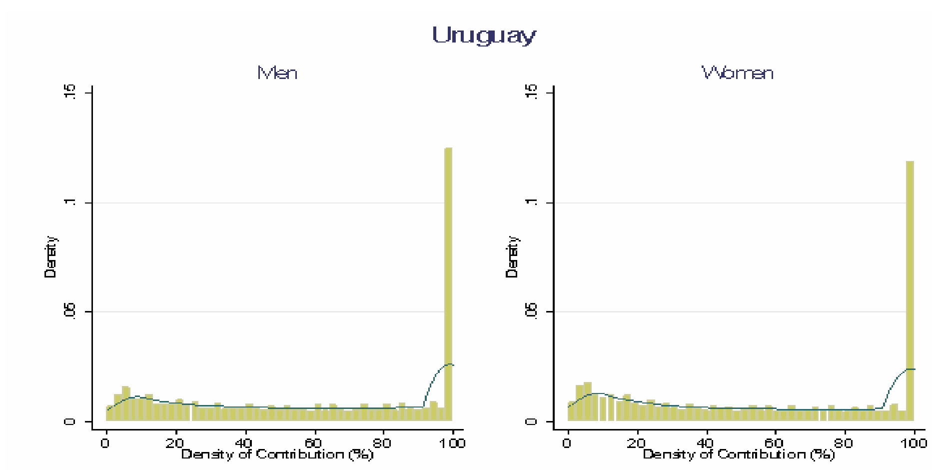
The average density of contributions is about 60 percent (Table 8). As in Argentina, the distribution of the density of contributions has two modes and is strongly asymmetric (Figure 6), characteristics also pointed out by Lagomarsino and Lanzilotta (2004). A similar pattern has been reported before in Argentina (Farall et al. 2003; Bertranou and Sánchez 2003) and Chile (Bravo et al. 2006; Berstein et al. 2006). In our sample, 26 percent of the workers have full contribution density. This was the most frequent density of contributions in the database. Over 40 percent do not register contributions for at least half of the potential months of contribution. Men present slightly higher densities of contribution than women (59.6 and 57.0 percent respectively).

Table 8: Uruguay—BPS Contribution Densities, Several Groups of Workers

Characteristics	*Mean*	*Median*	*Percentage of Contributors with Densities (d) …*				
			d<25%	*25% ≤ d < 50%*	*50% ≤ d < 75%*	*75% ≤ d < 100%*	*d=100%*
Total	58.4	61.0	26.6	16.6	15.0	16.1	25.6
Sex							
Men	59.6	63.0	24.9	16.9	15.4	16.9	25.9
Women	57.0	58.1	28.6	16.3	14.6	15.2	25.3
Sector							
Public	79.7	100.0	11.8	8.0	9.8	11.5	59.0
Private	54.6	53.3	29.2	18.2	16.0	17.0	19.6
Income Bracket							
Poorest Quintile	43.5	32.4	44.4	17.1	11.9	10.8	15.8
2nd Quintile	56.1	55.2	27.3	19.1	16.0	15.8	21.8
3rd Quintile	60.3	62.9	21.8	19.2	17.4	17.2	24.4
4th Quintile	63.9	70.5	19.2	16.6	17.4	20.1	26.8
Richest Quintile	68.4	85.7	20.1	11.2	12.4	16.8	39.5
Age							
20	37.8	16.7	51.8	10.4	7.9	6.6	23.3
35	66.3	100.0	29.1	4.4	3.8	3.9	58.9
50	71.6	100.0	24.7	3.2	3.2	3.1	65.8

Note: The sample window is 105 months. Each income bracket includes the minimum of the interval.
Source: Authors' computations.

Figure 6: Uruguay—Distribution of the Density of Contribution (April 1996-December 2004)



Source: Authors' Computations.

As expected, public sector workers have significantly higher densities of contribution than private sector workers.[16] While public workers contributed about 80 percent of the time, private sector workers contributed about 55 percent of the time. Contributions spanning the entire period ranged from 60 percent for public employees to less than 20 percent for private employees. A considerable number of individuals classified as public employees, nevertheless, present low densities of contribution. This is partly because workers classified as public employees may actually spend as much as half of their working life as private employees. Also several groups of public workers have special regimes by which they compute more than one year of contribution per year of effective contribution (teachers, workers handling radioactive products, etc.), so they need less than 35 years of contribution to access to pensions (i.e., they have shorter working careers).

Additionally, we grouped the individuals in the sample in quintiles of the earnings distribution. The average density of contribution consistently rises with the quintile of these distributions: it is almost 44 percent for the poorest quintile and more than 68 percent for the richest quintile.

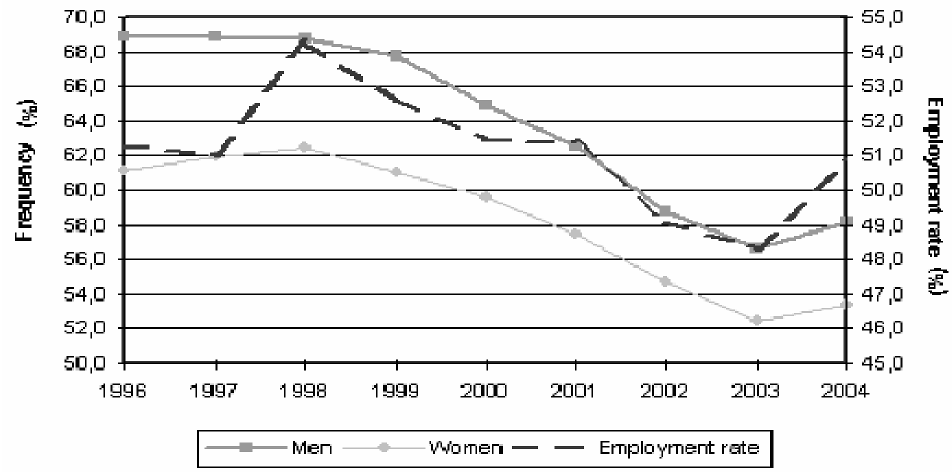
There are also significant differences according to age. At 20 the density of contribution is approximately 49 percent on average and it continuously increases with age, exceeding 71 percent when workers are in their 50s. However, there is an important dispersion among individuals of each age.

The business cycle seems to have had a significant impact on the density of contribution. Uruguay entered into a recession in 1999, followed by the most severe economic crisis in its history. A recovery that began slowly in 2003 accelerated to significant growth by 2004. The unemployment and employment rates observed between 1996 and 2004 capture this period of significant macroeconomic volatility: The density of contribution mirrored the rate of employment (

Figure 7).

[16] We consider a public worker to be anyone who worked in the public sector at least half of the total time during which contributions were made. According to this criterion, we identify 58,617 (85 percent) as private and 10,380 (15 percent) as public sector workers in the database.

Figure 7: Uruguay—Density of Contribution to BPS and Rate of Employment



Source: Bucheli et al. (2008).

Table 9 and Table 10 summarize the information about the duration of the spells of contribution and no contribution in the Uruguayan sample. On average, the spells of contribution last 33 months and the spells of no contribution last 20 months. Notice that 105 is the maximum duration that can be observed in this sample, being the length of the observation window (April 1996 to December 2004). Women have longer spells of contribution on average than men, but they also have longer spells of no contribution. Therefore, men seem to have higher turnover in Uruguay. Public sector workers have much longer spells of contribution and shorter spells of no contribution than private sector workers.

Like in Argentina and Chile, the distribution of duration is skewed to the right in the Uruguayan BPS. For contribution, the median duration is only 13 months (while the mean is 33) and for no contribution, it is 10 (while the mean is 20). About one third of all spells of contribution and no contribution lasted less than six months.

Table 9: Uruguay—Duration of Spells of Contribution in BPS Sample

Characteristics	*Mean*	*Median*	*Percentage of Spells in Sample with Durations...*				
			<6 Months	*6-12 Months*	*12-24 Months*	*24-36 Months*	*>36 Months*
Total	33.2	13	31.5	15.3	13.0	6.8	32.8
Sex							
Men	31.4	12	33.5	15.7	13.0	6.5	30.7
Women	35.7	16	28.6	14.7	13.0	7.2	35.8
Sector							
Public	67.8	96	11.7	7.6	7.0	3.4	67.9
Private	29.1	11	33.8	16.2	13.7	7.2	28.7
Income Bracket							
Poorest Quintile	24.4	8	42.6	15.8	12.1	5.8	23.2
2nd Quintile	30.4	13	30.7	17.0	14.8	7.8	29.2
3rd Quintile	31.8	14	28.9	17.1	14.9	7.4	31.2
4th Quintile	35.3	17	27.2	15.3	13.7	7.5	35.6
Richest Quintile	46.8	30	27.4	10.1	8.6	4.9	47.9

Note: The sample window is 105 months. Only spells within the observation window were computed. Each income bracket includes the minimum of the interval.

Source: Authors' computations.

Table 10: Uruguay—Duration of Spells of No Contribution in BPS Sample

Characteristics	*Mean*	*Median*	*Percentage of Spells in Sample with Durations...*				
			<6 Months	*6-12 Months*	*12-24 Months*	*24-36 Months*	*>36 Months*
Total	20.2	10	33.7	21.4	15.7	9.2	20.3
Sex							
Men	18.4	9	36.4	21.7	15.9	8.8	17.6
Women	22.9	11	29.8	20.9	15.5	9.8	24.4
Sector							
Public	16.5	7	44.9	20.5	12.8	8.0	15.1
Private	20.4	10	33.1	21.5	15.9	9.2	20.6
Income Bracket							
Poorest Quintile	25.0	13	29.1	18.2	15.7	9.9	27.5
2nd Quintile	20.1	10	34.2	20.0	16.5	9.3	20.2
3rd Quintile	18.1	9	36.1	21.6	16.1	9.2	17.3
4th Quintile	17.5	8	36.8	23.2	15.3	8.7	16.4
Richest Quintile	19.5	9	32.5	26.2	14.4	8.5	19.2

Note: The sample window is 105 months. Only spells within the observation window were computed. Each income bracket includes the minimum of the interval.

Source: Authors' computations.

IV. Methodology and Results

In order to predict the distribution of pension entitlements, we proceed by forecasting both labor income and contribution status along the life cycle of individuals. Projections are based on the estimation of econometric models that benefit from the longitudinal structure of the different datasets.

4.1 Projections of the labor income [17]

We estimate two different equations in order to predict each individual future wage stream. Wages in the second and following months of a spell of contribution are modeled using a dynamic equation. Wages in the first month of a contribution spell are modeled with a static equation. We chose the most parsimonious specifications in all cases.

Our goal is to simulate the future stream of labor income for these workers, so we are particularly interested in the impact on labor income of age, time trends and other observable covariates which are time invariant or evolve in a deterministic way. Typical covariates are education and cohort. Data availability will determine the exact set of controls we will be able to include in regressions for each country. Because of lack of data, we could not estimate labor income equations for Argentina.

