The Impact of Private Participation and Countervailing Information on Disability Costs: Evidence from Chile

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Abstract

Many social security systems face high and escalating disability costs. In Chile's new system, the disability assessment procedure includes participation by private pension funds (AFPs) and insurance companies, who finance the benefit, have a direct pecuniary interest in controlling costs and are able to pursue this objective by helping to set criteria and providing countervailing information. We hypothesize that these procedures and incentives will keep costs low, by cutting the incidence of successful claims. Using the Cox proportional hazard model and a retrospective sample of new and old system affiliates (EPS 2002), we find that disability hazard rates are only 20-35% as high in the new system as in the old traditional system. Analysis of mortality rates suggests that the new system has accurately targeted individuals with more severe medical problems.

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Social security systems in many countries face problems of high and escalating disability costs. This has been ascribed, alternatively, to demand-side responses to unemployment rate and generous benefits, supply-side factors such as easy eligibility rules and exogenous forces such as population aging.. The impact of alternative assessment procedures has also been a much-debated issue (see, for example, Rupp and Stapleton 1995 and 1998, Gruber 2000, Autor and Duggan 2003, Duggan and Imberman 2006, Autor and Duggan 2006, von Wachter, Song and Manchester 2007). This paper analyzes the impact of a radical change in disability assessment procedures and incentives implemented in Chile as part of its new pension system.

When Chile's new retirement system was introduced in 1981, workers then in the labor force had a choice between staying in the old system or switching to the new system. New entrants to the labor force had to join the new system. Thus, after 1981 the two systems co-existed, although the old system is gradually being phased out. Those who switched were subject to a new disability system as well as a new pension system. The disability part is less well-known than the old age pension part, but it is equally innovative. It differs from traditional public disability insurance in two ways: 1) it is largely pre-funded--through the accumulation in the retirement account and later through an insurance payment made when the person becomes permanently disabled, sufficient to cover a lifetime defined benefit annuity; and 2) the disability assessment procedure includes participation by private pension funds (AFPs) and life insurance companies, who help define and apply the criteria for disability, finance the benefit and have a direct pecuniary interest in controlling costs. Survivors' insurance for working affiliates is handled in the same way—it provides a pre-funded defined benefit for survivors—and both are financed through a combined disability and survivors' (D&S) insurance fee.

These fees are strikingly lower in Chile's new system than in countries with pure public PAYG systems. The insurance fee is currently about 1% of wages, with 2/3 of this—approximately .7% of wages--for lifetime disability benefits. D&S insurance fees are .9%-1.7% of wages in other Latin American countries that adopted features of the Chilean model (AIOS 2005). For comparison, the disability cost is 1.8% of wages and running into financial difficulties in the U.S., over 3% in most other industrialized countries and up to 10% in some European countries (Andrews 1999; U.S. Social Security Advisory Board 2001).[1] Costs and age-specific inflow of newly disabled are also much lower than in Chile's old system. In this paper we focus on differences between the new and old systems, mainly during the period when they co-existed.

In this paper we focus on differences between the new and old systems, mainly during the period when they co-existed. Part I describes the Chilean procedures. We hypothesize that the participation of private organizations in the assessment procedure of the new system, with incentives to keep costs low and the power to provide countervailing information and appeal, has cut the incidence of successful disability claims. Part II tests this hypothesis, applying the Cox proportional hazard model to a recent retrospective sample of old and new system affiliates (EPS 2002). We find that the hazard of workers becoming disability pensioners in the new system is only 20-35% that in the old system, after controlling for age, gender and unemployment rates. Tests for selection bias rule that out as the major explanation. Further, the new system accurately targets the disabled with more severe medical problems, as measured by their higher mortality rates, while the old system does not sort between those with high versus low mortality rates. The Conclusion considers social cost-benefit trade-offs and policy implications for other countries.

Designing social policy regarding disability is especially complex because of the impossibility of avoiding subjective evaluations and value judgments. Insurable disability can be defined as an objective medically determined condition or as a more subjective inability to work at one's normal job or at any job. The latter definition may be the raison d'etre for insurance but the former is more clear-cut to apply and less subject to moral hazard. Given the ambiguities, both type 1 errors (false positives) and type 2 errors (false negative) are inevitable; the latter fails to protect claimants who are truly disabled while the former imposes potentially high costs on workers. For efficiency, systems should strive to be on the frontier giving the least false negatives for any given number of false positives, but value judgments must be made to determine the preferred point on this frontier. The Chilean system places a heavy weight on objective medical conditions and on minimizing false positives, thereby economizing on costs. This paper analyzes the procedures used to accomplish this and presents some outcomes. We do not take a normative position on whether this is the "right" choice of objectives and criteria since, in fact, the right choice probably varies according to a country's values, capacities and resources.

I. Cost Controls by Private Pension Funds in the Chilean Scheme[2]

Procedures in Chile's old vs. new systems

Prior to 1981 Chile's disability system was like traditional publicly managed schemes in other countries. Multiple programs existed and they were run on a pay-as-you-go basis, although with a projected long term deficit. This is still the case today in the old system. Disability claims are filed by individuals with the help of their doctors. They are assessed by regional Medical Commissions that are part of the public health system and have no financial incentive to limit successful claims. Each regional Commission acts independently, at the provincial level, without a technical commission to set medical protocols or a central commission to assure uniform standards. Although each Commission applies its own criteria, typically the Commissions define disability as a condition that prevents the worker from doing his or her *habitual work*. The pension system, which pays the bills, does not participate in this procedure. Technically it can appeal approved claims, but it virtually never does. Data are not publicly available on acceptance and rejection rates on claims, although anecdotal evidence indicates that a large majority are approved.

In contrast, Chile's new disability system defines disability as a medical inability to do *any work.* Regional Medical Boards are set up specifically to assess disability claims but criteria are spelled out in a detailed technical manual that is used by each Medical Board. AFPs and insurance companies regularly present information and frequently appeal approved claims. The Association of AFPs keeps track of decisions and their variations over time and regions. Thus, decision-making power remains vested in public gatekeepers but this is balanced with countervailing information from private companies that have strong financial incentives to contain costs.

How disability insurance works: incentives to contain costs

Disability insurance in Chile's new system starts with the mandatory retirement accounts, to which each individual must contribute 10% of wages. This contribution is invested in the pension fund company (AFP) of his choice. Old age pensions depend on this defined contribution plus investment earnings. In contrast, if a worker becomes disabled before retiring, he receives a defined benefit. This is accomplished through the private insurance market, with government providing detailed regulations and guarantees.

Specifically: Each insured worker is guaranteed a benefit that is 70% of his average wage if he is totally disabled, 50% if partially disabled, indexed to inflation. During an initial three-year period of temporary disability, this benefit is paid by the AFP. After the provisional period, if the worker is certified as permanently disabled, the entire lifetime benefit is fully pre-funded. Part of this benefit is covered by his or her own retirement account. The remainder is covered by a top-up (the "additional payment") needed to finance an annuity that equals the specified defined benefit. The disabled worker uses these funds to purchase a lifetime annuity or a programmed withdrawal pension that follows a formula set by the regulator.[3]

Each AFP is required to purchase a term group insurance policy that covers the additional payment for its affiliates. The typical contract shares the risk: the AFP pays a provisional rate that covers costs up to a maximum and keeps most of the savings beneath that maximum, while the insurance company takes over after the maximum rate has been reached and keeps most of the investment earnings on the provisional premiums. The provisional and maximum rates for future contracts, of course, depend on the experience in past contracts. Survivors' insurance for workers is covered in the same way, in exchange for a combined D&S insurance fee that is passed on to workers by the AFP.

The D&S insurance fee is included in the general administrative charge that each worker pays the AFP. Each AFP sets its own fees and, apart from a small flat component, is required to charge all its affiliates the same percentage of their wage—regardless of age, gender, occupation, health status or account size.[4] For any given total fee the AFP charges, part is spent on disability and survivors insurance and part on administration, so lower disability costs mean greater profits. AFP fees currently average around 2.4% of wages, of which the disability insurance portion is about .7% and the survivors' insurance portion is .3% (authors' estimates). Suppose that profits comprise .4% for a typical AFP. Then if it cuts the combined insurance cost from 1% to .8% and continues charging the same market-determined fee, its profits increase by .2% of wages, or 50% of the initial profit amount. AFPs are therefore highly motivated to keep disability probabilities low, and they are given a role in the evaluation procedure that allows them to pursue this goal. We hypothesize that these procedures and incentives will produce lower age-specific disability rates than would have obtained in the old system.

Assessment for temporary and permanent disability. Initial claims are evaluated by 21 Regional Medical Boards, each made up of three doctors hired by the public Superintendencia of AFPs (SAFP). The member may present his/her own medical tests and invite his/her personal doctor to take part in the discussions (but not the vote). The AFPs and life insurance companies also have a non-voting representative--their Associations have organized a group of medical observers, who regularly attend Board meetings, raise questions and monitor its work. Thus, the decision-makers are still medical and public but information and arguments are provided by private representatives. According to data from the Association of AFPs, about 60% of all claims are approved at the first assessment, for a temporary disability benefit.

Three years later, the member is re-assessed. AFPs also participate in this second assessment. Approximately 70% of those who are provisionally disabled eventually come up for a second reassessment. About 94% of these are approved as permanently disabled (Association of AFPs 2004, 2005 and 2006). The additional payment to cover the cost of the defined benefit is put into the account of the individual at the point when he is certified as permanently disabled. He receives a lifetime benefit even if he returns to work; hence work by disabled beneficiaries is not penalized, as it is in many countries that base benefits on incapacity to hold a job. (In this sense, the Chilean disability system rewards work, as does the old age system; see Edwards and James 2005).

Appeals by AFPs and life insurance companies. Traditional public systems usually do not allow agencies to appeal against approved claims; they only allow workers to appeal denials of disability status. And workers who appeal are, in some countries, allowed to be represented by attorneys. As a result, the appeals process invariably raises successful claims and costs. In the U.S., for example, appeals raise successful claims by 15-20 percentage points.[5] In Chile, the process is more symmetrical—AFPs, life insurance companies and workers can all appeal the decisions of the Regional Boards to a Central Board. The Central Medical Board is also made up of three physicians appointed and paid by the SAFP. Some AFPs and insurance companies hire their own doctors to try to build strong appeals. Over-all, in 2004-06, AFPs and life insurance companies appealed 17-20% of insured approved claims and one third of these appeals were successful. This roughly offset the successful appeals by workers, so the net impact of appeals was close to 0 (Association of AFPs 2004, 2005 and 2006).

