Job Security and the Age-Composition of Employment: Evidence from Chile

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Abstract

This paper develops and tests a mechanism by which job security affects the age-composition of employment. This mechanism is based on the relative costs of dismissing young versus older workers resulting from job security provisions that are related to tenure. Using 39 consecutive annual household-surveys from Chile, we find that job security is associated with a substantial decline in the wage employment-to-population rate of young workers. In contrast, we do not find such a decline in young self-employment rates or in the wage employment rates of older workers. Our results also indicate that the negative effect on youth wage employment is driven by the slope of the severance pay-tenure profile. Regarding aggregate employment rates, we find that a raise in tenure-based severance pay reduces long-run employment, while a raise in a flat severance pay would marginally increase it.

Journal of Economic Literature classification: E24, J23, J65

1 Introduction

Economists have often criticized job security provisions for their negative impact on labor market performance. Although the majority of studies have focused on the impact of such policies on employment and turnover rates, some evidence suggests that job security might also have a large impact on the age-composition of employment[1]. Lazear [1990], for instance, finds that across the OECD, high job security countries like Spain, Italy, and Greece have lower youth employment-to-population rates than low job security countries like the U.K. and the U.S. More recently, Nickell [1997] finds some additional support for a prime-age employment bias in high job security countries. Although middle-age and older workers may benefit, reduced employment opportunities for the young are a source of concern. Young unemployed or discouraged workers may feel excluded from a society that does not allow them to participate in a gainful way. At worse, they may turn their energy into illegal or violent activities that impose large costs on society as a whole.

In this paper we explore and test the channels by which job security may bias employment against young workers. Lazear [1990] proposed an explanation based on a "barrier-to-entry " effect; by reducing hiring and firing rates, job security protects workers that are already employed, but it reduces the employment opportunities of new entrants. However, lower firing and hiring rates may have undetermined effects on young employment rates, depending on which effect dominates. Therefore, a barrier-to-entry effect, by itself, may not explain the association between job security and the age-composition of employment.

We propose an alternative mechanism based on the explicit link between severance pay and tenure observed in the majority of countries in which job security is mandatory; as severance pay increases in tenure, and tenure tends to increase with age, older workers become more costly to dismiss than younger ones. If wages and other labor costs do not adjust accordingly, negative shocks may result in a disproportionate share of layoffs among young workers. By reducing hiring rates and increasing youth layoff rates –relative to other age groups–, tenure-based job security may result in lower young to older workers

[1] At the theoretical level see, among others, Bertola [1990], Hamermesh[1993] and Nickell [1986] for a partial equilibrium analysis and Hopenhayn and Rogerson [1993] for a general equilibrium analysis of the impact of job security on employment and turnover. At the empirical level, see Anderson [1993], Grubbs and Wells [1993], Hamermesh [1995, 1996a, 1996b], Lazear [1990] and Nickel [1997].

employment rates.

We develop this argument in the context of a simple model and test these predictions using the remarkable time variability experienced by the Chilean labor legislation. Chile constitutes an ideal social laboratory in which to test the impact of job security for a variety of reasons: First, during the period 1960-1998, Chile went through six labor market reforms, evolving from a situation of dismissal-at-will to a highly regulated labor market by OECD standards.[2] Second, the Chilean labor market reforms have involved shifts in the relative costs of dismissing short versus long-tenure workers providing the variability required to test the mechanisms developed in the theoretical model. Third, job security provisions in Chile are very similar to the ones observed in most OECD countries. This implies the Chilean case could offer valuable lessons for other countries. Fourth, some of the labor reforms were associated to periods of drastic changes in political regimes –i.e. from democracy to dictatorship and vice-versa– increasing the likelihood that labor reforms were exogenous to contemporaneous labor market conditions. Lastly, Chile has a large formal sector by Latin American standards which facilitates the task of measuring the impact of regulations on labor market performance.

We use the wage and employment information contained in 39 consecutive household surveys spanning the period 1960-1998, and find a substantially different impact of job security across young, middle-age and older workers. As predicted, an increase in job security leads to a decline in the wage employment-to-population rates of young workers. In contrast, we do not find such a decline in young self-employment –where job security regulations do not apply– or in the wage-employment rates of older workers. The effect on the wage-employment rates of young workers is high; a 100% increase in severance-pay is associated with a contemporaneous 1.6 percentage point decline in youth employment-to-population rate and a 1.3 percentage point decline in participation rates. After 10 years, the employment decline would have surpassed 15 percentage points. As indicated by these magnitudes, the decline in youth employment rates is mostly absorbed by a decline in

[2] Marquez (1997) computes a cross-country ordinal index of labor market rigidity for Latin America, based on a methodology used by the OECD. He finds that most Latin American countries are more "rigid" than the majority of OECD countries. Thus for example, while Chile stands as an intermediate country by Latin American standards, it ranks on a par with Greece or Spain, the most rigid countries in the OECD sample.

participation, with few or no effect on youth unemployment rates.

When we decompose the effect of job security between the effect of changes in severance pay –keeping the slope of the severance pay-tenure profile constant– and the effect of changes in the slope, we find that job security affects young employment rates through its link to tenure. Thus, a high relative cost of dismissing long versus short-tenure workers is associated with lower youth employment-to-population rates. In contrast, our results indicate that a raise in a flat severance pay (i.e. not linked to tenure) would have no effect on young employment rates.

We also find interesting results regarding the impact of job security on aggregate employment rates: while a raise in tenure-based severance pay reduces long-run employment a raise in a flat pay would marginally increase it.

The rest of the paper is organized as follows: In section 2, we develop a partial equilibrium model to assess the impact of tenure-based job security on firms' hiring and firing decisions. In Section 3, we discuss labor reforms in Chile and construct a summary index of labor legislation. In addition, we describe the data, the empirical methodology and present our empirical results. Finally, in Section 4, we conclude.

2 The Model

In this section we present a simple model to analyze firms' firing and hiring decisions in environments where mandated severance pay increases with a worker' tenure at a firm.

Assume a continuum of firms operating a technology that has constant returns to scale and use labor as the only factor of production. Each worker has a per-period productivity equal to θ_i where the subscript i identifies whether the economy is in a good ($i = g$) or a bad ($i = b$) state and ρ is the probability that the economy remains in a given state in the following period. Consequently, we assume that $\theta_g > \theta_b$.

In this economy, workers have a finite working life that lasts during three periods, after which they retire from the labor market. The timing of firms' decisions is as follows: In each period, and conditional on θ_i, firms decide how many new workers to hire and how many of the existing workers to keep. For simplicity, we assume that firms only hire among workers that are at the beginning of their working careers; Older workers that have lost their jobs or never found one, remain unemployed or leave the labor force.

In the event of dismissal, firms have to pay a severance that increases with a workers' tenure at the firm. The mandated schedule is as follows: After one period, the severance pay is F, which increases to αF, $\alpha > 1$ after two periods at the firm. After three periods, the worker retires at zero cost.

Given these assumptions, the expected value of continuing a match with a worker that has been j periods at the firm, given state i, (SW_{ji}) is:

$$SW_{2i} = Max\{\theta_i - w, -\alpha F\} \tag{1}$$

$$SW_{1i} = Max\{\theta_i - w + \beta E_i SW_{2i}, -F\} \tag{2}$$

$$SW_{0i} = \theta_i - w + \beta E_i SW_{1i} \tag{3}$$

where $E_i SW_{ji}$ is the expected future value of a match conditional on state i, β denotes the discount rate and w is the wage, where $w \in [\theta_g, \theta_b]$. In this model, wages are given and assumed to be constant across a worker's life and across states for simplicity[3]. Expression (1) indicates that the value of continuing a match with a worker that has been two years at a firm (from now on a two-period worker) is the maximum between, what a employer obtains if the worker remains employed and what the firm has to pay if the worker is dismissed. If given the state of nature θ_i, the continuation value of a worker is below the threshold given by the firing cost, the match is terminated and $SW_{2i} = -\alpha F$. Condition (2) is equivalent to condition (1) , however the threshold for workers that have been only one period at a firm (one-period workers) is lower, reflecting their lower firing costs.

Our last assumption is that hiring a worker implies a cost that depends on how tight is the labor market. In periods of low unemployment, it requires more time and effort to fill a vacancy. Let $H(U)$ be the cost of hiring a worker depending on the state of the labor

[3] Assuming that wages and productivity increase over a worker's life would not alter our results as long as wages increase in direct relation to productivity. In the same manner, assuming $w_g > w_b$, but not so much as to imply lower employment in good states, would not alter our results but add extra notation to the model.

market and let $H'(.) < 0$. This assumption ensures that if $SW_{0i} > 0$, new jobs are created until the value of hiring a new worker is driven to zero.[4] That is;

$$SW_{0i} - H(.) \leq 0$$

In this simple set up, a firm operate "units" of production, namely workers, as long as the expected value of the unit is larger than the value of shutting the unit down. Future higher costs of dismissals are internalized because they tend to reduce the expected value of keeping a worker, therefore making it more likely to push him or her over the threshold at the present. Our next question is, given bad times, which workers are more likely to be dismissed first?

2.1 Which workers are more likely to be fired?

Proposition 1 *Given a severance pay schedule with slope α, there is a value ρ^* such that: i) For values of $1 > \rho \geq \rho^*$, one-period workers will be more likely to be dismissed than two-period workers. ii)For values of $0 < \rho < \rho^*$, two-period workers will be more likely to be fired, and iii) ρ^* is increasing in $\theta_g - w$ and decreasing in α.*

Proof: See Appendix A

Why does the likelihood of dismissing a one-period or a two-period worker depends on the persistency of a shock?. Assume a firms falls in a bad state and the persistency of the shock is expected to be large (that is, ρ is close to 1). Given this state of affairs, existing workers are not good investments since a employer is likely to incur in many periods of looses before emerging into a good state. Moreover, young workers are worst investments than older ones since old workers will retire soon. In this case, firms will dismiss young workers first. As the probability of remaining in a bad state declines, young workers become more attractive because they might become good values in the future. An employer then, will be more likely to cash loses with an old worker and keep the option value with a younger one. Moreover, as the slope of the severance pay-tenure profile increases, the likelihood that young workers are dismissed first increases.

