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7

Who Benefits from Labor Market Regulations?
Chile, 1960–1998

Claudio E. Montenegro and Carmen Pagés

Introduction

The economic literature has devoted considerable attention to studying the impact of labor market regulations on labor market outcomes. However, the issue of whether some subgroups of workers bear the brunt or enjoy the benefits of such regulations has been much less studied.[1] One notable exception has been the burgeoning literature studying the effect of statutory minimum wages on youth employment. Although this subject remains controversial, many studies have found negative effects of minimum wages on teenagers and young workers.[2] Less attention has been paid to the issue of whether minimum wages particularly affect women versus men or unskilled versus skilled workers. One exception is the study by Lang and Kahn (1998) for the United States, which finds that a rise in the minimum wage shifts the composition of employment in the eating and drinking sector from adults to teenagers and students. Neumark, Schweitzer, and Wascher (2000) also examine the effect of minimum wages across

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1. One reference in this literature is the paper by Bertola, Blau, and Kahn (2002) on the effect of unions' involvement in wage setting on the relative employment of youth, women, and older individuals.

2. Among the most recent studies, Williams and Mills (2001), Partridge and Partridge (1998), and Bazen and Skourias (1997) find a negative relation between minimum wages and youth employment, while Katz and Krueger (1992), Card, Katz, and Krueger (1994), and Card and Krueger (2000) find no evidence of such an effect.

different individuals by focusing on differential impacts of workers at different points in the wage distribution. They find that although wages of low-wage workers increase, hours worked and employment levels decline, reducing earnings for these workers.

Similarly, relatively little attention has been paid to the effect that job security provisions may have on particular subgroups of the labor force. Two recent exceptions are the Organization for Economic Cooperation and Development (OECD) (1999) and Bertola, Blau, and Kahn (2002). The OECD (1999) reports negative, but not statistically significant, effects of job security provisions on youth and prime-age females. Bertola, Blau, and Kahn (2002) find evidence that job security provisions increase the employment rates of male prime-age workers relative to the employment rates of male older workers. They also find evidence that job security provisions are associated with higher employment rates for prime-age women relative to women aged fifteen to twenty-four. Instead, they do not find statistically significant effects on youth relative to prime-age employment rates for male workers or in the distribution of employment across women and men.

In this chapter, we take advantage of the unusual variance in labor market policies in Chile to examine how minimum wages and job security provisions affect different types of workers. We look at the effects of regulations on the distribution of employment by age, and also, by skill, which to our knowledge has not been examined before. To this effect, we use a sample of repeated household surveys spanning the period 1960–1998 and several measures of labor market regulations across time. We make use of cross-section and time series methods to estimate the effect that these policies have on the distribution of employment and on particular subgroups' employment rates. We are able to control for time effects that affect all workers in a similar manner, as well as demographic groups-specific effects of business cycles and labor market institutions. In addition, to assess whether our estimates are reflecting the effect of regulations instead of the effect of some unobservable correlates, we also estimate the effect of labor policy on sectors not covered by regulations. We find large and statistically significant effects on the covered sectors and no effects, or effects going in the opposite direction, on the uncovered sectors.

Our results indicate that labor market regulations are far from neutral. We find that job security provisions and minimum wages reduce the employment rates of youth and the unskilled at the benefit of older and skilled workers. We also find opposite effects of these policies on women's and men's employment shares and rates. Job security provisions tend to benefit men at the expense of women, while the reverse seems to be true for an increase in the minimum wage.

We then explore some explanations for these regularities and, while we cannot fully discriminate among all of them, we are at least able to reject some hypotheses. There is little evidence that the differential effects of job security are driven by differences in labor supply elasticities or wage adjustments across subgroups. Instead, our findings suggest that job security regulations produce unequal shifts in labor demand across groups of workers. Regarding minimum wages, our results tend to fit the predictions of the competitive model for age and skill but not for gender. Contrary to our results, the competitive model predicts higher effects of minimum wages for women because they tend to earn lower wages than men.

The rest of the chapter is organized as follows. Section 7.1 reviews the arguments that predict nonneutral effects of regulations. Section 7.2 describes the evolution of job security and minimum wage regulations in Chile. Section 7.3 describes the data used in our empirical section. Section 7.4 describes the methodology implemented to estimate the effects of regulations on the distribution of employment. Section 7.5 describes our results for both the distribution of employment and the overall effect on employment rates. Finally, section 7.6 concludes.

7.1 Why Regulations May Affect Some Workers Differently

There are a number of reasons to suspect that labor market regulations alter the distribution of employment across subgroups. In the next two subsections, we review the theoretical arguments that predict differential effects of job security provisions and minimum wages across workers of different age, skill level, and gender.

7.1.1 Job Security

Job security provisions are introduced to discourage firms from adjusting their labor forces in the face of adverse economic conditions. However, job security provisions also alter hiring decisions. In good times, firms hire fewer workers because they take into account that these workers may have to be laid off in the future, and that is costly. The overall impact of job security provisions on employment rates is undetermined because it depends on whether the negative effect on layoffs is offset by the reduction in hiring rates.[3]

Job security provisions will have differential effects across subgroups of workers if changes in legislation bring changes in hiring and layoff rates that have a larger impact on some subpopulations than on others. Lazear (1990) conjectured that an increase in job security might act as a barrier, preventing the entry of young workers into the labor market. This is because job security reduces job creation, and entry rates are especially high among youth. This argument, however, does not consider that the effect of

3. See Bertola (1990), Bentolila and Bertola (1990), Bertola (1991), Bentolila and Saint-Paul (1994), Hopenhayn and Rogerson (1993), and Risager and Sorensen (1997), among others, for a theoretical discussion of the effects of job security on employment rates.

lower job creation rates can be offset by lower job destruction rates—which also tend to be large among youth. Pagés and Montenegro (1999) suggest an argument whereby job security provisions may actually *increase* young workers' layoff rates. Their argument is related to the regularity that, across countries, job security is positively related with a worker's tenure. Mandatory severance payments that increase with tenure change the cost of dismissing workers with short tenures relative to workers with more seniority at the firm. In this context, it is expected that job security concentrates layoffs among youth because, other things being equal, young workers tend to have lower average tenures than older workers. If severance pay increases substantially with tenure, and this effect is important, job security simultaneously reduces entry and increases layoffs among youth, resulting in a lower employment share and lower employment rates for this group of workers. Instead, the share of older workers in employment tends to increase due to their relatively lower layoff rates.

Similar reasoning can be used to predict the effect of job security provisions across gender. To the extent that women experience higher rotation and, therefore, have lower average tenure than males at every age, high job security will tend to concentrate layoffs among women. This effect will tend to reduce their employment share relative to men. However, higher turnover rates also imply that stringent job security may be less of an issue when hiring female workers because employers expect them to quit prior to attaining high job security.[4] In this case, employers might be more willing to hire women relative to men, but also more likely to lay them off should bad times arise. The overall effect on female versus male employment rates is undetermined and remains an empirical issue.

It is tempting to extend the former argument to unskilled and skilled workers. If unskilled workers have higher rotation and lower tenures than skilled workers, the same reasoning applies. However, while higher female turnover rates may be motivated by life-cycle decisions exogenous to the employer, such exogeneity is more difficult to claim when explaining the higher rotation of unskilled workers.

The insider-outsider literature provides further arguments for why job security may have a differential effect on the employment rates of different subpopulations.[5] According to this literature, more stringent job security reduces the elasticity of wages to changes in the unemployment rate. When employed workers know their jobs are insured against demand fluctuations, they may be less willing to accept the wage adjustments necessary to reduce unemployment rates. This situation may help to create two kinds of workers: insiders, who hold their jobs and have high wages; and outsiders,

4. See Pagés and Montenegro (1999) for a more formal development of this argument in the context of a partial equilibrium model.
5. See, for instance, Lindbeck and Snower (1988).

who either are unemployed or hold temporary, part-time or fixed-terms jobs without job security.[6] If women, the young, and the unskilled are more likely to be outsiders, then job security (through this wage effect) will bias employment against these groups.

Finally, differences in labor supply elasticity may contribute to differential effects across subpopulations, even if job security brings a uniform change in labor demand across groups. Let us assume that an increase in job security reduces labor demand. If women, the young, and the unskilled have higher labor supply elasticity than the average worker, higher job security would bring a higher decline in employment for these workers than for other groups with a lower elasticity of labor supply.[7]

In summary, the arguments put forth in this section suggest that youth, and possibly women and the unskilled, bear the brunt of job security regulations.