4.1.1 *Projection of labor income for Chile*

The results of estimating the labor income equations with the Chilean data are presented in Table 11. Panel A presents the results for income in the second and following months of the contribution spells and panel B presents the results for income in the first month of the contribution spell. All the explanatory variables included in the regressions for income in the second and following months of the contribution spells are strongly significant at a 1 percent level. As expected, the lagged wage is a strong predictor of current wage. The tenure variable (log of duration) shows how wages tend to increase along a labor relationship. This effect is much stronger in the case of men than in the case of women, probably related with the type of jobs in which women are participating, with wage discrimination practices among Chilean firms or the interrupted careers of Chilean women. The specification also includes a linear trend (month) to capture general trends in real wages in Chile.

[17] See section A of the methodological appendix for the details about the methodology.

Finally, the age variables exhibit the usual concave shape with a positive return to experience but at a decreasing rate. Once again, these returns are higher for men than for women.

Table 11 also includes the standard deviations of the two error components involved in the individuals wage stream: a time-invariant individual effect and an idiosyncratic shock. The individual effect accounts for as much as 31 and 44 percent of total variance for men and women, respectively. The R-squared of the within-group estimator is 0.54 for both men and women.

The explanatory variables in the regressions for income in the first month of the contribution spell include the individual effects estimated in the previous equation. This is a non-standard practice. However, the results look sensible both from a theoretical and an empirical point of view. First, we cannot observe the education level of the individual and this term could be capturing the joint effect of education and ability. Second, as expected, the coefficient associated to the estimated individual effect is significant and positive (see panel B). Third, the R-squared of the regression increases significantly when we introduce it.

As expected, the goodness-of-fit is lower for this equation (0.24 and 0.20 for men and women, respectively) than for the previous one.

Table 11: Chile—Estimation of Labor Income

A) Equation (1): $\ln w_{it} = \rho \ \ln w_{it-1} + \beta_1 \ln dur_{it} + \beta_2 a_{it} + \beta_3 a^2_{it} + \delta_t + v_i + e_{it}$

Independent Variables	*Men*	*Women*
$\ln w_{it-1}$	0.629***	0.588***
Log of Duration	1.756***	0.981***
Month	0.001***	0.002***
Age	0.227***	0.052***
Age^2	-0.026***	-0.009***
Constant	3.468***	3.909***
Nº of observations	694858	389951
Nº of individuals	6083	4645
R-squared	0.54	0.54
Standard Deviation of v_i	0.2473	0.3050
Standard Deviation of e_{it}	0.3685	0.3428

Note: * significant at 10%, ** significant at 5%, *** significant at 1%. w_{it} is the real wage of individual i at period t, expressed in pesos of December 2004. Duration is divided by 100. Age is measured in years and is divided by 10. Age^2 is divided by 100. Monthly dummies were included.

Source: Authors' computations.

B) Equation (9): $\ln b_i = \alpha_1 + \alpha_2 a_i + \alpha_3 a^2_i + \alpha_4 \hat{v}_i + \varepsilon_i$

Independent Variables	*Men*	*Women*
v	1.640 ***	1.257***
Age	0.647***	0.274***
Age^2	-0.075***	-0.037***
Constant	10.527***	11.136***
Observations	23533	13555
R-squared	0.24	0.2

Note: * significant at 10%, ** significant at 5%, *** significant at 1%. b_i is the average real wage of individual i in the first 12 months of the contribution spell. Age is measured in years and is divided by 10. Age^2 is divided by 100. v is the individual effect computed in the wage equation presented in panel A (see equation (1)).

Source: Authors' computations.

4.1.2 *Projection of labor income for Uruguay*

We split the Uruguayan dataset in four groups, based on gender and sector of activity (private and public). Unlike in the case of Chile, where the variable to be explained was the real wage, we modeled in the Uruguayan case the behavior of the ratio of the nominal wage of individual i at period t respect to the nominal wage index of the economy at period t. The results of the estimations are presented in Table 12.

Table 12: Uruguay—Estimation of Labor Income

A) Equation (1): $\ln w_{it} = \rho \ \ln w_{it-1} + \beta_1 \ln dur_{it} + \beta_2 a_{it} + \beta_3 a^2_{it} + \delta_t + v_i + e_{it}$

Independent Variables	*Men*		*Women*	
	Private Sector	*Public Sector*	*Private Sector*	*Public Sector*
$\ln w_{it-1}$	0.652***	0.511***	0.686***	0.563***
Log of Duration	1.060***	6.313***	-0.116**	5.211***
Age	0.093***	0.155***	0.044***	0.130***
Age^2	-0.016***	-0.016***	-0.006***	-0.013***
Constant	0.787***	1.116***	0.635***	0.877***
Nº of Observations	1572014	391141	1164871	416175
Nº of Individuals	31693	4977	24883	5212
R-squared	0.48	0.37	0.52	0.41
Standard Deviation of v_i	0.38	0.35	0.41	0.28
Standard Deviation of e_{it}	0.32	0.28	0.29	0.28

Note: * significant at 10%, ** significant at 5%, *** significant at 1%. w_{it} is the ratio of the nominal wage of individual i at period t respect to the nominal wage index of the economy at period t. Duration is divided by 100. Age is measured in years and is divided by 10. Age^2 is divided by 100. Monthly dummies were included. * significant at 10% ** significant at 5% *** significant at 1%.

Source: Authors' computations.

B) Equation (9): $\ln b_i = \alpha_1 + \alpha_2 a_i + \alpha_3 a^2_i + \alpha_4 \hat{v}_i + \varepsilon_i$

Independent Variables	*Men*		*Women*	
	Private Sector	*Public Sector*	*Private Sector*	*Public Sector*
v	1.214***	1.571***	1.324***	1.577***
Age	0.304***	0.386***	0.110***	0.409***
Age^2	-0.042***	-0.041**	-0.016***	-0.044***
Constant	2.320***	2.440***	2.368***	2.217***
Nº of Observations	34986	1105	24209	1799
R-squared	0.24	0.37	0.32	0.29

Note: * significant at 10%, ** significant at 5%, *** significant at 1%. b_i is the average of the nominal wage of individual i at period t relative to the average wage index of the economy in the first 12 months of the contribution spell. Age is measured in years and is divided by 10. Age^2 is divided by 100. v is the individual effect computed in equation (1).

Source: Authors' computations.

As in the case of Chile, the persistence coefficient (ρ) is significant and positive for the four categories. The estimated coefficients are slightly greater for women than for men within each sector (private and public). In turn, the coefficients for the private sector are approximately 20 percent greater than those for the public sector.

Tenure is also significant and positive for three groups: men working in the private sector and men and women working in the public sector. However, it is significant but negative for women working in the private sector. This result is counterintuitive but the magnitude of the coefficient is very small. Notice also that the effect of tenure appears to be stronger for men than for women in the public sector, and very well above the effect for men working in the private sector. The latter results are in line with our previous expectations because we know that pay for tenure is one of the deterministic rules that are used to determine wages in the public sector but not in the private sector.

The coefficients of Age and Age Squared are both significant and present the expected sign in all categories. The estimated polynomial for men and women in the public sector are very similar. Moreover we can conclude that the effect of age (after controlling for tenure) is higher in the public sector, which is in line with the behavior we expected.

The coefficients of the month dummies (not reported in Table 12) capture the seasonality of the series in a flexible manner and behave as expected. In particular, those of June and December capture the fact that in these months employees receive a Christmas Bonus that attains approximately half of their monthly wage. Obviously the Wage Indexes smooth this effect and thus the month dummies appear significant, positive and with a coefficient very close to 0.50 in all categories.

The individual effects account for a larger share of total variance in the Uruguayan than in the Chilean database. For most categories, the variance of the individual effects represents about 60 percent of the variance of the combined time-invariant individual effects and idiosyncratic shocks. For women working in the public sector it is in the order of 50 percent.

The goodness of fit can be analyzed by using different measures. First, the R-squared of the within-group estimator is between 0.37 and 0.52. Second, we regressed the observed on the predicted wage ratio. The R-squared of this regression attains: 0.88, 0.93, 0.86 and 0.85 for men and women in the private sector and men and women in the public sector respectively.

Table 12 also summarizes the results of the estimation of the initial wage of a contribution spell. Like in the case of Chile, the individual effects estimated with the first equation were used as explanatory variables in the initial wage equation. The coefficient associated to the estimated individual effects resulted significant and positive in all categories (see panel B) and the R-squared of the regressions increased from less than 0.01 to more than 0.24 (depending on the categories) when we introduced the effects. Also, the variables Age and Age Squared are significant and present the expected sign (positive and negative respectively), for all categories.

4.2 Projections of the contribution status

We estimate the probabilities of contributing to the system using two alternative procedures: linear probability and duration models. We then perform Monte Carlo simulations. We present and compare the results with different procedures to assess the robustness of the simulations to the estimation procedure. In appendix B, we present the details and discuss the pros and cons of each procedure.

We are not interested in the pension rights of a representative worker but on the distribution of pension rights that arises from the existing heterogeneity. In the regression models, we try to capture the heterogeneity that is relevant for pensions (i) using covariates that capture the observed heterogeneity, (ii) computing the individual unobserved effects when this is possible, and (iii) characterizing the distribution of the individual effects, otherwise.

Analysis similar to ours has been done in the past for Chile and Uruguay. Berstein et al. (2006) run probit models to estimate the probability of contributing in Chile. Using the probit model coupled with the labor income equation, they predict the history of contributions of workers in their sample beyond the period of observation. This methodology has not been used with either the Argentinean or Uruguayan data. Bucheli et al. (2008) used survival analysis to model the probabilities of transition between the states contributing and not contributing in Uruguay. They then run Monte Carlo simulations to generate a whole distribution of work histories. This method has not been used with the Argentinean or Chilean data.

4.2.1 *Contribution status projected with the linear probability model*

The estimated linear probability models are summarized in Table 13 to Table 15. The explanatory variables include a polynomial of degree three in age (degree two in the case of Chile), a dummy “elderly” for individuals aged 60 and over, and the rate of unemployment. We also included the individual effects computed in the wage equations in Chile and Uruguay, and dummies for ages 14 through 17 for some categories of workers and run different regressions for workers aged 18 and less and workers aged 19 and more in Uruguay.

Table 13: Argentina—Linear Probability Models

A) Equation (15): $Contribute\, s_{it} = x_{it}'\beta^{LPM} + \eta_i + \theta_{it} = X_{it}'\beta^{LPM} + \varsigma_{it} \quad , \quad t \geq 1$

Independent Variables	*Men*	*Women*
Age	0.006***	0.011***
Age^2	-0.001***	-0.002***
Age^3	0.000***	0.000***
Elderly	0.304***	-0.193***
Constant	-0.679***	-1.382***
Nº of Observations	537146	221731
R-squared	0.03	0.04

Note: * significant at 10%, ** significant at 5%, *** significant at 1%. Age is measured in months. Age^2 is divided by 1,000, and Age^3 is divided by 100,000. Elderly is a dummy equal to 1 if the individual is 60 years or older.

