Estudios de Economía. Vol. 34 - Nº 2, Diciembre 2007. Págs. 109-139

JOB SECURITY AND THE AGE-COMPOSITION OF EMPLOYMENT: EVIDENCE FROM CHILE

COSTOS DE DESPIDO Y LA COMPOSICIÓN ETARIA DEL EMPLEO: EVIDENCIA PARA CHILE

Carmen Pagés
Claudio E. Montenegro*, **

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. Comparing results for men and women and using measures of relative dismissal costs, we find that the adverse effect of job security on youth employment is driven by the link between severance payments and tenure. We also find that job security does not have a significant impact on overall aggregate employment, participation or unemployment rates.

Key words: *job security, employment composition, severance pay, Chile, job tenure.*

Resumen

Este trabajo desarrolla y prueba un mecanismo por el cual los costos de despido afectan la composición etaria del empleo. Este mecanismo se basa en el costo relativo de despedir trabajadores jóvenes versus trabajadores viejos, resultando de regulaciones laborales asociadas a la antigüedad en el empleo. Usando 39 encuestas anuales de hogares consecutivas para Chile, encontramos que

---

* Inter-American Development Bank (IDB) and the World Bank, respectively.

** For comments, please contact Carmen Pagés at carmenpag@iadb.org or Claudio E. Montenegro at cmontenegro@worldbank.org. We gratefully acknowledge the assistance of Leticia Cuevas, Yolanda Marshall and the Department of Economics at the Universidad de Chile for invaluable information on Chilean labor codes and the data used in this study. We are also very grateful to F. Alvarez, P. Beaudry, G. Becker, J.J. Dolado, S. Edwards, D. Hamermesh, A. Kugler, E. Lazear, and the participants at various seminars at the World Bank and IDB for their invaluable comments. This paper reflects the opinions of the authors and not necessarily those of the IDB or the World Bank.

los costos de despido están asociados con una baja sustancial en el cuociente empleo-población de los trabajadores asalariados jóvenes. En contraste, no encontramos esta caída en las tasas de empleo de trabajadores jóvenes por cuenta propia o de los asalariados de mayor edad. Comparando los resultados para hombres y mujeres usando mediciones de los costos relativos de despido, encontramos que el efecto negativo de los costos de despido sobre el empleo juvenil está dado por la relación entre las indemnizaciones por despido y la antigüedad en el empleo. También encontramos que los costos de despido no tiene un impacto significativo a nivel agregado en el empleo, en la participación laboral o en las tasas de desempleo.

Palabras clave: *costos de despido, composición del empleo, indemnización por despido, Chile, antigüedad en el empleo.*

JEL Classification: *E24, J23, J65.*

1. Introduction

In most countries, job security provisions were set up with the aim of protecting workers against economic fluctuations. However, economic theory and empirical evidence suggest that restrictions on firing workers decrease job creation and job destruction flows, reduce the speed at which employment adjusts to shocks, and increase the incidence of long-term unemployment.[1] Thus, even when some workers may benefit from these provisions, others, less fortunate, may be hurt.

Restricted inflows to employment and long-term unemployment will affect some workers more than others depending on the characteristics of workers and the design of job security policy. In this paper, we explore and test a mechanism by which job security may bias employment against young workers and in favor of older ones. Our mechanism is based on the explicit link between severance pay and tenure found in most countries in which job security is mandatory. As severance pay increases with 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. Therefore, by reducing hiring rates and increasing youth layoff rates, tenure-based job security may result in lower young-to-older-workers employment rates.

We are not the first examining the effect of job security on youth employment. Across a panel of OECD countries, Lazear (1990) finds some evidence that young-to-older-workers employment rates are negatively associated with job security provisions. Yet, his results are not robust to alternative econometric specifications. Nickell (1997) finds some additional support for a prime-age

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), Grubb and Wells (1993), Hamermesh (1995), Hamermesh and Pfann (1996a), Hamermesh and Pfann (1996b), and Nickell (1997).

employment bias in high job security countries. Bertola, Blau, and Kahn (2002) find that job security increases the unemployment rate of young workers relative to other groups. More recently, Kahn (2007) finds that higher job security (as measured by the OECD) increases the relative incidence of joblessness for women, immigrants and young workers as well as the relative incidence of temporary employment among these groups. In these studies, higher unemployment or lower relative employment for the youth is expected because job security reduces hiring, which should particularly affect the youth. Most rely on cross country differences in job security to identify its effects.

In this paper we take a different route. We develop our main argument in the context of a simple model and then we test its 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] By relying on this time-variability and using the wealth of information contained in cross-section data, we reduce the scope for omitted variable bias. Second, the Chilean labor market reforms have involved shifts in the relative costs of dismissing short versus long-tenure workers, which provides 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 many OECD and Latin American countries. This implies that these results can offer valuable lessons for other countries. 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. This study also fills a gap in the literature, since relative to OECD countries, there are few studies analyzing the effects of employment protection in developing countries.[3]

We use the wage and employment information regarding male workers 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 by our model, 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 rates and a 1.3 percentage point decline in participation rates. After 10 years, the estimated decline in employment surpasses 15 percentage points. This decline is mostly absorbed by a fall in participation, with little or no effect, on youth unemployment rates.

We then explore the channels by which job security alters the age-composition of employment. We distinguish between a "barrier-to-entry" effect and the relative firing cost mechanism developed in this paper. Our findings strongly suggest

---

[2] See Heckman and Pagés (2004).

[3] Important exceptions are the studies analyzing the effects of job security in India. See for example Fallon and Lucas (1991), Fallon and Lucas (1993), Besley and Burgess (2004), Dutta Roy (2004), and Ahsan and Pagés (2007).

that job security affects young employment rates through its link to tenure. We obtain this conclusion out of two separate findings. First, as suggested by our model, young and older female workers are less affected by job security than their male counterparts. Second, youth employment rates for male workers are more sensitive to relative than to absolute firing costs.

Finally, we provide results for the overall sample. Our findings suggest that job security provisions do not have a significant effect on total employment, participation or unemployment rates.

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 build 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's tenure at a firm.

We assume a continuum of firms operating with a technology that has constant returns to scale and uses 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 a total of three periods, after which they retire from the labor market. The timing of firms' decisions is as follows: In each period, and depending 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 a dismissal, firms have to pay a severance penalty that increases with a worker's 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 θ_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 θ_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, for simplicity, assumed to be constant across a worker's life and across states.[4] Expression (1) indicates that the value of continuing a match with a worker that has been at firm for two years (from now on a two-period worker) is the maximum between what an employer obtains if the worker remains employed and what the firm has to pay if the worker is dismissed. If the per-period productivity is given to be 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 at a firm for only one period (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 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.[5] That is:

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

In this simple set up, a firm operates "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 that the worker is fired at 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 that a firm falls into 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 the employer is likely to incur many periods of losses before emerging into a

4 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 an extra notation to the model.