7.1.2 Minimum Wages

The effect of minimum wages on employment remains a controversial topic. In the competitive model, workers are paid their marginal product, and any artificial increase in the price of labor above the marginal product therefore prices the worker out of the labor market. Conversely, models that allow for employers' monopsony power predict wages lower than the marginal product, and, thus, an increase in minimum wages can increase wages without reducing employment rates.[8] In the Lang and Kahn (1998) model of bilateral search, the effects of minimum wages also differ from the expected effects in the competitive model. In their model, minimum wages affect the quality of the pool of applicants to jobs. Higher minimum wages allow firms to get better applicants for jobs, while reducing the employment prospects of less-productive workers.

On average, youth, women, and the unskilled tend to have lower wages than older, male, or skilled workers. Therefore, because minimum wages are more likely to be binding among these workers, the competitive model predicts larger unemployment effects for the first group. In the imperfect competition model, however, the effects are less clear-cut. In principle, the magnitude and sign of the minimum wage effect will depend on how far wages are from their respective marginal products in each subpopulation. If that gap is larger in some groups than in others, an increase in minimum wages may have "competitive" effects on some groups and "noncompeti-

6. The insider-outsider argument requires a strong union fixing wages for new entrants. Otherwise, firms could always pay very low wages at the beginning of the employment relationship to compensate for higher wages in the future. See Bertola (1990) for an analytical study of this issue.

7. See Hamermesh (1993).

8. There are many situations that give rise to imperfect competition in the labor market, such as monopolistic power on the part of employees, incomplete information, or imperfectly mobile workers.

tive" effects on others. Given this ambiguity, the sign and magnitude of the effects become an empirical question.

7.2 Labor Market Regulations in Chile

Chile has experienced a very wide range in labor market policies, providing a privileged case scenario for analyzing the impact of regulations on labor market outcomes. We distinguish between job security provisions and statutory minimum wages.[9]

7.2.1 Job Security Provisions

Among the most interesting aspects of the Chilean experience is that, in the thirty-nine years covered by our sample, Chile has gone from a situation of dismissal at will to a rigid labor market by OECD standards (Heckman and Pagés 2000). Since their inception in 1966, job security provisions have favored full-time indefinite employment over part-time, fixed-term, or temporary contractual relationships. To this end, in case of a firm-initiated separation, labor codes regulate the following: (1) compulsory advance notice periods; (2) the causes for which a dismissal is considered justified or unjustified; and (3) severance pay related to the tenure of a worker and the cause of dismissal. While the minimum period of advance notice has always been kept constant and equal to one month, the formula for computing severance pay and the causes for just or unjust dismissal have varied widely over the years. This is the variance that we exploit in our empirical work.

Table 7.1 summarizes the changes in legislation that took place in the 1960–1998 period. From 1960 to mid-1966, firms had to provide a one-month advance notice (or pay the equivalent of one month of salary), but, otherwise, "employment at will" was the norm. In 1966, the congress approved a new law under which firms had to pay compensation equal to one month's wage per year of work to all workers dismissed without just cause. The economic needs of the firm were considered a just cause in the law, and, therefore, a worker dismissed for this reason would not qualify for severance pay. In practice, however, workers would appeal to courts, and judges tended to consider these dismissals unjustified (Romaguera, Echevarría, and González 1995). In that event, the employer could choose between paying the mandatory compensation—plus wages foregone during trial—or reinstate the worker in his or her old post. This reform substantially increased the difficulty and the cost of labor force adjustments.

After 1973, a violent change in political regime brought about a de facto liberalization. Although job security provisions were not modified in the

9. See Edwards and Cox-Edwards (2000) for an excellent summary of labor market reforms in Chile during the 1960–2000 period.

Table 7.1 Employment Protection Provisions in Chile: 1960–1998

Period	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 forgone wages during trial. Trials could last at most 6 months. There is no maximum in the amount to be awarded.	All workers
1973–1978	1 month	Labor courts were much more pro-firm. Workers' claims were weaker.	Same as previous period	Same as previous period	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 workers hired after June 1978
1981–1984 (August 14, 1981): Law 18,018	1 month	Economic needs were considered just cause.	zero	1 month's wage per year of work with a maximum of 150 days	Only workers hired after August 1981
1984–1990 (Dec. 1984): Law 18,372	1 month	Economic needs were no longer considered just cause for dismissal.	zero	1 month's 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's 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

Source: Pagés and Montenegro (1999).

law, in practice, it was more likely that judges ruled against workers, effectively reducing dismissal costs. In 1989 and 1981, successive modifications reduced the cost of dismissal under the law. In 1981, the maximum amount to be awarded to a worker dismissed without just cause was reduced to the equivalent of five months' pay. This reform substantially reduced the cost of dismissal, particularly for workers with long tenures, although it only applied to newly hired workers.

After 1984, the tide shifted and job security provisions became progressively stricter. In December of that year, the law was modified to exclude economic needs of the firm as a justified cause of dismissal. However, the maximum amount payable to a worker was kept at five months of pay. In 1990, after the return of democracy, a new labor reform still in force further increased the cost of dismissal. This law considers dismissals motivated by the economic needs of the firm justified, but employers are still liable to pay compensation equal to one month's pay per year of work, with a maximum amount of eleven months of pay. It is the responsibility of the firm to prove just cause. If such causality cannot be demonstrated, there is a 20 percent surcharge in the amount of compensation.

We summarize this variance in law and court practice by means of a job security measure derived in Pagés and Montenegro (1999).[10] This measure is computed as follows:

$$\mathrm{JS}_t = \sum_{i=1}^{T} \beta^i \delta^{i-1}(1-\delta)[b_{t+i} + a_t \mathrm{SP}^{\mathrm{jc}}_{t+i} + (1-a_t)\mathrm{SP}^{\mathrm{uc}}_{t+i}],$$

where δ is the probability of remaining in a job, β is the discount factor, T is the maximum tenure that a worker can attain in a firm, b_{t+i} is the advance notice to a worker that has been i years with a firm, a_t is the probability that the economic difficulties of the firm are considered a justified cause of dismissal, $\mathrm{SP}^{\mathrm{jc}}_{t+1}$ is the mandated severance pay in that event to a worker that has been i years at the firm, and finally, $\mathrm{SP}^{\mathrm{uc}}_{t+1}$ denotes the payment to be awarded to a worker with tenure i in case of unjustified dismissal.

This measure computes the expected cost, at the time a worker is hired, of dismissing this worker in the future. This cost is measured in terms of monthly wages. The advantage of this measure in respect to other measures that compute the cost conditional on having achieved a certain tenure is that our job security measure captures the whole profile of severance pay at each level of tenure. The assumption is that firms evaluate future dismissal costs based on current law. Higher values of this variable indicate periods of relatively high job security, whereas lower values characterize periods in which dismissals were less costly.

10. See the mentioned paper and Heckman and Pagés (2000) for a complete description of the methodology used, how it is applied across time and countries, and the relative advantages and costs of using this measure versus other measures of job security.

Table 7.2 **Parameters Used to Compute the Job Security Index**

	β	δ	*b*	*a*	SP^{fc}	SP^{uc}
1960–1965	0.92	0.88	1	1	0	0
1966–1973	0.92	0.88	1	0.2	0	(1)
1974–1977	0.92	0.88	1	0.5	0	(2)
1978–1980	0.92	0.88	1	0.8	0	(2)
1981–1984	0.92	0.88	1	0.8	0	(3)
1985–1990	0.92	0.88	1	0	0	(3)
1991–	0.92	0.88	1	0.9	(4)	(5)

Notes: To compute β we use the fact that the average real interest from 1960–1998 was 8.4 percent. To compute δ we assume that the average Chilean turnover rate *without* employment protection would be similar to the U.S. rate. According to Davis and Haltiwanger (1995), average turnover rates average 12 percent a year in the United States. (1) corresponds to one month's pay per year of work augmented by three months to capture the average payments in forgone 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 an upper limit of five months' pay. (4) = one month's pay per year of work with an upper limit of eleven months' pay. (5) = 1.2 months of pay per year of work with eleven months upper limit. We assume the maximum tenure a worker can attain at a firm is twenty-five years.

Based on the legal information summarized in table 7.1 and assumptions regarding β, δ, *a*, and *T*, we obtain a measure of job security (JS). We take β to be a constant value such that the average real interest is equal to 8.4 percent, which corresponds to the average real interest rate in Chile during the 1960–1998 period. The discount rate is computed based on the assumption that without job security, turnover rates in Chile would be comparable to those observed in the United States.[11] Davis and Haltiwanger (1992) report an average annual turnover rate of 12 percent. The probability that a dismissal originated by the economic needs of the firm will be considered just depends on whether the law says so and whether labor judges rule so if workers take firms to court. For the period 1966–1984, although economic needs of the firm were considered just cause in the law, we assume *a* to be larger than zero and determined by the position taken by labor courts. Finally, we assume $T = 25$. See table 7.2 for a complete description of the parameters used in the computation of the job security measure.