AFP role in shaping criteria for total and partial disability. A Technical Commission meets periodically to determine the medical criteria for granting partial and total disability. These criteria are spelled out in a lengthy protocol manual that is used by all decision-making groups. Representatives of the AFPs and the insurance companies, as well as three public representatives, sit on this commission, with a vote. For each medically-defined handicap presented by the member, the rules allot a certain percentage of disability, which are summed to produce the total degree of disability. The Medical Boards may increase this percentage discretionarily according to specified "complementary factors" in the case of older members with low levels of education, which make it difficult to retrain for a new line of work. If the degree of disability exceeds 67% the member is considered totally disabled, whether or not he has continued to work, and is granted a 70% defined benefit. If the degree of disability is 50-67% he is partially disabled and gets a 50% defined benefit. If degree of disability is less than 50%, he is not considered disabled. During 2004-06, almost 25% of approved claims were for partial disability. This proportion has been increasing over time.

Eligibility for insurance: avoiding adverse selection. Adverse selection could potentially be a big problem in an economy like Chile's, with a high degree of informality and self-employment. Self-employed individuals and independent contractors are not required to contribute to the system but may voluntarily do so. A healthy individual could work as an independent contractor or in the informal sector, thereby avoiding contributions, but start contributing or move to formal sector employment if he develops a complaint and anticipates filing a disability claim. Individuals who are approaching old age with small accounts, because they have worked in the informal sector or self-employment for part of their lives, also have a strong incentive to enter the system and apply for disability, because their defined benefit would be greater than the old age benefit based on their own accumulation. This would require a large "additional payment" by the AFP or insurance company. Thus, adverse selection based on size of contingent top-up as well as probability of disability might raise the rate of disability applications among contributors and insurance costs for the system as a whole.

However, insurance eligibility rules, monitoring and marketing strategies by AFPs reduce the likelihood that this will happen in Chile. AFPs market aggressively to individuals who fall into low-risk groups but pay small or zero commissions to sales agents for new members who are high risk or become disabled shortly after joining (Estelle James and Augusto Iglesias 2006). More important, eligibility for the defined benefit and the insurance top-up depends on recent work history. In general, an individual must have worked and contributed within the past 12 months in order to be insured and get the additional payment.[6] AFPs keep the contribution records of affiliated workers and thus can check whether they have contributed long and recently enough to be eligible.[7] In 2004-06 about 40% of all approved claims at the first stage and 30% at the second stage were deemed ineligible for insurance (Association of AFPs 2004). They get access to the balance in their own accounts but they do not get the additional payment that would cover a 70% replacement rate. (Some of these are eligible for the government's minimum pension guarantee—see below). The AFPs then concentrate their medical challenges on the insured.

Combating strategic behavior by monitoring the reference wage. Another way to discourage strategic behavior by workers with irregular contribution histories is to set a low reference wage for such people. The reference wage used to determine the defined benefit is the simple average of earnings during the prior ten years, expressed in the price-indexed Chilean currency, the UF (Unidad de Fomento), and with a ceiling. Workers who have been in the informal sector, unemployed, or out of the labor force for part of the last ten years will have 0's averaged in and will therefore have a low reference wage and benefit, even if they are insured.[8] AFPs use their records to ensure that these rules for defining the reference wage are strictly applied, thereby diminishing adverse selection and disability costs.

Consequently, during 2004-06, only 28% of original claimants ended up permanently disabled and insured, a proportion that is quite low by international standards (in the US, for example, acceptance rates are more than double). One quarter of these were considered partially disabled and many get a benefit that is far less than 70% (or 50%) of their full working wage (calculations by authors based on data in Association of AFPs, 2004, 2005, 2006). We argue that a major role in containing these costs was played by the AFPs and life insurance companies who actively participate in the assessment procedure and have a vested interest in containing costs.

The government's minimum pension guarantee

Underpinning these privately financed arrangements is the government's minimum pension guarantee (MPG), which sets a floor—currently 25-29% of the average wage—to disability and old age pensions (see James, Martinez and Iglesias 2006, Edwards and James 2006). Twenty years of contributions are needed for MPG eligibility among old age pensioners, but only ten years (or less) are needed among disability pensioners.[9] If the pensioner's accumulation (including the additional payment) is not large enough to cover an annuity above the MPG level, he or she must keep the money in the account and withdraw monthly benefits equal to the MPG. When the account balance becomes zero, the government steps in to pay the pension, if the pensioner is eligible for the MPG. Presently, about half of all disabled pensioners have benefits at the MPG level and in one third of these cases the benefits are already financed by the public treasury. This proportion will probably grow over time as more accounts are depleted. Most likely to receive the MPG are the disabled who are not insured, insured individuals with a small reference wage, those who are only partially disabled, and surviving widows of disabled men. For each of these categories, policy choices reduce the cost of the private insurance but increase the cost of the public contingent liability. Thus, the MPG serves as an escape valve for a cost-conscious private disability insurance system but its own future costs are uncertain.

Disability inflow rates in Chile vs. other countries

We expect these procedures to lead to a low inflow of age-specific newly disabled beneficiaries relative to insured population in Chile compared with other countries and, indeed, this is the case. For example, in 1999, for age group 45-54, 2.9 per thousand insured members were accepted to newly disabled status in Chile, compared with 7.8 per thousand people in that age group in the US and 8.6 in Organization for Economic Cooperation and Development (OECD) countries as a whole (OECD 2003 and Table 1). Over all ages, 1 per thousand insured members were accepted to newly disabled status in Chile in 2004, compared with 3 to 5 per thousand in the U.S. over the past two decades

(U.S. Social Security Board of Trustees 2005; Social Security Bulletin, various years). These low incidence rates lead to low insurance fees. Of course, many factors besides system incentives enter into these incidence differentials--in particular, the definition of disability, the existence of other programs that cover certain groups of disabled, and the safety conditions in occupations covered by the system (also see endnote 1). However, it seems likely that the participation of private pension funds and insurance companies in the assessment procedure and the incentives facing these organizations is an important part of the story. We test this hypothesis, while holding country-specific factors constant, by comparing disability hazard rates in the new Chilean system with the old one, which was broadly similar to public systems in other countries.

II. Testing the New System Impact

The sample

To isolate the impact of system procedures and incentives, we apply the Cox proportional hazard model to a recent sample survey of new and old system affiliates (EPS2002). EPS is a large retrospective survey, with about 17,000 observations, that was conducted in 2002 and is representative of the universe of people who were affiliated with the new or old systems at some point between 1982 and 2001. We know each individual's system affiliation and age in 2002, work history, age of death, disability or old age pension (if relevant), schooling, sex and marital status. We are interested in the propensity of these affiliates to become disabled pensioners and whether this propensity is different under the new and old systems, after controlling for other factors that might affect the disability hazard.

To carry out this analysis, from this large sample we constructed a sub-sample of individuals who were born 1932-1962. Individuals born before 1932 pose potential survival bias and memory problems, discussed later. We focus our analysis on the hazard of disability pensioning between ages 40 and 64, because disability pensions below age 40 are a very low probability occurrence, and eligibility stops at age 65 for men, 60 for women. Effectively, this means that we omit cohorts who were born after 1962, who were under 40 by 2002.

Applying these cohort cut-off criteria, we create a sub-sample consisting of 1840 individuals born 1932-62 who were old system affiliates throughout their working lives, 2988 who were new system affiliates throughout their working lives and 2890 who were old system affiliates initially but switched to new system affiliation. We construct the retrospective work histories of these individuals, going back to age 40 in the old system, age 40 (or age in 1982, if greater than 40) in the new system. Of course, old system members come disproportionately from earlier cohorts and are observed disproportionately at older ages than new-system members. Nevertheless, there is substantial overlap in cohorts, ages and years observed in the old and new systems (Tables 2, 3 and 4).

Workers who switched systems posed a special problem--they are treated as if they were in the old system sample before 1982 and the new system sample afterwards.[10] By definition, they all appear in our analysis of new system hazards. However, they appear in our old system analysis only if they were at least 40 years old in 1982. In reality, most switchers were younger than 40 in 1982, leaving just 606 switchers to be included in our old-system sub-sample. We create a double identity for these 606 switchers—one in the old system prior to 1982 and one in the new system afterwards; they are identical except for system affiliation, ages and years observed.

Therefore we end up with 8324 identities —1840 + 2988 + 2890 + 606--of whom 70% are in the new system (Table 2). Disability pensioners comprise about 2% of the total, or 172 individuals, one-third of whom are in the new system (Table 4). [11] For our analysis of selection bias in a later section, we divide this sample into groups that had choice of system versus those who were assigned and had no choice. (Tables 5 and 6).

Kaplan-Meier survival functions for new and old systems

We start by examining Kaplan-Meier survival functions, separately for the new and old systems (Figure 1), for ages 40-64. This shows us the cumulative probability ($CumS_{aget}$) that an individual who is a member of the "at risk group" will remain without a disability pension up to a given age, t. It is obtained by multiplying the single-year survival rates (S_{aget}) for all preceding ages, up to and including age t. We define the "at risk" group at a given age as all individuals who reach that age without a disability or old age pension and the hazard of dropping out as a disability pensioner is the proportion of the at-risk group who become a new disabled pensioner at that age. These are the "failures" and the rest are "survivors."[12]

(1) $$CumS_{aget} = CumS_{aget-1} * S_{aget} = S_{age40} * S_{age41} * \ldots S_{aget-1} * S_{aget}$$

where:

S_{aget} = % $atrisk_t$ *who survived through age t* = $1-(failures_t)/atrisk_t)$

$atrisk_t = atrisk_{t-1} - failures_{t-1} - dropouts_t + newentrants_t$

$failures_t$ = *number of newly disabled during age t*

$dropouts_t$ = *number* $atrisk_{t-1}$ *who died or became old age pensioners by t*

$newentrants_t$ = number of new entrants to system between age t-1 and t

a*$trisk_{40}$ = all non-pensioner affiliates at start of age 40*

S_{age40} *= % $atrisk_{40}$ who survived through age 40 = 1 – ($failures_{40}$/$atrisk_{40}$)*

It is immediately apparent from the K-M curve in Figure 1 that new-system affiliates have a substantially higher probability of surviving as non-disability-pensioners than old-system affiliates. For example, among old-system affiliates at age 40, 4.9% were likely to become disability pensioners by age 55, against just 1.6% of new system affiliates. We proceed to estimate the Cox proportional hazard model, which allows us to control for other factors and to establish whether these differences are statistically significant.