5 The assumption that hiring costs depends on market tightness is found in matching models, such like Mortensen and Pissarides (1997).

good state. Moreover, young workers are worse investments than older ones since old workers will retire sooner at no cost for the firm. 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, the likelihood that young workers are dismissed first increases with the slope of the severance pay-tenure profile.

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 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.

2.2. Impact of a Labor Reform

Proposition 2. *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 $SW_{2b} = -\alpha F$ but $\theta_b - w > -\alpha' F$. In such case, $SW_{2b}' = \theta_b - w < SW_{2b}$ 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 a bad state of nature leads to higher future firing costs for these workers. It is clear then, 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 benefit by an increasing severance pay profile. 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.

2.3. Higher Exogenous Turnover and Labor Reforms

Assume that certain workers have higher quit rates than others and that this characteristic is unrelated to job security policy. We denote the quit rate as $1 - q$ and assume the existence of two groups of workers $h = m, f$, for which $q^f < q^m$. The corollary of proposition (2) is that the impact of a higher α will be stronger for the group with lower exogenous turnover rates. In order to see this, it is useful to write conditions (1) -(3) including the quit rate. Thus, $SW_{2i}^{m} = SW_{2i}^{f} = Max\{\theta_i - w, -\alpha F\}$, $SW_{1i}^{f} = Max\{\theta_i - w + \beta q^{f} E_i SW_{2i}^{f}, -F\}$ and $SW_{1i}^{m} = Max\{\theta_i - w + \beta q^{m} E_i SW_{2i}^{m}, -F\}$ whereas $SW_{0i}^{f} = \theta_i - w + \beta q^{f} E_i SW_{1i}^{f}$ and $SW_{0i}^{m} = \theta_i - w + \beta q^{m} E_i SW_{1i}^{m}$. Assume $\alpha' > \alpha$, $SW_{2b}^{m} = SW_{2b}^{f} = -\alpha F$ and $SW_{2b}^{m'} = SW_{2b}^{f'} = \theta_b - w > -\alpha' F$. In addition, assume $SW_{2g}^{m} = SW_{2g}^{f} = SW_{2g}^{m'} = SW_{2g}^{f'} = \theta_g - w$. Then $\Delta SW_{2b}^{f} = \Delta SW_{2b}^{m} = \alpha F - (\theta_b - w) < 0$, $\Delta SW_{1b}^{h} = \beta q^{h} \rho \Delta SW_{2b}^{h}$ and $-SW_{1b}^{h} - F \leq -\Delta SW_{1b}^{h} \leq 0$, for $h = m, f$. And since $q^{f} < q^{m}$ it follows that $0 \geq \Delta SW_{1b}^{f} \geq \Delta SW_{1b}^{m}$ and $0 \geq \Delta SW_{0i}^{f} \geq \Delta SW_{0i}^{m}$.

Thus, an increase in the rate at which severance pay increases with tenure implies a higher decline in the continuation value of the workers with lower quit rates. Consequently, workers in group *m* would have a lower probability of being hired and a higher probability of being laid off after one period than workers in group *f*. We will make use of this distinction later in our empirical work.

It is easy to see that the predictions of this model hold in the case that workers 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 a policy change takes the form of reducing the maximum, say to $S' < S$. Those workers whose continuation values prior to the policy change were above *–SF*, but below $-S'F$, can be affected 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 present and future lower dismissal costs in increasing the continuation value of a worker is lower for workers that are closer to retirement age, and higher for those who have more years ahead. Therefore, less tenured workers benefit more from the reduction in dismissal costs, and their expected continuation 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, as he also states, the presence of borrowing constrains or lack of trust is likely to limit the extent of such transfers, and therefore the extent 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 focus 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 (Heckman and Pagés, 2004). These 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 modeled 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 of 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.

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 especially 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 position of labor courts as mediators between firms and workers. For example, in the period from 1966 to 1981, a dismissal originated by 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, Echeverría, and González (2007)). In such cases,

6 A few exceptions are Rama (1995) and Heckman and Pagés (2004) who use a cross-country sample of developing and developed countries to study this issue. See also Fallon and Lucas (1993).

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 do 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 in 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, 1978 Decree 2,200	1 month	Economic needs were 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 were 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 were not considered just cause for dismissal any more	zero	1 month' wage per year of work with a maximum of 150 days	All workers
1990-1998 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

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 on the outcome of the judiciary process.

Our approach is to compute an index combining information on notice periods, compensation for dismissal, the likelihood that a 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:

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

where δ is the probability of remaining in a job, δ^{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_{t+i}^{jc} is the mandated severance pay in such event to a worker that has been i years at the firm, and finally, SP_{t+i}^{uc} 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 the 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 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 a firms' economic difficulties are 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 turnover rates in the United States. 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 Haltinwanger (1992) estimate that U.S. job destruction rates average 12% a year. We also assume that the maximum tenure at a firm is 25 years. Finally, we compute the discount rate based on the fact that Chilean real interest rates averaged 8.4% during the 1960-1998 period. The resulting index series is plotted in Figure 1.

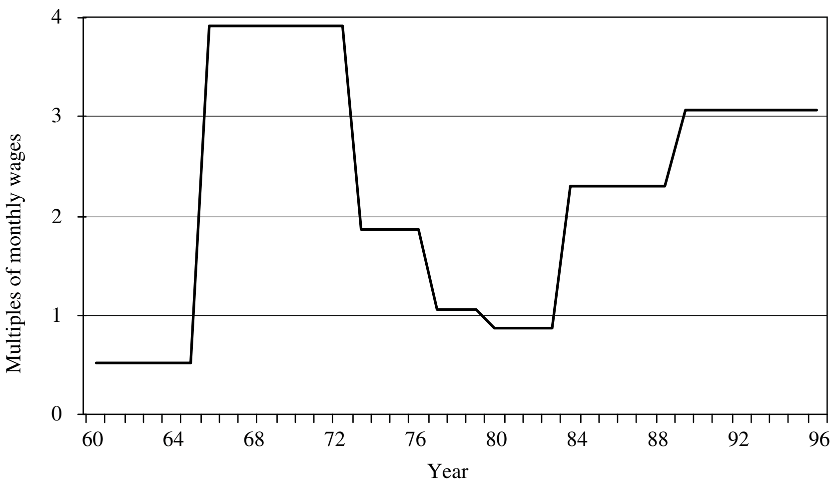
The index exhibits a maximum during the 1966-1973 period and a minimum during 1981 to 1984. After 1985, and especially after 1990, job security increased

TABLE 2
PARAMETERS USED TO COMPUTE *INDEX*

	β	δ	b	a	Sp_{fc}	Sp_{uc}
1960-65	.92	.88	1	1	0	0
1966-73	.92	.88	1	.2	0	(1)
1974-77	.92	.88	1	.5	0	(2)
1978-80	.92	.88	1	.8	0	(2)
1981-84	.92	.88	1	.8	0	(3)
1985-90	.92	.88	1	0	0	(3)
1991-98	.92	.88	1	.9	(4)	(5)

Notes to Table 2: To compute β we use the fact that the average real interest from 1960-1998 was 8.4%. To compute δ we assume that the average Chilean turnover rate without employment protection would be similar to the US one. According to Davis and Haltinwanger (1992) average turnover rates average 12% a year in the United States. (1) Corresponds to one month's pay per year of work augmented in three months to capture the average payments in 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 11 months upper limit. (5) 1.2 month's pay per year of work with 11 months upper limit. We assume the maximum tenure a worker can attain at a firm is 25 years.