The evolution of this variable over time is depicted in figure 7.1. After some years of relatively low employment protection, job security increases eightfold after the introduction of compulsory severance pay in the law. Expected dismissal costs decline markedly in 1973 and then successively in 1978 and 1981. Subsequently, employment protection increases again, but without reaching the levels attained during the late 1960s.

11. Although turnover rates can be measured, this measure is itself affected by labor law. Given this endogeneity, we choose instead to use the U.S. turnover rate, because it is well established that dismissal costs in the United States are very small.

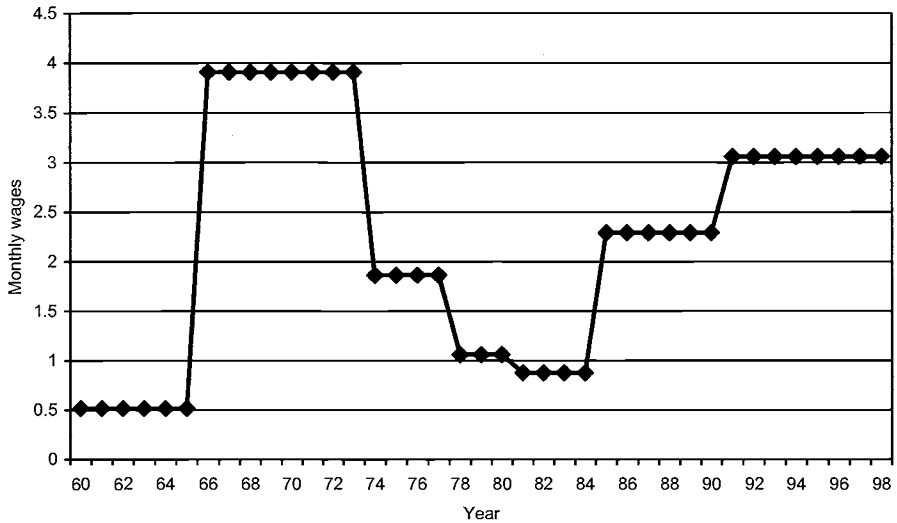

Fig. 7.1 Job security (in monthly wages)
Source: Pagés and Montenegro (1999).

7.2.2 Minimum Wages

Columns (2) and (3) in table 7.3 present the hourly real minimum wage in 1998 pesos; these indices were constructed using Chile's Central Bank Bulletins.[12] It is interesting to note that since 1989 there has been a lower minimum wage for workers eighteen years old or younger. This wage has been fixed at a level between 15 and 20 percent of the adult wage. Figure 7.2 summarizes the evolution of the minimum wage in relation to the average wage for teen and adult workers. The figure shows that minimum wages are much higher, relative to each group average rate, for teens than for adult workers. It also shows that the level of teen minimum wages has been quite volatile relative to the average wage.

Between 1960 and 1998, adult real minimum wages increased by 186 percent and teen minimum wages by 104 percent. However, because average ages rose more than the increase in the minimum wages, minimum wages lost ground in relation to the average wage. Despite this long-term secular trend, Chile experienced a wide range of fluctuations in minimum wages, both in its rate of growth (in real terms) and in its level in relation to the average wage. During the 1960s, the real value of minimum wages was held constant, but since real wages increased, the ratio of the minimum to the average real wage declined. In the early 1970s, minimum wages increased

12. Per hour minimum wages are constructed as monthly minimum wages divided by 4.2 · 40 hours.

Table 7.3 **Basic Statistics of the Sample**

		Minimum Wage				Average Wage											
				Bargaining Index		By Sex		By Skill Level		By Age Group						
Year	Job Security Index (1)	Age Under 18 (2)	Age Over 18 (3)	Original (4)	Smoothed (5)	Male (6)	Female (7)	Low (8)	High (9)	15–24 (10)	25–49 (11)	50–65 (12)	GDP Deviation from Trend (%) (13)	Employment Rate (%) (14)	Wage Employment Rate (%) (15)	Self-Employment Rate (%) (16)
1960	0.5199	119	119	3.3333	3.3333	302	152	157	475	133	283	306	–0.86	52.5	39.8	12.7
1961	0.5199	114	114	3.3333	3.3333	370	179	171	554	164	331	435	–1.41	52.2	41.1	11.1
1962	0.5199	126	126	3.3333	3.3333	373	203	181	615	162	361	418	–1.37	53.2	41.2	11.9
1963	0.5199	109	109	3.3333	3.3333	376	206	n.a.	311	219	342	395	0.20	53.0	41.4	11.5
1964	0.5199	107	107	3.3333	3.3333	268	160	n.a.	230	133	272	296	–2.15	52.9	42.3	10.6
1965	0.5199	114	114	3.3333	3.3333	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	–5.23	54.4	43.3	11.2
1966	3.9090	118	118	3.3333	3.3333	380	211	187	591	179	376	434	1.50	53.0	42.2	10.8
1967	3.9090	116	116	3.3333	3.34724	427	268	222	648	217	420	539	1.50	54.0	43.2	10.8
1968	3.9090	111	111	3.3333	3.39543	466	278	224	699	251	450	502	1.79	53.2	41.9	11.4
1969	3.9090	107	107	3.3333	3.46403	475	279	231	709	218	470	560	2.79	52.4	41.2	11.2
1970	3.9090	133	133	3.66667	3.53596	549	351	256	804	248	536	693	2.97	52.3	41.4	10.9
1971	3.9090	183	183	3.66667	3.57675	689	437	302	957	307	660	779	9.67	53.7	42.1	11.5
1972	3.9090	195	195	3.66667	3.52856	712	457	342	929	359	698	729	7.28	52.7	41.3	11.4
1973	3.9090	108	108	3.66667	3.40525	525	332	279	671	280	512	553	0.37	51.4	39.6	11.8
1974	1.8642	204	204	3	3.26140	435	310	275	561	255	436	496	0.12	49.0	37.1	11.8
1975	1.8642	245	245	3	3.12419	376	277	225	483	214	376	420	–14.58	45.0	34.7	10.4
1976	1.8642	259	259	3	3.01390	486	352	249	635	280	474	542	–12.67	45.8	34.5	11.2
1977	1.8642	269	269	3	2.88227	692	512	320	953	357	696	786	–5.01	48.3	38.1	10.1
1978	1.0599	346	346	3	2.62090	868	517	360	1,090	400	799	1,072	0.87	48.0	37.1	10.9
1979	1.0599	345	345	2.66667	2.27455	913	640	432	1,150	496	904	1,009	6.66	47.8	36.8	10.9
1980	1.0599	354	354	1.33333	1.90434	890	611	424	1,120	476	881	932	11.83	47.4	36.6	10.7
1981	0.8772	334	334	1.33333	1.53353	1,057	799	510	1,338	590	1,099	1,016	15.64	50.9	39.3	11.6
1982	0.8772	365	365	1.33333	1.25825	1,235	852	508	1,499	618	1,206	1,295	–1.15	41.8	33.0	8.8
1983	0.8772	276	276	1	1.13070	842	622	345	1,056	416	872	721	–6.79	43.5	34.4	9.1
1984	0.8772	243	243	1	1.06209	843	573	355	1,028	371	845	780	–4.19	46.1	35.8	10.3

(*continued*)

Table 7.3 (continued)

Year	Job Security Index (1)	Minimum Wage: Age Under 18 (2)	Minimum Wage: Age Over 18 (3)	Bargaining Index: Original (4)	Bargaining Index: Smoothed (5)	Average Wage: By Sex: Male (6)	Average Wage: By Sex: Female (7)	Average Wage: By Skill Level: Low (8)	Average Wage: By Skill Level: High (9)	Average Wage: By Age Group: 15–24 (10)	Average Wage: By Age Group: 25–49 (11)	Average Wage: By Age Group: 50–65 (12)	GDP Deviation from Trend (%) (13)	Employment Rate (%) (14)	Wage Employment Rate (%) (15)	Self-Employment Rate (%) (16)
1985	2.2915	220	220	1	1.01390	699	480	312	808	323	683	725	–6.19	46.4	36.6	9.8
1986	2.2915	215	215	1	1	653	471	301	742	314	634	731	–5.35	47.0	37.3	9.7
1987	2.2915	199	199	1	1	796	539	288	932	355	764	907	–4.05	50.1	39.5	10.5
1988	2.2915	222	222	1	1.02781	766	542	316	902	376	751	799	–2.93	50.9	38.6	12.2
1989	2.2915	293	340	1	1.12419	869	679	376	981	434	868	973	0.41	53.1	41.6	11.5
1990	2.2915	298	346	1	1.26140	1,003	682	390	1,074	462	960	1,011	–2.83	52.0	40.5	11.4
1991	3.0598	278	327	1.66667	1.40525	971	694	401	1,046	470	951	949	–2.47	53.2	41.2	11.9
1992	3.0598	293	340	1.66667	1.54247	904	726	455	998	503	914	900	1.47	55.7	43.6	12.1
1993	3.0598	294	341	1.66667	1.63885	1,072	832	496	1,158	627	1,054	1,093	0.98	55.9	44.0	11.9
1994	3.0598	294	342	1.66667	1.66667	1,141	840	535	1,194	624	1,101	1,163	–1.22	55.4	42.5	12.9
1995	3.0598	302	351	1.66667	1.66667	1,230	919	566	1,310	657	1,215	1,199	0.81	55.5	42.8	12.7
1996	3.0598	279	324	1.66667	1.66667	1,329	1,047	621	1,412	725	1,283	1,465	1.59	55.8	43.7	12.0
1997	3.0598	248	333	1.66667	1.66667	1,392	1,100	613	1,505	775	1,380	1,335	2.79	56.7	44.1	12.6
1998	3.0598	243	341	1.66667	1.66667	1,356	1,136	759	1,427	792	1,325	1,500	0.70	56.8	43.6	13.2

Source: Authors' calculations (see data section) and Banco Central de Chile (2001).