Cox proportional hazard model

The Cox proportional hazard model estimates a baseline hazard function (in which the values of co-variates are fixed) and the impact of co-variate variation on this hazard. It enables us to estimate age-specific hazards and the statistical significance of differences due to co-variates. It is based on the assumption that the hazard, *h(t)* (i.e. the proportion of the population at risk that becomes a newly disabled pensioner at age t), is:

(2) $h(t) = e^{\mathbf{Xi\beta}} h_0(t)$

where:

h(t) is the hazard, given the values of co-variates

$h_0(t)$ is the baseline hazard, setting co-variates to zero or other fixed value

$\mathbf{X_i}$ is a vector of covariates

$\boldsymbol{\beta}$ is a vector of parameters to be estimated.

Dividing both sides of the equation by $h_0(t)$ we obtain

(3) $h(t)/h_0(t) = e^{\mathbf{Xi\beta}}$.

That is, the model assumes that the effect of $\mathbf{Xi\beta}$ is proportional over all ages covered by the baseline hazard, hence the term proportional hazard function. Our main object is to measure the impact of the co-variate "new system," which we represent by a dummy variable equal to 1 for those affiliated to the new system. The baseline hazard applies to old-system affiliates. A $\boldsymbol{\beta}$ different from zero (or an exponentiated $\boldsymbol{\beta}$ different from 1) would lead us to reject the hypothesis that individuals of the same age have the same disability hazard in the two systems. In that case, estimated coefficients give the amount by which the baseline hazard must be multiplied to obtain the new-system hazard. We include a second co-variate, unemployment rate, which may have direct effects on the hazard of disability that might affect the measured system impact. In the baseline this is set at 10%, which is about average for the period. In some specifications we include education as a co-variate but this did not significantly affect the baseline hazard or the reform effect, so we present most of our results without education—see below.

A proportional effect over all ages would imply that the ratio of hazards with and without co-variates is constant for all ages. Initial estimates for our data indicated that the proportionality assumption does not hold for the entire baseline. Therefore, we split the baseline into five sections (40-44, 45-49, 50-54, 55-59 and 60-64) to make the proportionality rule applicable over pre-established sections. Instead of estimating one coefficient for the effect of new system and unemployment rate, we estimate five coefficients, one for each age group.

Tests further showed that baseline disability hazards vary by individual characteristics such as sex and marital status, so we estimate the Cox model stratified by these two variables. In other words, we allow for differences in the baseline hazard between men and women, married and singles, while constraining the proportionate impact of the co-variates to be the same for both genders and marriage groups. The estimated baseline hazards are lower for women than men and for married over single individuals (Figure 2). Baseline hazards are zero for women in the age range 60-64, since disability pensioning is ruled out after normal retirement age, which is 60 for women, 65 for men. Baseline hazards generally increase with age and peak at 5-year intervals (ages 50, 55, 60).

The new system impact

Applying the Cox model, the exponentiated coefficients for the "new system" dummies are far smaller than one and significant at the 1% level in all the age ranges between 40 and 59. Over most ages, workers in the new system are only 21-35% as likely to start a disability pension as they were in the old system. The largest new-system decline occurs at ages 45-49 while the smallest is for ages 60-64 (the latter was tested for men only and was only marginally significant) (Table 7, col. 1). These differences in hazard rates are similar to the difference in stock of disability pensioners (relative to retirement pensioners) in the new and old systems in Chile.[13] They are also roughly similar to the observed differences in incidence between Chile and other countries with traditional systems. Figure 3 compares the old system (baseline) hazard and the new system hazard (baseline*estimated coefficients for each age range) for married men.

To obtain the Cox model survival function we calculate (1-hazard rate) cumulatively for the old system and (1-baseline hazard*estimated coefficient) for the new system, separately for married and single men and women, holding unemployment constant at 10% (Figure 4). As expected, in each case the estimated new-system survival functions are far above the old system estimates. Although the single-age hazard rates become closer after age 60, the cumulative survival function remains higher in the new system through age 65, because of the larger number of survivals from earlier ages. These results are consistent with the hypothesis that the assessment procedure in the new system has had a strong negative impact on the rate of successful disability claims, compared with the old public disability insurance system.[14]

Effect of changing unemployment rates on disability hazard and new system effect

It is often observed, in the U.S. and other countries, that disability claims rise during periods of unemployment, as individuals lose their jobs, can't find new ones and may try to avail themselves of disability benefits (see Rupp and Stapleton 1995 and 1998, Gruber 2000, Autor and Duggan 2003, Duggan and Imberman 2006, Autor and Duggan 2006, von Wachter, Song and Manchester 2007; also see Castro 2005). Then, if unemployment rates were higher during earlier periods in Chile, this could help explain the higher disability rates observed for the old system.

We were able to disentangle the effects of new system and unemployment because there is a fair degree of overlap in the years at risk and years of pensioning of old and new system affiliates (see Tables 3 and 4). Additionally, the Chilean economy went through cyclical upturns and downturns both before and after the reform. Unemployment was low through the 1960's, reaching 3% in 1973, then rose sharply to 23% by 1982, during a period of economic chaos and hyper-inflation. Post-reform, the economy went through a period of prolonged growth during which unemployment fell to 6% by 1995 but then rose during a cyclical downturn, reaching 14% by 2002.

In earlier work on old age pension probabilities, we found that age-specific probabilities of pensioning rose during periods of high unemployment—but this did not change the reform effect (Edwards and James 2005). That is exactly what we found here, for disability pensioning. Individuals at risk in periods of higher unemployment have a larger hazard of becoming a disability pensioner. Specifically, in our main specification this probability rises 5-9 percentage points for each one percentage point increase in the unemployment rate for individuals in their 50's (but in the specification that includes earlier cohorts the unemployment rate becomes insignificant; apparently this variable is sensitive to choice of cohorts and time periods.) However, the new system effect barely changes when unemployment is in the equation--in fact, it become stronger for some ages (compare columns 1, 2 and 4 in Table 7).

Impact of education on hazards and system effects

Low income and education levels may also increase disability risk. Blue collar workers may have a higher probability of injury than white collar workers and high earners may have better access to preventative and remedial medical care that prevents long-lasting disability. Additionally, high earners may be less likely to file for disability benefits as their wage rates are higher, and their accounts may be larger due to more sustained work, hence gains from the insurance top-up are smaller. If high earners with white collar jobs were more likely to switch to the new system, this could result in an over-estimate of the new-system effect on hazard rates.

We do not know the previous wage levels of pensioners, most of whom no longer work. However, we use education as a proxy for permanent earning capacity. We experimented with several specifications, to test if differences in schooling levels are driving the new-system effect. First, we ran the hazard model with years of education as a co-variate and this turned out to be insignificant. Second, we estimated the model with a dummy for those who had a secondary degree or higher. This depresses hazard rates, as expected, but the effect is not significant and the other coefficients do not change. Third, we added an interaction term between secondary and system, and found that the new system effect is stronger for those with secondary degrees. However, this effect, too, is insignificant, while the main new-system effect remains strong and significant (col. 3, Table 7). Fourth, given that most disability pensioners in our sample (143 out of 172) do not have a secondary degree, we reran the model for the sub-sample with lower education, and again obtained the same results—a highly significant reform impact. We conclude that education, and selection based on education, are not driving these results.

Survivor bias and results with enlarged sub-sample

To see if these results are sensitive to our definition of sub-sample, we add cohorts born 1922-31, most of whom were in the old system (see Tables 3 and 4). The trade-off for the gains from sample enlargement is greater potential memory error by individuals about events from the distant past, missing time-specific variables that may be responsible for some of the change in disability rates, and survivor bias, which may lead to an underestimate of old system hazards and hazard reductions in the new system. If underlying health and safety conditions were worse in earlier years, the new-system advantage might appear to grow as earlier cohorts are added, but this would be an overestimate of the true advantage. However, this potential overestimate is offset by the fact that medical reasons for disability changed over time to include broader diagnoses such as psychiatric problems and back pains, which constitute an increasing proportion of the total and would be expected to add to total disability hazards, ceteris paribus.[15]

On survivor bias: The EPS is a retrospective survey, representative of the universe of affiliates in the new and old systems between 1982 and 2001. But affiliates who died before 1982 were not included. Disabled pensioners are likely to have higher death rates than non-disabled. In the U.S., for example, mortality rates for disabled pensioners ages 35-55 after 10 years of disability status are 3 times higher than for those who applied but were denied disability status (30% vs. 10%, Von Wachter, Song and Manchester 2007). If that is the case in Chile, the proportionate exclusion of the disabled due to death before 1982 would far exceed the proportionate exclusion of the at-risk group. All of these individuals were in the old system. Therefore, survivor bias would understate the size of the disability hazard in the old system and the hazard reduction in the new system. This understatement would apply particularly to cohorts born before 1932, whose members were more likely to die before 1982. Survivor bias is further compounded by the possibility that individuals who died after 1982 may also have been undercounted, as a result of practical difficulties in getting information about them in 2002.[16] To estimate these effects, we use mortality tables (RV98) from Chile to calculate the proportion of affiliates from various cohorts, alive at age 40, who were expected to die before 1982 and therefore never appeared in the sample frame, as well as those who died after 1982 but were unreported. The numbers are consistent with our expectation that understatement of disability hazards in the old system and hazard reduction from the new system are likely to be small in the main sample but much larger in the expanded sample.[17]

Indeed, when we ran the Cox model we found that baseline hazards obtained from the enlarged sample are lower than those obtained from the main sample. The impact of the reform also falls slightly. Nevertheless, it remains highly significant. The new system reduces disability hazards by 60-70% at the 1% significance level for ages 45-59 (marginally significant for ages 40-44) in the expanded sample, compared with a 65-80% reduction in our main analysis (Table 7, col. 4).