FIGURE 1
INDEX OF EMPLOYMENT PROTECTION IN CHILE



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

In our empirical estimates we use the period 1960-1998 for a number of reasons. First, since that 1998, there have not been any changes in severance pay. In addition, at least since the year 2000, there has been an ongoing policy debate on labor reform, without much real reform. The anticipation of reforms, however, could have changed employment dynamics in ways difficult to predict since during the discussion it was not clear whether reforms would end up reducing or increasing job security. Finally, in 2001 the parliament approved a new unemployment insurance law based on a mixed scheme of individual saving accounts and some subsidies from the government. No country has implemented an unemployment insurance system with such design and therefore the effects of this new law on the employment rates across age groups are hard to predict and control for.

We use data obtained from 39 consecutive household surveys to construct annual time series on employment, unemployment, participation rates and a wage index during the period 1960-1998. The data was collected by the University of Chile. 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, the Grand Santiago area was home of about one third of Chile's total population, and a higher proportion of GDP.[7]

In our model, job security affects the age-composition of employment because severance pay is linked to tenure and tenure is related to age. Complementary information obtained from the CASEN National Household Surveys indicates that average tenure is monotonically increasing in age both for women and men, although mean tenure for women increases at a slower rate (See Figure 2 plotting mean tenure by age for women and men).[8, 9]

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 and gender. Complementary data from the 1987 CASEN national surveys indicate that our age split offers a good approximation of expected time at the firm and expected dismissal costs for each age group (see Figure 3 for the distribution of tenure of male workers by age group). In 1987, for instance, approximately 70% of male workers 15-25 years old had been less than 2 years in their current job. This proportion was only 35% among male workers 26-50, and 20% among male older workers (51-65).

7 The use of metropolitan instead of national data may introduce some biases induced by migration flows following policy changes. Yet, the magnitude of the bias is likely to be small since inflows and outflows will tend to offset each other.

8 The CASEN (Encuesta de Caracterización Socioeconómica Nacional) is a nationally and regionally representative Chilean household survey carried out by the Ministerio de Planificación y Cooperación, through the Department of Economics of the University of Chile. Questions are asked at the household and the individual level. The sample sizes are quite large. For instance, in the year 1996, the sample included information on 134,262 individuals.

9 We also have information for 1996. Mean tenure in this latter year looks very similar, although slightly lower than mean tenure in 1987.

FIGURE 2
MEAN TENURE BY AGE
(Moving average)

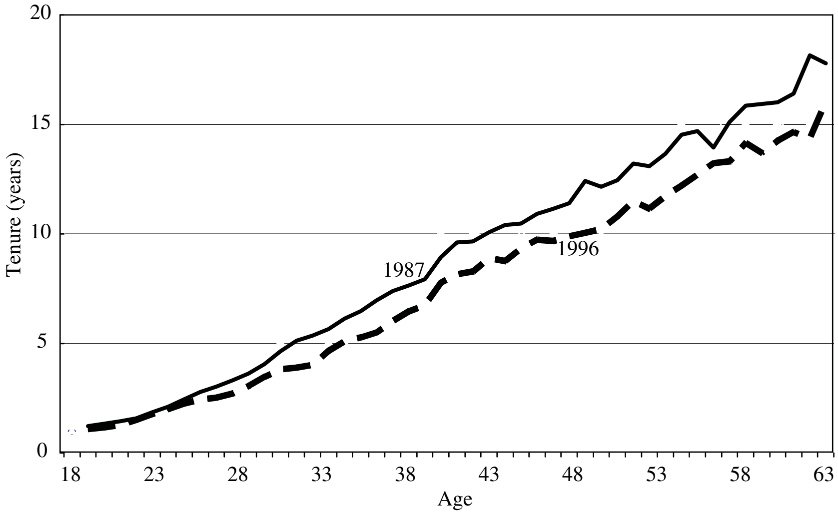
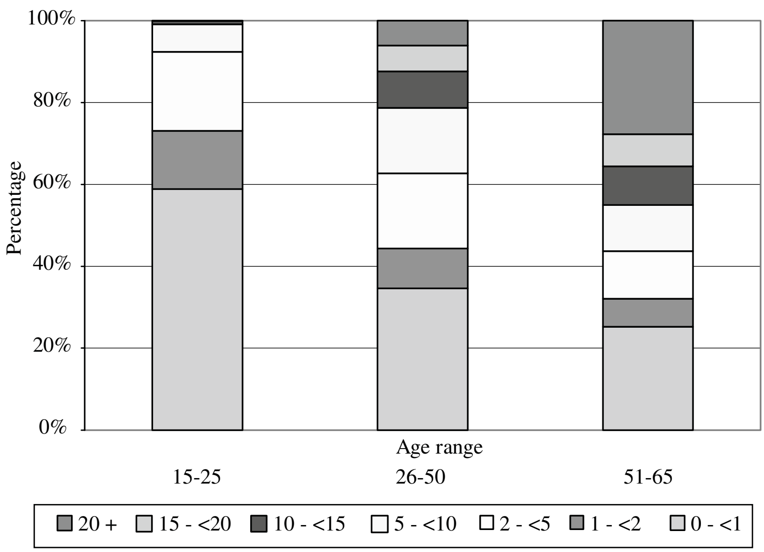


FIGURE 3
TENURE BY AGE RANGE IN 1996



We further split employment data between self-employment, in which regulations do not hold, and wage-employment, in which regulations may or may not hold. The University of Chile surveys do not contain information on whether wage workers hold formal or informal jobs. Although, this is a serious shortcoming, data from the CASEN Household Survey indicates that about 70% of wage workers are formal. Therefore if regulations are to have any effect, this effect should be reflected in the wage-employment series.

Table 3 provides summary statistics for our data. Approximately 45% of the population 15-25 years old was employed (40% in the wage-employment sector and 5% in the self-employment sector). Employment-to-population rates averaged 89% for prime-age workers and 67% among the older age group. Unemployment rates for the overall male population averaged 10%, but this number was much higher for the young (18%). Mean wages increased with age, whereas female participation, employment rates and wages were much lower for women than for men.

Figure 3 highlights 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 during a bad period. 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 during bad times, and (ii), increased job security reduces hiring rates, which, as shown in Figure 3, are especially 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, although higher job security will tend to reduce hiring rates, since new entries 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 Figure 4 seems to confirm these predictions. The shaded areas correspond to periods in which labor legislation was stricter, 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 during this period. 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 (Figure 5); 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.