Note: n.a. = not available.

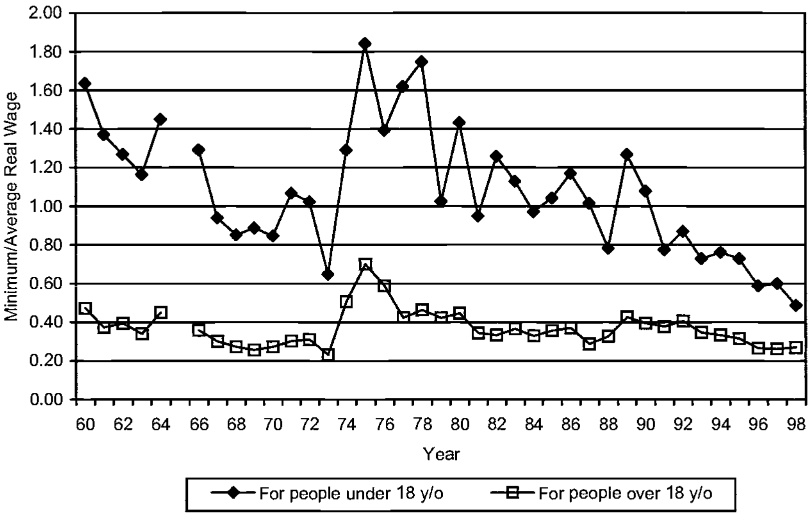

Fig. 7.2 Minimum to average real wages
Source: Authors' calculations (see data section).

substantially, surpassing the growth rate of average wages. In consequence, the ratio of the minimum to the average real wage increased sharply in that period. From 1975 to 1980, minimum wages lost ground relative to the average wage. After the return to democracy in 1990, real minimum wages increased steadily, but they continued declining relative to the average wage. The decline was particularly sharp for the teen group, whose minimum to average real wage rate fell from 1.80 in 1975 to 0.50 in 1998. It is interesting to note that while there are several studies in the Chilean case that suggest that the minimum wage is binding, others such as Bravo and Vial (1997) suggest that it is not.[13]

7.3 Data

The household surveys used in this study were obtained from the University of Chile's economics department. The economics department's survey monitors the employment-unemployment status in the metropolitan area of Santiago, Chile four times a year. Unfortunately, only the surveys taken in June of each year contain information about wages and other employment status variables. Therefore, these are the surveys used in this study. The format of the survey and the definition of the variables have been kept constant since 1957, when the survey started, and so the infor-

13. See, for instance, Castañeda (1983), Paredes and Riveros (1989), Montenegro (2002), and Cowan et al. (2003). An excellent review of the impact of minimum wages in the case of the United States can be found in Kosters (1996). A more recent survey on the international evidence of minimum wages can be found in Dowrick and Quiggin (2003).

mation contained in them is comparable across years.[14] During the period from 1960 to 1998, the surveys interviewed between 10,000 and 16,000 people and around 3,700 and 5,400 active labor force participants each year. During this period, the metropolitan area of Santiago, Chile represented about one-third of Chile's total population and a higher proportion of gross domestic product (GDP).[15] The data set is formed by stacked cross-sectional data sets, which means that individuals are not followed over time. The only restriction applied to our sample is that the people included in the estimates must be at least fifteen years old and no older than sixty-five.

We merge labor policy and macrovariables taken at the annual frequency with our individual-level annual data. We include the job security index and the minimum wage data described in Section 2. We also include a measure of wage bargaining to control for changes in union activity that can be correlated to our variables and to employment. While perhaps the best measure of the influence of unions on wage determination is union coverage, that is, the share of workers whose wages are affected by collective bargaining, a time series of this nature does not exist in Chile. Because union membership is also not available for all years covered in our sample, we measure unions' bargaining power by means of an index that reflects the degree of centralization of collective bargaining constructed by Edwards and Cox-Edwards (2000). This variable takes values from 1 (total decentralization) to 4 (total centralization). The use of this measure is based on the observation that union coverage tends to be larger in countries where collective bargaining is centralized. Finally, we include as a measure of economic activity deviations, with respect to potential GDP. To obtain this variable, we use GDP data from the World Bank and apply a Hodrick-Prescott filter to obtain trend GDP.

Table 7.3 summarizes some basic statistics of our sample, by year. The first three columns display the value of the job security index and the real minimum wage for people eighteen or younger and for adult workers. The next two columns summarize the index of bargaining (column [4] presents the original index, and column [5] presents the smoothed index). The evolution of these variables over time is depicted in figure 7.3. Higher values of this measure, like those registered from 1960 to 1970, reflect periods of higher union centralization.[16] The next seven columns summarize the average hourly wage broken down by sex (columns [6] and [7]); skill level (col-

14. In this study we use data from 1960 on, because the previous years (1957–1959) do not have reliable data.

15. According to the 1992 census, the metropolitan area accounted for 39 percent of the total population.

16. Although not shown in the results, we checked the robustness of our results using the strikes index constructed by Edwards and Cox-Edwards (2000) instead of the centralization index. The results were invariant to different specifications.

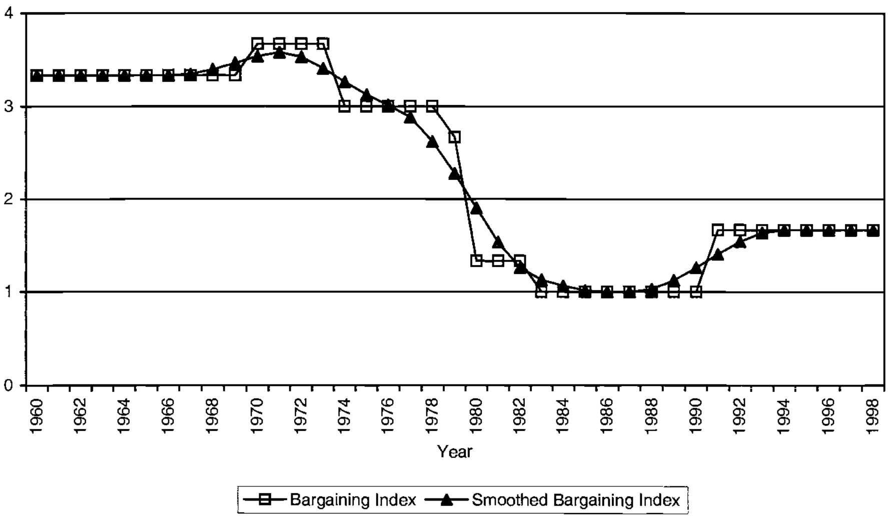

Fig. 7.3 Bargaining index

Source: Edwards and Cox-Edwards (2000).

Notes: Bargaining index measures the degree of centralization of wage bargaining. It takes values from 1 to 4. Higher values indicate higher centralization of collective bargaining.

umns [8] and [9]); and age group (columns [10], [11] and [12]). Column (13) summarizes the deviation of the GDP from its potential or trend value. Finally, columns (14), (15), and (16) present the percentage of total people employed, the percentage of people that work for someone else (wage employment), and the percentage of people self-employed as a proportion of total population between fifteen and sixty-five years old. These three rates are also depicted in figure 7.4, which, jointly with figure 7.5 (which shows GDP deviations from its trend), illustrates the violent swings experienced by the Chilean economy during the 1960–1998 period, and, in particular, between 1970 and 1985.[17] Some additional indicators describing the performance of the Chilean economy are summarized in table 7.4.

7.4 Methodology

To estimate the differential impact of labor market regulations across subpopulations we assume that the employment status of an individual is characterized by

17. Chilean economic performance has been extensively documented by Edwards and Cox-Edwards (1991, 2000), de la Cuadra and Hachette (1992), Wisecarver (1992), Bosworth, Dornbusch, and Laban (1994), Hudson (1994), Soto (1995), and Cortazar and Vial (1998).