Testing for selection bias: choice versus no-choice groups

The biggest challenge to these results is the possibility of selection bias. Since many members of our sample chose their system, rather than being randomly assigned, it is possible that selection bias is present. Those who thought they had a higher probability of applying for disability benefits may have stayed in the old system because of the greater likelihood that their claim would be approved there and the greater uncertainty about how the new system would operate. Additionally, AFPs may have discouraged from membership those who they considered to be more prone to disability (James and Iglesias 2006). In that case, the observed reduction in disability hazards might be due to the fact that the new system attracted individuals who had a low risk a priori rather than to new-system procedures that cut hazards.[18]

We carried out several tests to investigate this possibility. Our negative findings for selection based on education were already discussed above. In this section, we draw upon the fact that our sample includes groups who had choice as well as some (pre-1982 observations and new entrants to the labor market post-1982) who had no choice. We derive hypotheses about what the relative disability hazards would be between the choice and no-choice groups in the old versus new systems if selection bias were driving our results. Tests of these hypotheses indicate that selection bias between systems is not a major factor; it does not invalidate our initial results.

We start by dividing our subjects into four groups: no-choice in old system (observed 1972-81), with choice in old system (observed in old system 1982-2002 but with some prior employment before 1982), with choice in new system (observed in new system 1982-2002 but with some prior employment before 1982) and no-choice in new system (entered labor market after 1982 and observed up to 2002) (see Tables 5 and 6).[19] Members of the no-choice groups were assigned to their respective systems, therefore pose no selection issues. By contrast, members of the choice groups chose their systems and therefore pose potential selection bias.

This disaggregation of our sample presents some problems. As we disaggregate by choice, we end up with a small number of disability pensioners in each group, especially in the no-choice group, which has only 28 pensioners between the two systems. In addition, we only observe ages 40-50 for the old system no-choice group; by the time they exceed age 50 it was already post-1982 and they had choice. Furthermore, members of the new-system no-choice group, who entered the labor market after 1982, are relatively young; this helps account for the small number of disability pensioners—most disability occurs after age 50. Finally, the no-choice people in the old system were born 1932-41 and are observed in the years 1972-81, in contrast to the no-choice people in the new system who were (almost) all born after 1941 and are observed after 1982. The absence of overlap in birth and observation years of the two systems means that a comparison of the two no-choice groups is more subject to survivor bias and other time-specific forces than is a comparison of the two systems as a whole. Indeed, these are the reasons why we didn't confine our entire analysis to the no-choice groups.

We proceed with hypotheses about the impact of selection bias. If individuals with higher disability propensities chose to stay in the old system after 1982, we should observe differences in disability hazards for the choice vs. no-choice groups in both systems, as well as differences in the new system impact on these groups and over time.

1. In the new system, age-specific disability hazards will be higher for individuals who entered the labor market after 1982 (no choice), compared with individuals who voluntarily switched to the new system in 1982. To test this hypothesis we apply the Cox model to new system observations, derive the baseline hazard for the no-choice group and estimate the impact of the co-variate, membership in the choice group. As in our previous regressions, we stratify by gender and marital status and control for unemployment rate. Contrary to the selection hypothesis, we find no difference in disability hazards between the choice and no-choice groups in the new system (Table 8, column 1).

2. In the old system, age-specific disability hazards will be lower for no-choice individuals who were in the old system before 1982, compared with those who voluntarily stayed in the old system after 1982. To test this, we apply the Cox model to old system observations, derive the baseline hazard for the no-choice group and estimate the impact of the co-variate, membership in the choice group. We carry out this analysis only for ages 40-50, since we have no observations beyond age 50. Although the hazard is higher for the choice group, this difference is not statistically significant (Table 8, column 2).

3. Differential new system effects for choice vs. no-choice groups. If selection bias is important, the reduction in disability hazards for new system members should appear greater for the choice groups than the no-choice groups—but this would be, in part, a consequence of selection rather than real cost-reduction. If the entire result is due to selection, the new system effect would disappear for the no-choice groups.

The results just depicted for tests #1 and #2 suggest that we will not find this differentiated new system effect, but we test this hypothesis directly, using the Cox model to estimate the old system baseline and new system effect, interacting new system with choice and thereby allowing the new system effect to vary between the two groups. Since we don't have observations in the old system no-choice group beyond age 50, the main terms after that age apply to the choice group only. The main term for ages 40-50 applies to both groups, and the interaction term modifies this for the choice group.

Results from test #2 suggest that we should not stratify the baseline by choice. In this case the main new system effect is highly significant, reducing disability hazards by 40-65% for both groups for ages 40-59, and the interaction term is insignificant—contrary to the selection hypothesis. If, in contrast, we do stratify the baseline by choice, the main terms remain significant and of roughly the same magnitude, but the interaction coefficient also becomes significant (apparently due to the insignificantly higher disability hazard for the choice group in the old system). For the choice group, disability hazards are reduced by 76% (=1-.57*.43) for ages 40-50 and by 60% for ages 51-59. The new system coefficient is smaller but still significant for the no-choice group, reducing disability hazards by 43% in the observed age range of 40-50. This suggests that some selection may be present, but selection does not account for the entire new system effect, which remains strong both in the choice and no-choice groups (Table 8, col. 3 and 4).

4. New system effect in early versus late periods for choice group. Workers are likely to have more information about their near-future disability propensities than their far-future propensities. Then, if selection created the reduced disability hazards, we would expect the new system effect to be concentrated in years closer to 1982. To test this hypothesis, we divided the period 1982-2002 into two sub-periods, 1982-91 and 1992-2002 and compared the old system baseline and new system coefficient for the choice group for these two sub-periods. For this analysis we exclude all observations prior to 1982 so the two systems co-exist throughout the years observed, and we include all individuals who were age 40 or over at some point within their respectively periods. We ran the Cox model for these observations, stratifying by period to allow the baseline to vary and interacting with period to allow the new system effect to vary. The main terms were highly significant for ages 40-59, indicating hazard reductions of 56-73% for both periods, and the interaction terms were insignificant—contrary to the selection hypothesis (Table 8, col. 5). We also re-ran our main regression omitting all observations during years 1982-87, which were most likely to be sensitive to selection bias. We got very similar results to those for the full set of years (Table 7, column 5).

Over-all, these tests indicate that some selection bias may be present but it does not account for most of the main new system effect.

Mortality rates as an indicator of system stringency and accuracy

To further determine whether differential assessment standards and countervailing information are responsible for the lower hazard rates, we sought direct evidence on severity of approved disability claims. If the new system got its lower hazard rates though selection of healthier workers as members, disabled pensioners in both systems would have a similar severity of disability but would simply be scarcer in the new system. In contrast, if the reduction in hazard rates were due to more stringent assessment procedures, disabled pensioners in the new system would tend to have more severe disabilities.

We use comparative mortality rates among pensioners in the new versus old systems as a test of whether the new procedures are causing the lower hazard rates and also as a test of the accuracy of the assessment standards (see von Wachter, Song and Manchester 2007 for use of mortality rates as a test of accuracy in the U.S. system). Of course, mortality rates are only one dimension of disability severity. If a constant standard of severity were being applied, we might find lower aggregate mortality rates among new system pensioners because of the shift in diagnoses toward psychiatric and other chronic conditions and away from cardiac conditions.[20] If, instead, mortality rates are higher in the new system, this suggests that the severity bar has been raised substantially, that benefits have been targeted toward the more medically disabled and hazard rates have been cut by rejecting the lesser end of the medically disabled distribution.

We start by using probit models to estimate the probability of death by 2002 of disabled pensioners and non-disabled affiliates in the old and new systems, controlling for age, gender, years of education and period of exposure to disabled status. Because those who died before 1982 are not included in the sample frame, we confine the analysis to those who pensioned after 1982. (This means that we omit members of the old system no-choice group from the mortality analysis and that years observed in the new and old system years are all overlapping). We caution that these results should be regarded as preliminary since the number of dead disability pensioners is small.

Not surprisingly, the probits show that the disabled have a (9 percentage point) higher probability of death by 2002 than the non-disabled. Every year of age increases the probability of death, women have lower mortality rates than men and more education means lower death rates. Those who pensioned after 1986 are less likely to have died by 2002 than the others because they have fewer years of exposure.[21] These effects are all significant at the 5% level or less (Table 9). Most important for our purposes, the dummy variable for new system has a negative (almost 4 percentage point) effect on mortality rates but the interaction between new system and disability pensioners has a large (13 percentage point) positive coefficient, significant at the 1% level, after controlling for these other factors. In other words, life survival for the non-disabled increases in the new system, possibly due to health improvements for later cohorts. However, life survival for disabled pensioners drops dramatically. The new-system coefficient for disability deaths is even larger (18 percentage points) when we run the same regression for men only.

To investigate this issue further, we develop Kaplan-Meier life-survival curves by age and Cox proportional hazard regressions of the hazard of dying, by year of exposure to the disability pensioner state. In this case, those "at risk' are all disability pensioners (disabled) or all other affiliates (nondisabled), the "failures" are those who die and the "survivors" are those who live. It turns out that, in the old system, age-specific Kaplan-Meier survival probabilities are practically the same for disability pensioners and other affiliates (Figure 5). Apparently "disability" does not target those with higher expected mortality rates in the old system. In contrast, a wide disparity opens up between survival rates of disabled and other affiliates in the new system. Apparently the new system does an effective job of sorting between disabled and non-disabled. The Cox model further shows that the hazard of dying within 3 years of disability pensioning is the same for the two systems, but the hazard of dying within the next 12 years is much greater for the new system, implying far lower life-survival probabilities (Table 10 and Figure 6).