[4] The assumption that hiring costs depends on market tighness is found in matching models, such like Mortensen and Pissaridies [1997].

The next question is whether the threat of increasing dismissal costs leads firms to over-rotate workers. That is, hiring and dismissing them as soon as possible (regardless of the state of the economy) to avoid paying high dismissal costs in the future.

2.2 Will firms over-rotate workers?

We examine a simple case where firms hire workers and then dismiss them after one period —regardless of the state of nature—, to avoid paying high dismissal costs.

Proposition 2 *If* $\alpha < \frac{1+\beta(1+\beta\rho)}{\beta(1-p)}$ *then over-rotation (as defined above) does not occur.*

Proof: See Appendix A.

This condition states that over-rotation depends on the value of α, which determines the steepness of the severance pay schedule. If severance pay increases very rapidly with tenure, firms will dismiss workers after a short period at the firm. Yet if α is sufficiently low, rotating workers every period is very costly and firms will keep workers in good states.[5]

In the following subsection we use this simple model to assess the impact of a labor codes reform on firms' optimal hiring and firing decisions. Studies like Bertola [1990] and others, analyze the impact of altering severance payments on hiring and firing decisions when severance pay does not vary by tenure. Their predictions are well-known: lower job security leads to more workers hired in good times and also more workers fired in bad times with undetermined effects on average employment. In this paper we perform a different exercise: We analyze the impact of a change in the *slope* at which severance pay increases with tenure.

2.3 Impact of a Labor Reform

Proposition 3 *An increase in severance pay, in the form of a higher* α*, leads to a reduction in new hirings and a Last-in First-out (LIFO) firing policy.*

Proof: Assume $\alpha' > \alpha$. If $SW_{2b} > -\alpha F$ then the change is not binding and nothing occurs. Assume instead that

[5] If we take the period to be a year, imputing some reasonable values for β and ρ, yields that the severance payment profile has to be very steep in order to induce over-rotation. For example, if $\beta = .9$ and $\rho = .5$,then $\alpha = 5.12$,whereas if $\beta = .95$ and $\rho = .8$, then $\alpha = 14.15$. This indicates that severance payments have to be multiplied by, at least, a factor of 5 over the scope of one year to lead to over-rotation. This rate of increase is much larger than the typical increase of one month pay per year of tenure.

$SW_{2b} = -\alpha F$ but $\theta_g - w > -\alpha' F$. In such case, $SW'_{2b} = \theta_g - w < SW_{2b}$ and two-period workers, that were dismissed before the policy change, will not be fired afterwards. This change implies that $SW'_{1g} = \theta_g - w + \beta\rho SW_{2g} + \beta(1-\rho)SW'_{2b} < SW_{1g}$ and in the same manner, $SW'_{1b} < SW_{1b}$ thus, first-year workers are now more likely to be dismissed than before. Finally, because $SW'_{1i} > SW_{1i}$ then $SW'_{0i} > SW_{0i}$ and firms will want to hire less young workers ■

Thus, the continuation value of one-period workers decreases after the change in policy since bad states of nature lead to higher future firing costs for these workers. It is then clear, that one-period workers would be *more* likely to be dismissed, while more tenured workers will be *less* likely to be fired after an increase in α. Therefore, an increase in the slope at which severance pay increase over time does not affect all workers in the same manner. In particular, more tenured workers will be over-protected by an increasing severance pay profile. Finally, it should be fairly clear that an opposite type of reform, that is, one that reduces the slope of the severance pay profile, would induce an increase in new hirings and a First-in, First out (FIFO) firing policy.

It is easy to see that the predictions of this model hold in the case that workers have longer working lives and may potentially remain more periods at the firm. Assume a severance pay profile tF, where t is tenure at the firm, and assume that workers can obtain payments up to a maximum SF. Assume further that the policy change takes the form of reducing the maximum, say to $S' < S$. Those workers whose values, prior to the policy change, were above $-SF$ but below $-S'F$ can be hit by the reform. Out of this group, firms would start firing from the more tenured workers since, in the context of this model, they have less periods to go until retirement. This is because the impact of future lower dismissal costs in increasing the continuation value of a worker is lower in workers that are closer to retirement age. Instead, younger workers can benefit from future lower costs of dismissal and therefore, their expected value (after the change in policy) will increase with the potential horizon at the firm making them less likely to be pushed over the threshold..

The above discussion relies on wages not adjusting to the existence of or changes in severance pay legislation. Lazear[1990] shows that wages can offset the effect of state-mandated severance pay as long as firms are able to charge a fee to workers upon starting a job. Yet, he also acknowledges that the presence of borrowing constrains or lack of trust is likely to limit the extend of such transfers, and therefore the extend at which wages can adjust to fully compensate for mandated pay. Consequently, it is expected that changes in the severance pay legislation will result in changes in employment allocations.

3 Labor Market Reforms in Chile

In recent years, a substantial body of literature has examined the effects of employment protection on employment and unemployment rates. The large majority of these studies focuses on developed countries, with few studies examining the impact of job security provisions in developing countries[6]. Yet, most Latin American countries experience higher rates of protection than any industrial country (Marquez [1997]). This high rates of protection combined with significant changes in labor codes imply that Latin American countries offer excellent laboratories in which to test the impact of these provisions. In this study we turn to Chile to examine the impact of job security on the age-composition of employment.

3.1 An index of labor legislation

As in many other Latin American countries, Chilean labor codes were modelled after those existing in the countries of the South of Europe. Therefore, it is not surprising that –as occurs in Spain, Italy, France or Greece– Chilean codes favored full-time, permanent employment over other part-time or more temporary contractual relationships. Since its inception in 1966, labor laws mandated that in the case of a firm-initiated separation, a worker had to be compensated with a payment equivalent to one month's pay per year of work at the firm. However, throughout the years, the law was modified in a number of ways altering the maximum amount that a worker could receive or the instances under which a worker could be dismissed without severance pay. Table 1 summarizes the changes in legislation that took place in the 1960-1998 period, focusing on (i) advance-notice periods, (ii) the amount of compensation in case of justified dismissal, (iii) the amount of compensation in case the unjustified dismissal, (iv) whether financial or economic needs of the firm where just cause for dismissal and finally, (v) to which workers the reforms applied.

[6] A few exceptions are Rama [1995] and Marquez [1997] who use a cross-country sample of developing and developed countries to study this issue. See also Fallon & Lucas [1993].

< *Insert* Table 1: Labor Legislation in Chile>

One of the most interesting aspects of the Chilean experience is that in a span of 39 years, Chile has gone from a situation of dismissal at will (up to 1966) to a very rigid labor market by OECD standards. In assessing the cost of dismissing a worker, two factors are specially relevant. The first one is the severance pay profile. Over the years, the existence of different caps in the maximum severance pay has substantially changed the cost of dismissing high versus low tenure workers. Whereas in 1966, to dismiss a worker with 20 years of tenure implied a severance pay equivalent to 20 months of pay, this amount was reduced to 5 months in 1984, and raised back to 11 months in 1990.

The second determinant of the cost is related to the labor codes' definition of justified dismissal as well as the stand of labor courts as mediators between firms and workers. For example, in the period from 1966 to 1981, a dismissal originated by the economic difficulties of a firm was considered justified and no compensation had to be awarded. Yet, it was also the firm's responsibility to prove the economic causality. In practice, workers would appeal and take the case to courts. It would then be up to a judge to decide whether severance payments had to be awarded. Casual observation and conversations with labor judges indicate that in the period 1966-1973 the majority of cases were ruled in favor of workers (Romaguera et al [1995]) In such cases, a firm could choose between paying a compensation –plus any wages foregone during the legal process– or reinstate the worker. After the 1973 military coup, there was a de-facto liberalization and it was much easier to dismiss a worker at no cost.

Our goal is to elaborate a synthetic index that measures the relative rigidity of Chilean codes over the years. This is not an easy task since dismissal costs are not given by a number, but instead by a profile that changes with a worker's tenure. It is therefore necessary to come up with synthetic measures that summarize the shape of the schedule and are sensitive to changes in upper limits, since that is the margin that suffered more changes and that is the most relevant for our exercise. Another difficulty comes from the fact that an important part of the cost is related to whether firms end up paying any compensation or not depending of the outcome of the judiciary process.

Our approach is to compute an index combining information on notice periods, compensation for dismissal, the likelihood that firms' difficulties be considered as justified cause of dismissal, and the severance pay that is due in that event. The formula to compute the cost of job security in period t is the following:

$$Index_t = \sum_{i=1}^{T} \beta^i \delta^{i-1}(1-\delta)(b + aSP^{jc}_{t+i} + (1-a)SP^{uc}_{t+i})$$

where δ is the probability of remaining in a job, $\delta^{i-1}(1-\delta)$ is the probability of dismissal after i periods at the firm and β is the annual discount factor. In addition, b denotes the cost of advance notice, a is the probability that the economic difficulties of a firm are considered a justified cause of dismissal, SP^{jc}_{t+i} is the mandated severance pay in such event to a worker that has been i years at the firm, and finally, SP^{uc}_{t+i} denotes the payment to be awarded to a worker with tenure i in case of unjustified dismissal.

The constructed index measures the expected cost, at a time the worker is hired, of dismissing a worker in the future. The advantage of this measure is that it captures the whole profile of severance pay. The assumption is that firms evaluate future costs based on current labor law. Higher values of the index indicate periods of relative high job security, whereas lower values characterize periods in which dismissals were less expensive.

Based on the legal information summarized in Table 1 we feed the parameters summarized in Table 2.A (See appendix), into the index formula. The schedule of the severance pay is clearly stated in the labor codes and therefore could be readily used in our formula. In contrast, the probability that a dismissal was considered justified or not is difficult to determine. We performed educated guesses based on two pieces of information: (1) whether firms' economic difficulties were considered a justified cause for dismissal according to the labor codes and (2) information on the stand of labor courts in each period.