Source: Authors' computations.

B) Equation **(16)**: $\theta_{it} = \rho\theta_{it-1} + \varepsilon_{it} \quad , \quad t \geq 1.$

Independent Variables	*Men*	*Women*
θ_{it-1}	0.794***	0.822***
Constant	-0.001***	-0.001***
Nº of Observations	530858	219188
R-squared	0.63	0.70

Note: * significant at 10%, ** significant at 5%, *** significant at 1%.

Source: Authors' computations.

Table 14: Chile—Linear Probability Models

A) Equation (15): $Contribute\ s_{it} = x_{it}'\beta^{LPM} + \eta_i + \theta_{it} = X_{it}'\beta^{LPM} + \varsigma_{it} \quad , \quad t \geq 1$

Independent Variables	Men	Women
Age	0.004***	0.003***
Age^2	-0.000***	-0.000***
Unemployment	-0.010***	-0.006***
ν	0.389***	0.447***
Constant	-0.243***	-0.088***
Nº of Observations	1,303,354	995,651
R-squared	0.10	0.12

Note: * significant at 10%, ** significant at 5%, *** significant at 1%. Age is measured in months. Age^2 is divided by 1,000, and Age^3 is divided by 100,000. Unemployement is the country's unemployment rate. ν is the individual effect computed in the wage equation (see Equation (1)).

Source: Authors' computations.

B) Equation (16): $\theta_{it} = \rho\theta_{it-1} + \varepsilon_{it} \quad , \quad t \geq 1$

Independent Variables	Men	Women
θ_{it-1}	0.869***	0.899***
Constant	-0.002***	-0.001***
Nº of Observations	1,297,271	991,006
R-squared	0.75	0.80

Note: * significant at 10%, ** significant at 5%, *** significant at 1%.

Source: Authors' computations.

Table 15: Uruguay—Linear Probability Models

A) Equation **(15)**: $Contribute\ s_{it} = x_{it}'\beta^{LPM} + \eta_i + \theta_{it} = X_{it}'\beta^{LPM} + \varsigma_{it} \quad , \quad t \geq 1$.

Independent Variables	*Men*			
	Private Sector		*Public Sector*	
	≤18 Years Old	*≥19 Years Old*	*≤18 Years Old*	*≥19 Years Old*
Age	0.280***	0.005***	-4.284	0.011***
Age^2	-1.515***	-0.008***	19.596	-0.017***
Age^3	0.274***	0.000***	-2.983	0.001***
D14	-0.051***			
D15	-0.045***			
D16	-0.031***			
D17	-0.040***		0.007	
Elderly		-0.038***		-0.021***
Unemployment	-0.005***	-0.013***	0.010***	-0.001***
ν	0.016***	0.279***	0.014	0.147***
Constant	-17.210***	-0.169***	311.541	-1.322***
Nº of Observations	265407	2884624	3153	459540
R-squared	0.09	0.08	0.02	0.14
Independent Variables	*Women*			
	Private Sector		*Public Sector*	
	≤18 years old	*≥19 years old*	*≤18 years old*	*≥19 years old*
Age	0.532***	0.006***	-2.109	0.019 ***
Age2	-2.748***	-0.010***	9.505	-0.034 ***
Age3	0.473***	0.001***	-1.424	0.002 ***
D14				
D15	-0.023***			
D16	-0.017***			
D17	-0.016***		0.003	
Elderly		-0.086***		-0.121 ***
Unemployment	-0.001 ***	-0.010***	0.013 ***	0.002 ***
ν	0.012 ***	0.268***	0.011	0.176 ***
Constant	-34.381 ***	-0.399***	155.434	-2.641 ***
Nº of Observations	162337	2321848	6339	500437
Adjusted R-squared	0.06	0.09	0.02	0.22

Note: * significant at 10%, ** significant at 5%, *** significant at 1%. Age is measured in months. Age^2 is divided by 1,000, and Age^3 is divided by 100,000. D14-D17 are dummies for ages 14-17. Elderly is a dummy equal to 1 if the individual is 60 years or older. Unemployment is the country's unemployment rate. ν is the individual effect computed in the wage equation (see Equation (1))).

Source: Authors' computations.

B) Equation (16): $\theta_{it} = \rho\theta_{it-1} + \varepsilon_{it} \quad , \quad t \geq 1$

Independent Variables	*Men*			
	Private Sector		*Public Sector*	
	≤18 Years Old	*≥19 Years Old*	*≤18 years Old*	*≥19 Years Old*
θ_{it-1}	0.804 ***	0.863 ***	0.760 ***	0.906 ***
Constant	-0.000	-0.001 ***	0.001	-0.001 ***
Nº of Observations	258512	2859826	2981	454735
R-squared	0.62	0.74	0.54	0.82
Independent Variables	*Women*			
	Private Sector		*Public Sector*	
	≤18 Years Old	*≥19 Years Old*	*≤18 Years Old*	*≥19 Years Old*
θ_{it-1}	0.803 ***	0.893 ***	0.789 ***	0.895 ***
Constant	0.000	-0.000 ***	0.001	-0.001 ***
Nº of Observations	157334	2301968	5991	495573
R-squared	0.60	0.80	0.61	0.80

Note: * significant at 10%, ** significant at 5%, *** significant at 1%.
Source: Authors' computations.

The explanatory variables were significant at 1 percent in most regressions. The unemployment rate has the expected negative impact on the probability of contributing.[18] The individual effect computed from the wage equation (ν) is meant to capture characteristics of individuals that impact on wages and could not be observed in our estimations, like education and ability. Whenever it could be included, this effect turned out to be highly significant and positive. The error terms in the contribution status equations show considerable persistence: the estimated coefficients of the lagged errors lie between 0.76 and 0.90 and are significant at 1 percent in all cases, as shown in panel B of tables 13 to 15. These results indicate that the probability that a worker contributes is substantially higher if he contributed the previous month than otherwise.

The linear probability model fits the data reasonably well in the three countries: the percentage of correct predictions is in all cases above 65 percent (Table 16).

[18] The significance of this coefficient should be taken with caution though, because of the aggregate variable problem first identified by Moulton (1990).

Table 16: Goodness-of-Fit of Linear Probability Models

	Percentage of Correct Predictions for...	
	$\tilde{C}_{it} = 0$	$\tilde{C}_{it} = 1$
Argentina		
Men	77.0	75.5
Women	76.7	75.4
Chile		
Men	64.5	75.1
Women	75.1	66.6
Uruguay		
Men private sector	74.0	80.7
Men public sector	67.2	95.8
Women private sector	75.7	78.7
Women public sector	65.0	95.0

Source: Authors' computations.

The main results for the Monte Carlo simulations with the linear probability models are summarized in Table 17. This table provides information about the proportion of the population in the simulated database that would accumulate at least 20, 30 and 35 years of contribution at 65 years old (I20, I30 and I35).

Table 17: Argentina, Chile and Uruguay—Proportion of Workers with at Least 20, 30 and 35 Years of Contribution at Age 65 (Linear Probability Model)

	Men			*Women*		
	I_{20}	I_{30}	I_{35}	I_{20}	I_{30}	I_{35}
Argentina	0.54	0.40	0.33	0.57	0.40	0.31
Chile	0.64	0.37	0.25	0.47	0.25	0.16
Uruguay						
Private Sector	0.72	0.54	0.47	0.68	0.48	0.40
Public Sector	0.92	0.87	0.82	0.93	0.87	0.82

Source: Authors' computations.

According to these simulations, significant swaths of the population would not reach 20, 30 or 35 years of contribution at 65 years of age. The required number of years of contribution to access a pension is 30 in Argentina and 35 in Uruguay. In the case of Chile, workers with less than 20 years of contribution receive a pension but they are not covered by the minimum pension guarantee. Neither Argentina nor Uruguay show significant differences between genders, but Chile does. In Chile, women have significantly lower periods of contribution than men.

Part of the population that lies below these thresholds could be "rescued" by a reduction of the required number of periods of contribution. Governments have recently started parametric reforms moving in that direction. The reform passed in the Chilean Congress in January 2008 replaced the minimum pension guarantee with a Solidarity Pension that does not require a minimum number of contributions (available to the elderly among the 60 percent poorest individuals in the population). The Uruguayan parliament passed a reform in the last quarter of 2008 that reduced the number of years of contribution required to access an ordinary pension from 35 to 30 and reduced the minimum age to access to the advanced-age pension from 70 to 65. The advanced-age pension requires 25 years of contribution rather than 30, but the benefit is smaller than the ordinary pension. These changes will be effective in 2009 and 2010.

4.2.2 *Contribution status projected with the duration model*

The covariates include a constant, a polynomial in age, duration, two variables with the interaction of duration and age, and the national rate of unemployment. In the case of Chile and Uruguay, we also include the individual effects estimated in the wage equation as a proxy of education and/or ability.[19] The results of estimating the duration models are summarized in Table 18, Table 19 and Table 20.

[19] As mentioned before, we do not have information in the Argentinean database to compute a labor income equation and identify individual effects.

Table 18: Argentina—Duration Models

		Transitions Out of Contributing		*Transitions Out of Not Contributing*	
		Men	*Women*	*Men*	*Women*
Duration		-0.293 ***	-0.369 ***	-0.784***	-0.989***
Duration*D18-		-0.011	0.468 ***		
D18				-0.547**	-0.978**
D19				-0.130	-0.315
D20				-0.019	-0.078
Age		-0.489**	-0.105	0.421**	0.525
Age^2		1.827**	0.167	-1.463**	-2.018
Age^3		-0.300**	0.002	0.230*	0.353
Age^4		0.181**	-0.017	-0.137*	-0.228*
Elderly		-0.286**	0.223		
Self-Employed		-0.881***	-0.967 ***	0.783***	0.971***
Constant		3.545**	0.351	-6.479***	-7.041**
Likelihood-ratio test of rho=0 [A]	Chi squared	1258.13	774.83	5059.38	2786.92
	Prob >= Chi squared	0.000	0.000	0.000	0.000

Note: [A] rho is the ratio of the individual effects variance to total variance.

Source: Authors' computations.

Table 19: Chile—Duration Models

Estimation of equation (22): $h_{it} = 1 - \exp\left(-\exp\left(x_{it}^{'}\ \beta^D + \gamma_t + u_i\right)\right)$

		Transitions Out of Contributing		*Transitions Out of Not Contributing*	
		Men	*Women*	*Men*	*Women*
Duration		-0.469***	-0.436***	0.085***	-0.105***
Duration*D18-		-0.012	0.142**		
D18				-0.288***	-0.381**
D19				-0.04	-0.078
D20				-0.019	0.019
Age		-0.671***	-0.693***	0.400***	-0.221
Age^2		2.580***	2.980***	-1.748***	0.466
Age^3		-0.433***	-0.554***	0.309***	-0.022
Age^4		0.272***	0.378***	-0.199***	-0.022
Elderly		-0.209**	-0.174		
Unemployment		0.027***	0.006**	-0.025***	-0.018***
ν		-1.380***	-1.514***	0.472***	0.516***
Constant		3.701***	3.384***	-5.608***	0.254
Likelihood-ratio test of rho=0 [A]	Chi squared	1846.95	1137.76	2752.66	1121.17
	Prob >= Chi squared	0.000	0.000	0.000	0.000

Note: [A] rho is the ratio of the individual effects variance to total variance.