Thus, the new system seems to reduce the incidence of successful disability claims by targeting those individuals with the most severe medical problems, who have much higher mortality rates than disabled pensioners in the old system or non-disabled affiliates in the new system.[22]

III. What Can Other Countries Learn from the Chilean Experiment?

Countries around the world are faced with rising costs of old age security programs. In many countries, disability expenditures are a high proportion of total social security costs and have been rising even faster than old age expenditures. The Chilean experience shows that costs can be contained, quite drastically, by adopting procedures and incentives that favor this goal. But it also raises questions about the degree to which this is socially desirable.

The Chilean system for disability insurance has two innovative features that help contain costs: it is pre-funded (which reduces fees in the long run; see James and Iglesias 2006) and it utilizes private pecuniary incentives to dampen successful claims (on which this paper has focused). Pre-funding takes place in two stages: first, building the retirement accounts through the worker's career and second, using an "additional payment" when he becomes disabled to enable the purchase of a lifetime defined benefit.

AFPs charge an insurance fee that covers the "additional payment" and have an incentive to keep costs low by controlling successful claims.

The assessment process includes participation by private AFPs and insurance companies and enables them to pursue this objective. We hypothesize that the pressure they create toward strict medical criteria and tight application of these criteria, as well as their right to present countervailing information and appeal initial evaluations, have the effect of reducing the incidence of approved disability cases. Our calculations of hazard and survival rates, using the Cox proportional hazard model and a retrospective data set of new and old system affiliates (EPS 2002) show significantly lower hazards of becoming disability pensioners in the new system. These hazards are cut by 65-80% compared with the old system, after controlling for other variables that might affect outcomes. Tests for selection bias, largely based on analysis of choice versus no-choice groups in the two systems, rule this out as the major explanation. Comparisons of mortality rates among new and old system disability pensioners suggest that this reduction in disability hazard is achieved by targeting those with the most severe medical conditions: life survival rises in the new system for non-disabled affiliates, but falls for the disabled.

A reduction in approved disability claims would not be socially desirable if those who are rejected are indeed disabled. Similarly, an increase in approved claims would not be socially desirable if those who receive benefit are not disabled. Our analysis indicates that the new system has cut costs by effectively sorting between those who are more versus less disabled, as measured by relative mortality rates, whereas the old system did not. However, in reality disability is hard to define and probably consists of a continuum with many dimensions other than expected mortality rates. Value judgments are involved in where to draw the line and by which definition of disability. The very low hazard rates in Chile's new system suggest it has chosen to emphasize hard conditions and to minimize type 1 errors (false positives) at the possible expense of more type 2 errors (false negatives), while the old system made the opposite choices.

To avoid the false negatives, some societies might wish to grant disability insurance benefits more liberally, even though this will cost more and may involve more false positives. In fact, Chile has made a move in that direction in its 2008 reforms.[23] Other countries, however, consider their current disability hazard rates and costs excessive, in the sense that they impose heavy insurance fees on non-disabled workers. The main lesson they might learn from Chile is that costs can be contained by introducing stringent procedures, countervailing information and incentives with this goal. These include the development of careful definitions of disability that are uniformly followed, the possible participation of private organizations, allowing private organizations and public agencies to appeal, and permitting them to be represented by third parties (such as lawyers) who have a financial interest in winning.[24] This would increase the probability that information favoring both sides would be forcefully presented, which might save money and lead to greater accuracy at the same time.

Table 1: Inflow to disability benefit status, Chile vs. U.S. and OECD, 1999
(new inflow, per thousand in insured population)

Age group	20-34	35-44	45-54	55-59	60-64
Chile	.2	.9	2.9	7.2	12.3
US	2.7	4.5	7.8	13.9	12.8
OECD	2.3	4.2	8.6	14.9	14.1

Source: OECD data from OECD (2003), p. 81
Chilean data calculated by authors from claims and assessment data supplied by Association of AFPs, contributor and member data supplied by SAFP. Only disabled who are insured are included here—in 1999 this was about 70% of those who were granted disabled status in Chile. Inflow to temporary disability status is given; inflow to permanent disability status would be about 75% as large, depending on age. Ratios are given as % of [(members + contributors)/2] since insured population includes some affiliates who are not currently contributing.
OECD numbers are newly disabled beneficiaries as % of (population in the relevant age group, minus the stock of people in that age group who are already on disability benefits). The denominator includes some people who are not eligible for insurance. If this definition were used for Chile, Chile's disability rate would be lower than given here.

Table 2: Main sample composition by old vs. new systems, 1932-62 cohorts > age 40

	Number of identities			**Median age at exit**[a]
Affiliation	Non-switcher	Switcher[b]	All[b]	All
Old System	1,840	606	2,446	52.5
New System	2,988	2,890	5,878	48
Total	4,828	3,496	8,324	49

[a] in 2002 or year of pension or death if sooner, or 1982 for old-system switchers
[b] includes duplicate identities for 606 individuals > age 40 who were in old system pre-1982 but switched to new system in 1982; they are considered switchers in both systems.

Table 3: Sample composition by birth cohorts, 1922-62[a]

Year of birth	Age-1982	Age-2002	Ages observed	Years included	Old-system	New-system	Total[b]
1922-31	51-60	71-80	40-64	1962-1995	1079	220	1299
1932-36	46-50	66-70	40-64	1972-2000	671	272	943
1937-41	41-45	61-65	40-64	1977-2001	768	464	1,232
1942-46	36-40	56-60	40-60	1982-2002	426	687	1,113
1947-51	31-35	51-55	40-55	1987-2002	263	985	1,248
1952-56	26-30	46-50	40-50	1992-2002	174	1,287	1,461
1957-62	20-25	40-45	40-45	1997-2002	144	2,183	2,327
1932-62	20-50	40-70	40-64	1972-2002	2,446	5,878	8,324
1922-62	20-60	40-80	40-64	1962-2002	3525	6098	9623

[a] Main sample contains 1932-62 cohorts. Expanded sample adds 1922-31 cohorts.
[b] includes duplicate identities for individuals > age 40 who were in old system pre-1982 but switched to new system—606 for 1932-62 sample and 765 for 1922-62 sample.

Table 4: Sample of disability pensioners[a]

Birth year	Old	New	Total	Pension year	Old	New	Total	Pension age	Old	New	Total
1922-31	40	5	45	**Pre-1979**		0					
1932-36	35	5	40	**1979-81**	10	0	10	**40-44**	16	12	28
1937-41	34	13	47	**1982-86**	17	0	17	**45-49**	24	10	34
1942-46	28	14	42	**1987-91**	21	4	25	**50-54**	30	15	45
1947-51	14	11	25	**1992-96**	24	15	39	**55-59**	30	11	41
1952-56	4	12	16	**1997-01**	38	31	69	**60-64**	16	8	24
1957-62	1	1	2	**2002**	6	6	12				
1932-62	116	56	172	**1979-02**	116	56	172	**40-64**	116	56	172
1922-62	156	61	217								

[a] Numbers by pension year and pension age are given for 1932-62 cohorts (main sample)

Table 5: Main sample composition by choice vs. no-choice, 1932-62 cohorts > age 40

	Old system	New system	Total
No-choice	1570[a]	2988[c]	4558
Choice	1788[b]	2890[d]	4678
Total	3358	5878	9236[e]
Dual identities			1518

[a] Includes all old system members > age 40 before 1982--964 who remained in old system plus 606 who switched to new system after 1982. Dual identities were created in choice groups post-1982 for these individuals minus 52 who remained in old system because they became pensioners before 1982. (An additional 876+2284 = 3160 old system members pre-1982 were < age 40 in 1982 and appear only in the post-82 choice group sample; see notes b and d).

[b] Includes 964 members > 40 in 1982 who remained in old system post-1982 + 876 members < age 40 in 1982 who remained in old system – 52 who pensioned pre-1982.

[c] Corresponds to all lifetime new system members (entered labor force after 1982).

[d] Corresponds to all who switched out of old system into new system in 1982; this includes 606 who were > 40 in 1982 + 2284 who were < 40 in 1982.

[e] This total is greater than total in Table 2 because dual identity is created for 912 old system members who were in no-choice group before 1982, choice group after 1982 (912 = 962 members – 52 pensioners pre-1982)

Table 6: New and old system disability pensioners by choice vs. no-choice, 1932-2002 cohorts > age 40

	Choice	No-choice	Total
Old system	102	14	116
New system	42	14	56
Total	144	28	172

Table 7: Determinants of Hazard of Disability Pension for 1932-62 cohorts
(Cox Proportional Hazard Stratified by Sex and Marital Status; numbers given are hazard ratios relative to baseline)

Covariate and segment of hazard over which effect applies	1932-62 cohorts (1)	1932-62 cohorts (2)	1932-62 cohorts (3)	1922-62 cohorts (4)	Exclude 1982-87 (5)
New System: Age 40 to 44	0.355 (-2.71)*	0.337 (-2.88)*	.425 (-2.07)**	0.497 (-1.83)***	0.579 (-1.24)
Age 45 to 49	0.21 (-4.12)*	0.213 (-4.1)*	.157 (-3.77)*	0.300 (-3.24)*	0.145 (-4.41)*
Age 50 to 54	0.325 (-3.5)*	0.307 (-3.71)*	.369 (-2.82)*	0.383 (-3.14)*	0.368 (-2.64)*
Age 55 to 59	0.296 (-3.23)*	0.317 (-3.25)*	.326 (-2.78)*	0.397 (-2.87)*	0.299 (-3.18)*
Age 60 to 64 (males only)	0.493 (-1.63)***	0.553 (-1.33)	.546 (-1.34)	0.638 (-1.24)	0.493 (-1.63)***
Unemployment: Age 50-54	1.054 (1.84)***		1.055 (1.86)***	1.010 (0.42)	1.111 (1.99)**
Age 55-59	1.093 (1.98)**		1.095 (2.02)**	0.980 (-0.87)	1.094 (1.98)**
Secondary degree			.795 (-0.83)		
Sec*newsys40-44			.633 (-0.64)		
Sec*newsys45-49			2.360 (1.24)		
Sec*newsys50-54			.677 (-0.46)		
Sec*newsys55-59			.721 (-0.39)		
Sec*newsys60-64			.602 (-0.46)		
# observations	8324	8324	8324	9623	8217

* significant at 1% level or less
** significant at greater than 1% but less than 5% level
*** marginally significant at less than 10% level
Notes: Baseline differs by sex and marital status, but impact of co-variates relative to baseline is same for all. Numbers in parentheses are z-statistics.