Regarding the turnover rate we assumed that in absence of job security, average Chilean turnover rates would be similar to the U.S. ones. This choice was based on the fact that the probability of dismissal is itself affected by severance pay legislation, and that turnover information prior to the inception of the Chilean labor codes is difficult to obtain. We use Davis and Haltinwager [1995] estimate that U.S. job destruction rates average 12% a year.

Finally, we compute the discount rate based on the fact that Chilean real interest rates averaged 8.4% during the 1960-1996 period. The resulting index series is plotted in Graph 1.

< *Insert* Graph 1: Index of employment protection in Chile>

The index exhibits a maximum during the 1966-1973 period and a minimum during 1981 to 1984. After 1985, and specially after 1990, job security increased again but to levels below those attained during 1966-1973. The fact that the index is not monotonically increasing or decreasing is helpful in identifying its potential impact on employment or other labor market variables.

3.2 Empirical Strategy

We use data obtained from 39 consecutive University of Chile Household Surveys, to construct annual time series on employment, unemployment and participation rates during the period 1960-1998. These are comparable and representative annual surveys for the metropolitan area of Santiago, designed to monitor labor market conditions in the capital area. Each survey contains individual and labor market information covering between 10,000 and 16,000 people and 3,700 and 5,400 labor force participants. During this period, Santiago represented about one third of Chile's total population, and a higher proportion of GDP.

In our model, job security affects the age-composition of employment because severance pay is linked to tenure and tenure is related to age. Since men tend to have a higher attachment to the labor force, a stronger link between tenure and age is expected for this group. Therefore, we restrict our analysis to male workers. Unfortunately, the University of Chile household surveys do not contain information on a worker's tenure as to allow us to examine the association between tenure and age in our sample. However, complementary information obtained from the CASEN National Household Surveys indicates that average tenure is monotonically increasing in age (See Graph 2 plotting mean tenure by age in 1987 and 1996).[7]

[7] The CASEN (Encuesta de Caracterizacion Socioeconomica Nacional) is a nationally and regionally representative Chilean household survey carried out by the Ministerio de Planificacion y Cooperacion, through

< *Insert Graph* 2 : *Mean Tenure by Age* >

We split the University of Chile Household data into three age groups; young (15-25 years old workers) middle-age (ages 26-50), and older workers (51-65) and construct employment, participation and unemployment series for each age group. Table 3 reports the summary statistics of our data. Complementary data from the 1987 and 1996 CASEN national surveys indicates that our age split offers a good approximation of expected time at the firm and expected dismissal costs for each group (See Graphs 3 and 4). In 1987, for instance, approximately 70% of workers 15-25 years old had been less than 2 years in their current job. This proportion was only 35% among workers 26-50, and 20% among older workers (51-65). In 1996, after a relatively long period of sustained growth and increased hirings, the percentage of low-tenure workers was slightly higher among all age groups.

< *Insert Graph* 3 *and* 4: *Distribution of tenure by age* >

Using these data we perform three separate exercises: First, we test the implications of our model by assessing whether employment, unemployment and participation rates exhibit an age-specific response to labor reforms and, whether this response corresponds to what is predicted by our model. Regrettably, the University of Chile survey does not include information on whether a worker is covered by labor laws. However, the surveys allow for a separation between wage and self-employment. Since self-employed workers are not covered by job security provisions a differential response between these groups to changes in legislation provides additional information on the nature of the effects captured by our estimations.

Second, we elaborate a more direct measure of the relative costs of dismissal for high and low tenure workers and examine whether this relative cost is related to young and old workers employment rates. Finally, we assess the impact of labor reforms on overall employment, unemployment and participation rates in Chile.

the Department of Economics of the Universidad de Chile, with the dual objectives of generating a reliable portrait of socioeconomic conditions across the country, and of monitoring the incidence and effectiveness of the government 's social programs and expenditures. Questions are asked at the household and the individual level. The sample size are quite generous for international standards. For instance, in the year 1996, the sample included information on 134262 individuals.

3.3 Age-Specific Response to Job Security Provisions

In this section, we examine whether employment, participation and unemployment rates exhibit a differential response to labor market reforms by age group. To do so, we estimate the following expression:

$$\Delta Y_{jt} = c + \gamma Y_{jt-1} + B_1.\Delta \log GDP_t + B_2 \Delta \log(wage_{jt}) + B_3 \log(index_t) + B_{4j}(L)\Delta Y_{jt-L} + \epsilon_t \quad (4)$$

where Y_{jt} represents, depending on the specification, employment, wage-employment, self-employment or participation as a percentage of total population, as well as unemployment rates –as a % of total labor force— for age group j. The lagged employment variable captures the fact that firms may choose not to adjust immediately their labor force given changes in the forcing variables. In particular, in presence of linear adjustment costs, $\gamma + 1$ reflects the fraction of employment that remains constant after a shock in the forcing variables (Anderson [1993]). Aggregation over units with different adjustment costs may lead to further lags. Bearing this in mind we allow for further lags in the endogenous variable. We include lags in the participation and unemployment specifications as well, if its inclusion is not rejected by the data. Finally, we include GDP and the age specific wage growth as forcing variables[8].

The inclusion of wages and GDP in growth rates instead of levels is justified by our definition of the endogenous variables as ratios that can only vary between zero and one. This choice is also justified by the presence of unit roots in all Y_t and the fact that none of the Y_t variables were cointegrated with GDP and wages in levels. Table 4 reports the results of testing for unit roots in all our variables according to an augmented Dickey-Fuller test. In all cases the values of the test were below the critical values and the unit root hypothesis could not be rejected. Yet, Johansen Cointegration tests did not reject cointegration between Y_t and the forcing variables in expression (4).

[8] Age-specific real wages series were constructed out of the University of Santiago surveys and are deflated by the CPI. They capture real hourly earnings for male workers employed full-time in the wage-employment sector. (See Table 3 for summary statistics of these variables)

Graphs 3 and 4 highlight that even when mean tenure and age are positively correlated, workers of different tenures coexist within each age group Thus, in going from our tenure-based model to our age-based empirical exercise, direct parallelisms with the three theoretical tenure groups cannot be drawn. First, while in the theoretical model, the young workers group had only new entrants, the 15-25 empirical group will also include workers that have been more than one period at the firm. According to our model, for sufficiently persistent shocks these workers are the most likely to be dismissed in a bad state. Thus, one implication of our model is that the employment rates of young workers are likely to decrease after a labor market reform that increases severance pay for high-tenure workers. This is due to the confluence of two effects working in the same direction: (i) Increased job security for high-tenure workers reduces the continuation value of young workers making them more likely to be dismissed in bad times, and (ii), increased job security reduces hirings rates, which, as shown in graphs 3 and 4 are specially relevant for young workers. Thus, the expected sign on *index* is negative for the young workers group.

Regarding the prime-age and the 50–65 age group, our model implies a different response to a labor market reform. In particular, a labor market reform that increases job security for high-tenure workers should decrease dismissals among prime-age and older workers. In addition, since new hirings are not as relevant in these age groups, an increase in job security may well increase prime-age and older workers employment rates. Thus, the sign on *index* can be either positive or negative, but is expected to be larger for the older group.

The employment series by age group plotted in Graph 5 seem to confirm these predictions. The shaded areas correspond to periods in which the labor legislation was more strict, whereas the light areas correspond to times of lower job security. Whereas wage-employment rates for young workers fell abruptly from 1966 to 1973, employment rates for prime-age workers and older workers increased. In the same manner, from 1985 onwards, and corresponding with a period of sustained growth, employment rates for all groups tended to increase, however, employment rates for older and middle-age groups increased at a much faster rate. The evolution of the young-to-old wage-employment rate is even more telling (Graph 6); the proportion of young workers (adjusted by population) fell in both periods of high job security and increased in the period 1973-1985, which was characterized by a relative flexibilization of the labor market.

<Graph 5: Male Wage-Employment-to-Population Rates by age group>

<Graph 6: Young to Old Male Wage-Employment-to-Population Rates by age group>

We now provide a more formal test of differences in behavior across age groups.

3.3.1 Results for Workers 15-25 Years Old.

The first group of estimations correspond to the group of workers 15-25 years old. The results are summarized in Table 5. To reduce the likely endogeneity of wages we lagged them one period in all the specifications. In addition, the inclusion of the endogenous variable lagged one period was not rejected in any of the specifications.

Column (1) confirms that an increase in job security measured by *index* reduces employment rates for young workers. Moreover, our results indicate that this effect is entirely due to a decline in wage-employment rates. The decline in employment that can be attributed to a change in job security is large. Consider, for example, the 1966 reform. In that year a new law increased the rigidity of the labor code. Our legislation index captures this reform with a 200% increase over its value in 1965. Column 2 implies that within a year, employment-to-population rates had declined by approximately 2.8%. Moreover, our estimates suggest that this decline in young employment rates is likely to persist as long as the severance pay provisions remain high[9].

The estimates in column (3) suggest that self-employment rates were not affected by changes in job security provisions. This differential response of wage and self-employment indicates that our legislation variable is not capturing contemporaneous environmental changes, such like changes in schooling policy, that may also affect employment[10]. It also

[9] Several econometric problems were encountered when estimating the long run response of employment to legislation changes. While Johansen test indicated that employment was cointegrated to the forcing variables, Dickey-Fuller test on the residuals of the cointegration relation indicated that it was not. Due to this inconsistency, often encountered in small samples, we could not estimate the long run relationship. Moreover, this possible lack of cointegration means that the coefficient on Yt-1 is not different from zero (since it cointains a unit root). Therefore, the coefficient on the legislation index should be interpreted as permanently affecting the change in the endogenous variable in the context of our short period sample.

[10] Since there have been substantial changes in secondary schooling policies and enrollment rates we conducted an additional test to ensure that changes in schooling were not biasing our results. We re-estimated the specifications reported in Table 5 with a subsample of people 18-25 years old. The estimates were very similar to the ones reported here.

suggest that in the face of reduced wage-employment opportunities, self-employment is not an option for young workers. These results confirm the findings of Maloney [1996] suggesting that workers without capital or know-how are less likely to move to self-employment.