TABLE 3
DATA SUMMARY STATISTICS

Variable	Obs	Mean	Std. Dev.	Min	Max
GDP Growth	38	0.040917	0.054946	–0.134450	0.111345
Emp-to-pop. Males 15-25	39	0.455138	0.070880	0.330357	0.600946
Wage-Emp-to-pop Males 15-25	39	0.398883	0.070428	0.280907	0.541010
Self-Emp-to-pop Males 15-25	39	0.056224	0.008574	0.035922	0.074627
Participation Rate Males 15-25	39	0.552113	0.054872	0.492523	0.667981
Unemp. Rate Males 15-25	39	0.177950	0.073366	0.080529	0.388818
Emp-to-pop. Males 26-50	39	0.887651	0.046998	0.745203	0.944797
Wage-Emp-to-pop Males 26-50	39	0.669857	0.043902	0.552239	0.732780
Self-Emp-to-pop Males 26-50	39	0.217638	0.015585	0.184017	0.256764
Participation Rate Males 26-50	39	0.961154	0.007800	0.942044	0.975224
Unemp. Rate Males 26-50	39	0.076616	0.045619	0.019071	0.222037
Emp-to-pop. Males 51-65	39	0.672282	0.070536	0.498084	0.790816
Wage-Emp-to-pop Males 51-65	39	0.397385	0.048641	0.272884	0.475836
Self-Emp-to-pop Males 51-65	39	0.274853	0.030732	0.191564	0.338435
Participation Rate Males 51-65	39	0.734261	0.043938	0.637306	0.826531
Unemp. Rate Males 51-65	39	0.086440	0.054143	0.015385	0.259259
Emp-to-pop. Males 15-65	39	0.707278	0.057683	0.557678	0.791482
Wage-Emp-to-pop Males 15-65	39	0.535722	0.050761	0.422460	0.616675
Self-Emp-to-pop Males 15-65	39	0.171458	0.013356	0.135218	0.194512
Participation Rate Males 15-65	39	0.787080	0.026029	0.743418	0.839300
Unemp. Rate Males 15-65	39	0.102346	0.053293	0.033914	0.263618
Wage-Emp-to-pop. Females 15-25	39	0.283574	0.044800	0.208000	0.365500
Wage-Emp-to-pop. Females 26-50	39	0.3387	0.0593	0.2467	0.4565
Wage-Emp-to-pop. Females 51-65	39	0.1490	0.0326	0.10937	0.2478
Wage Males 15-25	39	261.68	127.159	96.72	513.41
Wage Males 26-50	39	532.94	240.68	203.10	1004.01
Wage Males 51-65	39	550.97	239.23	223.40	1065.17
Wage Males 15-65	39	470	217.58	173.66	890.62
Wage Females 15-25	39	265.36	132.98	92.77	543.22
Wage Females 26-50	39	413.23	203.95	143.20	837.91
Wage Females 51-65	39	422.82	233.30	126.05	1070.89
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

We now provide a more formal test of differences in behavior across age groups. First, we test the implications of our model by assessing whether employment, unemployment, and participation rates exhibit an age-specific response that corresponds to what is predicted by our model. We focus our study on male workers because we expect a stronger impact of job security on these workers. Second, we examine the channels by which job security may alter the age-composition of employment. We do so in two alternative ways: firstly, we assess whether the age-response to labor reforms is different for women than for men as predicted by our model. Secondly, 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.

FIGURE 4
WAGE-EMPLOYMENT TO POPULATION RATES MALES

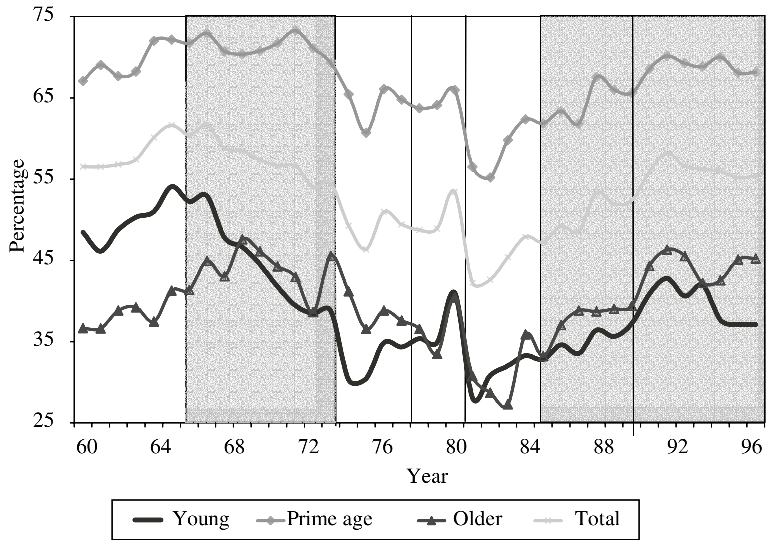
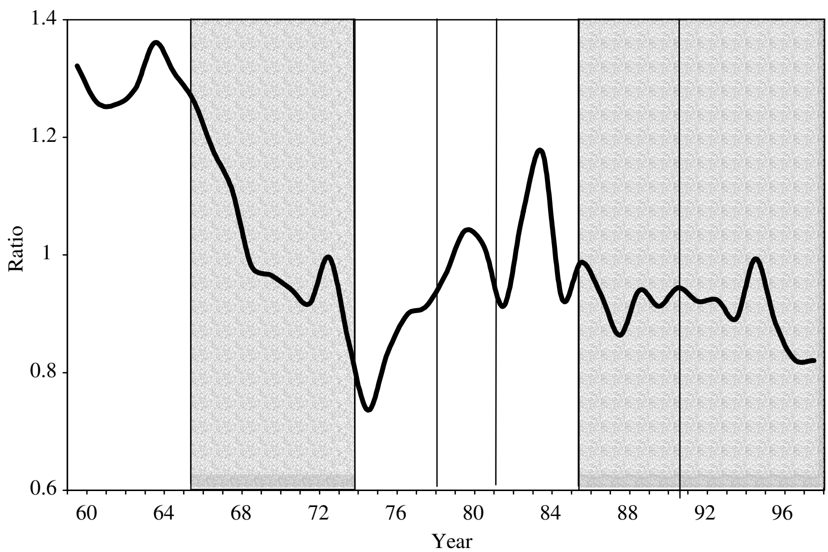


FIGURE 5
YOUNG-TO-OLD WAGE-EMPLOYMENT-TO-POPULATION RATES



3.3 Age-Specific Response to Job Security Provisions

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

$$\text{(eq1)} \qquad \Delta Y_{jt} = c_j + \lambda_j (Y^*_{jt} - Y_{jt-1}) + B_j(L)\Delta Y_{jt-L} + \varepsilon_{jt}$$

where:

$$\text{(eq2)} \quad Y^*_{jt} = B_{1j}\Delta \log GDP_t + B_{2j}\Delta \log(wage_{jt}) + B_{3j}\log(index_t)$$

and 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 (as % of total labor force) for age group j in period t. This specification allows for inertia in the adjustment of variables to their long-run equilibrium levels. If, for example, $Y^*_{jt} - Y_{jt-1}$ is positive, the desired rate, Y^*_{jt} is larger than the last period's rate and the endogenous variable increases by a fraction of this deviation. The larger is λ_j, the faster the speed of adjustment of the variables. Aggregation over units with different adjustment costs may lead to further lags of the endogenous variable. We therefore allow for these additional lags and test whether their inclusion is rejected by the data.