Table 8: Disability hazards: Selection bias tests based on choice and no-choice groups

Age segment of hazard	Effect on hazard of being in choice group		Effect of hazard in new system relative to old system baseline, for various groups		
	New system	Old system	Choice vs. no-choice, not stratified	Choice vs. no-choice, stratified	1982-91 vs. 1992-2002, choice group
40-50	1.27 (.58)	1.55 (1.29)	0.58 (-3.28)*	0.57 (-2.65)*	0.27 (-3.53)*
51-59	1.93 (1.16)	na	0.35 (-3.91)*	0.41 (-3.12)*	0.44 (-2.65)*
60-64	na	na	0.5 (-1.61)***	0.64 (-1.04)	0.63 (-1.06)
40-50* choice	na	na	0.67 (-1.23)	0.43 (-2.34)**	na
40-50* 1982-91	na	na	na	na	0.59 (-.8)
51-60* 1982-91	na	na	na	na	0.32 (-1.03)
#individuals observed	5878	3358	9236	9236	4678
#pensioners observed	56	116	172	172	144

Column 1 gives effect of co-variate—being in choice group--relative to baseline hazards for no-choice group (for new system members only).
Column 2 gives same for old system members only.
Column 3 gives effect of co-variate—being in new system—relative to baseline hazards for old system (with interaction term for choice group, not stratified by choice). Ages 51-64 applies to choice group only, since we do not have old-system no-choice observations in this age range. Ages 40-50 apply to both groups. 40-50*choice interacts age with choice.
Column 4 gives same, stratified by choice. Since interaction terms is significant, coefficient for choice group ages 40-50 = .57*.43 = .25.
Column 5 gives effect of co-variate—being in new system—relative to baseline hazards for old system (with interaction term for early period, 1982-91, stratified by period)
* significant at 1% level or less
** significant at greater than 1% but less than 5% level
*** marginally significant at less than 11% level

Note: All regressions include unemployment rates. Baseline differs by sex and marital status, but impact of co-variate relative to baseline does not vary by gender or marital status. Numbers in parentheses are z-statistics.

Table 9: Probit Analysis of Mortality Rates in New vs. Old Systems (change in probability of dying by 2002 in percentage points)

Probit regression, reporting marginal effects

	Men + women		Men only	
	dF/dx	z	dF/dx	z
Age-2002	0.06	2.65*	0.06	1.58
Yrs ed	-0.14	-3.38*	-0.27	-4.01*
female	-2.22	-6.06*		
Disability	9.02	2.08**	16.56	2.4**
Dis86	-2.65	-2.03**	-3.81	-2.09**
Dis90	7.6	1.15	13.43	1.33
Dis94	-2.1	-1.24	-3.26	-1.48
Dis98	6.11	1.5	5.06	1.03
New sys	-3.73	-7.08*	-5.10	-6.07*
Dis*newsys	12.78	3.15*	17.79	3.06*
# observations	7674		4429	
Pseudo R2	0.085		0.078	
Prob>chi2	0.00		.00	
Obs P	3.47		4.54	
Pred P (at x-bar)	2.58		3.63	

* significant at 1% level or less

** significant at greater than 1% but less than 5% level

Dis86 = started disability pension 1986 or after, relative to 1982-85

Dis90 = incremental effect of starting disability pension 1990 or after; this is added to Dis86 to get total effect for pensioning 1990 or after relative to 1982-85. And so on.

New sys = new system affiliate

Dis*newsys = disability pensioner in new system

Table 10: Estimates of Determinants of Hazard of Death after Disability Pension

(Cox Proportional Hazard Model; hazard ratios relative to old-system baseline)

Covariate: Segment of the hazard over which effect applies	1932-62 cohorts[a] Full Sample Baseline Differs by Sex	1932-62 cohorts[a] Men Only
New System: 1-3 years after pension	0.72 (-0.40)	1.04 (0.04)
4-7 years after pension	10.21 (2.07)**	7.82 (1.78)***
8-15 years after pension	6.75 (1.90)***	6.75 (1.90)***
#observations	172	120

* significant at 1% level or less

** significant at greater than 1% but less than 5% level

*** marginally significant at less than 10% level

Numbers in parentheses are z-statistics.