Instead, columns (4) and (5) indicate that a decline in wage-employment is compensated by a reduction in youth participation rates, without any significant effect on unemployment. Thus, our results imply that employment protection is detrimental for young workers, yet a researcher that only analyzed unemployment rates would not capture this effect.

Since residual autocorrelation leads to inconsistent estimates in equations were lagged endogenous variables are included, the last two rows in table 5 report tests on residual autocorrelation. In presence of the lagged endogenous variables the Durbin-Watson statistic (DW) is likely to be biased towards finding no autocorrelation, therefore we also provide critical values for the Breusch-Godfrey serial correlation test. In all cases, the null hypothesis of no autocorrelation was not rejected.

3.3.2 Results for Prime-Age Workers (26-50 Years Old).

The results for workers in this age group are summarized in Table 6. Series for prime-age workers and in particular, the total and wage employment series, displayed a substantial amount of serial correlation. In that context, the specification used for young workers was rejected according to the Breusch-Godfrey test. To correct this problem, we included two lags of the endogenous variable in all especifications. In addition, in the specifications for wage and total employment we included two lags of GDP and wage growth. Once these variables are included, we cannot reject the hypothesis of no serial correlation.

Our results indicate a substantial difference in the impact of job protection for prime-age workers. In particular, the coefficient on the index in total and wage employment indicates a positive, albeit not statistically significant, effect of job protection on employment. These results coincide with Nickell's [1997] findings that higher job protection has a positive but not significant impact on prime-age employment. The lack of impact is not surprising since, an increase in job protection affects differently high and low tenure workers within this age group. Whereas, low tenure workers will be more likely to be dismissed in recessions, high tenure workers will be less likely to be so. The resulting balance of these two effects offsets lower hirings in expansions.

Columns (3), (4) and (5) indicates that the low response of employment rates to labor market reforms is matched by a comparatively low response in participation and unemployment rates.

3.3.3 Results for Older Workers (51-65 Years Old).

Results for workers 51-65 years old are summarized in Table 7. As in the prime-age case, series for older workers displayed considerable serial correlation. This lead us to reject the specification used for young workers. The correction applied in the prime-age case was not rejected. Therefore the results reported in Table 7 include as additional regressors two lags of the endogenous variable and, for the employment and wage employment specifications, two lags of GDP and wage growth. The impact of job security has the same sign than for prime-age workers but, as expected, the coefficient is larger in magnitude. In particular, the coefficient on *index* is now positive and statistically significant for total employment. Thus, an increase in job security increases overall employment rates for older workers. The size of the coefficient suggest a large change: A 100% increase in the index variable is associated with a 2.9% increase in pre-retirement age employment rates. Estimates for wage and self-employment confirm that *both* types of employment increase with increased job security, albeit wage employment increases more. It is unclear why self-employment increases with job security. Perhaps, severance pay provide the start-up capital necessary to move into self-employment. Finally, columns (4) and (5) suggest that older workers participation rates increase and unemployment rates decline in period of increased job security.

The results presented in Tables 5-8 are robust to the introduction of time trends and further lags in the endogenous and exogenous variables. Overall, our findings suggest that older workers are the ones that benefit the most from job security provisions. Instead, younger workers, bear the burn of increased job security because they endure lower hirings and a disproportionate share of firings in recessions..

3.4 Alternative Measures of Job Security

Our estimates in the former section exhibit a differential response by age groups to changes in job security provisions. Our model predicts that the bulk of these differences should arise from the relative differences in dismissal costs across age groups. In this section we construct two additional measures of job security to isolate the channels by which job security affects young employment rates.

The overall profile of severance pay measured by *index* can be decomposed into the relative cost of dismissing long versus short tenure workers, namely α , and the cost of dismissing workers in the latter age group, namely F. We proxy F with the cost of dismissing a worker after two years of tenure ($index2$). In addition, we proxy α, with the relative cost of dismissing a worker after twenty ($index20$) and two years ($index2$) at a firm.[11] That is,

$$\hat{F} = index2$$

$$\overset{\wedge}{\alpha} = index20/index2$$

Tables 8 reports the results of re-estimating expression (4) substituting $log(index)$ for $log(index2)$ and $log(index20) - \log(index2)$. Since there is substantial correlation between these two measures the coefficients are now estimated with larger standard errors. Nonetheless, this exercise yields some interesting results.

Table 8 shows that, as predicted by our model, an increase in the relative cost of dismissal of old workers lead to a decline in youth total employment and wage employment-to-population rates. In contrast, an increase in $index2$, has a much smaller (and less statistically significant) effect on youth total employment and wage employment rates. This is because an increase in the absolute severance pay reduces both firings and hiring rates whereas an increase in the relative cost reduces hiring while increasing firing rates. As before, none of these effects seem to be present in the self-employment data. Results for participation and unemployment rates show an interesting pattern: Higher dismissal costs for short-tenure workers induce lower participation rates and smaller unemployment rates. In contrast, higher relative cost of dismissal are associated with higher youth unemployment rates and no changes in participation. These results suggest that increased firings (as a consequence of higher α) result in higher unemployment rates whereas lower hiring and

[11] $index2$ and $index20$ have been computed as the expected dismissal costs at the time of dismissal of a worker that has been 2 and 20 years at the firm, respectively. In particular, the formula to compute $indexi$ is the following:

$indexi_t = b + aSP_i^{jc} + (1-a)SP_i^{uc}$ where SP_i^{jc} and $SP_i^{uc}$are the costs for a justified and an unjustified dismissal after i years of tenure in a firm.

firing rates (as a consequence of higher F) result in a withdrawal of youth workers from the labor force. They also indicate that, as suggested by our model, the relevant parameter for youth employment rates is the *slope*, more than the *level* of the severance pay.

To summarize, increased job security, in the form of higher relative costs of dismissing short versus long-tenure workers, increases firing and reduces hiring rates for the youth. The resulting effect is a progressive decline in their employment rates in favor of higher employment rates for older workers. The composition of employment is biased towards older workers. It is unclear, however, whether this effect results in lower overall employment and higher total unemployment rates. The following and last section of this paper examines this question.

3.5 Results for the overall sample

In this section we provide results for the overall sample (males, 15-65 years old). This exercise is relevant at least for two reasons: First, when evaluating the impact of a policy, the overall impact is as important as its composition effects. Second, since most of the literature examines the impact of job security on overall employment and unemployment rates, it is useful to determine whether our measured impact coincides with the results obtained in other individual country or cross-country studies.

Table 9 summarizes our results for the overall sample of males. Judging from the coefficient on *index*, an impact on overall job security has a negative —albeit, not statistically significant— impact on overall wage employment and a positive impact on self employment rates. In addition, we find that employment security provisions have a negative –but not significant- effect on participation and unemployment. These findings suggest that job security provisions affect the composition but no the average level of employment or unemployment.

Columns (4) and (10) however, suggests that the lack of effect is due to conflicting signs on $index2$ and $index20/index2$. Whereas an increase in the cost of dismissing short tenure workers has a positive, but not statistically significant, effect on overall wage employment, an increase in the relative cost of dismissal reduces overall employment rates. In the same manner, when we split the severance pay measure into relative and absolute costs we find that increased absolute job security reduces unemployment, whereas higher relative costs tend to increase it.

Overall, our estimates for absolute firing costs (F), confirm the thorethical findings of Bertola [1990] that a flat rate (i.e. not linked to tenure) would have opposite effects on hiring and firing that will tend to offset each other. In contrast, changes in the relative cost of dismissing short versus long-tenure workers, have a substantial negative effect on total employment. This effect is driven by the decline in the youth wage-employment-to-population rate, not compensated by neither higher self-employment rates nor by higher employment rates for the old.

4 Conclusion

In this paper, we show that tenure-based job security biases employment in favor of middle-age and older workers. Our results also indicate that tenure-based job security reduces long-run aggregate employment rates. In contrasts, a flat severance pay would have little effect on youth-employment or on aggregate employment and unemployment rates. We believe these results have two important implications for the design of future labor market reforms:

First, labor market reforms have important redistributive effects. In anticipation of such redistribution, some groups may try to block the process of reforms. Thus, while young workers could benefit from the measure, they are less likely to vote or to organize themselves supporting the reforms. In contrast, middle-age and older workers are more likely to be unionized or to exert pressure on policy-makers to block any attempt of reform that undermines their status in the labor market. Indeed, this relative higher political power is likely to explain why job security provisions are tied to tenure in almost all OECD and Latin American countries. In this context, understanding the political economy of the reforms may help policy-makers to design compensation packages aimed at attaining overall employment gains.

Second, our results support reforms aimed at reducing the link between severance pay and tenure. This effect could be achieved by: mandating a flat severance pay; reducing the maximum amount a worker can receive as severance pay; or reducing the rate at which severance pay increases with tenure. Such reforms would bring an expansion in youth and overall employment rates, although they could come at a cost of lower older workers employment rates. Since in many countries, retirement incentives have already pushed many older workers into retirement, reforms like the ones described above may require additional policies to bring older workers back to work.