We posit that the long-run equilibrium rates of employment, participation and unemployment depend on Gross Domestic Product (GDP) growth, age-specific real wage growth and on the level of employment regulations. The inclusion of wages and GDP in growth rates rather than in levels is justified by our definition of the endogenous variable as a ratio that can only vary between zero and one. This choice is also justified by the presence of unit roots in Y_{jt} and the fact that none of the Y_{jt} variables were cointegrated with GDP or wages in levels. The coefficient B_{3j} denotes the long-term elasticity of age-specific employment, unemployment or participation to changes in job security. Age specific real wages were constructed out of the same University of Chile survey data that was used to construct the employment, unemployment and participation series and are deflated by the Consumer Price Index, CPI. The wage series measure changes in real hourly earnings of male workers employed full-time in the wage-employment sector. The GDP series was obtained from the World Bank Development Indicators and the CPI data from the Statistical Bulletin of the Central Bank of Chile (see Table 3 for summary statistics of the variables).

Prior to estimating expressions (eq1) and (eq2) we pretest the time series for their order of cointegration performing an augmented Dickey-Fuller (*ADF*) test for all the variables used in this study both in levels and first differences. In all cases, we assume three lags and a non-zero intercept. The results summarized in Table 4 suggest that almost all variables are integrated of order one, I(1), that is, they contain a unit root in levels but their first differences are stationary. Although it is well known that unit root test have low power, the possible existence of unit roots requires being careful with the estimation procedure. The *ECM* described above will be well specified if Y^*_{jt} is cointegrated with Y_{jt-1}. In order for this to be the case there must exist at least one cointegration vector β such that the linear combination $Y_{jt-1} + \beta Y^*_{jt}$ is stationary. Table 5 reports the results of performing the Johansen Test for the cointegration of Y_{jt-1} and Y^*_{jt}. We reject the hypothesis of no-cointegration for all the employment, unemployment and participation variables used in this study since in all cases, there is at least one cointegration relation (CR). Therefore, the ECM is well defined and can be estimated once expression (eq2) is substituted in expression (eq1).

TABLE 4
AUGMENTED DICKEY-FULLER UNIT ROOT TEST

Critical value 1%	–3.682	
Critical value 5%	–2.972	
Critical value 10%	–2.618	
Variable label	ADF test Variable in Levels	ADF test Variable in First Diff.
employment rate males 15-25	–1.642	–2.487
wage-employment rate males 15-25	–1.743	–2.428
self-employment rate males 15-25	–2.004	–3.417 (**)
participation rate males 15-25	–1.507	–3.082 (**)
unemployment rate males 15-25	–1.728	–3.390 (**)
employment rate males 26-50	–1.783	–3.331 (**)
wage employment rate males 26-50	–1.462	–3.043 (**)
self-employment rate males 26-50	–2.975	–3.441 (**)
participation rate males 26-50	–1.943	–3.234 (**)
unemployment rate males 26-50	–1.783	–3.525 (**)
employment rate males 51-65	–0.929	–2.696 (*)
wage employment rate males 51-65	–1.213	–2.655 (*)
self-employment rate males 51-65	–0.967	–3.075 (**)
participation rate males 51-65	–0.970	–3.236 (**)
unemployment rate males 51-65	–1.637	–3.05 (**)
employment rate males 15-65	–1.380	–2.749 (*)
wage-employment rate males 15-65	–1.308	–2.55
self-employment rate males 15-65	–2.455	–3.35 (**)
participation rate males 15-65	–1.3511	–3.281 (**)
unemployment rate males 15-65	–1.64	–3.211 (**)
Wage employment rate females 15-25	–1.953	–2.566
Wage employment rate females 26-50	0.378	–4.516(**)
Wage employment rate females 51-65	–1.214	–3.689(**)
real wage males 15-25	–1.059	–3.701 (**)
real wage males 26-50	–1.268	–4.294 (**)
real wage males 51-65	–1.219	–3.643 (**)
real wage males 15-65	–1.181	–4.163 (**)
real GDP	0.638	–3.295 (**)

Notes to Table 4:
(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

A source of concern is endogeneity, which comes from the inclusion of endogenous lagged variables and wages in the specification. In the case of the lagged endogenous variable, endogeneity is an issue if there is serial correlation in the error term. To minimize this problem we test for serial correlation in all our specifications. 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 some cases the null hypothesis of serial correlation is rejected,

TABLE 5
COINTEGRATION ANALYSIS

		Johansen Cointegration Test			
	# of CR	Likelihood Ratio test for no Cointegration *Critical value 5%=39.89*	Likelihoood Ratio test for at most 1 CR *Critical value 5%=24.31*	Likelihood Ratio test for at most 2 CR *Critical value 5%=12.53*	Likelihood Ratio test for at most 3 CR *Critical value 5%=3.84*
Total Emp. Male 15-25	3 at 5%	64.17	35.03	13.75	3.33
Wage Emp. Male 15-25	3 at 5%	61.29	34.39	13.58	2.53
Self. Emp. Male 15-25	1 at 5%	42.43	23.34	5.68	0.41
Participation Male 15-25	2 at 5%	74.49	33.57	9.78	2.41
Unemp. Male 15-25	1 at 5%	46.56	23.30	5.82	0.006
Total Emp. Male 26-50	2 at 5%	51.41	25.01	5.33	0.06
Wage Emp. Male 26-50	3 at 5%	66.25	36.31	15.92	0.25
Self. Emp. Male 26-50	1 at 5%	41.89	19.10	3.97	0.27
Participation Male 26-50	2 at 5%	58.44	31.51	10.34	0.22
Unemp. Male 26-50	1 at 5%	50.83	23.19	3.56	0.00
Total Emp. Male 51-65	2 at 5%	67.48	30.06	11.06	0.07
Wage Emp. Male 51-65	2 at 5%	68.76	29.79	11.53	0.03
Self. Emp. Male 51-65	2 at 5%	58.71	33.80	10.86	0.17
Participation Male 51-65	2 at 5%	82.62	31.23	8.61	0.11
Unemp. Male 51-65	1 at 5%	43.38	16.04	2.32	0.09
Total Emp. Male 15-65	2 at 5%	63.86	37.00	12.10	0.36
Wage Emp. Male 15-65	3 at 5%	71.08	41.08	16.24	0.87
Self. Emp. Male 15-65	2 at 5%	51.09	26.03	7.33	0.16
Participation Male 15-65	2 at 5%	78.06	38.34	9.54	0.52
Unemp. Male 15-65	1 at 5%	49.62	22.27	3.81	0.01
Total Emp. 15-65	3 at 5%	79.47	36.10	15.62	0.02
Wage Emp. 15-65	3 at 5%	90.79	36.54	15.55	0.001
Self. Emp. 15-65	2 at 5%	51.84	25.37	9.77	0.34
Participation 15-65	2 at 5%	67.20	34.87	11.29	0.57
Unemp. 15-65	1 at 5%	52.63	23.67	4.12	0.001

Note to Table 5: Johansen cointegration tests assume that there is no deterministic trend in data and three lags interval.

in which case we re-estimate the specification including additional lags of the endogenous and forcing variables. With this change, the null hypothesis of no serial correlation is not rejected in all cases. To deal with the endogeneity of wages, we lag them one period. Given that the null hypothesis of no serial correlation is not rejected, lagged wages are a predetermined variable not correlated with the contemporaneous error term.