Source: Authors' computations.

Table 20: Uruguay—Duration Models

Estimation of Equation (22): $h_{it} = 1 - \exp\left(-\exp\left(x_{it}^{'}\ \beta^{D} + \gamma_t + u_i\right)\right)$

Men					
		Transitions Out of Contributing		*Transitions Out of Not Contributing*	
		Private	Public	Private	Public
Duration		-0.345***	-0.549***	-0.266***	-0.362***
Duration*D18-		0.193***	0.973***		
D14				-0.644	
D15				-0.827***	
D16				-0.877***	
D17				-0.461***	-2.146**
D18				0.014	-0.861
D19				0.023	-0.491
D20				0.009	-0.287
Age		0.113**	0.574*	0.259***	-0.117
Age^2		-0.616***	-2.661**	-1.286***	-0.171
Age^3		0.111***	0.438**	0.256***	0.115
Age^4		-0.065***	-0.227**	-0.184***	-0.128
Elderly		0.410***	0.981***		
Unemployment		0.024***	-0.034***	-0.069***	-0.036**
ν		-0.956***	-1.135***	0.437***	-0.108
Constant		-3.441***	-6.630**	-3.231***	0.896
Likelihood-ratio test of rho=0	Chi squared	2812.27	303.78	2880.10	66.42
	Prob >= Chi squared	0.000	0.000	0.000	0.000
Women					
		Transitions Out of Contributing		*Transitions Out of Not Contributing*	
		Private	Public	Private	Public
Duration		-0.331***	-0.566***	-0.292***	-0.386***
Duration*D18-		0.181***	1.447***		
D15				-1.715***	
D16				-1.434***	
D17				-1.244***	-36.316
D18				-0.217**	-1.263
D19				-0.130**	-1.855*
D20				-0.152***	-0.932*
Age		0.134**	0.742	-0.300**	1.386
Age^2		-0.855***	-3.062	0.877*	-6.264
Age^3		0.166***	0.472	-0.104	1.161
Age^4		-0.103***	-0.229	0.033	-0.768
Elderly		0.312***	-0.326		
Unemployment		0.015***	-0.064***	-0.068***	0.048*
ν		-0.747***	-1.558***	0.574***	0.086
Constant		-3.090***	-8.215	1.615	-13.191
Likelihood-ratio test of rho=0 [A]	Chi squared	1752.76	49.59	1398.92	17.15
	Prob >= Chi squared	0.000	0.000	0.000	0.000

Note: [A] rho is the ratio of the individual effects variance to total variance.

Source: Authors' computations.

In all cases, we reject the hypothesis that there are no individual effects. Hence, as expected, we have mixed populations, reflecting the existing heterogeneity in terms of the probabilities of leaving each state: contributing and not contributing.

The coefficient of duration, where duration is the number of months the individual has spent in the state, is significant at 1 percent in all regressions. As expected, it is negative in most equations, indicating that the probability of making a transition decreases as individuals spend time in the state. Only in the case of Chilean men leaving the state of not contributing do we find a positive coefficient associated with duration.

In Argentina, we could distinguish self-employed and salaried workers. Surprisingly, the probability of making a transition out of contributing is smaller and the probability of making a transition out of not contributing is higher for the self-employed.

In the cases of Chile and Uruguay, we could control for the rate of unemployment. In general, the probability of making a transition out of contributing rises and the probability of making a transition out of not contributing decreases with the rate of unemployment. In the case of Uruguay, where it was possible to distinguish public and private workers, we find this pattern only for private workers. Unlike private workers, public workers are less likely to make a transition out of contributing when unemployment is high. These contrasting patterns suggest that while private workers are forced to stop contributing in downturns (probably because they lose jobs), public workers might be choosing to leave public employment when conditions in the labor market are comparatively good.

The individual effect of the wage equation (ν) has a statistically significant negative impact on the probability of making a transition out of contributing and, in most cases, a significant positive impact on the probability of making a transition out of not contributing.[20] Therefore, the higher the individuals' "ability", the lower the probability that they will stop contributing if they are doing so and the higher the probability that they start contributing if they were not contributing. This is consistent with the finding in the linear probability model that individuals' "ability" positively impact on the probability of contributing.

[20] Only in the case of Uruguayan public workers, ν has a non-significant impact on the probability of leaving the state not contributing.

We present in Table 21 measures of the goodness of fit of the duration models. The percentage of correct predictions is well above 60 percent in most cases. Only in the case of the Chilean men do we get a lower figure (47 percent).

Table 21: Goodness of Fit of Duration Models

	Percentage of Correct Predictions for...	
	$\tilde{C}_{it} = 0$	$\tilde{C}_{it} = 1$
Argentina		
Men	72.0	64.2
Women	77.0	66.2
Chile		
Men	46.57	66.00
Women	66.79	65.72
Uruguay		
Men Private Sector	77.0	68.0
Men Public Sector	82.3	87.0
Women Private Sector	84.5	65.4
Women Public Sector	81.1	86.2

Source: Authors' computations.

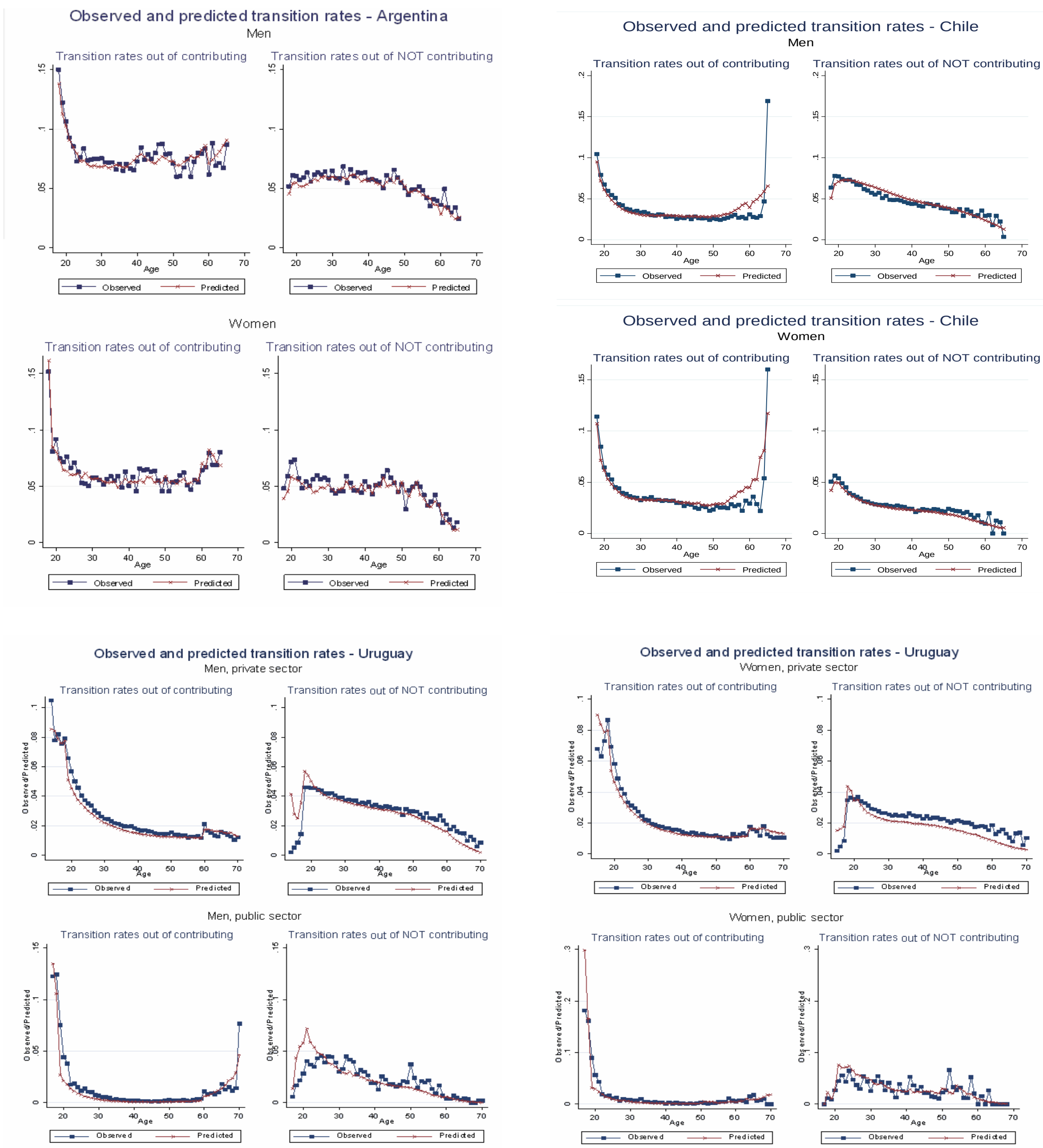
We present in Figure 8 the observed and predicted average transition rates by age in the three countries. In all cases, the predicted contribution profiles obtained by using the estimated model and just the first observation of each individual in the sample reproduce with high accuracy the observed transition rates. The transition rates out of contributing peak at early ages. Workers below 20 have monthly probabilities of leaving the state contributing as high as 15 percent in Argentina and about 10 percent in Chile and Uruguay.[21] These probabilities decrease with age up to 55 to 60. Nevertheless, the transition rates continue being relatively high, particularly so in the case of Argentina, where average transition rates out of contributing never fall below 5 percent per month. The rates of transition out of contributing tend to be smaller in Uruguay than in Argentina and Chile. As expected, the transition rates out of public employment in Uruguay are particularly low, but they are also low in the private sector compared to the rates we found in Argentina and Chile. This finding seems consistent with the comparatively high labor rigidity found in Uruguay by Forteza and Rama (2006).

The transition rates out of contributing start to increase at ages 55 to 60, probably due to retirement. This result may look natural, but it is somehow unexpected in our framework. Recall that we want to estimate transition rates conditional on not retiring or dying, because our ultimate goal is to predict

[21] Uruguayan public employees show higher transition rates, but these figures should be taken with caution for the number of public employees in these ages is very small.

the probability that a worker who has not retired or died continues contributing. Because of this, we treated transitions into retirement as censored observations. With this treatment, the reason for the rise in transition rates out of contributing at advanced ages is not clear. One possibility is that some workers stop contributing in order to retire, but they only receive the pension a few months later. These transitions would have been treated as censored observations had we been able to identify the motive, but they show up as transitions to the not contributing state because we cannot link them to retirement.

Figure 8: Transition Probabilities in Argentina, Chile and Uruguay



Source: Authors' Computations.