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Appendix A

Proof of Proposition 1:

Let $0 < \rho^* < 1$ be such that $-\alpha F - \rho^* \beta\alpha F + \beta(1-\rho^*)(\theta_g - w) = -F$. In this case, if $\theta_b - w = -\alpha F$,then $SW_{b1} = -F$ and both one period and two period workers are dismissed. For larger values of $\theta_b - w$ neither one or two period workers are dismissed. For lower values of $\theta_b - w$, both one and two year workers are dismissed. Therefore, when $\rho = \rho^*$, both types of workers are equally likely to be dismissed..

i)Assume $\rho \geq \rho^*$ and $\theta_b - w = -\alpha F$. Then $-F = -\alpha F - \rho^* \beta\alpha F + \beta(1-\rho^*)(\theta_g - w) > (\theta_b - w) - \rho\beta\alpha F + \beta(1-\rho)(\theta_g - w)$ and slightly larger values of $\theta_b - w > -\alpha F$ will not alter the fact that one period workers are dismissed, but second workers would not be. Therefore, for all $\rho \geq \rho^*$, there is a range of values of $\theta_b - w$, such that $-\alpha F < \theta_b - w \leq -\frac{F}{1+\beta\rho} - \frac{\beta(1-\rho)}{(1+\beta\rho)}(\theta_g - w)$ in which $SW_{b1} = -F$ and $SW_{b2} = \theta_b - w > -\alpha F$ and only first-year workers are dismissed. For values of $\theta_b - w \leq -\alpha F$, both first and second-period workers are dismissed. Hence, when $\rho \geq \rho^*$ first-period workers are more likely to be dismissed.

ii) Assume $\rho < \rho^*$and $\theta_b - w = -\alpha F$. In this case $-\alpha F - \rho\beta\alpha F + \beta(1-\rho)(\theta_g - w) > -\alpha F - \rho^* \beta\alpha F + \beta(1-\rho^*)(\theta_g - w) = -F$ and therefore one-period workers would not be dismissed even when $SW_{b2} = -\alpha F$ and two-period workers would go. A marginal decrease of $(\theta_b - w)$ would not alter the fact that second-period workers are dismissed, but first-period ones will be not. Therefore for or all $0 < \rho < \rho^*$, there is a range of values of $\theta_b - w$, such that,

$-F(1-\beta\rho\alpha) - \beta(1-p)(\theta_g - w) < \theta_b - w \leq -\alpha F$ in which $SW_{b1} < -F$ and $SW_{b2} = -\alpha F$ and only second-period workers are dismissed. For values of $\theta_b - w \leq -F(1-\beta\rho\alpha) - \beta(1-p)(\theta_g - w)$, both first and second-period workers are dismissed. Hence, when $\rho < \rho^*$ second-period workers are more likely to be dismissed.

iii) Since ρ^* is such that $-\alpha F - \rho^* \beta\alpha F + \beta(1-\rho^*)(\theta_g - w) = -F$,

$$\rho^*(\alpha, \theta_g - w) = \frac{F(1-\alpha) + \beta(\theta_g - w)}{\beta(\alpha F + \theta_g - w)}$$

and therefore$\frac{\partial\rho^*(\alpha,\theta_g - w)}{\partial\theta_g - w} > 0$ and $\frac{\partial\rho^*(\alpha,\theta_g - w)}{\partial\alpha} < 0$. $\blacksquare$

Proof of Proposition 2

If firms hire workers and fire them after one period, independently of the state of nature, it has to be that:

$$SW_{0g} = H(.) > 0 \tag{5}$$

$$SW_{0b} \leq H(.) \tag{6}$$

$$SW_{1g} = -F \tag{7}$$

and

$$SW_{1b} = -F \tag{8}$$

Conditions (6) and (7) imply that

$SW_{0g} = \theta_g - w + \beta\rho SW_{1g} + \beta(1-\rho)SW_{1b} = \theta_g - w - \beta F > 0$ and $\theta_g - w > 0$ so $SW_{2g} = \theta_g - w > 0$. Assume that $SW_{2g} = -\alpha F$, that is, the firm would incur in large future firing costs if first-period workers remain at the firm, then $SW_{1g} = (\theta_g - w)(1+\rho\beta) - \beta(1-\rho)\alpha F$ $>\beta(1+\rho\beta)F - \beta(1-\rho)\alpha F$ but if $\alpha < \frac{1+\beta(1+\beta\rho)}{\beta(1-p)}$ then condition (7) does not hold and firms will not dismiss workers in expansions. ■

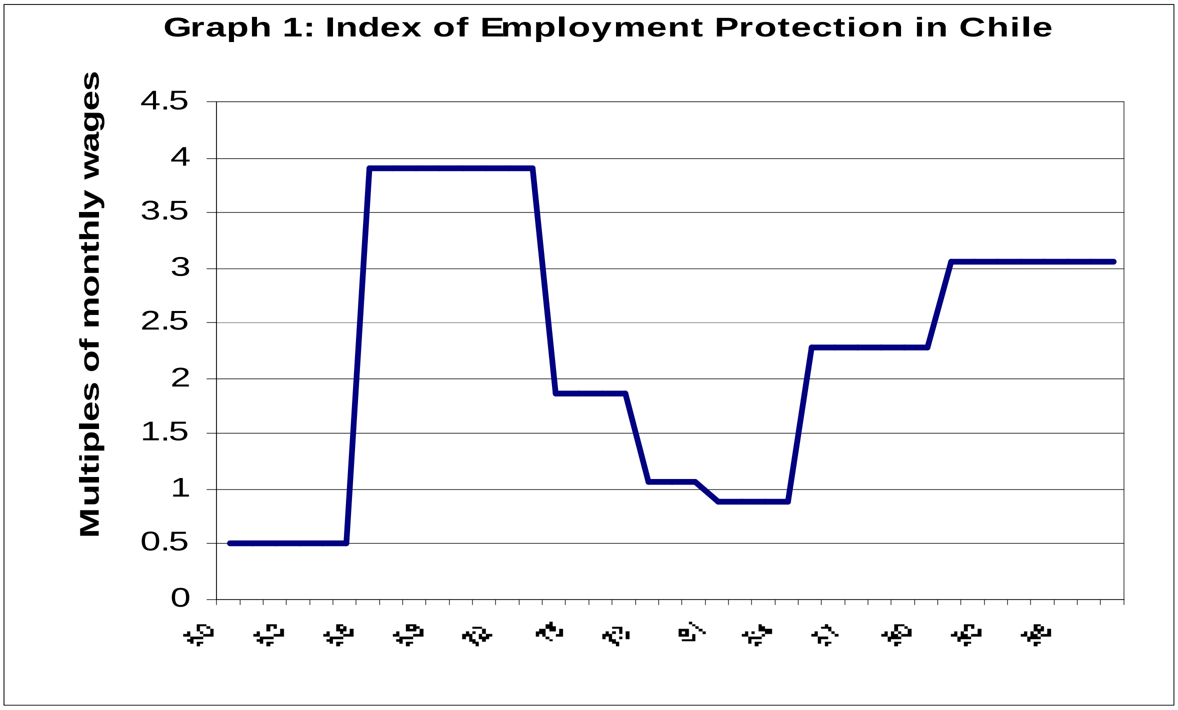
Appendix B

Table 2: Parameters used to compute *Index*

Periods[12]	β	δ	b	a	SP^{fc}	SP^{uc}
60-65	.92	.88	1	1	0	-
66-73	.92	.88	1	.2	0	(1)
74-77	.92	.88	1	.5	0	(2)
78-80	.92	.88	1	.8	0	(2)
81-84	.92	.88	1	.8	0	(3)
85-90	.92	.88	1	0	0	(3)
91–	.92	.88	1	.9	(4)	(5)

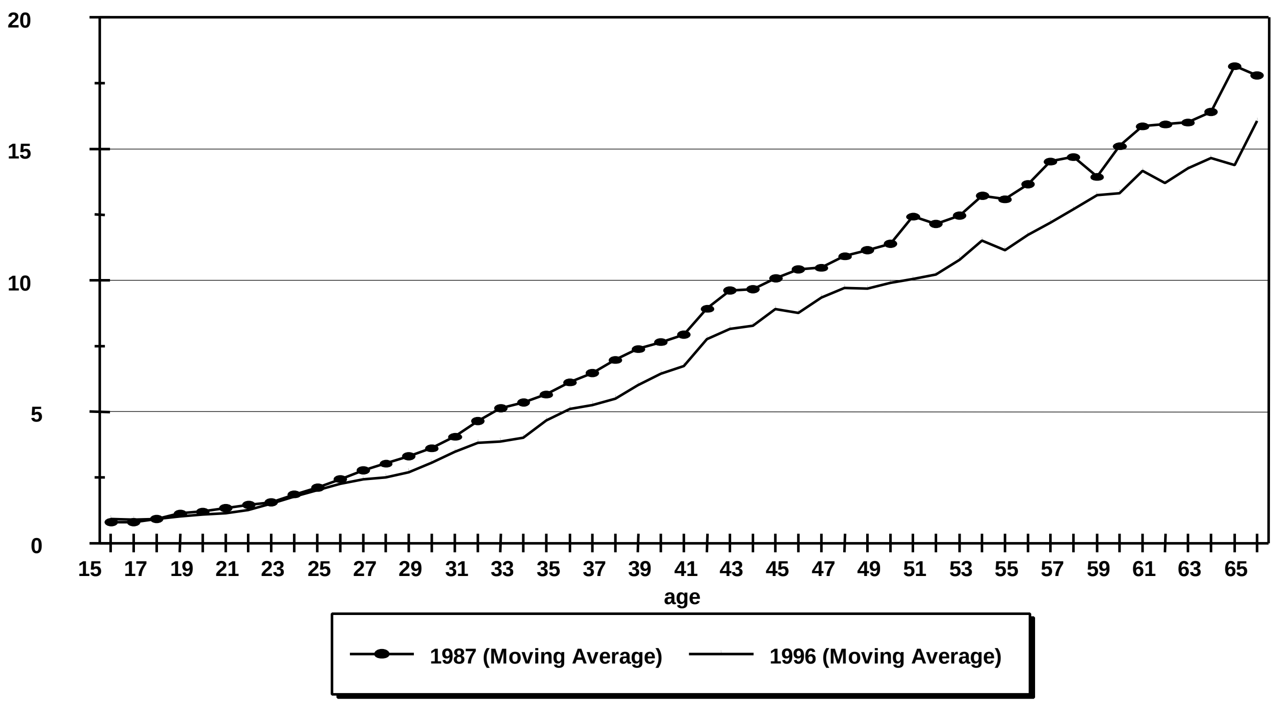
[12] To compute β we use the fact that the average real interest from 1960-1996 was 8.4%. To compute δ we assume that the Chilean turnover rates *without* employment protection would be similar to US ones. According to Davis & Haltinwager (1995) average turnover rates average 12% a year in the US. (1) Corresponds to one month's pay per year of work augmented in three months. This three months capture average payments for foregone wages during trial. (2) One month's pay per year of work without upper limit. (3) One month's pay per year of work with a five months upper limit. (4) One month's pay per year of work with a 11 months upper limit. The maximum tenure that a worker can attain at a firm is assumed to be 25 years.