Our specification includes wages to ensure that the impact of exogenous changes in the price of labor of different groups of workers are not confounded with the effects of job security. This implies that all our estimates hold wages constant and in particular, that we are not considering the possibility that the costs of firing is absorbed via changes in wages, as suggested by Gruber (1997). However, in a related paper (Montenegro and Pagés, 2004) we find little evidence that in Chile wages adjust to changes in job security.

3.3.1. *Results for Workers 15-25 Years Old*

The first set of estimates corresponds to the group of workers 15-25 years old. The results are summarized in Table 6. 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, youth employment-to-population rates would have declined by approximately 3 percentage points. The long run elasticity of job security is -0.25, that is, a 100% increase in job security would wipe out 25 percent of the youth employment relative to the working age population. Nonetheless, some caution should be exercised in interpreting the LR response as several econometric problems, common when dealing with relatively small samples, were encountered. 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. This possible lack of cointegration means that the coefficient on Y_{t-1} is not different from zero (since it contains a unit root). Therefore, the coefficient on the legislation index can be interpreted as permanently affecting the change in the endogenous variable in the context of our short period sample.

The estimates in column (3) suggest that self-employment rates are 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 suggests 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, and perhaps because there is no unemployment insurance in Chile, column (4) and (5) indicate that a decline in wage-employment is compensated by a reduction in youth participation rates, without any significant effects on unemployment. Thus, our results imply that employment protection is detrimental for young workers, yet a researcher that only analyzed youth unemployment rates would not capture this effect. The last two rows in Table 5 report tests on residual autocorrelation. In all cases, the null hypothesis of no autocorrelation is not rejected.

[10] Since there have been substantial changes in secondary schooling policies and enrollment rates we conducted various additional test to ensure that changes in schooling were not biasing our results. First, we re-estimated the specifications reported in Table 6 with a subsample of people 18-25 years old for whom the expansion in primary and secondary education should have had a lower impact on employment rates. The results obtained with this age group were very similar to the ones reported here. Second, and in order to control for slow and gradual changes in education enrollment, we estimated our model including time trends. Again, the results did not change.

TABLE 6
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	.013	.023	.074	.061
	(.028)	(.023)	(.011)	(.042)	(.021)
Y(t–1)	–.0599	–.0571	–.43	–.128	–.182
	(.06)	(.056)	(.19)	(.075)	(.103)
Log Index	–.0148	–.0167	.001	–.0131	.001
	(–.006)	(–.005)	(.002)	(.006)	(.010)
ΔLog(wages(t–1))	–.022	–.013	–.007	–.033	–.030
	(.025)	(.023)	(.007)	(.023)	(.040)
GDP Growth	.434	.428	–.012	.042	–.692
	(0.076)	(0.070)	(.023)	(.074)	(.131)
Adj. R2	.560	.559	.40	.20	.499
DW	2.28	2.21	2.09	1.99	2.08
T*R2	4.48	3.14	3.97	2.21	0.75
(Prob.)	(.21)	(.37)	(.26)	(.53)	(.86)
LR elasticity	–0.25	–0.29	0.00	–0.10	0.01

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.

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

The results for workers in this age group are summarized in Table 7. In this case, 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 specifications. 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 (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.

TABLE 7
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	.06	.147	.28	.03
	(.085)	(.064)	(.044)	(.15)	(.01)
Y(t-1)	–.187	–.120	–.68	–.298	–.22
	(.097)	(.097)	(.199)	(.163)	(.101)
Log Index	.003	.001	–.002	3.7E–05	–.0039
	(.006)	(.006)	(.003)	(.001)	(.006)
ΔLog(wages(t-1))	–.027	–.040	–.006	.003	.03
	(.028)	(.025)	(.017)	(.007)	(.027)
GDP Growth	.345	.30	.087	–.006	–.35
	(.079)	(.071)	(.048)	(.02)	(.07)
Adj. R2	.38	.37	.30	.124	.42
DW	2.22	2.45	1.91	1.88	2.15
T*R2	2.45	6.12	1.74	1.92	2.25
	(.48)	(.10)	(.62)	(.58)	(.52)
LR elasticity	0.016	0.008	–0.003	0.000	–0.018

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 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.

Columns (3), (4) and (5) indicate 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 8. As in the prime-age case, the specification includes two lags of the endogenous variable and, for the employment and wage employment specifications, two lags of GDP and wage growth to insure no serial correlation of the error term. In this case, the impact of job security has the same sign as 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 indicating that an increase in job security increases overall employment rates for older workers. The size of the coefficient suggests a large change: A 100% increase in the index variable is associated with a 2.9% short run increase in pre-retirement age employment rates. In the long run, the effects of such and increase are much bigger, as we estimate a long run elasticity of 14 %. Estimates for wage and self-employment confirm that *both* types of employment increase with higher job security, albeit wage employment increases more. It is unclear why self-employment increases with job security. Perhaps, severance pay provides

TABLE 8

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	.059	.049	.20	.04
	(.066)	(.052)	(.036)	(.13)	(.01)
Y(t-1)	–.205	–.206	–.23	–.295	–.22
	(.106)	(.139)	(.135)	(.187)	(.10)
Log *Index*	.029	.0131	.009	.014	–.017
	(.011)	(.009)	(.005)	(.010)	(.008)
ΔLog(wages(t-1))	–.09	–.041	–.049	.0268	.06
	(.039)	(.027)	(.017)	(.031)	(.027)
GDP Growth	.454	.19	.24	.174	–.35
	(.112)	(.10)	(.063)	(.104	(.09)
Adj. R2	.52	.41	.50	.19	.46
DW	2.01	2.09	2.12	2.16	2.39
T*R2	7.54	6.39	3.06	5.62	4.29
	(.056)	(.09)	(.38)	(.13)	(.23)
LR elasticity	0.141	0.064	0.039	0.047	–0.077

Notes to Table 8: 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.

the start-up capital necessary to move into self-employment. Finally, columns (4) and (5) suggest that older workers participation rates increase in periods of increased job security (although the coefficient is not statistically significant at conventional levels) while unemployment rates for older workers decline.