The average transition rates out of not contributing do not show the same age patterns across countries. Argentinean workers aged between 18 and 55 years have probabilities of making a transition from not contributing to contributing in the order of 5 percent per month. These probabilities fall at more advanced ages. Chilean workers show maximum transition rates out of not contributing at 20. At this age, the probability that a Chilean worker starts contributing in any month is about 7 to 8 percent for men and about 5 percent for women. The transition rates fall at higher ages. Uruguayan workers also show maximum transition rates out of not contributing at around 20 in the private sector and a bit later in the public sector. The maximum average transition rates out of not contributing by age are close to 5 percent per month in Uruguay, save for women in the public sector where the rates are a bit higher.

We present in Table 22 the number of transitions workers are expected to make in their working life. The table is based on the micro-simulations of work histories done with the two models. We compute the number of transitions for each simulated worker and report in the table the average number in each category. The expected number of transitions fluctuates between a minimum of 2, for men working in the Uruguayan public sector, to 18 for men in Argentina. In the three countries, men tend to present more transitions than women. In the only country in which we could distinguish private and public workers, we found that the difference between sexes shows only in the private sector. As expected, private workers show more transitions than public workers.

Table 22: Argentina, Chile and Uruguay—Lifetime Expected Number of Transitions

	Men		*Women*	
	Transitions Out of Contribution	*Transitions into Contribution*	*Transitions Out of Contribution*	*Transitions into Contribution*
Argentina				
Duration Model	7	8	3	3
Linear Probability Model	18	18	16	16
Chile				
Duration Model	9	8	6	5
Linear Probability Model	12	11	9	9
Uruguay				
Duration Model				
Private Sector	6	6	4	5
Public Sector	2	2	2	2
Total	5	5	3	4
Linear Probability Model				
Private Sector	13	13	11	11
Public Sector	4	5	4	5
Total	11	11	9	10

Note: Workers are assumed to begin in the no contribution state at initial ages (14-18, depending on the country, sex and activity sector).

Source: Authors' computations.

The linear probability models yield more transitions than the duration models. One possible explanation is that while by construction the probability of making a transition does not vary with duration in the linear probability model, it does vary in the duration model. In all the estimations done with the duration models, the probability of making a transition decreases significantly as workers spend time in the state. Hence, the linear probability model might overestimate the probability of making a transition for spells that are long enough.

As we did with the linear probability models, we used the duration models to simulate the work histories of a hypothetical generation that was born in 2005. In Table 23 we report the proportion of workers who, according to these simulations, would accumulate 20, 30 and 35 years of contribution at 65. In almost all the simulations, lower figures were produced with the duration than with the linear probability models. Only in the case of the Uruguayan private sector workers do we find similar figures with the two models. In all other cases, the differences are large.

Table 23: Proportion of Workers Who Would Accumulate 20, 30 and 35 Years of Contribution At 65 Argentina, Chile and Uruguay (Duration Model)

	Men			*Women*		
	I_{20}	I_{30}	I_{35}	I_{20}	I_{30}	I_{35}
Argentina	0.45	0.27	0.18	0.40	0.28	0.20
Chile	0.54	0.15	0.05	0.26	0.06	0.02
Uruguay						
Private Sector	0.73	0.56	0.45	0.61	0.44	0.34
Public Sector	0.82	0.73	0.65	0.79	0.68	0.59

Source: Authors' computations.

4.3 Projection of pension rights

4.3.1 *Conceptual issues*

Pension systems aim at providing protection against the risk of poverty in old age and helping individuals to smooth consumption along the life cycle. Usual indicators of performance that shed light on these two dimensions are: (i) the proportion of contributors that would be eligible for a pension at the usual retirement ages; (ii) the pension levels (relative to some normative lines); and (iii) the replacement rates. We compute these three indicators with the simulated projections.

Like most contributory pension programs with a defined benefit component, the Argentinean and Uruguayan public pension programs condition access to a pension to the accumulation of a minimum number of periods of contribution. This condition can vary with the retirement age. For example, in the case of Uruguay, workers aged between 60 and 69 years accrue rights to an ordinary pension served by the PAYG pillar after having contributed no less than 35 years, but they access to a (smaller) pension with only 15 years of contribution when they turn 70 years. We use our projections of contribution status to compute the proportion of workers who would comply with these eligibility conditions at the usual retirement ages.

The level of protection can be assessed by computing the distribution of pension levels and replacement rates. We describe the simulated distribution of pension rights relative to the minimum pension in each country at 65 that is, we report the proportion of contributors who, at age 65, would get (i) no pension, (ii) a pension below the minimum, (iii) the minimum pension, (iv) a pension larger than the minimum and lower than two minimums, and (v) a pension larger than two minimum pensions. We define the (apparent) replacement rate as the ratio of the total pension to the last salary.

Our simulations are designed to represent the pension programs in the long run (i.e., we do not discuss transition issues in detail). We also provide estimations of the distribution of pensions and replacement rates in the absence of minimum pensions and subsidies.

In principle, those workers who do not comply with the access conditions are not protected by the public pension programs. However, in the case of Uruguay we also provide estimations of the pensions they would get if these conditions were not enforced. The social security administrations do not have, in this case, work histories long enough to verify the requirement of years of contribution. In turn, according to our simulations, most contributors to the public pension systems in Argentina and Uruguay do not reach the required number of years of contribution to access to an ordinary pension. However, elderly coverage is comparatively high in these countries, which suggests that the procedures the administrations have adopted to circumvent the lack of enough records has in practice implied a loosening of the requisite. Because of this issue, in the case of Uruguay, we also provide estimations of pension rights generated at 65 years old, conditional on an in-practice no-binding years-of-contribution condition. Since we have loosened the years-of-contribution completely in these estimations, the results yield an upper bound to the rights the pension administration could recognize in the current circumstances.

4.3.2 *Estimation of pension rights*

We summarize in Table 24 the simulated distribution of pension rights in Chile under the conditions prevailing before the reform passed in January 2008, when the only source of assured protection was the minimum pension guarantee (MPG) for individuals with low pensions but at least 20 years of contribution.[22] To present the distribution of pension rights and make it comparable across countries, we define a reference point, considered to be the minimum pension level accepted by society. In the Chilean case, we define this point at Ch$75,000, which corresponds to the current level of the Basic Solidarity Pension (BSP). This is the minimum benefit guaranteed for all individuals in the 60 percent poorest portion of the population under the 2008 reform. Even if we are not calculating benefits

[22] In the Chilean case, we calculate pensions as single annuities bought at the legal retirement age (60 years for women, 65 for men), at the prevailing implicit interest rate on annuities (assuming 3.5% for future periods). To do so, we use the program "epension" developed by the Chilean Pension Supervising Authority, which is implemented as a Stata ADO file. For more information, see Pino (2005).

under the new reform, we use the BSP level as a reference point, as this is the most recent definition made by policymakers in Chile.

The BSP is smaller than the level that was guaranteed under the MPG scheme. This explains why nobody is receiving, in this scenario, a pension exactly equivalent to the BSP. The results suggest that almost 40 percent of the population would receive a pension smaller than the BSP and would not fulfill the 20 years requirement for the MPG.[23] The proportion of contributors who would receive less than the BSP is much higher among women (between 53 and 62 percent, depending on the estimation method) than men (between 21 and 23 percent).

As the system is a full defined contribution system and only affiliates are included in the sample, practically everyone would initially receive a pension, even if, in some cases, it was small and eliminated in due course. Approximately one third (almost half of all men and 15 percent of all women) of all retirees would receive a pension higher than two BSP. Median replacement rates would be relatively low under this scenario (about 35 percent for men and 11 percent for women). In general, the results obtained with the linear probability and the duration models are roughly consistent.

[23] These results are consistent with the ones presented in Berstein, Larrain and Pino (2006, Figure 4.2).

Table 24: Chile—Distribution of Simulated Pension Rights at 65 With Minimum Pension Guarantee

	Men		*Women*		*ALL*	
	LPM	*Duration*	*LPM*	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…						
No pension	1%	0%	1%	2%	1%	1%
A pension lower than the Basic Solidarity Pension a/	23%	21%	53%	62%	36%	38%
A Basic Solidarity Pension	0%	0%	0%	0%	0%	0%
More than one and less than two Basic Solidarity Pensions	28%	32%	29%	23%	28%	28%
Two or more Basic Solidarity Pension	49%	48%	17%	14%	35%	33%
Replacement Rate (Median)	34%	39%	10%	11%	21%	24%

Note: a/ The Basic Solidarity Pension level is equal to Ch$75.000 (starting in 2009), equivalent to US$116 (based on exchange rate on January 6, 2009).

Source: Authors' computations.

¡Error! No se encuentra el origen de la referencia. Table 25 summarizes the simulated pension rights at 65 in Uruguay under the current rules.[24] As in Chile, most workers would get a pension at 65, but many would receive pensions below the minimum public pension. These are workers who are not entitled to the public pension because they have not accumulated the required 35 years of contribution to access this benefit. These workers would not end up empty handed only because they would receive an annuity from their individual saving accounts. Only workers who never contributed to the individual accounts pillar would receive no pension at 65. It is important in this result the assumption we made that low income workers opted to contribute to the individual accounts. Otherwise, most of them would end up with no pension at all at 65.

[24] It should be noted that a reform that reduces the number of years required for accessing a pension and the minimum age at which the advanced-age pension is granted was passed in 2008, effectively reducing the proportion of contributors who would not fulfill the conditions. Current simulations are based on previous rules.

Table 25: **Uruguay—Distribution of Simulated Pension Rights at 65**

	Men			
	Private		*Public*	
	LPM	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,2%	1,0%	0,2%	3,0%
A pension lower than minimum pension	52,8%	52,9%	16,3%	28,4%
A minimum pension	0,0%	0,0%	0,0%	0,0%
More than one and less than two minimum pensions	20,0%	14,3%	3,8%	3,7%
Two or more minimum pensions	26,9%	31,8%	79,7%	64,9%
Replacement rate (median) - Total	43,6%	42,3%	45,8%	45,8%
Replacement rate (median) - Compliers	72,3%	71,3%	46,7%	48,4%
Replacement rate (median) – No compliers	22,9%	24,1%	15,3%	14,1%
	Women			
	Private		*Public*	
	LPM	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,6%	3,1%	0,2%	4,0%
A pension lower than minimum pension	58,4%	61,7%	17,5%	34,8%
A minimum pension	0,0%	0,0%	0,0%	0,0%
More than one and less than two minimum pensions	22,9%	15,0%	6,0%	5,2%
Two or more minimum pensions	18,0%	20,2%	76,3%	56,0%
Replacement rate (median) - Total	26,0%	24,1%	46,1%	45,2%
Replacement rate (median) - Compliers	60,2%	58,7%	47,1%	48,3%
Replacement rate (median) – No compliers	15,7%	16,0%	17,0%	14,9%

Source: Authors' computations.