Graph 1

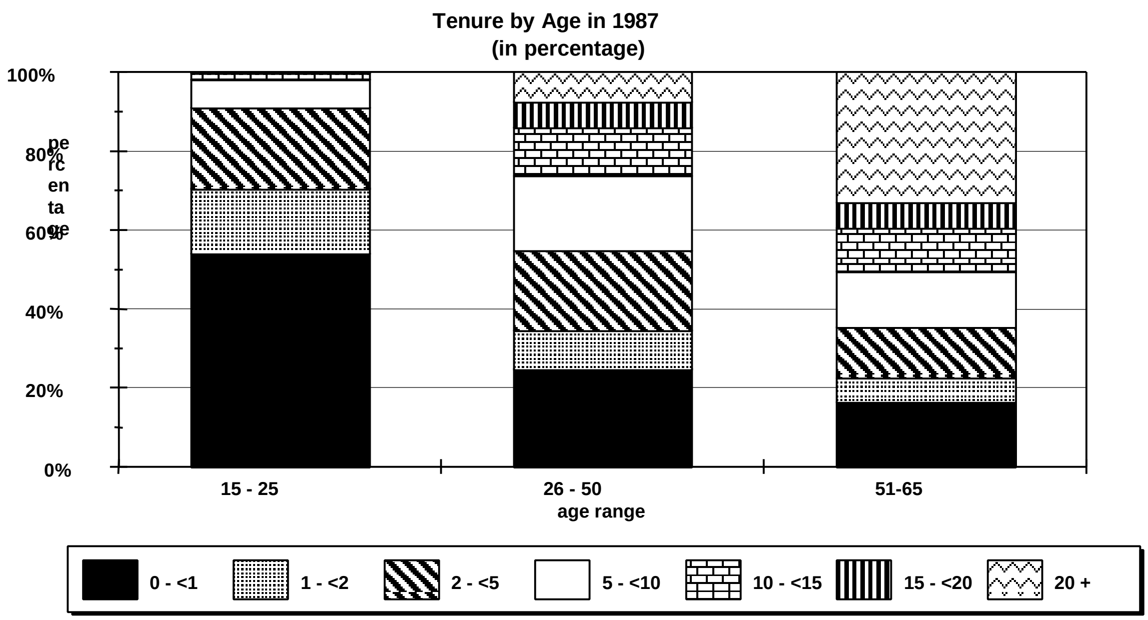


Graph 2

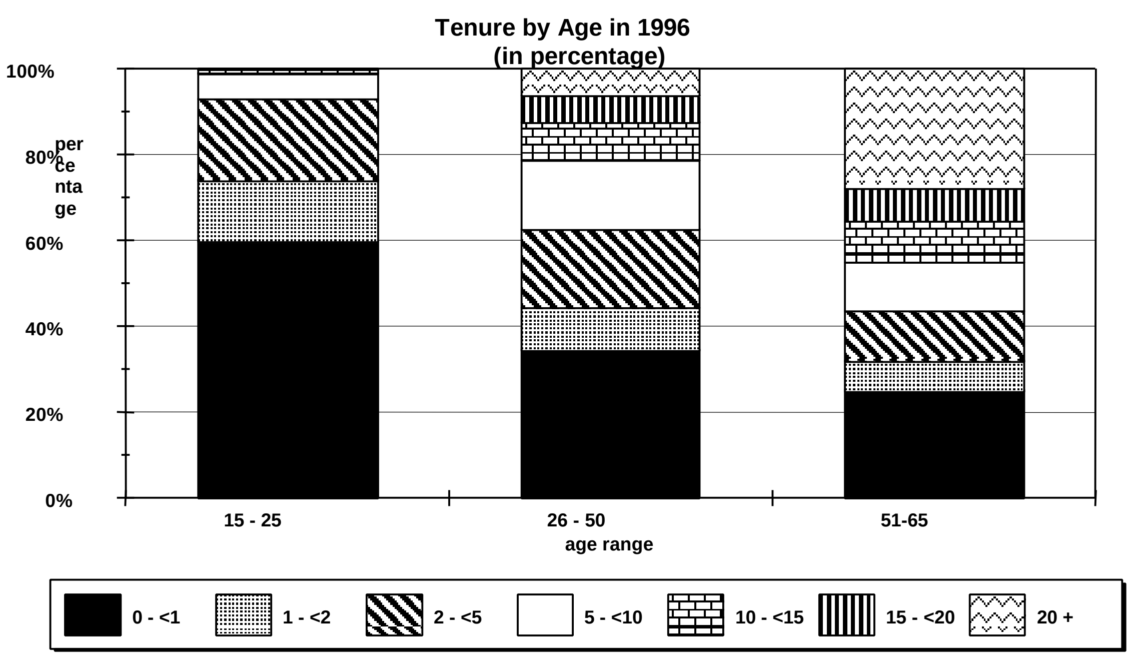
Mean Tenure by Age: 1987 and 1996



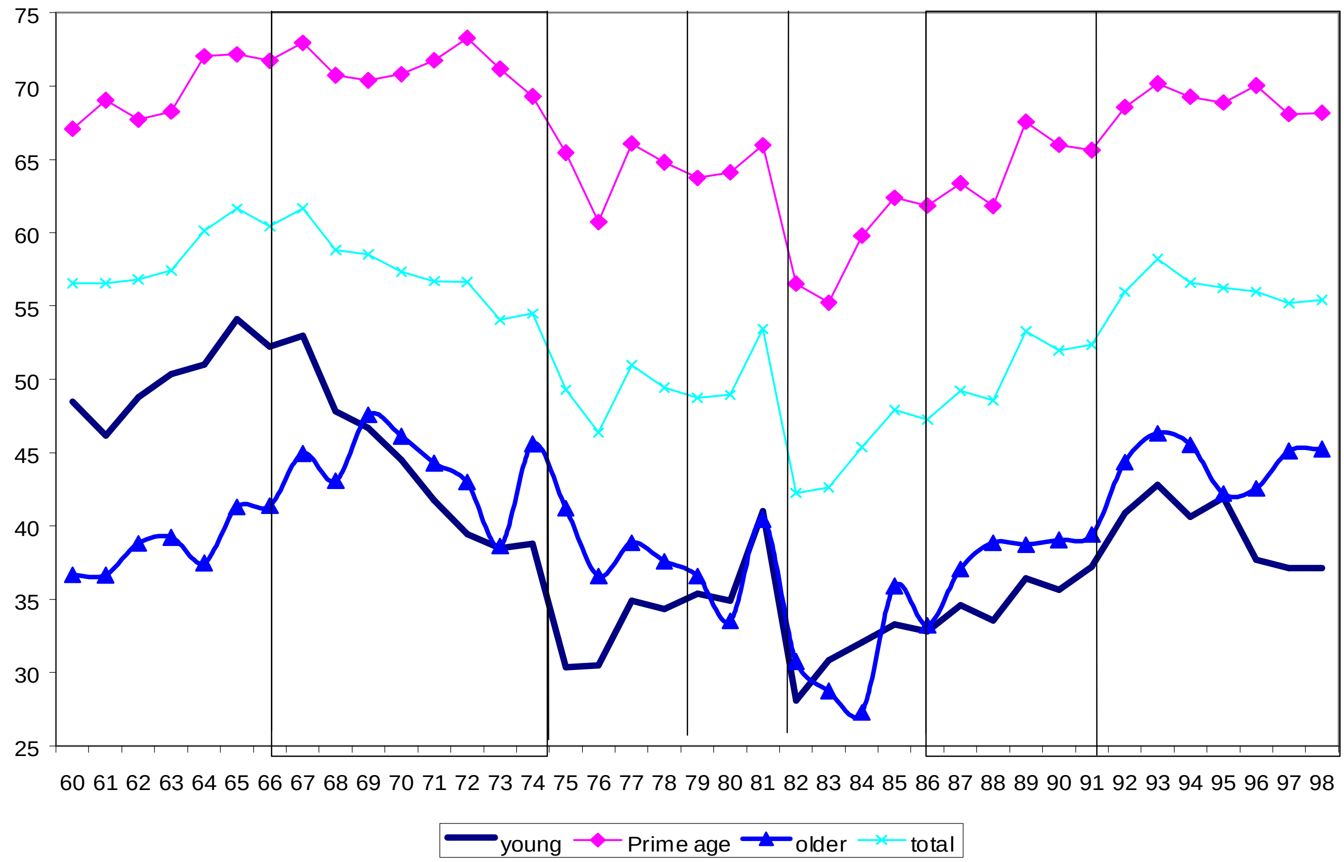
Graph 3



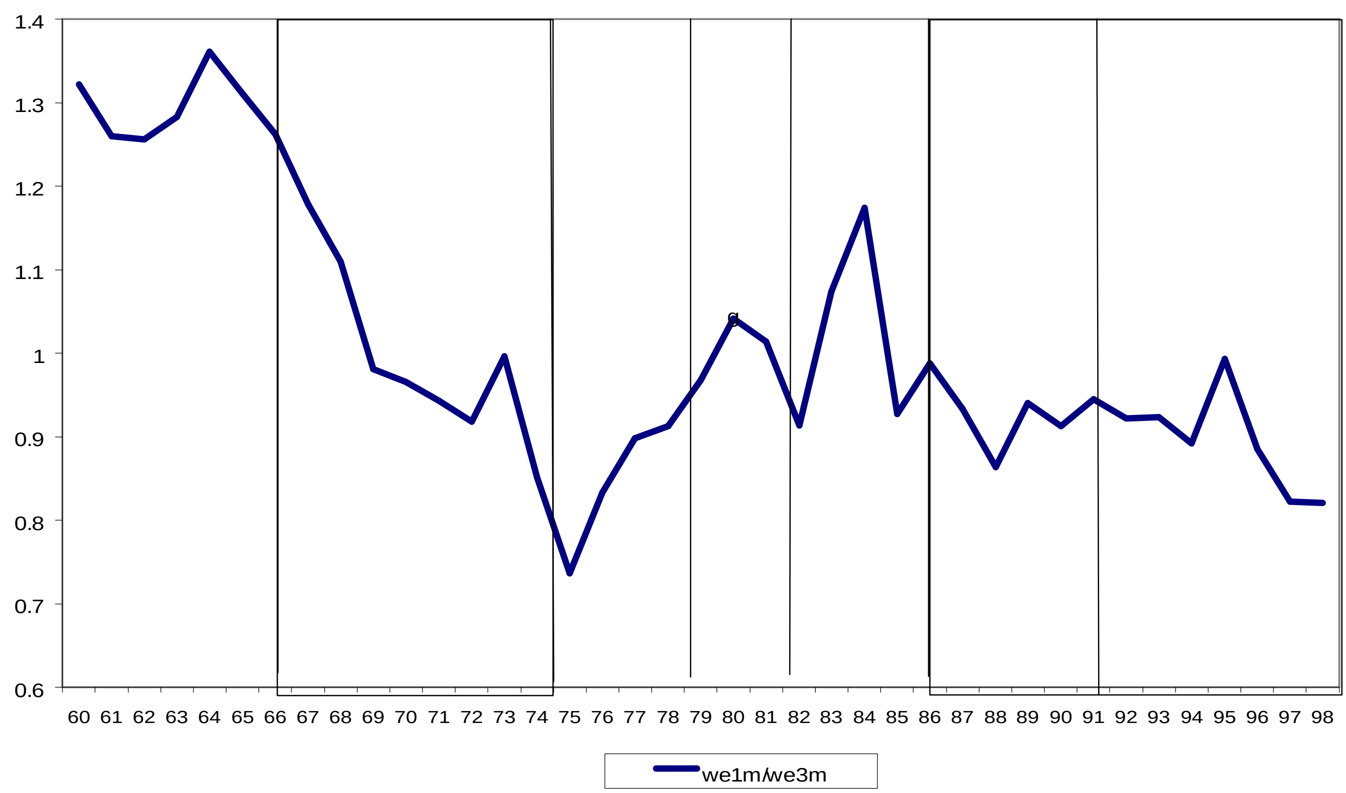
Graph 4



Wage-Employment to Population Rates
Graph 5



Graph 6: Young-to-Old Wage-Employment-to-Population Rates



Graph 7: Other Measures of Employment Protection (in logs)

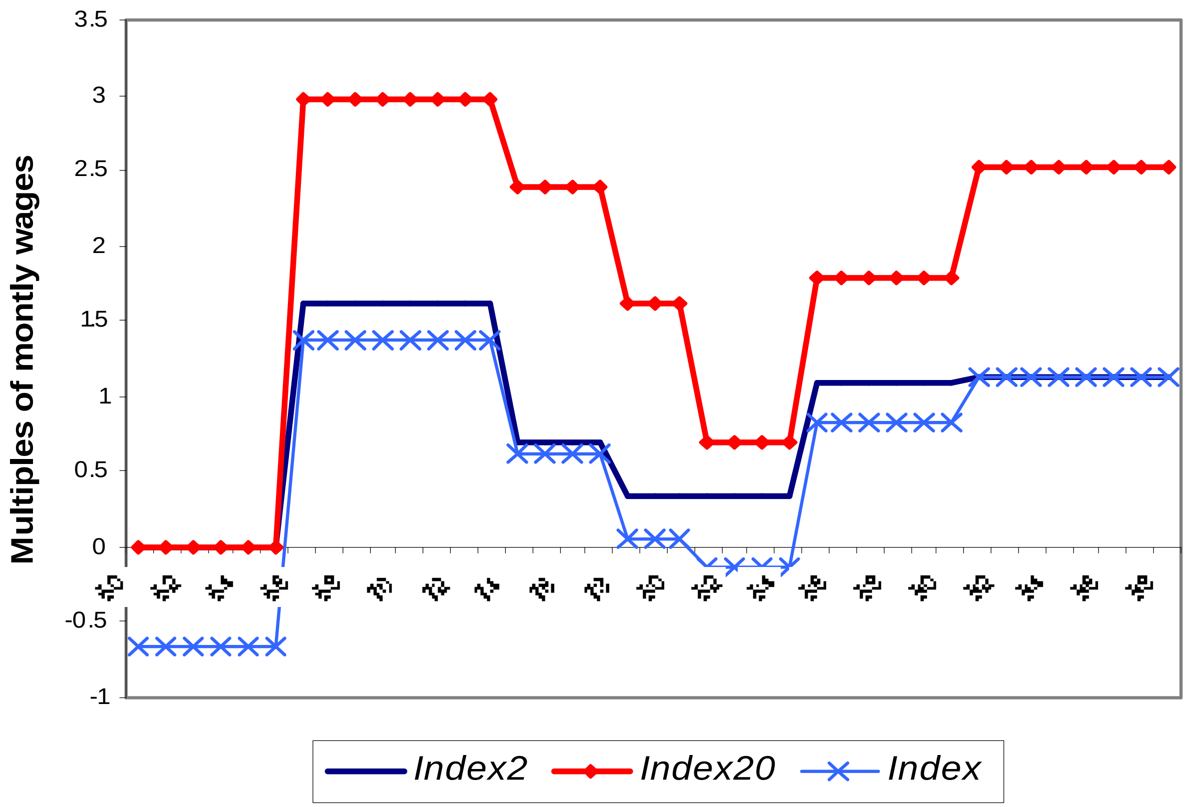


Table 1: Employment Protection Provisions in Chile

Periods	**Prior Notice Period**	**Economic reasons just cause for dismissal on the law? / in the courts?**	**Compensation for dismissal in case of just cause**	**Compensation for dismissal in case of unjust cause**	**To whom the changes apply?**
1960 –1966	1 month	Dismissals at will	Dismissals at will	Dismissals at will	Dismissals at will
1966-1973 Firms could not dismiss workers without a just cause.	1 month	Economic reasons were just cause on the law/ In practice labor courts considered most dismissals unjustified.	The law does not mandate any compensation in this case.	One month's pay per year of work at the firm plus foregone wages during trial. Trials could last at most 6 months. There is no maximum in the amount to be awarded	To all workers
1973-1978	1 month	Labor courts were much more pro-firms. Workers' claims were weaker.	Same than previous period	Same than previous period	To all workers
1978-1980 June 15, 78 Decree 2,200	1 month	Economic needs are considered just cause.	zero	1 month per year of work, without maximum limit.	Only to workers hired after June 1978
1981-1984 Law 18,018 (August,14, 1981)	1 month	Economic needs are considered just cause.	zero	1 month' wage per year of work *with a maximum of 150 days*	Only to workers hired after August 1981
1984-1990 Law 18,372 (Dec, 1984)	1 month	Economic needs are not considered just cause for dismissal any more	zero	1 month' wage per year of work *with a maximum of 150 days*	All workers