The results presented in Tables 6-8 are robust to the introduction of time trends and further lags of the endogenous and exogenous variables. In addition, differences in coefficients between young and older workers are not due to differences in the empirical specification used for the two groups.[11] Our results are also robust to the inclusion of the index variable without the logarithm and to alternative definitions of the regulation variable. In particular, we constructed an alternative index that measures the cost of dismissing a worker with 20 years of tenure along the different policy regimes.[12] The results using this variable are very similar to the ones reported here. Overall, our findings suggest that older workers are the ones that benefit the most from job security provisions. Instead,

[11] We re-estimated our results for young workers using the same empirical specifications we use for older ones (that is, including further lags of exogenous and endogenous variables). The results did not change.

[12] See the next section for a description of this alternative index.

younger workers bear the burden of increased job security because they endure lower hirings and a disproportionate share of firings in recessions.

3.4. Barrier-to-entry or Relative Costs?

The results obtained in the former section indicate that job security alters the age-composition of employment. In this section we investigate the source of this bias. Two alternative explanations can be proposed: The first explanation states that by reducing job creation, job security reduces the employment prospects of those entering in the labor market. We call this explanation ''barrier-to-entry'' effect and was first suggested by Lazear (1990). The second explanation is related to the relative cost of dismissing young versus old workers and has been fledged out in our theoretical section.

We rely on two different procedures to discriminate against these two explanations. We first compare the results for two groups of workers with different expected turnover rates and second, we use a direct measure of relative firing costs to test its impact on male youth employment rates.

3.4.1. *Gender-Age-Specific Response to Labor Reforms*

If job security affects young workers by reducing hires and slowing down entry in the labor market then young women and young men should be similarly affected by a change in regulations. Instead, if the mechanism is the link between severance pay and tenure, section 2.3 suggest a differential response between women and men. Thus, under the hypothesis that women have higher exogenous quitting rates (as suggested by their relatively lower tenure given a certain age), expectations of future cost may induce employers to prefer hiring women rather than men. However, as women and men become older, men acquire more tenure. This implies that in the face of an adverse economic shock, employers may tend to dismiss women first because their relative lower cost of dismissal. In terms of our empirical specifications, the model predicts that regulations will have a smaller effect on the age structure of employment for women than for men. This is because in the context of our model, again under the critical assumption of higher exogenous turnover rates for women-- tenure-based job security creates a larger negative effect on youth male employment and a larger positive effect on older male employment rates than for their female counterparts.

Table 9 reports the results of estimating our empirical model, defined by expressions (eq1) and (eq2), for women 15-25, 26-50, and 51-65 (columns 1-3), with the same specifications used for males in each group. As predicted by our model, job security, as measured by *index*, has a smaller effect on young females than on males. The coefficient is about one third of the one estimated for men and it is not statistically significant. In addition, and as predicted by the model, the effect of job security on females 51 to 65 is lower than for males of in the same age group.

Could these differences be the result of higher informality rates among women than among men, and therefore a lower effect of labor regulations in the female labor market? We found that the CASEN surveys for 1987 and 1996 reveal very similar levels of formalization across gender. For example, in 1996, 69.47% of all employed men and 71.67% of all employed women had a regular indefinite

TABLE 9
BARRIERS-TO-ENTRY VERSUS RELATIVE COSTS OF DISMISSAL:
GENDER DIFFERENCES

	Wage Employment Young Females (1)	Wage Employment Prime-Age Females (2)	Wage Employment Older female workers (3)
Log *Index*	–.0045	0.003	0.0054
	(.004)	(.005)	(0.004)
Adj. R2	.12	.11	.16
DW	2.05	1.83	1.86
T*R2	9.39	2.41	7.07
	(.024)	(.49)	(.069)

Notes to Table 9: Standard errors reported in parenthesis. Specifications in columns (1)-(3) are identical to the ones presented in tables 6-8 (col. 2). T*R2 reports the critical value of the Breusch-Godfrey Serial Correlation Test. In parenthesis, the probability of no-autocorrelation.

contract. Overall, we interpret these differential effects by gender as an important piece of evidence for the type of effects implied by our theoretical model.

3.4.2. *Relative versus Absolute Firing Costs*

Our estimates indicate 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 cost of dismissal across age groups. In this sub-section, we test this prediction constructing measures of relative and absolute dismissal costs.

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.[13] That is,

$$\hat{F} = index2$$

$$\hat{\alpha} = index20 \,/\, index2$$

[13] *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.

Figure 6 presents these indices and Table 10 reports the results of re-estimating our error correcting model 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. As predicted by our model, increasing the relative cost of dismissing older workers leads to a decline in youth total 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. These results suggest that an increase in absolute severance pay, keeping the tenure slope constant, reduces both firings and hiring rates. Instead, an increase in the relative cost reduces hiring while increasing firing rates.

Summarizing, higher 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 decline in youth 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 for at least 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

FIGURE 6
OTHER MEASURES OF JOB PROTECTION

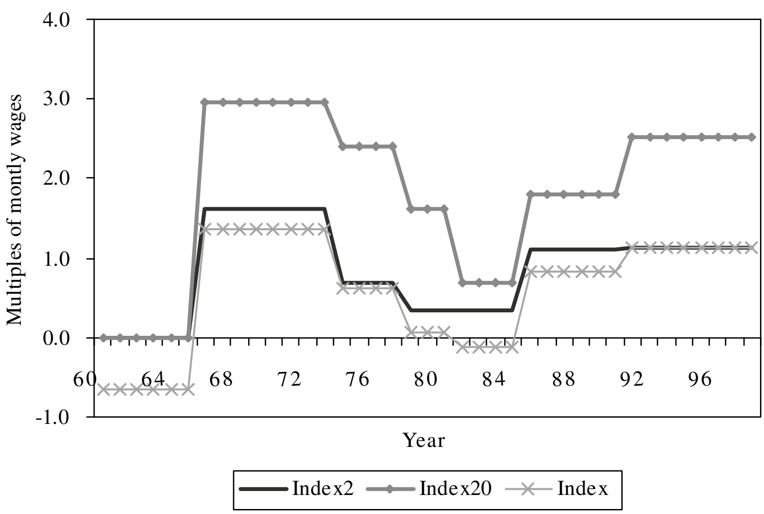


TABLE 10

RELATIVE AND ABSOLUTE DISMISSAL COSTS: MALES 15-25 YEARS OLD

15-25 years old Males	Total Employment (1)	Wage Employment (2)
Log Index2	–.006 (.010)	–.009 (.009)
Log *Index20* – Log *Index2*	–.0167 (.009)	–.0152 (.009)
Adj. R2	.57	.56
DW	2.20	2.15
T*R2 (Prob.)	4.83 (.18)	3.32 (.34)

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 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.

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 11 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 and not statistically significant 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 not the average level of employment or unemployment.

Columns (2) and (4) however, suggest 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 (not reported).