The proportion of workers receiving no pension would be much higher at 60 than at 65. The proportion of workers who do not access a public pension because they have not accumulated 35 years of contribution is of course higher at 60 than at 65 (Table 17 and Table 23). But also, workers are not entitled to the annuity at 60, unless they are also eligible for the public pension. So workers who did not accumulate 35 years of contribution at 60 are not entitled to either the public or the private pension. At 65, workers can claim the annuity regardless of whether they are eligible for the public pension or not.

No one in this simulation would receive exactly the minimum pension either. This is because the minimum operates only on the pay-as-you-go defined benefit (PAYG-DB) pillar and all workers in our simulations contribute to both pillars. Therefore, all workers who are entitled to a public pension at 65 would receive at least the minimum from the PAYG-DB pillar plus an annuity. Between 3 and 6 percent of Uruguayan workers in the private sector would receive a pension larger than the minimum but lower than two minimums, and between 30 and 45 percent would receive a pension above two minimum pensions. Men receive high pensions in higher proportion than women.

The results for the Uruguayan public workers are rather surprising: According to the duration model, as high as 41 percent of women and 35 percent of men working in the public sector would not access a public pension at 65 because they would not have accumulated 35 or more years of contribution. This is the counterpart of the unexpectedly low densities of contribution found among public workers mentioned in section 3.3. Thanks to the annuities, these workers would still get a pension at 65 and many of them would get a pension above the minimum, and yet the proportion of public sector workers who would receive a pension below the minimum public pension would be as high as 16 percent and 18 percent, for men and women respectively, according to the linear probability model. These percentages are even higher in the estimations done with the duration model.

We find two possible explanations for this result. First, a public sector worker in these computations is someone who made more than half of his/her contributions as a public employee, so some "public workers" might actually spend a significant part of their working life in the private sector where interruptions are much more common. The second reason is that some special groups of public employees are required to contribute fewer years. This is the case, for example, of workers dealing with radioactive substances and teachers. We have not been able to correct for the special regimes, and we do know that in at least some categories of public workers the proportion of individuals in special regimes can be quite significant. We nevertheless do not find these explanations sufficiently convincing: The proportion of public workers who**¡Error! No se encuentra el origen de la referencia.** would not get a public pension at 65 looks too high. Also the difference between the results obtained with the two models looks too big in the case of public sector workers.

In order to analyze the impact of minimum pensions and subsidies, we run simulations without these provisions for Chile and Uruguay (Table 26 and Table 27).

Chilean results without the MPG or NSP programs are actually very similar to the ones under the MPG (Table 24), reflecting the scarce protection that this program was projected to provide, as the combination of low pensions and 20 years of contribution would rarely be fulfilled by Chilean affiliates. These results confirm previous findings by Berstein et al. (2006) who were the first to show that the Chilean MPG program would protect few workers. Interestingly, our results are close to theirs even though the models we used are different.

Table 26: **Chile—Distribution of Simulated Pension Rights at 65 Without Minimum Pensions and Subsidies**

	Men		*Women*		*ALL*	
	LPM	*Duration*	*LPM*	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…						
No pension	1%	0%	1%	2%	1%	1%
A pension lower than Basic Solidarity Pension a/	28%	25%	62%	67%	43%	43%
A Basic Solidarity Pension	0%	0%	0%	0%	0%	0%
More than one and less than Basic Solidarity Pensions	23%	27%	19%	18%	21%	23%
Two or more Basic Solidarity Pension	49%	48%	17%	14%	35%	33%
Replacement Rate (Median)	33%	38%	10%	10%	20%	23%

Note: a/ The Basic Solidarity Pension level is equal to Ch$75.000 (starting in 2009), equivalent to US$116 (based on exchange rate on January 6, 2009).

Source: Authors' computations.

As in Chile, in Uruguay the results without the minimum pension look similar to the ones with the minimum pension (Table 25 and Table 27). As expected, the proportion of workers receiving pensions below the minimum would be higher if minimum pensions did not exist, but the changes are not dramatic. More than half of the men working in the private sector would still receive a pension below the minimum public pension and almost sixty percent of women would be in that case. This result is related to the fact that workers receiving less than the minimum pension do not comply with the 35 years of contribution required to access to a public pension and hence do not access to minimum pensions. So the reasons why minimum pensions do not work are very similar in Chile and Uruguay: most workers whose self-financed pension would fall below the minimum are not eligible for the minimum pension because they have not contributed the required number of years.

Table 27: Uruguay—Distribution of Simulated Pension Rights at 65 Without Minimum Pensions and Subsidies

	Men			
	Private		*Public*	
	LPM	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,2%	1,0%	0,2%	3,0%
A pension lower than minimum pension	57,9%	58,7%	16,6%	28,6%
A minimum pension	0,0%	0,0%	0,0%	0,0%
More than one and less than two minimum pensions	16,3%	9,2%	3,5%	3,5%
Two or more minimum pensions	25,6%	31,0%	79,7%	64,9%
Replacement rate (median) - Total	43,6%	42,2%	45,8%	45,7%
Replacement rate (median) - Compliers	69,9%	68,6%	46,7%	48,3%
Replacement rate (median) – No compliers	22,9%	24,1%	15,3%	14,1%
	Women			
	Private		*Public*	
	LPM	*Duration*	*LPM*	Duration
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,6%	3,1%	0,2%	4,0%
A pension lower than minimum pension	67,6%	69,7%	17,9%	35,0%
A minimum pension	0,0%	0,0%	0,0%	0,0%
More than one and less than two minimum pensions	13,8%	7,0%	5,6%	5,0%
Two or more minimum pensions	18,0%	20,2%	76,3%	56,0%
Replacement rate (median) - Total	25,9%	24,1%	46,1%	45,1%
Replacement rate (median) - Compliers	57,6%	56,4%	47,1%	48,2%
Replacement rate (median) – No compliers	15,7%	16,0%	17,0%	14,9%

Source: Authors' computations.

There is ample evidence that the requirement of having contributed at least 35 years to access a pension is not well enforced in Uruguay. In most cases, the social security administration is unable to control for this requirement because records have not been kept; and in the absence of records, the administration accepts testimonials as proof that someone claiming a pension has actually contributed. Finally, according to the household surveys, most of the Uruguayan elderly are covered by pensions (Table 1). Even if some are covered by non-contributory and survivors pensions, which are not included in our simulations, the gap between the results in our simulations and the surveys suggests that the degree of enforcement is extremely low.

Taking the low enforcement argument to an extreme, we ran simulations assuming that all workers who contributed at least once will receive a public pension. In order to compute pensions, we assumed that the administration granted workers who did not accumulate the required years of contribution the minimum public pension, which is added to the annuity they get from their individual accounts. The results in this scenario are summarized in Table 28. The impact of the minimum pension is now much greater. No one would receive less than the minimum pension. Between 50 and 74 percent of workers in the private sector would receive a pension between one and two minimum pensions. In the public sector, most workers would receive a pension above two minimum pensions.

Table 28: Uruguay—Distribution of Simulated Pension Rights at 65 Assuming Requirements Not Enforced

	Men			
	Private		*Public*	
	LPM	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,0%	0,0%	0,0%	0,0%
A pension lower than minimum pension	0,0%	0,0%	0,0%	0,0%
A minimum pension	0,2%	1,1%	0,2%	3,0%
More than one and less than two minimum pensions	64,2%	50,1%	9,9%	13,0%
Two or more minimum pensions	35,6%	48,8%	89,9%	84,0%
Replacement rate (median) - Total	71,0%	68,8%	46,4%	47,1%
	Women			
	Private		*Public*	
	LPM	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,0%	0,0%	0,0%	0,0%
A pension lower than minimum pension	0,0%	0,0%	0,0%	0,0%
A minimum pension	0,7%	3,2%	0,2%	4,1%
More than one and less than two minimum pensions	74,3%	60,9%	14,2%	18,1%
Two or more minimum pensions	25,1%	35,9%	85,7%	77,8%
Replacement rate (median) - Total	60,1%	58,9%	46,9%	46,9%

Source: Authors' computations.

These results show that the impact of minimum pensions is larger when the years-of-contribution requirement is not enforced than when it is. This is because workers with short contribution careers also tend to have smaller wages entitling lower pensions. The interaction between minimum pensions and enforcement can be farther analyzed simulating a scenario with no minimum pensions and no enforcement. Between 28 and 46 per cent of private workers would receive a pension below the minimum in this scenario (Table 29).

Table 29: Uruguay—Distribution of Simulated Pension Rights at 65
With No Minimum Pension or Subsidies
Assuming Contribution Requirements Not Enforced

	Men			
	Private		*Public*	
	LPM	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,0%	0,0%	0,0%	0,0%
A pension lower than minimum pension	31,9%	28,3%	2,3%	3,6%
A minimum pension	0,2%	1,1%	0,2%	3,0%
More than one and less than two minimum pensions	33,5%	22,5%	7,6%	9,4%
Two or more minimum pensions	34,3%	48,1%	89,9%	84,0%
Replacement rate (median)	60,9%	60,0%	46,3%	46,8%
	Women			
	Private		*Public*	
	LPM	*Duration*	*LPM*	*Duration*
Proportion of Contributors Who at 65 Would Receive…				
No pension	0,0%	0,0%	0,0%	0,0%
A pension lower than the minimum pension	46,3%	41,8%	3,5%	5,1%
A minimum pension	0,7%	3,2%	0,2%	4,1%
More than one and less than two minimum pensions	28,0%	19,1%	10,6%	13,0%
Two or more minimum pensions	25,0%	35,9%	85,7%	77,8%
Replacement rate (median)	47,2%	47,6%	46,6%	46,3%

Source: Authors' computations.

These results confirm that minimum pensions are effective under current Uruguayan laws and circumstances to a large extent because the eligibility conditions are not enforced. If they were, most workers who would need a minimum pension would not be entitled to it.

V. Concluding Remarks

In this document, we propose alternative methods to project pension rights and implement these methods in Chile and Uruguay and partially in Argentina. We use incomplete work histories databases from the social security administrations (spanning from eight years in Uruguay to 25 years in Chile) to project entire lifetime work histories. We first fit linear probability and duration models of the contribution status and dynamic linear models of the income level. We then run Monte Carlo simulations to project work histories and compute pension rights.

The transition rates between contributing and not contributing are on average quite large. In Argentina, the average monthly transition rates out of contributing are at no age smaller than 5 percent, and during most of the working life the transition rates into contributing are in the order of 5 percent per month. The average rates are smaller in Chile, but even there the probability that a contributor makes a transition out of contributing is on average 3 percent or more each month during most of his/her working life. The rates are smaller in Uruguay, particularly so in the public sector. Nevertheless, Uruguayan workers in the private sector transit from contributing to not contributing at rates no lower on average than 1.5 percent per month, with considerably higher rates at younger ages. These transition rates are comparable and often higher than the rates of job creation and destruction and the rates of firing and hiring reported in the labor literature that helped motivate the flow approach to labor markets (Davis et al. 2006).