Fig. 7.4 Employment and dependent rates
Source: Authors' calculations (see data section).

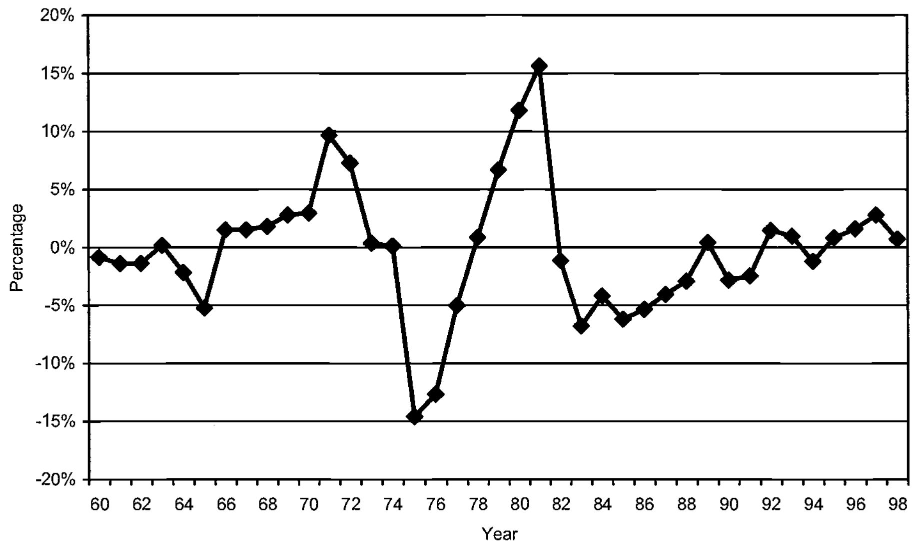

Fig. 7.5 GDP deviation from trend
Source: Authors' calculations (see data section).

Table 7.4 **General Economic Indicators: Chile 1960–1998**

Series Name	GDP per Capita Growth (annual %)	Inflation, Consumer Prices (annual %)	National Unemployment, Total (% of total labor force)	National Unemployment, Female (% of female labor force)	National Unemployment, Youth Total (% of total labor force ages 15–24)	Gran Santiago Unemployment, total (% of total labor force)	Gini Coefficient
1960	n.a.	n.a.	n.a.	n.a.	n.a.	8.0	42.5
1961	1.5	7.7	n.a.	n.a.	n.a.	7.1	45.2
1962	2.7	14.0	n.a.	n.a.	n.a.	5.7	45.5
1963	3.6	44.1	n.a.	n.a.	n.a.	5.2	n.a.
1964	0.3	46.0	n.a.	n.a.	n.a.	4.9	n.a.
1965	–1.8	28.8	n.a.	n.a.	n.a.	5.0	n.a.
1966	7.6	23.1	n.a.	n.a.	n.a.	6.0	45.2
1967	1.5	18.8	n.a.	n.a.	n.a.	5.9	45.8
1968	1.6	26.3	n.a.	n.a.	n.a.	6.4	48.1
1969	1.5	30.4	n.a.	n.a.	n.a.	7.1	48.0
1970	0.2	32.5	n.a.	n.a.	n.a.	7.0	47.5
1971	7.1	20.0	n.a.	n.a.	n.a.	5.2	47.7
1972	–2.5	74.8	n.a.	n.a.	n.a.	3.7	43.1
1973	–6.5	361.5	n.a.	n.a.	n.a.	3.1	44.1
1974	0.8	504.7	n.a.	n.a.	n.a.	10.3	40.7
1975	–12.8	374.7	n.a.	n.a.	n.a.	16.1	41.1
1976	1.8	211.8	n.a.	n.a.	n.a.	18.0	47.2
1977	7.1	91.9	n.a.	n.a.	n.a.	13.0	48.4
1978	5.9	40.1	n.a.	n.a.	n.a.	12.8	49.8
1979	7.1	33.4	n.a.	n.a.	n.a.	12.5	49.4

(*continued*)

Table 7.4 (continued)

Series Name	GDP per Capita Growth (annual %)	Inflation, Consumer Prices (annual %)	National Unemployment, Total (% of total labor force)	National Unemployment, Female (% of female labor force)	National Unemployment, Youth Total (% of total labor force ages 15–24)	Gran Santiago Unemployment, total (% of total labor force)	Gini Coefficient
1980	6.5	35.1	10.4	10.0	20.8	11.7	49.1
1981	3.2	19.7	11.3	9.9	21.5	9.0	47.3
1982	–11.7	9.9	19.6	18.3	30.5	23.2	51.2
1983	–5.3	27.3	14.6	14.7	24.7	22.7	52.7
1984	6.3	19.9	13.9	n.a.	25.2	18.4	54.2
1985	5.4	29.5	12.1	13.4	22.7	16.2	51.5
1986	3.9	20.6	8.8	9.7	17.3	15.4	48.7
1987	4.9	19.9	7.9	9.3	n.a.	13.5	57.6
1988	5.5	14.7	6.3	7.8	14.3	11.2	53.7
1989	8.7	17.0	5.3	6.1	13.2	9.3	50.8
1990	1.9	26.0	5.7	5.7	13.1	9.7	53.9
1991	6.2	21.8	5.3	5.8	12.7	8.3	52.4
1992	10.4	15.4	4.4	5.6	10.9	6.0	47.4
1993	5.2	12.7	4.5	5.1	11.0	6.4	45.4
1994	4.0	11.4	5.9	6.8	13.2	6.3	45.9
1995	8.9	8.2	4.7	5.3	11.5	6.1	46.3
1996	5.7	7.4	5.4	6.7	12.8	7.2	45.4
1997	6.0	6.1	5.3	6.6	13.0	6.7	n.a.
1998	2.5	5.1	7.2	7.6	16.7	6.9	n.a.

Sources: World BankWorld Development Indicators Data Base and Gini coefficient from background data, Montenegro (1998).

Note: n.a. = not available.

(1) $$y^*_{ijt} = \mathbf{X}_{it} \cdot \boldsymbol{\beta}_1 + \mathbf{X}'_{it} \cdot \mathbf{Z}_t \cdot \boldsymbol{\beta}_2 + \gamma_t + \varepsilon_{ijt},$$

where

$$y_{ijt} = 1 \text{ if } y^*_{ijt} > 0$$

$$y_{ijt} = 0 \text{ otherwise,}$$

and y^*_{ijt} is an unobservable variable that determines whether an individual i, in subpopulation j at time t will be employed or not, and y_{ijt} is the observable employment status of this individual. This variable takes a value of 1 if the individual is employed and zero if it is not. In some specifications, we focus only on wage employment (self-employment), and, therefore, this variable takes the value of 1 if an individual is wage (self-) employed and zero otherwise. The sample corresponds to the whole population between fifteen and sixty-five years old. In addition, $\mathbf{X}_{it}$ is a vector of variables that summarizes the personal characteristics of the individual i at time t, $\mathbf{Z}_t$ is a vector of variables that vary with t, γ_t is a year fixed effect, and ε_{ijt} is an error term. Among the personal characteristics, we include age, gender, skill level, number of children, and number of children interacted with gender. In some specifications, we also include age interacted with gender and age interacted with skill to capture differential effects of age across gender and skill groups. Given the number of observations available, we divided the data into three age groups (fifteen–twenty-four, twenty-five–fifty, and fifty-one–sixty-five) and two skill levels (nine years of education or less and more than nine years). Adding the skill and the age groups to the gender division, we have twelve different subpopulations, $j = 1, \dots 12$

In the vector of aggregate variables, $\mathbf{Z}_t$, we include the index of job security, deviations from GDP trend, and the union centralization variable (all in logarithms). We also include the minimum wage index (also in logarithms), but we let it change for individuals eighteen and younger. By construction, the vector of coefficients on the interaction of $\mathbf{X}_{it}$ and $\mathbf{Z}_t$, $\boldsymbol{\beta}_2$, gives the sign of the *differential* effect. In addition, assuming that the Prob($y^*_{ijt} > 0$) is distributed as a standard normal distribution, the size of the marginal differential effect is given by $\phi(\cdot)\mathbf{X}_{it}\boldsymbol{\beta}_2$, where $\phi(\cdot)$ is the normal density function.

Although specification (1) is a reduced form equation, in some cases it will be useful to add a measure of wages. To construct this variable, w_{ijt}, we assign to all workers $i \in j, j = 1, \dots, 12$, at period t, the average wage of all employed workers in group j at period t.