Overall, our estimates for absolute firing costs (F), confirm the theoretical 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

TABLE 11
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)	Part. (7)	Unem. (8)
Constant	.07 (.05)	.06 (.05)	.01 (.03)	.014 (.035)	.05 (.028)	.10 (.08)	.04 (.01)
Y(t-1)	–.139 (.08)	–.125 (.08)	–.05 (.07)	–.03 (.07)	–.382 (.164)	–.13 (.10)	–.19 (.10)
Log *Index*	.001 (.006)	.	–.006 (.005)	.	.003 (.002)	–.003 (.004)	–.007 (.008)
Log Index2	.	.012 (.009)	.	.004 (.007)	.	.	.
Log Index20- *Log Index2*	.	–.014 (.009)	.	–.016 (.007)	.	.	.
ΔLog(W(t-1))	–.078 (.03)	–.071 (.03)	–.070 (.024)	–.062 (.023)	–.005 (.011)	–.035 (.017)	.019 (.032)
GDP growth	.39 (.07)	.39 (.07)	.31 (.06)	.31 (.05)	.08 (.03)	.06 (.04)	–.42 (.09)
Adjusted R2	.52	.55	.58	.63	.33	.21	.45
DW	2.09	2.09	2.01	2.02	2.04	2.07	2.27
T*R2 (Prob.)	3.02 (.38)	4.29 (.23)	3.31 (.34)	3.75 (.28)	3.24 (.35)	2.10 (.55)	2.86 (.41)

Notes to Table 11: 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.

neither higher self-employment rates nor by higher employment rates of the older sections of the labor force.

4. Conclusion

In this paper, we show that job security biases employment in favor of middle-age and older workers. We also show that this bias is rooted in the lower cost of dismissing young workers relative to older ones. Our results suggest that job security provisions affect the composition but not the aggregate level of employment or unemployment.

We believe these results have two important implications for the design of future labor market reforms.

First, labor market reforms will 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 in order to support 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 may explain why job security provisions are tied to tenure in almost all OECD and Latin American countries.

Second, our results give support for 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 employment rates. However, 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.

APPENDIX

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 $\frac{\partial \rho^*(\alpha,\theta_g - w)}{\partial \alpha} < 0$. ■

References

Ahsan, Ahmad, and Carmen Pagés (2007). "Are All Labor Regulations Equal? Assessing the Effects of Job Security, Labor Dispute, and Contract Labor Laws in India". Washington, DC: World Bank Policy Research Working Paper Series 4259.

Anderson, Patricia M. (1993). "Linear Adjustment Costs and Seasonal Labor Demand: Evidence from Retail Trade Firms". *Quarterly Journal of Economics* 108 (4): 1015-42.

Bertola, Giuseppe (1990). "Job Security, Employment and Wages". *European Economic Review* 34: 851-86.

Bertola, Giuseppe, Francine D. Blau, and Lawrence M. Kahn (2002). "Labor Market Institutions and Demographic Employment Patterns". Cambridge, MA: National Bureau of Economic Research Working Paper Series 9043.

Besley, Timothy, and Robin Burgess (2004). "Can Labor Regulation Hinder Economic Performance? Evidence from India". *Quarterly Journal of Economics* 119 (1): 91-134.

Davis, Steven, and John C. Haltinwanger (1992). "Gross Job Creation, Gross Job Destruction, and Employment Reallocation". *Quarterly Journal of Economics* 107 (3): 819-63.

Dutta Roy, Sudipta (2004). "Employment Dynamics in Indian Industry: Adjustment Lags and the Impact of Job Security Regulations". *Journal of Development Economics* 73 (1): 233-56.

Fallon, Peter, and Robert Lucas (1991). "The Impacts of Changes in Job Security Regulations in India and Zimbabwe". *World Bank Economic Review* 5 (3): 395-413.

Fallon, Peter, and Robert Lucas (1993). "Job Security Regulations and the Dynamic Demand for Industrial Labor in India and Zimbabwe". *Journal of Development Economics* 40: 241-75.

Grubb, David, and William Wells (1993). "Employment Regulation and Patterns of Work in EC Countries". Paris: OECD Economic Studies 21.

Gruber, Jonathan (1997). "The Incidence of Payroll Taxation". *Journal of Labor Economics* 15 (3): 72-101.

Hamermesh, Daniel S. (1993). *Labor Demand*. Princeton, NJ: Princeton University Press.

Hamermesh, Daniel S. (1995). "Labour Demand and the Source of Adjustment Costs". *Economic Journal* 105 (430): 620-34.

Hamermesh, Daniel S., and Gerard A. Pfann (1996a). "Adjustment Costs in Factor Demand". *Journal of Economic Literature* 34 (3): 1264-92.

Hamermesh, Daniel S., and Gerard A. Pfann (1996b). "Turnover and the Dynamics of Labor Demand". *Economica* 63(3):359-67.

Heckman, James, and Carmen Pagés, eds. (2004). *Law and Employment: Lessons from Latin America and the Caribbean.* Chicago, IL: University of Chicago Press.

Hopenhayn, Hugo, and Richard Rogerson (1993). "Job Turnover and Policy Evaluation: A General Equilibrium Analysis". *Journal of Political Economics* 101 (5): 915-38.

Kahn, Lawrence M. (2007). "The Impact of Employment Protection Mandates on Demographic Temporary Employment Patterns: International Microeconomic Evidence". *Quarterly Journal of Economics* 117 (521): 333-56.

Lazear, Edward P. (1990). "Job Security Provisions and Employment". *Quarterly Journal of Economics* 58 (4): 757-82.

Maloney, William F. (1999). "Does Informality Imply Segmentation in Urban Labor Markets? Evidence from Sectoral Transitions in Mexico". World Bank Economic Review, vol. 13: 275-302.

Montenegro, Claudio E., and Carmen Pagés (2004). "Who Benefits from Labor Market Regulations? Chile 1960-1998". In: James Heckman and Carmen Pages, (eds.), *Law and Employment.* Chicago, IL: University of Chicago Press.

Mortensen, Dale, and Christopher Pissarides (1997). "Equilibrium, Unemployment Differences, and Fluctuations". Mimeo.

Nickell, Stephen (1986). "Dynamic Models of Labor Demand". In: O. Ashenfelter and Richard Layard, (eds.), *Handbook of Labor Economics.* Amsterdam: North-Holland.

Nickell, Stephen (1997). "Unemployment and Labor Market Rigidities: Europe versus North America". *Journal of Economic Perspectives* 11 (3): 55-74.

Rama, Martin (1995). "Do Labor Market Policies and Institutions Matter? The Adjustment Experience in Latin America and the Caribbean". *Labour* 1995: 243-69.

Romaguera, Pilar, Christian Echeverría, and Pablo González (2007). "Reforming the Labor Market in a Liberalized Economy: The Case of Chile". In: Gustavo Marquez (eds.), *Reforming the Labor Market in a Liberalized Economy.* Washington, DC: Inter-American Development Bank.