The transition rates out of contributing are higher at younger ages and decrease to reach a minimum at ages 40 to 50, approximately. This is especially worrying in defined contributions pension systems given the special importance of early contributions. In Chile and Uruguay, the transition rates out of not contributing also peak at younger ages and decrease later (the age pattern is less clear in Argentina). These age profiles of the transition rates suggest that, as expected, young workers are particularly mobile.

The transition rates from both contributing and not contributing are decreasing functions of the time spent in the state. Therefore what happens during the first months in the state seems to be a key determinant of work histories. Aggregate shocks do matter as the transition rates out of contributing are higher and into contributing are lower during downturns.

Higher income workers show lower transition rates out of and higher transition rates into contributing than low income workers. This is consistent with the finding that the densities of contribution increase with income level.

According to our results, if the current rules were strictly enforced, between 50 and 70 percent of contributors to the Argentinean SIJP and about 55 percent of private workers who contribute to the Uruguayan BPS would receive no public pension at age 65. This is because of the failure to satisfy the required years of contribution to access a pension. In Chile, contributors with short contribution histories do receive a pension, though they are not eligible for the minimum pension guarantee if they do not accumulate at least 20 years of contribution. Because of this condition, about 44 percent of contributors to the Chilean pension system would not get the minimum pension. Because of the correlation between densities of contribution and income levels, the percentage of workers who will not comply with the access conditions is higher the lower the average income of the worker.

The proportion of workers who would receive the minimum pension in Chile is surprisingly small. This result is driven by the positive correlation we find between the density of contribution and labor income. Indeed, most workers who comply with the years of contribution required to access a pension do have sufficient contributions to finance a pension above the minimum. In turn, most of the workers who would not reach the minimum pension because of their low wages and contributions would not be eligible for the minimum pension because they also have short contribution histories.

Under the MPG scheme and if the real interest rates in the pension funds were 4 percent per year, about 60 percent of contributors to the Chilean system would receive a pension above the basic solidarity pension. In Uruguay, the percentage of private workers getting above minimum pensions would be less than 50 percent.

These results highlight what seems to be a flaw in the design of pension programs with high eligibility requirements (as in many defined benefit PAYG schemes) and of the minimum pension provision in the three countries under study: The histories of contribution required to access pensions (in Argentina and Uruguay) or minimum pension benefits are apparently unattainable for large segments of the population, at least under the labor market conditions that have prevailed in the region. In the presence of widespread informality, these conditions were established to avoid abuse and provide incentives to contribute, but our findings suggest that these tough access conditions severely undermined the social protection function of the pension systems without being effective in providing incentives for continuous contribution.

In contrast, weak enforcement may have offset the stringent conditions, and the lack of recordkeeping made verifying contribution histories impossible. The social security administrations implemented alternative forms of recognizing years of contribution, including testimony of witnesses, which seem to have significantly loosened the eligibility conditions in practice. However, as the work history records accumulate, the administrations will be increasingly able to enforce the rules. The question is thus whether current rules will be appropriate when they are actually enforced.

Chile and Uruguay passed in 2008 reforms that will mitigate (or directly solve) the problems we have highlighted in this study. In January, the Chilean parliament passed a reform law that among other things eliminated the requirement of having to contribute for 20 years to receive any kind of support from the government. A solidarity pension that supplements the self-financed pension will substitute the minimum pension guarantee, and it has no years-of-contribution requirement to receive the solidarity pension. In turn, the Uruguayan parliament passed in September a reform law that reduced the number of years of contribution required to receive an ordinary pension from 35 to 30 years, and reduced the minimum age to access to the so-called advanced age pension from 70 to 65. With the new rules, a worker can retire at 60 with 30 years of contribution, at 65 with 25, and at 66 to 70 with two less years of contribution for each year retirement is postponed.

There are several reasons why the results we have presented in this document should be considered preliminary. First, we are not yet getting results with the linear probability and duration models similar enough to conclude that our results are robust (despite acceptable indicators of goodness of fit in both models). We hope the final results will be within the range of figures presented in this document, but this range remains too wide to be considered satisfactory. Second, we could not compute pension rights for Argentina because of data unavailability: Unfortunately, we do not expect the situation to improve in the near future. Third, the pension rights computed for Uruguay should be taken with caution, not only because of the problems mentioned in the first place, but also because there seems to be a big gap between *de jure* and *de facto* pension policies in the country. It is hard to believe that the comparatively high levels of coverage observed in Uruguay would have been possible, if the tough “on paper” conditions established to access to pensions were strictly enforced “in practice”. Because of this, our projections for Uruguay could be better-interpreted as the pension rights that would be observed under current conditions if the social security norms were strictly enforced.

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A. Methodological Appendix

We projected pension entitlements based on independent projections of labor income and contribution status. In turn, these projections are based on the estimation of econometric models that benefit from the longitudinal structure of the datasets.

A1. Projection of labor income

We estimate two wage equations. Wages in the second and following months of a spell of contribution are modeled using a dynamic equation. Wages in the first month of a contribution spell are modeled with a static equation. We chose the most parsimonious specifications in all cases.

Given that our main goal is to project income, we are particularly interested in exploring the impact on wages of time invariant and deterministic covariates, like age and time trends. Typical covariates in this type of regressions are education and cohort, but data availability will determine the exact set of controls we will be able to include in regressions for each country.

Wages in the second and following months of the spell of contribution are assumed to be governed by the following stochastic process,

$$\ln w_{it} = \rho \ \ln w_{it-1} + \beta_1 \ln dur_{it} + \beta_2 a_{it} + \beta_3 a^2_{it} + \delta_t + v_i + e_{it} \tag{1}$$

Where w_{it} is the real wage, in the case of Chile, and the ratio of the nominal wage of individual *i* at period *t* respect to the nominal wage index of the economy at period *t*, in the case of Uruguay (we considered separately the nominal wage index for the private and the public sector); $ldur_{it}$ is the tenure in the current job; a_{it} is the age; δ_t are month dummies; and v_i is a time invariant unobservable characteristic of individual i. The idiosyncratic shock e_{it} is assumed to be normally distributed with mean 0 and variance σ_i^2. As long as we expect w_{it} to be stationary we do not introduce any deterministic time trend in the equation. We do not observe the education level of the individuals. Therefore, the term v_i will capture, at least in part, the cross section heterogeneity that comes from education jointly with other time invariant unobservable characteristics like ability.

We estimate equation (1) using the Within Group estimator. It is important to notice that despite this estimator is biased in the case of a dynamic model that includes the lagged dependent variable as a regressor, the bias is of order 1/T and thus consistent in large samples. We computed the potential bias for the Uruguayan sample (T=105) and concluded that it is smaller than 2 percentage points.[25]

Furthermore, as we have long panels, we are able to obtain reliable estimates of the individual effects v_i. We compute the individual effects as:

$$\hat{v}_i = \frac{1}{T_i}\sum_{t=1}^{T_i} \left(\ln w_{it} - \left(\hat{\rho} \ \ln w_{it-1} + \hat{\beta}_1 ldur_{it} + \hat{\beta}_2 a_{it} + \hat{\beta}_3 a^2{}_{it} + \hat{\delta}_t\right)\right) \tag{2}$$

The model for the prediction of the labor income stream also includes the estimation of an individual variance through the following equation:

$$\hat{\sigma}^2{}_i = \frac{1}{T_i}\sum_{t=1}^{T_i} \hat{e}_{it}^2 \tag{3}$$

$$\hat{e}_{it} = \ln w_{it} - \left(\hat{\rho} \ \ln w_{it-1} + \hat{\beta}_1 ldur_{it} + \hat{\beta}_2 a_{it} + \hat{\beta}_3 a^2{}_{it} + \hat{\delta}_t + \hat{v}_i\right) \tag{4}$$

Finally, the non-financial-income stream predictions are calculated using the following equation:

$$\ln \tilde{w}_{is} = \hat{\rho} \ \ln \tilde{w}_{is-1} + \hat{\beta}_1 ld\tilde{u}r_{is} + \hat{\beta}_2 a_{is} + \hat{\beta}_3 a^2{}_{is} + \hat{\delta}_t + \hat{v}_i + \hat{\sigma}_i \tilde{z}_{is} \tag{5}$$

Where s = t + 1, ..., H; t is the last month of our sample; H is the previous month of the 71st anniversary of the individual; $\tilde{z}_{it}$ are pseudo-random draws from a Standard Normal distribution and

$$\tilde{w}_{it} = \begin{cases} w_{it} & \text{if wage of individual i in period t is observed} \\ \tilde{w}_{it} & \text{if wage of individual i in period t is not observed} \end{cases}$$

$$ld\tilde{u}r_{it} = \begin{cases} ldur_{it} & \text{if tenure of individual i in period t is observed} \\ ld\tilde{u}r_{it} & \text{if tenure of individual i in period t is not observed} \end{cases} \tag{6}$$

[25] The alternative is the GMM Arellano-Bond estimator that is unbiased for fixed and small T. However, we were not able to implement this estimator given the dimension of the panel, even when we tried to restrict the dimension of the instrument matrix using no more than 4 lags. In any case, it would be important to prove in the future the performance of the WG estimations with respect to the Arellano-Bond estimations by limiting the sample size (in N not in T) or trying to estimate the whole sample using STATA10+WINDOWS VISTA (we are currently facing memory constraints associated with a conflict between Stata 10 and Windows XP).

We also predict the Nominal Wage Index (NWI) for each sector using the following simple rule:

$$NWI_{j,s} = NWI_{j,s-1}(1+g) \quad ^{26} \tag{7}$$

with $s = t+1,.....H$ and j = private sector, public sector. Afterwards, we obtain the nominal covered wage for individual i period s (W_{is}) as:

$$\tilde{W}_{is} = NWI_s \tilde{w}_{is} \quad s = t+1,....H \tag{8}$$

Before continuing with the equation for wages in the first month of a contribution spell we notice that an exception is made in order to predict the future wage streams. This is the case when the individual enters the spell of non-contribution but stays in it for up to two months. In these cases we ignore that the Markov chain was broken and continue using equation (1) to predict wages paid in the months in which individuals are effectively contributing. That is to say, the wage stream in these months is predicted as if there had been no interruption in the contribution spell. Of course, we impute zero wages in the months the individual did not contribute.

The second equation is applied to the initial month of the contribution spells, save for the case mentioned in the previous paragraph of interruptions that lasted less than three months. The equation to estimate is as follows:

$$\ln b_i = \alpha_1 + \alpha_2 a_i + \alpha_3 a^2{}_i + \alpha_4 \hat{v}_i + \varepsilon_i \tag{9}$$

Where b_i is the average real wage in the first 12 months of the contribution spell, in the case of Chile, and the average of the nominal wage of individual i in the first 12 months of the contribution spell divided by the nominal wage index of the economy at period t, in the case of Uruguay. a_t is the age and $\hat{v}_i$ is the individual effect estimated with equation (1). We omit the index t in order to highlight the fact that here we are using a pooled cross-section. The model we estimate is static in nature. We used the OLS estimator with the White formula in order to obtain the standard errors. Notice that the introduction of the term $\hat{v}_i$ in equation (9) is a non-standard practice that should be considered with some caution. Indeed, we have not seen any work that handles the estimation in the initial period in

[26] To project the nominal wage index we assumed an annual inflation rate of 8% and an annual growth of the real wage of 1.5%.

the spell in this manner. We base our choice in two arguments. First, we expect that the individual effect estimation is roughly speaking a proxy for the education level and ability of the individual and thus help us to improve the wage predictions. Second, as we will show below this variable turns out to be significant and with the expected sign in all the estimations.