1990- today (Nov. 1990) Firms need to justify dismissals	1 month	Firms have to justify dismissals but economic needs are considered just cause for dismissal	Economic reasons: 1 month' wage per year of work with a maximum of 11 months' pay	1.2-1.5 months per year of work	All workers hired after August 1981

Table 3: Data Summary Statistics

Variable	Obs	Mean	Std. Dev.	Min	Max
GDP Growth	38	0.040917	0.054946	-0.13445	0.111345
Te1m	39	0.455138	0.07088	0.330357	0.600946
We1m	39	0.398883	0.070428	0.280907	0.54101
Se1m	39	0.056224	0.008574	0.035922	0.074627
Pa1m	39	0.552113	0.054872	0.492523	0.667981
Un1m	39	0.17795	0.073366	0.080529	0.388818
Te2m	39	0.887651	0.046998	0.745203	0.944797
We2m	39	0.669857	0.043902	0.552239	0.73278
Se2m	39	0.217638	0.015585	0.184017	0.256764
Pa2m	39	0.961154	0.0078	0.942044	0.975224
Un2m	39	0.076616	0.045619	0.019071	0.222037
Te3m	39	0.672282	0.070536	0.498084	0.790816
We3m	39	0.397385	0.048641	0.272884	0.475836
Se3m	39	0.274853	0.030732	0.191564	0.338435
Pa3m	39	0.734261	0.043938	0.637306	0.826531
Un3m	39	0.08644	0.054143	0.015385	0.259259
Te0m	39	0.707278	0.057683	0.557678	0.791482
We0m	39	0.535722	0.050761	0.42246	0.616675
Se0m	39	0.171458	0.013356	0.135218	0.194512
Pa0m	39	0.78708	0.026029	0.743418	0.8393
Un0m	39	0.102346	0.053293	0.033914	0.263618
Wm1	39	261.68	127.159	96.72	513.41
Wm2	39	532.94	240.68	203.10	1004.01
Wm3	39	550.97	239.23	223.40	1065.17
Wm	39	470	217.58	173.66	890.62
Lindex	39	0.590906	0.71781	-0.65412	1.363282
Lindex2	39	0.860731	0.558433	0	1.609438
Lindex20	39	1.840952	1.035985	0	2.965273

Note: Variables should be read as follows: te1m, for instance, means total employment (te) for age group 1 (15-25), male (m) The prefix te= total employment, we=wage employment, se=self-employment, pa=participation and un=unemployment. Age groups; 0=15-65, 1=15-25, 2=26-50, 3=51-65. Wm*i*= Real wage in age group i. For more information see Table 4 which displays variable labels.