[a] for group that started pension 1982 or afterwards

Figure 1: Kaplan-Meier survival as non-disability-pensioner, by system

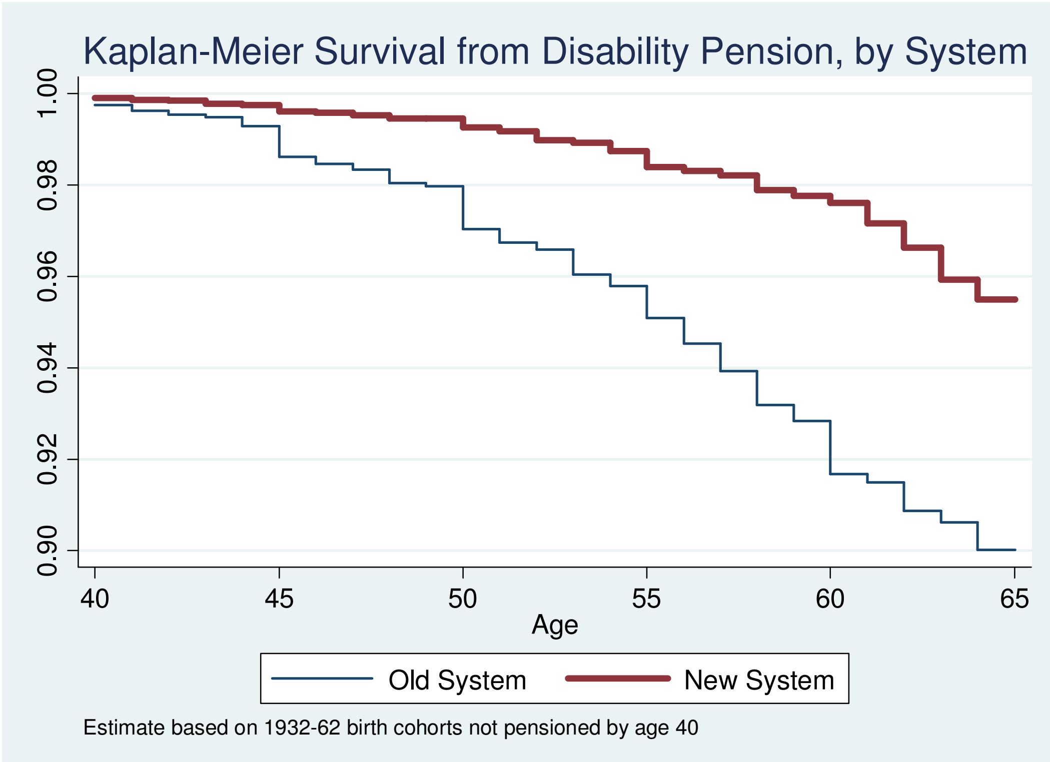


Figure 2: Baseline disability hazards by gender and marital status

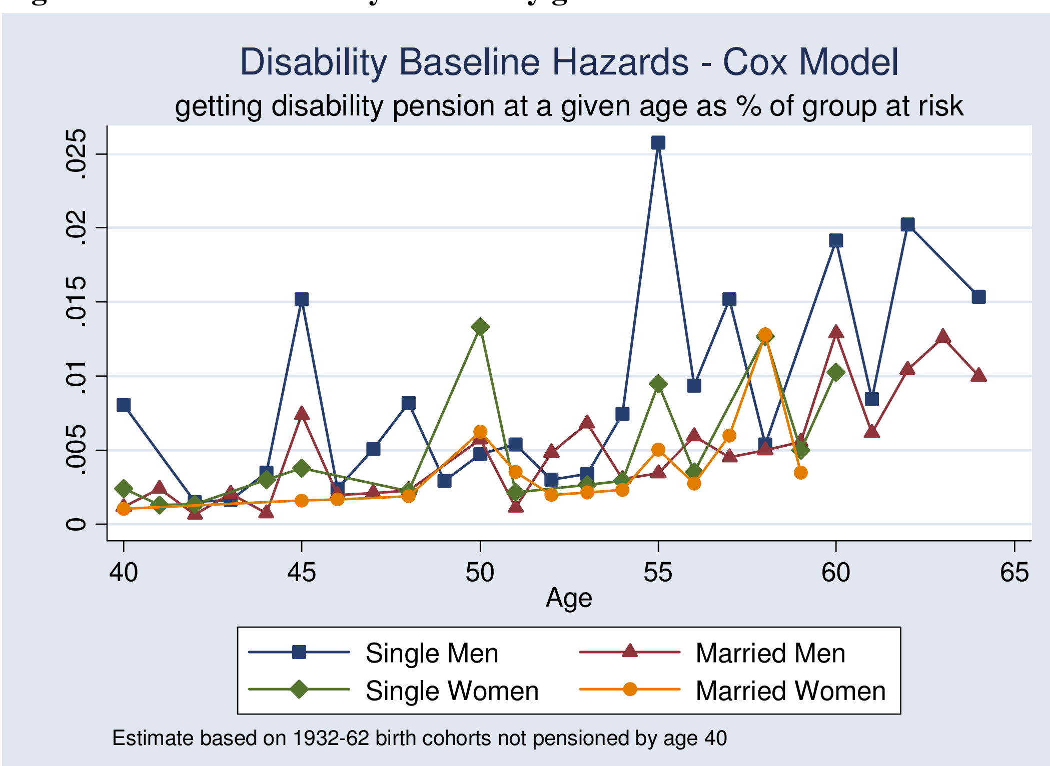


Figure 3: New vs. old system disability hazards for married men

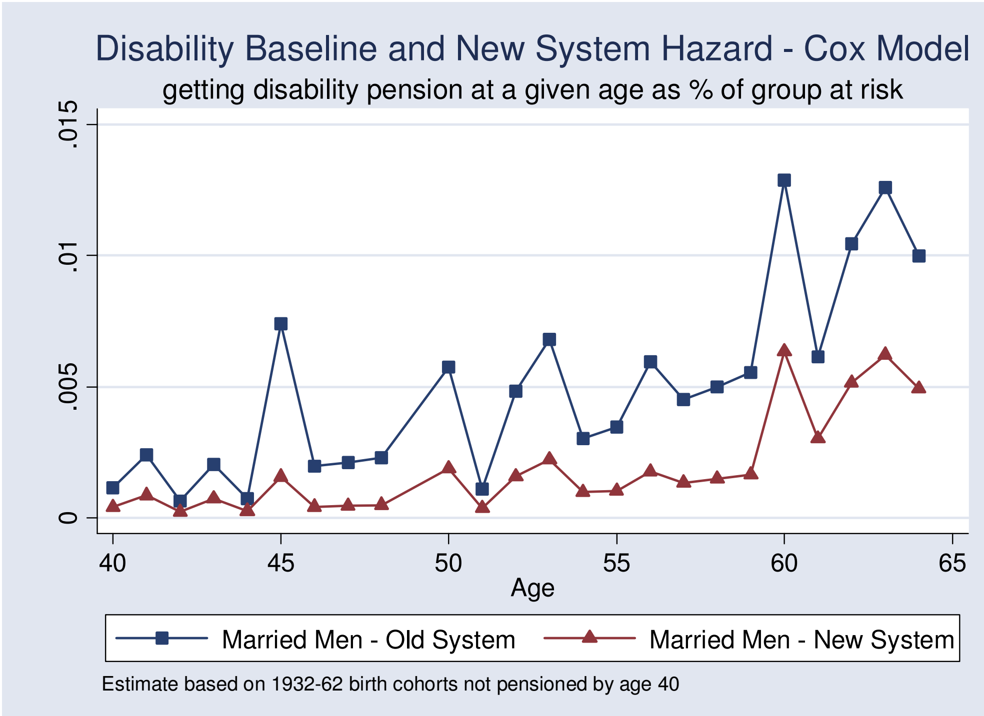


Figure 4: Survival as non-disability pensioner, married men—Cox model

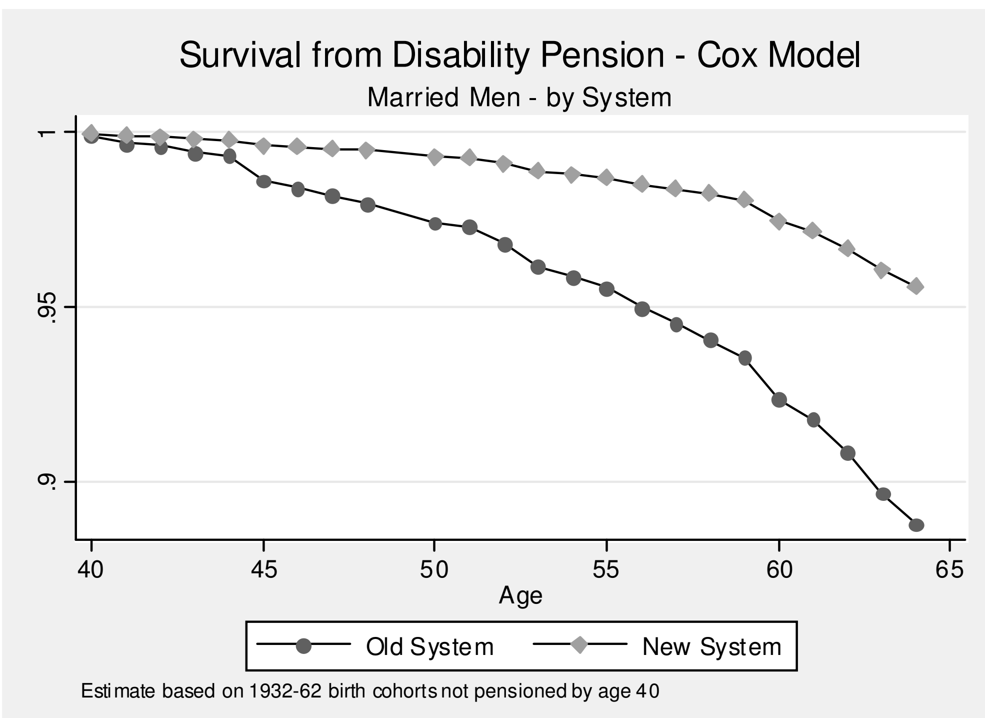


Figure 5: Kaplan-Meier: Survival rates (not dying) by disability status and system

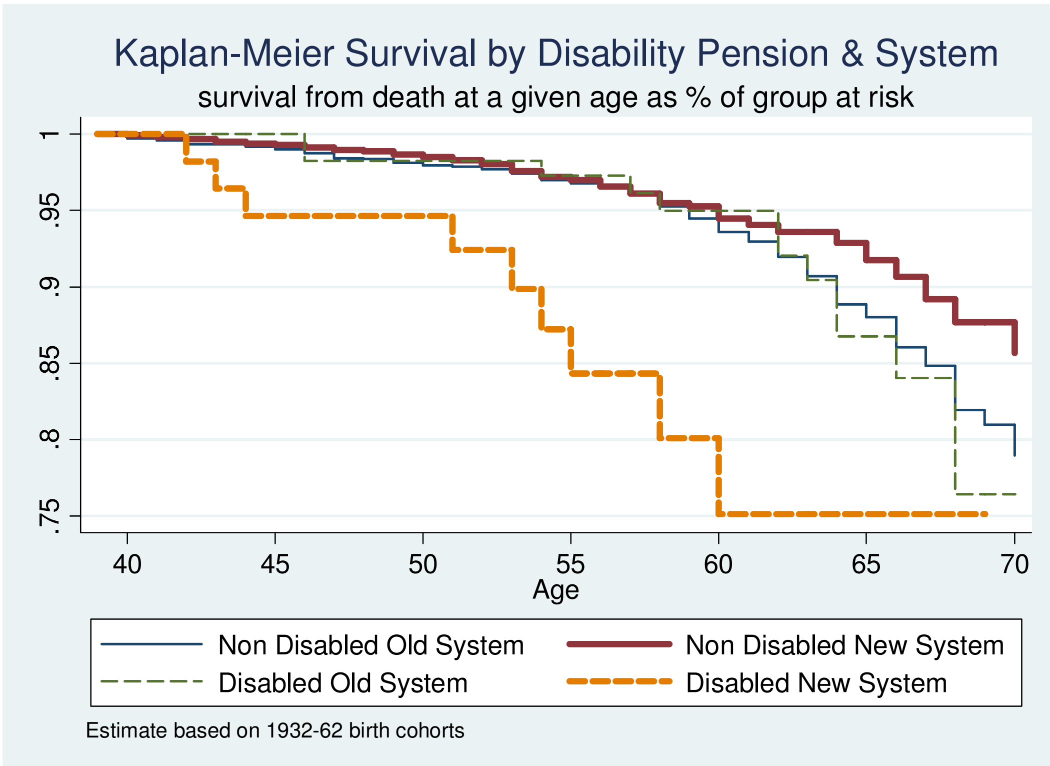
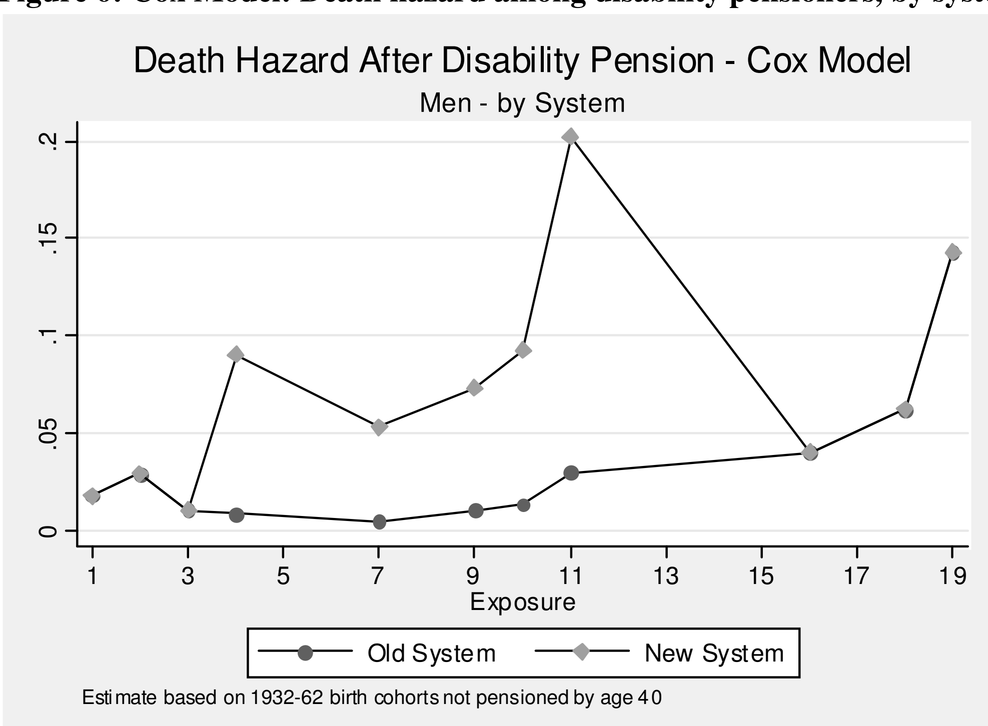


Figure 6: Cox Model: Death hazard among disability pensioners, by system



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Footnotes

[1] International comparisons of D&S costs are fraught with difficulties. D&S benefits in Chile and some other Latin American countries are financed partially by government payments through the minimum pension guarantee, in addition to the private insurance fees. Most European countries also provide minimum pensions, which are separately financed. The insurance top-up in Chile is supplemental to the money in the worker's own account, which helps cover disability pensions. Countries with pay-as-you-go systems do not have such account balances, which serve triple duty as part of old age, disability and survivors' insurance. In Chile the annuity financed by the disability insurance fee plus account balance covers the lifetime of the individual and his survivors, while in the U.S. and some other countries it only covers the individual until age 65, at which point old age insurance takes over.