Our original intention was to estimate

(1′) $$y^*_{ijt} = \mathbf{X}_{it} \cdot \boldsymbol{\beta}_1 + \mathbf{X}'_{it} \cdot \mathbf{Z}_t \cdot \boldsymbol{\beta}_2 + \mathbf{Z}_t \cdot \boldsymbol{\beta}_3 + \varepsilon_{ijt}.$$

With such a specification we could recover the *total* marginal effect of a labor policy on subpopulation j as $\phi(\cdot)(\mathbf{X}_{it}\boldsymbol{\beta}_2 + \boldsymbol{\beta}_3)$. However, despite finding robust estimates for the differential effects, our estimates for the level effect (β_3) proved to be extremely sensitive to the set of variables included in $\mathbf{Z}_t$, suggesting that our time variables did not properly account for the time variation of the series. In view of these results, we opted for estimating specification (1). This estimation still allows us to compute marginal effects, but the total effects are now absorbed by the constant term. Therefore, we can measure the impact of labor market regulations on the *distribution* but not on the *level* of employment. Nonetheless, estimating equation (1) instead of (1′) offers substantial advantages from an econometrics point of view. It allows controlling for macroeconomic trends and cycles as well as policy changes and other unobservable variables that are common to all individuals and that could be correlated to employment and labor market regulations and bias the estimation. In addition to the inclusion of time variables, we minimize the risk of omitted variable biases and spurious correlations in four additional ways.

First, by using individual data from a series of stacked household surveys to estimate specification (1), we can control for changes in the relative size of the population of each group and changes in fertility, which, if omitted, could bias our estimates. Second, by controlling for effects of changes in the business cycle (using GDP deviations from its trend) across individuals (that is, including $\boldsymbol{X}'_{it} \cdot \mathbf{Z}_t$, where $\mathbf{Z}_t$ contains the business-cycle variable), we can partially control for changes in policy and institutions that are endogenous to changes in relative employment. This is because such movements are likely to be correlated with changes in the business cycle. Third, by estimating the differential effect of policy while including contemporary labor market policies and institutions, we make sure that our measured effects are not biased by the correlation between these variables and the distribution of employment. Finally, by comparing the estimated effects on the probability of wage employment (which is covered by labor policy) with the results on self-employment (which is not covered) once appropriate pull-push factors from and to self-employment are accounted for, we assess whether we are capturing the effect of policy, or, instead, the effect of some unobservable correlate with group-specific employment.

7.5 Empirical Results

7.5.1 The Effect of Job Security on the Distribution of Employment

Our results indicate that job security provisions have a differential impact across demographic subgroups. In table 7.5, we report the results of estimating our empirical specification (1) assuming normality in the distribution of errors. The reported numbers correspond to the coefficients of the probit model, while the marginal effects for selected subpopulations of workers are reported in table 7.6. The *t*-tests, reported next to the coefficients, are robust to the presence of heteroskedasticity of unknown kind

Table 7.5 **The Effect of Job Security and Minimum Wages, Probit Results**

	(1) Employed		(2) Employed		(3) Wage Employment		(4) Self-Employment		(5) Employed		(6) Employed		(7) Employed	
Dependent Variable	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test
Dummy young	–0.8954	–104.2	0.4921	2.6	0.9189	5.0	–0.4202	–1.4	–1.1703	–6.1	–0.9651	–4.9	1.2757	9.1
Dummy old	–0.6709	–66.8	–1.6509	–7.3	–1.6967	–7.5	0.4176	1.7	–2.0996	–9.1	–2.1226	–9.0	–1.4101	–8.6
Dummy women	–0.5461	–66.7	–2.0260	–12.2	–1.8595	–11.6	–0.3632	–1.7	–2.4113	–14.2	–1.9622	–11.3	–2.7873	–22.7
Dummy unskilled	0.0007	0.1	1.8635	10.9	1.8843	11.2	–0.3281	–1.5	1.4867	8.6	1.8356	10.3	2.2867	18.1
Children per father	0.1570	45.0	0.1569	44.6	0.0594	25.7	0.0273	11.3	0.1152	32.0	0.1152	31.5	0.1562	44.6
Children per mother	–0.3931	–93.9	–0.3921	–92.7	–0.3147	–86.9	–0.0196	–5.4	–0.3179	–70.1	–0.3160	–68.5	–0.3919	–93.1
Interacted with logarithm of job security														
Dummy young	–0.0935	–10.8	–0.1112	–12.7	–0.0826	–9.7	–0.0161	–1.2	–0.0913	–5.6	–0.1163	–6.7		
Dummy old	0.0124	1.2	0.0196	1.8	0.0292	2.7	0.0173	1.5	0.0253	1.2	0.0123	0.6		
Dummy women	–0.0468	–6.1	–0.0266	–3.4	–0.0021	–0.3	0.0267	2.7	–0.546	–4.5	–0.0873	–6.8		
Dummy unskilled	–0.0334	–4.2	–0.0563	–7.0	–0.0733	–9.3	0.0344	3.4	–0.0382	–3.3	–0.0596	–4.8		
Dummy young · dummy women									0.0835	4.7	0.1033	5.4		
Dummy old · dummy women									–0.0035	–0.2	0.0064	0.3		
Dummy young · dummy unskilled									–0.0381	–2.2	–0.0164	–0.9		
Dummy old · dummy unskilled									0.0033	0.2	0.0146	0.6		
Interacted with logarithm of minimum wage														
Dummy young			–0.1406	–8.2	–0.1557	–9.3	–0.0366	–1.3	–0.0111	–0.6	–0.0215	–1.2	–0.2129	–16.0
Dummy old			0.0913	4.4	0.0911	4.4	–0.0286	–1.3	0.1301	6.2	0.1301	6.1	0.0715	4.6
Dummy women			0.1455	9.6	0.1551	10.7	–0.0299	–1.5	0.1677	10.8	0.1303	8.2	0.2097	18.0
Dummy unskilled			–0.1811	–11.6	–0.1811	–11.9	0.0304	1.5	–0.1587	–10.1	–0.1810	–11.2	–0.2196	–18.3
Dummy young · dummy women									0.0248	11.0	0.0223	9.8		
Dummy old · dummy women									–0.0035	–1.3	–0.0019	–0.7		
Dummy young · dummy unskilled									0.0393	17.4	0.0346	15.2		
Dummy old · dummy unskilled									0.0133	4.9	0.0145	5.3		

(*continued*)

Table 7.5 (continued)

	(1) Employed		(2) Employed		(3) Wage Employment		(4) Self-Employment		(5) Employed		(6) Employed		(7) Employed	
Dependent Variable	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test	β	*t*-test
Interacted with union centralization														
Dummy young			0.1320	8.2	0.1422	9.2	0.0800	3.0	–0.3006	–13.1	–0.2785	–11.9		
Dummy old			0.0272	1.4	0.0241	1.2	0.0152	0.7	–0.0966	–3.2	–0.0854	–2.8		
Dummy women			–0.0968	–6.8	–0.1222	–8.9	0.0802	4.2	–0.2494	–13.5	–0.2177	–11.6		
Dummy unskilled			0.0756	5.2	0.0480	3.4	0.0358	1.9	–0.0843	–4.6	–0.0599	–3.3		
Dummy young · dummy women							0.2957	12.3	0.2712	10.9				
Dummy old · dummy women							0.1530	5.2	0.1359	4.5				
Dummy young · dummy unskilled									0.3485	14.1	0.3306	13.0		
Dummy old · dummy unskilled									0.0265	0.9	0.0249	0.8		
Interacted with GDP deviation from path														
Dummy young			–0.0852	–0.9	0.2102	2.2	0.0208	0.1	–0.2928	–1.7	–0.3618	–2.1		
Dummy old			–0.3872	–3.1	–0.2161	–1.7	–0.0041	0.0	–0.7902	–3.4	–0.8027	–3.4		
Dummy women			–0.4917	–5.5	–0.3108	–3.6	0.3153	2.7	–0.8047	–6.0	–0.8958	–6.7		
Dummy unskilled			0.4345	4.8	0.3467	3.9	0.0777	0.7	0.4079	3.2	0.4152	3.2		
Dummy young · dummy women									0.3973	2.0	0.5022	2.5		
Dummy old · dummy women									0.3863	1.6	0.4749	1.9		
Dummy young · dummy unskilled									–0.2455	–1.3	–0.1571	–0.8		
Dummy old · dummy unskilled									0.1912	0.8	0.1761	0.7		
Logarithm of hourly wage											0.1520	16.9		
No. of observations	303,945		303,945		303,945		303,945		303,945		295,318		303,945	
Pseudo R^2	0.196		0.168		0.11		0.08		0.211		0.210		0.197	

Notes: Besides the control variables mentioned in the table, all specifications include yearly dummies (not reported). Standard errors are robust to the presence of heteroskedasticity. The employed dummy variable is defined as 1 if the person is employed and zero otherwise (unemployed or inactive). The wage employment dummy variable is defined as 1 if the person is a dependent employee and zero otherwise (independent, unemployed, or inactive). The self-employed dummy variable is defined as 1 if the person is an employer or if the person works as an independent worker and zero otherwise (dependent, unemployed, or inactive).