We use equation (9) to predict the covered wage of the first month of a new spell of contribution provided the individual has previously stayed in a spell of non-contribution at least three months. Thus the prediction is given by:

$$\ln \tilde{b}_i = \hat{\alpha}_1 + \hat{\alpha}_2 a_i + \hat{\alpha}_3 a^2{}_i + \hat{\alpha}_4 \hat{v}_i + \hat{\sigma}_\varepsilon \tilde{z}_i \tag{10}$$

where $\hat{\sigma}_\varepsilon$ is the estimated standard error of the regression (9) and $\tilde{z}_i$ are pseudo-random draws from a Standard Normal distribution. Notice that $\hat{v}_i$ is estimated from equation (1) while the remaining parameters come from equation (9).

After the prediction of b_i (using equation (10)) we proceed as in equation (8) to predict the first covered wage of a spell of contribution as:

$$\tilde{W}_{is} = NWI_s \tilde{b}_i \tag{11}$$

Then, wages for the second and subsequent months of contribution are predicted using equation (8).

Predictions according to equations (5) and (10) can only be computed for the individuals in the sample, i.e. individuals for which we can compute the individual effects. But the model is used in this document to predict the labor income flow of "new" individuals. In this case, we simulate the individual effects:[27]

$$\tilde{v}_i = \hat{\sigma}_v \tilde{z}_i \tag{12}$$

where $\hat{\sigma}_v$ is the squared root of the distribution of the individual effects in equation (1) and $\tilde{z}_i$ are pseudo-random draws from a Standard Normal distribution. The labor income stream of the newborn individuals is thus computed using the following two equations:

[27] The implicit assumption here is that the distribution of the individual effect does not vary with age or cohort.

$$\ln\tilde{b}_i = \hat{\alpha}_1 + \hat{\alpha}_2 a_i + \hat{\alpha}_3 a^2{}_i + \hat{\alpha}_4 \tilde{v}_i + \hat{\sigma}_e \tilde{z}_i \tag{13}$$

$$\ln\tilde{w}_{is} = \hat{\rho}\ \ln\tilde{w}_{is-1} + \hat{\beta}_1 ld\tilde{u}r_{is} + \hat{\beta}_2 a_{is} + \hat{\beta}_3 a^2{}_{is} + \hat{\delta}_t + \tilde{v}_i \tag{14}$$

Where $\tilde{v}_i$ is simulated from a standard normal distribution with zero mean and $\hat{\sigma}_v^2$ variance.

A2. Projection of the contribution status

As mentioned in the body of the text, we run two different types of models, the linear probability and the duration models.

A2.1 The linear probability model

A simple approach to estimating the probability of making contributions that directly exploits, for prediction purposes, the longitudinal nature of the data is to fit a fixed effect linear probability model. The main advantage of this type of models is that they allow using estimated individual fixed effects to make predictions for the entire lifetime. This is particularly relevant if the data does not allow including sufficiently rich control variables.

A2.1.1 *Specification and estimation of the linear probability model*

In the linear probability model, the dependent variable is equal to one if the individual makes a contribution during a particular month and zero otherwise $\left(Contributes_{it} \in \{0,1\}\right)$. We use a model with an autoregressive error term to capture the persistence of contribution spells. The model is as follows:

$$Contributes_{it} = x_{it}'\beta^{LPM} + \eta_i + \theta_{it} = x_{it}'\beta^{LPM} + \varsigma_{it} \quad , \quad t \geq 1 \tag{15}$$

$$\theta_{it} = \rho\theta_{it-1} + \varepsilon_{it} \quad , \quad t \geq 1 \tag{16}$$

The β^{LPM} can be estimated consistently in the first equation using OLS, if the regressors in x_{it} are exogenous, or using the within groups estimator otherwise. The individual effects can be computed as: $\hat{\eta}_i = \sum_{t=1}^{T_i} \hat{\varsigma}_{it} \Big/ T_i$, where $\hat{\varsigma}_{it}$ is the estimated residual of the first equation. Subtracting the individual effects from the residual of the first equation we compute the θ_{it} that we then use to estimate ρ and later to simulate the work histories.

A2.1.2 *Projection of work histories with the linear probability model*

We simulated the contribution status of workers across their lifetime up to 70 years old. Our simulations are thus conditional on the individual not retiring or dying before 70. For each group of workers, we determine in the database the first age at which workers start contributing and use this age as our starting point in the simulations. We then set the contribution status of each worker in each of the following periods using the estimated regressions to simulate the probability of contributing. More specifically we simulate the probability of contributing $\left(\tilde{P}_{it} = \Pr\left(Contributes_{it} = 1\right)\right)$, draw realizations from a uniform (0,1) distribution ($draw_{it}$) and set C_{it} as: $\tilde{C}_{it} = 1 \ if \ draw_{it} < \tilde{P}_{it} \ and \ 0 \ otherwise$. In turn, the simulated probability of contributing is computed as:

$$\tilde{P}_{it} = X_{it}'\hat{\beta} + \hat{\eta}_i + \tilde{\theta}_{it},$$
$$\tilde{\theta}_{it} = \hat{\rho}\tilde{\theta}_{it-1} = \hat{\rho}\left(\tilde{C}_{it-1} - X_{it-1}'\hat{\beta} - \hat{\eta}_i\right), \quad \textbf{(17)}$$

We compute the percentage of correct predictions in the sample to assess the goodness of fit of the models.

A2.2 Transition probabilities plus Monte Carlo simulations

Survival analysis is the standard methodology used to model transition probabilities.[28] In this project we have worked with two discrete-time models, the complementary-log-log (cloglog) and the logit models. Time is intrinsically discrete in our setting, because individuals can only be in one contribution status in each month that is, they cannot change the contribution status on a continuous basis. The complementary-log-log model assumes proportional hazard rates and the logit model assumes proportional odd rates. As it is often the case, both models yield very similar results in our study so we will only present in detail the cloglog model. We use these models to simulate the transition probabilities and the histories of contribution status.

[28] For surveys of survival analysis, see Jenkins (2005) and Van den Berg (2001).

A2.2.1 Specification and estimation of the duration model

Assuming proportional hazards, the discrete time model looks as follows:

$$\log\left(-\log\left(1-h_{it}\right)\right)= x_{it}'\beta^{D}+\gamma_t+u_i \tag{18}$$

Where h_{it} is the interval *t* hazard rate of individual *i; x_{it}* are the covariates; γ_t is the cloglog transformation of the baseline hazard: $\gamma_t=\log\left(-\log\left(1-h_{0t}\right)\right)$; and u_i is the unobserved individual effect ("frailty"). The baseline hazard can in turn be modeled in different ways. Lacking theoretical guidance, we first use duration dummies to estimate flexible piece-wise constant baseline hazards and then, depending on the results, we fit more parsimonious models.

The contribution of individual i to the sample likelihood is:[29]

$$\Gamma\left(u_i\right)=\begin{cases}\prod_{t=1}^{T_i}\left(1-h_{it}\right) \ \ if \ \ i \ \ makes \ no \ transition\\ \left[\dfrac{h_{iT_i}}{1-h_{iT_i}}\right]\prod_{t=1}^{T_i}\left(1-h_{it}\right) \ \ if \ \ i \ \ makes \ a \, transition\end{cases} \tag{19}$$

Let $y_{it}=1$ if individual *i* makes a transition in period *t* and 0 otherwise. Then, the contribution of individual *i* to the sample likelihood can be written as:

$$\Gamma\left(u_i\right)=\left[\frac{h_{iT_i}}{1-h_{iT_i}}\right]^{y_{iT_i}}\prod_{t=1}^{T_i}\left(1-h_{it}\right) \tag{20}$$

Assuming that u_i is normally distributed with zero mean and variance σ_u^2, the total likelihood is:

$$\Gamma=\int_{-\infty}^{\infty}\frac{e^{-u_i^2/2\sigma_u^2}}{\sqrt{2\pi}\sigma_u}\Gamma\left(u_i\right)du_i \tag{21}$$

[29] We are assuming here single spells, so that individuals can be identified with spells. This is a highly simplifying but clearly not very realistic assumption. Unfortunately, we have no simple way of getting rid off this assumption. The comparison of the results obtained with the two methods, the linear probability model and the duration model, can serve as a test of the robustness of our results despite of these shortcomings.

This can be seen as a latent variable model where $y_{it} \in \{0,1\}$, with y_{it} =1 if and only if $e_{it} < x_{it}^{'}\beta^{D} + \gamma_{t} + u_{i}$ and e_{it} is distributed according to the Gumbel(0,1) distribution, i.e. its cumulative distribution function is D(e)=1-exp(-exp(e)). With these assumptions, the hazard rate can be expressed as:

$$h_{it} = \Pr(y_{it} = 1) = \Pr(e_{it} < x_{it}^{'}\beta^{D} + \gamma_{t} + u_{i}) = 1 - \exp(-\exp(x_{it}^{'}\beta^{D} + \gamma_{t} + u_{i})) \tag{22}$$

In the current version of this paper, we separately estimate duration models for contribution and non-contribution spells. Because the assumption that the two processes are independent is admittedly strong, we plan in a future version to also explore more recent techniques in which the distribution of random effects is estimated jointly (Huff Stevens 1999).

A2.2.2 Projection of work histories

We first simulate the hazard rates of contributing and not contributing using the estimated cloglog model. We draw the individual effects from a normal distribution with zero mean and the standard deviation estimated before $(SD(u_i))$:

$$\tilde{u}_i = SD(u_i) * \tilde{z} \quad ; \quad \tilde{z} \sim \text{Normal}(0,1) \tag{23}$$

We then run Monte Carlo simulations with the following hazard rates:

$$\log(-\log(1 - \tilde{h}_{it})) = x_{it}^{'}\hat{\beta}^{D} + \tilde{\gamma}_{t} + \tilde{u}_{i} \tag{24}$$

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 $draw_{it}$ represent a draw from a uniform distribution in the $[0,1]$ interval. The individual contributes in *t* if $draw_{i,t-1} \geq \tilde{h}_{i,t-1}^{C}$ and he was contributing in *t-1*, or if $draw_{i,t-1} \leq \tilde{h}_{i,t-1}^{N}$ 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 $1 - \tilde{h}_{i,t-1}^{C}$, 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 $\tilde{h}_{i,t-1}^{N}$, which is the probability of leaving the state "not contributing". Simulated individuals start in the state not contributing and switch into contributing and back into not contributing according to these hazard rates.

We compute the percentage of correct predictions for the contribution status in the sample as our measure of the goodness of fit of the models.