Table 4: Augmented Dickey-Fuller Unit Root Test

Critical value 1%	-3.628	
Critical value 5%	-2.947	
Critical value 10%	-2.611	

Variable code	variable label	ADF test Variable in Levels	ADF test Variable in First Diff.
te1m	employment rate 18-25	-1.428	-2.833 (*)
we1m	wage employment rate 18-25	-1.533	-2.699 (*)
se1m	self-employment rate 18-25	-2.001	-3.368 (**)
pa1m	participation rate 18-25	-1.265	-3.238 (**)
un1m	unemployment rate 18-25	-1.672	-3.38 (**)
te2m	employment rate 26-50	-1.783	-3.331 (**)
we2m	wage employment rate 26-50	-1.462	-3.043 (**)
se2m	self-employment rate 26-50	-2.975	-3.441 (**)
pa2m	participation rate 26-50	-1.943	-3.234 (**)
un2m	unemployment rate 26-50	-1.783	-3.525 (**)
te3m	employment rate 51-65	-0.929	-2.696 (*)
we3m	wage employment rate 51-65	-1.213	-2.655 (*)
se3m	self-employment rate 51-65	-0.967	-3.075 (**)
pa3m	participation rate 51-65	-0.97	-3.236 (**)
un3m	unemployment rate 51-65	-1.637	-3.05 (**)
Tem	employment rate	-1.38	-2.749 (*)
Wem	wage-employment rate	-1.308	-2.55
Sem	self-employment rate	-2.455	-3.35 (**)
Pam	participation rate	-1.3511	-3.281 (**)
Unm	unemployment rate	-1.64	-3.211 (**)
wm1	real wage 18-25	-1.059	-3.701 (**)
wm2	real wage 26-50	-1.268	-4.294 (**)
wm3	real wage 51-65	-1.219	-3.643 (**)
Wm	real wage all males	-1.181	-4.163 (**)
GDP	real GDP	0.638	-3.295 (**)

Notes:
(1) All regressions included three lags and intercept
(**) The hypothesis of unit root is rejected at the 5% level.
(*)The hypothesis of unit root is rejected at the 10 % level

Table 5: Results for Male Population 15-25 years old. Sample: 1960-1998

15-25 years old Males	Total Employment (1)	Wage Employment (2)	Self-employment (3)	Participation (4)	Unemployment (5)
Constant	.015 (.028)	.013 (.023)	.023 (.011)	.074 (.042)	.061 (.021)
Y(t-1)	-.0599 (.06)	-.0571 (.056)	-.43 (.19)	-.128 (.075)	-.182 (.103)
Log *Index*	-.0148 (-.006)	-.0167 (-.005)	.001 (.002)	-.0131 (.006)	.001 (.010)
ΔLog(wages(t-1))	-.022 (.025)	-.013 (.023)	-.007 (.007)	-.033 (.023)	-.030 (.040)
GDP Growth	.434 (0.076)	.428 (0.070)	-.012 (.023)	.042 (.074)	-.692 (.131)
Adj. R2	.560	.559	.40	.20	.499
DW	2.28	2.21	2.09	1.99	2.08
T*R2 (Prob.)	4.48 (.21)	3.14 (.37)	3.97 (.26)	2.21 (.53)	0.75 (.86)

Notes to Table 5:All dependent variables with the exception of unemployment rates are measured as % of working age population and in first differences. Standard Errors reported in Parenthesis. In addition to the variables reported in this table, all specifications include ΔY(t-1) as a regressor. T*R2 reports the critical value of the Breusch-Godfrey Serial Correlation Test. In parenthesis, the probability of no-autocorrelation.

Table 6: Results for Male Population 26-50 years old. Sample: 1960-1998

26-50 years old Males	Total Employment (1)	Wage Employment (2)	Self-employment (3)	Participation (4)	Unemployment (5)
Constant	.151 (.085)	.06 (.064)	.147 (.044)	.28 (.15)	.03 (.01)
Y(t-1)	-.187 (.097)	-.120 (.097)	-.68 (.199)	-.298 (.163)	-.22 (.101)
Log *Index*	.003 (.006)	.001 (.006)	-.002 (.003)	3.7E-05 (.001)	-.0039 (.006)
ΔLog(wages(t-1))	-.027 (.028)	-.040 (.025)	-.006 (.017)	.003 (.007)	.03 (.027)
GDP Growth	.345 (.079)	.30 (.071)	.087 (.048)	-.006 (.02)	-.35 (.07)
Adj. R2	.38	.37	.30	.124	.42
DW	2.22	2.45	1.91	1.88	2.15
T*R2	2.45 (.48)	6.12 (.10)	1.74 (.62)	1.92 (.58)	2.25 (.52)

Notes to Table 6: All dependent variables with the exception of unemployment rates are measured as % of working age population and in first differences. Standard Errors reported in Parenthesis. In addition to the variables reported in this table, all specifications include ΔY(t-1) as a regressor. T*R2 reports the critical value of the Breusch-Godfrey Serial Correlation Test. In parenthesis, the probability of no-autocorrelation.

Table 7: Results for Male Population 51-65 years old. Sample: 1960-1998

51-65 years old Males	Total Employment (1)	Wage Employment (2)	Self-employment (3)	Participation (4)	Unemployment (5)
Constant	.090 (.066)	.07 (.056)	.049 (.036)	.20 (.13)	.04 (.01)
Y(t-1)	-.205 (.106)	-.248 (.152)	-.23 (.135)	-.295 (.187)	-.22 (.10)
Log *Index*	.029 (.011)	.0173 (.011)	.009 (.005)	.014 (.010)	-.017 (.008)
ΔLog(wages(t-1))	-.09 (.039)	-.039 (.034)	-.049 (.017)	.0268 (.031)	.06 (.027)
GDP Growth	.454 (.112)	.18 (.11)	.24 (.063)	.174 (.104	-.35 (.09)
Adj. R2	.52	.42	.50	.19	.46
DW	2.01	1.93	2.12	2.16	2.39
T*R2	7.54 (.056)	5.62 (.13)	3.06 (.38)	5.62 (.13)	4.29 (.23)

Notes to Table 7:All dependent variables with the exception of unemployment rates are measured as % of working age population and in first differences. Standard Errors reported in Parenthesis. In addition to the variables reported in this table, all specifications include two lags of the endogenous variable. In addition, in the specifications for total and wage employment two additional lags of GDP and wage growth were included to correct for residual autocorrelation. T*R2 reports the critical value of the Breusch-Godfrey Serial Correlation Test. In parenthesis, the probability of no-autocorrelation.

Table 8: Relative and Absolute Dismissal Costs: Male 15-25 years old

15-25 years old Males	Total Employment (1)	Wage Employment (2)	Self-employment (3)	Participation (4)	Unemployment (5)
Log *Index2*	-.006 (.010)	-.009 (.009)	.002 (.003)	-.016 (.010)	-.023 (.015)
Log *Index20* – Log *Index2*	-.0167 (.009)	-.0152 (.009)	-.0003 (.0029)	-.0005 (.01)	.031 (.014)
Adj. R2	.57	.56	.38	.18	.55
DW	2.20	2.15	2.12	1.99	2.03
T*R2 (Prob.)	4.83 (.18)	3.32 (.34)	3.83 (.27)	2.68 (.44)	1.82 (.61)

Notes to Table8: All dependent variables with the exception of unemployment rates are measured as % of working age population and in first differences. Standard Errors reported in Parenthesis. In addition to the variables reported in this table, all specifications include a constant, Y(t-1) , GDP growth, Real Wage growth and ΔY(t-1) as regressors. T*R2 reports the critical value of the Breusch-Godfrey Serial Correlation Test. In parenthesis, the probability of no-autocorrelation.

Table 9: Results for the Overall Population: 15-65 years old. Sample: 1960-1998

51-65 years old Males	Tot. Emp. (1)	Tot. Emp. (2)	Wage Emp. (3)	Wage Emp. (4)	Self-Emp. (5)	Self-Emp. (6)	Part. (7)	Part. (8)	Unem. (9)	Unem. (10)
Constant	.07 (.05)	.06 (.05)	.01 (.03)	.014 (.035)	.05 (.028)	.05 (.029)	.10 (.08)	.09 (.08)	.04 (.01)	.034 (.01)
Y(t-1)	-.139 (.08)	-.125 (.08)	-.05 (.07)	-.03 (.07)	-.382 (.164)	-.377 (.171)	-.13 (.10)	-.11 (.11)	-.19 (.10)	-.15 (.10)
Log *Index*	.001 (.016)	.	-.006 (.005)	.	.003 (.002)	.	-.003 (.004)	.	-.007 (.008)	.
Log Index2	.	.012 (.009)	.	.004 (.007)	.	.002 (.004)	.	-.005 (.006)	.	-.022 (.011)
Log Index20-Log Index2	.	-.014 (.009)	.	-.016 (.007)	.	.0007 (.0041)	.	-.0001 (.005)	.	.018 (0.011)
ΔLog(W(t-1))	-.078 (.03)	-.071 (.03)	-.070 (.024)	-.062 (.023)	-.005 (.011)	-.005 (.012)	-.035 (.017)	-.034 (.017)	.019 (.032)	.015 (.031)
GDP growth	.39 (.07)	.39 (.07)	.31 (.06)	.31 (.05)	.08 (.03)	.08 (.03)	.06 (.04)	.06 (.04)	-.42 (.09)	-.43 (.08)
Adjusted R2	.52	.55	.58	.63	.33	.30	.21	.19	.45	.49
DW	2.09	2.09	2.01	2.02	2.04	2.01	2.07	2.06	2.27	2.25
T*R2 (Prob.)	3.02 (.38)	4.29 (.23)	3.31 (.34)	3.75 (.28)	3.24 (.35)	3.41 (.33)	2.10 (.55)	2.09 (.55)	2.86 (.41)	3.82 (.28)

Notes to Table 10: All dependent variables with the exception of unemployment rates are measured as % of working age population and in first differences. Standard Errors reported in Parenthesis. In addition to the variables reported in this table, all specifications include two lags of the endogenous variable. In addition, in the specifications for total and wage employment two additional lags of GDP and wage growth were included to correct for residual autocorrelation. T*R2 reports the critical value of the Breusch-Godfrey Serial Correlation Test. In parenthesis, the probability of no-autocorrelation.