Table 7.6 **Marginal and Total Effects of Labor Market Regulations**

	Marginal Effects		Total Effects	
	Job Security (1)	Minimum Wage (2)	Job Security (3)	Minimum Wage (4)
Men, 15–25, unskilled	–0.066 [0.000]	–0.0516 [0.000]	–0.049	–0.0516
Men, 15–25, skilled	–0.0351 [0.000]	–0.004 [0.52]	–0.0181	–0.004
Men, 26–50, unskilled	–0.008 [0.001]	–0.036 [0.000]	0.009	–0.036
Men, 51–65, unskilled	–0.0035 [0.620]	–0.005 [0.54]	0.0135	–0.005
Men, 51–65, skilled	0.008 [0.22]	0.045 [0.000]	0.025	0.045
Unskilled	–0.0343 [0.000]	–0.012 [0.09]	–0.0173	–0.012
Skilled	–0.015 [0.000]	0.044 [0.000]	0.002	0.044
Women	–0.0278 [0.000]	0.0463 [0.000]	–0.0108	0.0463
Men	–0.0151 [0.000]	–0.017 [0.000]	0.0019	–0.017
Young	–0.0394 [0.000]	0.0134 [0.08]	–0.0224	0.0134
Older	–0.008 [0.14]	0.0596 [0.0000]	0.009	0.0596

Note: P-values of the test that the marginal effects are equal to zero are reported in square brackets.

using the White (1980) method. Most coefficients on the individual characteristic variables exhibit the expected patterns: Female and older workers are less likely to be employed than prime-age (twenty-six–fifty) men. Additionally, the number of children per father increases the probability of being employed, and the number of children per mother decreases the probability of being employed. Instead, the coefficients on the variable young and unskilled change signs across specifications.

In column (1) we report the results of interacting the job security measure with dummies for age (young and older), gender (women), and skill level. A negative (positive) sign indicates that periods of more stringent job security provisions are associated with a decline (increase) in the probability of employment of a particular subpopulation, relative to the omitted category. We find strong age effects. The coefficient on the young-job security interaction is negative and statistically significant, while the coefficient on the older-job security interaction is positive although not statistically significant. Our results suggest that high job security tends to bias the distribution of employment against younger workers. We also find significant effects across the skill divide. The coefficient on the unskilled-job security interaction is negative and statistically significant, suggesting that job security provisions reduce the probability of employment of unskilled workers relative to skilled ones. Last, the coefficient on the female-job security interaction suggests a negative effect of job security on the probability of employment of women relative to men.

Column (2) shows the results once we control for the evolution of the minimum wage, union activity, and deviations of GDP with respect to its trend, as well as interaction of these variables with age, gender, and skill dummies. The only difference with respect to column (2) is that the coefficient on the dummy for older workers is now somewhat larger and statistically significant at the 10 percent level, suggesting that job security provisions benefit the employment prospects of older workers relative to prime-age ones. In columns (3) and (4) we report the coefficients resulting from estimating the same specification for wage employment and self-employment separately. Our results are encouraging because they suggest that our findings are driven by policy changes instead of by some unobservable factors correlated with labor policy and employment. The signs and magnitudes of the coefficients for total and wage employment are very similar, except for the coefficients on women. Instead, for self-employment, the coefficients are either not statistically different from zero or going in the opposite direction than for wage employment. This is the case with the coefficients on the gender and unskilled variables, which suggests that more stringent job security regulations increase the probability that women and the unskilled are employed in the self-employment sector relative to men and the skilled.

Column (5) exhibits the results once we allow for further interactions between age, skill, and gender groups. With this finer level of disaggregation we can examine whether the impact of job security is the same across young men and young women, or across young skilled and unskilled workers. These additional variables not only provide a more complete description of the effects of job security on the distribution of employment, but also help to infer the channels through which job security affects that distribution. The coefficients for these additional interaction variables are all statistically significant, and a test for their joint significance strongly rejects the null hypothesis of all the coefficients being zero.

The estimates in column (5) contain some interesting additional information relative to the estimates in columns (1) to (4). We find that an increase in job security tends to reduce the employment probabilities of young men relative to those of young women. However, we also find that this effect is reversed at older ages. Thus, job security provisions seemingly reduce the probabilities of employment of middle-aged and older women relative to those of men in that same age group. Our estimates also suggest that an increase in job security provisions reduces the probability of employment of both skilled and unskilled youth, but the effect is larger for unskilled youth.

Finally, column (6) reports the results of estimating the same specification as in column (5), but in addition controls by the average wage of each subpopulation group in period *t*. Controlling for the wage level of each group allows us to assess whether some of the observed effects are driven by differences in wage adjustment across subpopulations. Yet the results should be taken with caution because some wage movements may be endogenous to the probability of employment. Overall we find that holding wages constant does not affect our main results. The only coefficient that changes size and significance is the interaction between the young unskilled and job security. Holding wages constant reduces the coefficient and the significance of the effect on unskilled youth (relative to more skilled youth). Instead, most of the other coefficients become larger (in absolute value) than the ones reported in column (5). This suggests that more stringent regulations are partly paid by workers in the form of lower wages.

The marginal effects reported in table 7.6 correspond to the specification reported in column (5) of table 7.5. They are computed for different combinations of the dummies for gender, age, and skill.

The results indicate that the largest adverse effects are on unskilled youth. However, the effects on skilled youth are also substantial; an increase of 100 percent in job security reduces the probability of employment by 0.066 points (or 6.6 percentage points) for unskilled youth and by 0.0351 for skilled youth workers, relative to prime-age skilled workers. The results in table 7.6 suggest that skilled prime age male workers gain relative to all other groups with the exception of older workers. In addition, the marginal effects suggest that job security policies tend to have more adverse effects on women than on men.

In light of the different theories described in section 7.2, how do we explain the results presented previously? Although we cannot totally discriminate among different theories, we are at least able to reject some hypotheses. The fact that most of our results remain unchanged when wages are included suggests that the differential effects presented previously cannot be explained by differences in the elasticity of labor supply across demographic groups. The only exception is the larger effect on young unskilled workers, which seems to be driven by a higher labor supply elasticity of this group.[18] Our results also suggest that these differential effects cannot be explained by insider-outsider theories, because in that case the effect would also be through wages. Instead, our results suggest that the differ-

18. Cowan et al. (2003) find that, in Chile, seemingly high transitions between schooling and the labor market lead to a very elastic labor supply for the young unskilled.

ential effects on employment are demand driven: Changes in job security provisions bring about changes in hiring and firing rates that selectively affect different types of workers.

A barrier-of-entry effect can explain the negative impact of job security on the employment rates of young workers relative to other demographic groups. However, it cannot account for the estimated differences in impact between young women and young men. One possible way to explain these findings is to consider differences in turnover rates across groups. As discussed in section 7.2, a higher exogenous turnover rate can bring about two effects. On the one hand, workers with a higher propensity to rotate have lower average tenures and, therefore, are more likely to be laid off in bad times. On the other hand, higher rotation reduces expected severance payments and, therefore, increases the incentives to hire these workers. Consequently, higher rotation among women can explain why job security provisions affect young women less than young men. It can also explain why middle-aged and older women benefit less from job security than men of the same age.

Differences among turnover rates could also partially explain the results for skilled and unskilled workers. Higher rotation among the unskilled would imply lower tenure rates and higher probabilities of dismissal for middle-aged and older unskilled workers, relative to more skilled ones. This is consistent with the deleterious effect of job security on the employment rates of middle-aged and older unskilled workers, relative to skilled ones. Of course, the higher turnover rates among unskilled workers are less likely to be exogenous to the decisions of employers than female turnover rates. In consequence, a complete discussion of this effect requires a model that explains why turnover rates are different in the first place. This model does not seem to be able to explain why the effect on employment appears more negative on the unskilled than on skilled youth, but as we have seen, this effect seems to be driven by a relatively more inelastic labor supply of the latter.

7.5.2 Distribution of the Effect of Minimum Wages

Table 7.5 also reports the results of interacting personal characteristic dummies with the evolution of minimum wages over time. An increase in the statutory wage has qualitative effects on the distribution of employment across age and skill that are similar to the qualitative effects of stricter job security provisions. To account for contemporary employment policies and economic conditions, we include measures of union activity, job security provisions, and GDP deviations, interacted with demographic dummies in all specifications in columns (2) to (6), but not in column (7). As in other studies for developed countries, the results in column (7) suggest that an increase in the minimum wage reduces the employment prospects of young workers relative to older ones. We also find a negative effect on the unskilled. Instead, our results also indicate that minimum wage hikes may increase the probability of employment for women relative to men.