[2] For previous discussions of disability insurance in Chile and other countries with individual accounts see James and Iglesias 2006, James, Iglesias and Edwards 2008, Grushka and Demarco 2003, Castro 2005 Wiese 2006, Valdes and Navarro 1992.

[3] Annuities last the entire lifetime, thereby providing longevity insurance. Programmed withdrawals have the same expected present value as annuities. They do not provide longevity insurance but do give the worker bequest rights over the balance in the account when he dies. Currently, 40 percent of disabled pensioners have annuitized, compared with 65 percent of retirement pensioners. See James, Martinez and Iglesias 2006, Edwards and James 2006. For an analysis of the impact of pre-funding on long run vs. short run costs and on sensitivity to changes in interest rates and demographic structure of the work force, see James and Iglesias 2006, James, Iglesias and Edwards 2008.

[4] Cross-subsidization and incentives for selection by AFP's that result from the mandate of uniform percentage insurance fees are discussed in James and Iglesias 2006 and James, Iglesias and Edwards 2008. As part of the 2008 reforms, a portion of this fee will be rebated to the accounts of women, who have lower disability and survivors' risks. This same fee also covers benefits paid by AFPs during the three years of temporary disability.

[5] In the U.S., appeals can only be brought by workers whose initial claims have been denied, so appeals inevitably increase approved cases. Attorneys who specialize in disability cases often represent applicants in appeals. In 2000, only 38 percent of claims were approved initially, but the majority of those denied benefits appealed and more than half of all appeals eventually won. Therefore, 55 percent of all claims were eventually accepted. (Social Security Advisory Board 2001, pp. 8, 18, 19; Autor and Duggan 2006).

[6] The individual must 1) be working and contributing at the time of the claim, or 2) have contributed during the last 12 months and also paid at least 6 contributions in the year immediately preceding the last registered contribution. 3) Self-employed workers must have paid at least one contribution in the calendar month before the date of the claim. 4) In addition, he must not be a pensioner or be over the normal retirement age (65M/60W). These conditions are lighter than those in other countries with contributory schemes. For example, 3 years of contributions are typically required in Latin America, 5 years in

OECD countries (OECD 2003, Grushka and Demarco 2003, Andrews 1999). In the U.S. the applicant must have worked in 5 of the last 10 years and cannot be working currently.

[7] Poor record-keeping by public agencies in many countries has made it difficult to enforce insurance eligibility requirements. This was true of the old system in Chile, which allowed workers to become eligible based on contributions in the distant past. For example, in the SSS (the largest program), workers needed to have at least 50 weeks of contributions in the past and a 50% density of contributions over the entire membership period. Young workers could qualify more readily in the new system, but older workers who fulfilled the requirements and then dropped out of employment could qualify more readily in the old system. Assertions by older workers of long-past service were difficult to refute. The new system's emphasis on the recent past made it easier for AFPs to prove that workers were ineligible for insurance. Workers who are certified as disabled based on medical criteria but do not meet the eligibility conditions for the insurance top-up can withdraw their own money as an annuity or programmed withdrawal.

[8] The average density of contributions in Chile—that is, the portion of his working life that an average worker contributes—is about 60% (Alberto Arenas, Jere Behrman and David Bravo 2004; Solange Berstein, Guillermo Larrain and Francisco Pino 2005; Alberto Arenas et al 2007). A person with this average density will therefore get a defined benefit that is 60% of 70% = 42% This de facto 42 percent replacement rate is comparable with the disability replacement rate in many OECD countries, although lower than in the Netherlands or Sweden and higher than the U.S. or U.K. See Andrews 1999 and U.S. Social Security Advisory Board 2001 for numbers in other countries.

[9] To be eligible for the MPG, the disabled pensioner must have: 1) at least 10 years' contributions in the social security system, or 2) at least two years' contributions in the last 5 years prior to the disability claim, or 3) 16 months contributions if he has joined the labor force within the last 2 years, or 4) been contributing at the date of disability, if this was caused by an accident. If the individual has other sources of income this may invalidate his eligibility for the MPG; but we don't know if this means-test is enforced.

[10] The survey tells us if new system affiliates in 2002 had previously belonged to the old system, but it does not record years of switching. We know from other sources that most switching occurred between 1982 and 1985. For purposes of this analysis, we assume that all switching took place at the end of 1982. The survey also doesn't distinguish between the partially versus fully disabled or between those with or without insurance.

[11] Respondents were asked whether they received a pension and, if so, what kind (old age, disability, etc.). They were also asked who paid the pension. Payment by AFPs or insurance companies comprised the new system, by INP comprised the old system. This analysis focuses on "normal" disability pensioners rather than those disabled by labor accidents, most of whom are covered by separate insurance payers (Mutuales) with different procedures. However, some victims of labor accidents are covered by INP and we could not separate them from normal disability pensioners in EPS. According to aggregate data, in 2006 1.8% of all disabled pensions awarded by the INP were for labor

accidents; on a stock basis this ratio was 4.8% (Association of AFPs; SAFP; Boletin Estadistico 2006, Superintendencia de Seguridad Social). It is also possible that some additional disabled pensioners should have been categorized as labor accidents but were instead mis-categorized as "normal." This implies a possible overestimate of 3-6% of normal disability pensioners in the old system, due to the inclusion of labor accidents. This possible overestimate is less than the likely underestimate of old-system disabled due to survival bias (see below) and small relative to the estimated new-system effects.

[12] The initial "at risk" group at the start of age 40 is the total number of individuals who were in the system at 40 and had not yet pensioned. Individuals who take a new disability pension during age 40 are considered a "failure" and the survival probability for age 40 = $1 - (failures_{40}/atrisk_{40})$. The "at risk" group at the start of 41 is the total number of individuals who were in the system at 41 and had not yet pensioned. This equals the number at risk at 40 - failures at 40 - dropouts for an old age pension or death between 40 and 41 + new labor market entrants between 40 and 41. The survival rate at 41 = $1 - (failures_{41}/atrisk_{41})$, the cumulative survival rate at 41 = $CumS_{40} * S_{41}$ and so on.

[13] The ratio of stock of disabled pensioners to old age pensioners was 30 percent in the old system in 2005 but only one-third of that--10 percent--in the new system in 2003 (INP 2006 and SAFP 2003).

[14] Neither differential eligibility conditions nor differential benefits could be producing these differences. As noted earlier, eligibility conditions in the old system favored older workers with more contributory years in the distant past, while the new system favored younger workers with more recent contributions. This was also true of the benefit formula. In the largest old system, the SSS, disability benefits were "equal to the retirement pension". The retirement pension, in turn, was based on number of working years—50 percent for first ten years + 1 percent for every year after ten up to 70 percent. This old-system formula is more generous than the new-system for workers who have 30 years of service but have not contributed regularly during the past ten years, while the new system formula is more generous for workers who have less than 30 years' service but have contributed regularly during the past ten years. If these eligibility and benefit differences were the driving force behind the fall in disability hazards, we would expect the decline to occur mainly in the older age groups. But this was not the case.

[15] Between 1995 and 2004, the percentage of permanent disabilities accounted for by psychiatric disorders increased from 12 to 20 percent of the total in Chile, while cardiac disorders moved in the opposite direction—from 19 percent to 12 percent. (Data provided by Association of AFP's). A similar change has occurred in the U.S. and other countries.

[16] Information in EPS about individuals who died between 1982 and 2002 was obtained from system records and relatives. Undercounting might have occurred if records were incomplete and relatives unavailable for some of these individuals.

[17] We estimate that missing dead affiliates (as a proportion of those alive in 1982) add up to 3 percent for the 1932-41 cohorts vs. 8 percent for the 1922-31 cohorts. Chile does not

have up-to-date mortality tables for the disabled. If the expected mortality rate among disabled is double that of the average affiliate, then the missing disabled (as a proportion of those alive in 1982) would be 9 percent for the 1932-41 cohorts vs. 24 percent for the 1922-31 cohorts. The missing dead percentage are much greater for the disabled than for the at risk group as a whole, and for the 1922-31 cohorts as compared with later cohorts.

[18] As a counteracting force: while collar workers (such as government employees) had generous programs in the old system, which may have made them less likely to switch than blue collar workers.

[19] For the 912 individuals who were observed working in the old system both before and after 1982 we create dual identities in order to distinguish their behavior when they were there by choice versus mandate. Thus, Table 5 displays 9236 identities, although Table 2, which did not distinguish between choice and no-choice groups, had only 8324 identities.

[20] In 1985 the 4-year mortality rate in the US for disability pensioners as a whole was 22 percent, but for mental disorders it was only 5 percent (Duggan and Imberman 2006).

[21] To proxy years of exposure, we include 4 dummy variables interacted with disability pension, indicating the years when the pension started: Dis86, Dis90, Dis94 and Dis98, with 1982-85 as the omitted period. We expect a negative sign (fewer deaths) for later periods because they indicate fewer years of exposure to disability status. These dummy interactions measure marginal effects. The first term turns out to be negative and highly significant; those who pensioned after 1985 are less likely to have died by 2002 than those who pensioned before. However, the coefficients on subsequent dummies are insignificant—no further incremental distinctions after 1986. This suggests a possible offset by less survival bias or by the increasing share of new system pensioners.

[22] It would have been useful to compare mortality rates of new-system disability pensioners versus those who applied for disability benefits and were rejected. This would have allowed us to determine whether the sorting occurred in the application or adjudication procedure. It also would have been useful to observe the employment and income histories of rejected applicants, to see whether they were capable of continued work. Unfortunately, this was not possible since ESP2002 did not inquire whether individuals had applied for disability benefits and were rejected and the pension system also does not keep track of the identity of such individuals.

[23] Changes introduced to the disability system as part of Chile's 2008 pension reforms indicate a greater concern about false negatives and may shift the balance away from the current emphasis on avoiding false positives. For example, disability claimants in the future may be represented by a doctor who is publicly financed. Disability insurance policies will be pooled across AFPs, which might reduce their incentive to oppose claims.

[24] Recommendations along these lines have been made by the U.S. Social Security Advisory Board 2001 and Autor and Duggan 2006.