Controlling for the subgroup effects of contemporary changes in policy and the business cycle does not alter the results reported in column (7).[19] The comparison between the results obtained from the wage employment and the self-employment specifications (columns [3] and [4]) is also encouraging. As with the coefficients associated with job security provisions, we find that the coefficients on wage employment are very similar to the ones obtained for total employment, while the coefficients on self-employment are not statistically significant. All in all, these results suggest that the effects we are capturing are indeed associated with changes in policy rather than with some unobservable correlate of employment across demographic groups.

In column (5) we present our results once we allow for differential effects across age-skill and age-gender categories and control for contemporaneous changes in policy and economic conditions. As in column (7), we find a negative effect of minimum wages on the employment probabilities of unskilled workers. The effect of minimum wages is negative for young unskilled workers and not statistically significant for young skilled ones. Instead, higher minimum wages tend to shift employment toward older workers. Finally, we find that women, and in particular the young, tend to benefit from minimum wage policies.

The former specification assumes that the effect of raising the minimum wage is unrelated to the level of the going wage. However, it is plausible that the effect may be positively related to the distance between the statutory and the going wage. To account for this possibility, we include average wages, computed as described in section 7.5.[20] The results reported in column (6) indicate that controlling for the time evolution of the average wage of subpopulation $j = 1, \ldots, 12$ does not alter the results reported in columns (3) to (5).

Column (2) in table 7.6 summarizes the marginal effects, which give an estimate of the magnitude of the effects on different demographic groups. A 10 percent rise in the minimum wage reduces the employment probability of young unskilled workers by 0.005 (0.5 percentage points). While the effects on youth skilled workers are insignificant, the results indicate an adverse effect on prime-age unskilled workers. This is an interesting result in the context of a literature that almost exclusively focuses on the effects on youth workers.

While most of our findings are consistent with the competitive model,

19. See column (3) as well.

20. Including such variables is tantamount to including a set of noncoverage adjusted, demographic group-specific Kaitz ratios. However, we are not imposing the constraint that the coefficient on the minimum wage is the same as the coefficient on the group-specific average wage.

some are difficult to explain with this paradigm. For instance, this model cannot explain why minimum wages tend to shift employment toward women. One possible interpretation is that while men are able to obtain wages that are close to the competitive ones, women's wages are below their marginal products. This would be consistent with the systematic wage gaps found between observationally identical men and women and with the asymmetric gender effects of minimum wages. If wage gaps are explained by imperfect competition in female labor markets, employers are supply constrained when hiring women. Therefore, an increase in minimum wages reduces the demand for male workers and increases the supply of labor for women.

7.5.3 Total Effects

In our previous results, all the estimated coefficients measured the effects of labor regulations on each particular subpopulation *relative* to the omitted category, but they did not provide information on whether the employment probabilities of the different subgroups increased or declined in absolute terms after changes in policy. In this section, we attempt to gauge the total effects of labor market policies on the probability of employment by estimating their effect on the aggregate employment rates of prime-age skilled men (the omitted category in the specifications reported in table 7.5). To do so, we estimate the following error correction specification:

$$(2) \quad \Delta N_t = c - \lambda(N_{t-1} - N^*) + B_1(y_t - y_t^*) + B_2\Delta \log(W_t) + B_{3\Lambda}\Delta N_{\tau-L} + \varepsilon_t,$$

$$(3) \quad \text{where } N_\tau^* = \gamma_0 + \gamma_1 \log(\mathrm{JS}_t) + \gamma_2 \log(\mathrm{MW}_t) + \gamma_3 \log(\mathrm{Union}_t),$$

and where N_t denotes the employment rate—that is, the employment to population ratio—of prime-age male skilled workers in period t, N_τ^*denotes long-run equilibrium employment, $y_t - y_t^*$denotes GDP deviations from its trend (in logs), W_t denotes average wages for prime-age skilled male workers, JS_t denotes the measure of job security, MW_t denotes minimum wages, Union_t denotes the index of wage bargaining, and L is the length of the maximum lag. In expression (2), employment changes in function of previous period deviations from long-run equilibrium employment, GDP deviations from its trend, and changes in wages and short-run dynamics. Expression (3) assumes that, in the long run, employment rates are a function of labor market policies and the structure of wage bargaining.

Using aggregate time series techniques to estimate the effect of policies on the reference group allows us to model short- and long-run employment dynamics. The first step in the estimation of expression (2) and (3) is to test whether the variables are stationary. The first panel in table 7.7 reports the results of testing for the presence of unit roots using the Augmented Dickey-Fuller test (ADF). The tests are specified with three lags. In those cases in which the plot of the series indicated the presence of a time trend,

Table 7.7 **Unit Root and Cointegration Tests**

Names of the Series	Symbol	Specification	ADF Test Statistic	5% Critical Value
GDP deviation from its trend	$y - y^*$	Constant	–4.8412	–2.9472
Wage growth	$\Delta(\log W)$	Constant	–3.8514	–2.9705
Logarithm minimum wage	L(Minwage)	Trend	–1.4709	–3.5426
Logarithm job security	L(JS)	Constant	–2.43	–2.9472
Logarithm union centralization	L(Union)	Trend	–2.7568	–3.5426
Lagged employment rate	N_{t-1}	Constant	–1.6736	–2.9472
First diff. lagged emp. rate	ΔN_{t-1}	Constant	–3.0433	–2.9499
Change in log minimum wage	ΔL(Minwage)	Constant	–2.5591	–2.9499
Change in log job security	ΔL(Index)	Constant	–2.655	–2.9499
Change in log union	ΔL(Union)	Constant	–2.3443	–2.9499

Likelihood Ratio	5% Critical Value	Hypothesized Number of Cointegrating Equations
Johansen Cointegration Test: Series: N_{t-1} L(Minwage) L(JS) L(Union)		
108.64	53.12	None***
60.35	34.91	At most 1***
24.64	19.96	At most 2**
5.26	9.24	At most 3

***Denotes rejection of the hypothesis at the 1 percent significance level.

**Denotes rejection of the hypothesis at the 5 percent significance level.

we included a constant and a time trend in the specification; in the other cases, we included only a constant. While we can reject the unit root hypothesis for GDP deviations from its trend and for changes in hourly wages, we cannot reject nonstationarity for the lagged employment rate, the logarithm of minimum wages, the logarithm of the job security index, and the logarithm of union centralization. However, ADF tests on the first differences of these four series indicate that the hypothesis that these series are integrated of order one, $I(1)$, is not rejected.

Given the nonstationarity of the employment rate, expression (2) is well defined only if lagged employment deviations, with respect to the long-run equilibrium rate, are stationary. This is equivalent to saying that the series N_t^* has to cointegrate with N_{t-1}. The second panel in table 7.7 reports the results of the Johansen cointegration test between N^* and N_{t-1}. The likelihood ratio test indicates the presence of three cointegrating equations, indicating that the error correction model is well defined.

Table 7.8 presents the results of estimating the error correction model (ECM) once expression (3) has been substituted into expression (2). We use the results of the Akaike's Information Criteria (AIC) test to determine the optimal length of the lagged endogenous variable and determine that

Table 7.8 Level Effects on Male Prime-Age Employment

Independent Variable	(1)	(2)
N_{t-1}	–0.63	–0.66
	(–3.05)	(–3.24)
Deviations GDP_t	0.08	0.10
	(1.21)	(1.48)
Δ log Wt	—	0.018
		(0.84)
Log(JS)	0.011	0.015
	(1.80)	(2.23)
Log(Minwage)	–0.01	–0.014
	(–0.93)	(–1.13)
Log(Union)	0.03	0.029
	(1.54)	(1.45)
Constant	0.61	0.651
	(3.55)	(3.92)
ΔN_{t-1}	0.277	0.239
	(1.48)	(1.30)
No. of observations	37	35
Adj. R^2	0.16	0.23
Long-term effect of JS	0.017	0.023
Long-term effect of Minwage	0	0

Note: t-statistics shown in parentheses.

$L = 1$. We estimate the ECM with and without wages to see whether introducing wages alters our results, and we find the results to be very similar in both cases. Essentially, we find that job security provisions increase the long-run equilibrium rate of prime-age skilled male employment. This is not totally surprising. As mentioned in section 7.2, job security provisions increase the cost of dismissing workers with long tenure relative to the costs of dismissing less-tenured workers, reducing the layoff rate of the first relative to the layoff rate of the latter. Because prime-age skilled workers tend to have longer tenures than other, younger, less-skilled workers, job security provisions reduce the layoff rates of prime-age skilled workers relative to the layoff rate of other demographic groups. The positive sign in the ECM suggests that this effect on the layoff rate more than compensates for the negative effect of job security on employment creation. Instead, we do not reject the hypothesis that an increase in the minimum wage does not affect the employment rate of prime-age, skilled male workers, regardless of whether we control for the evolution of wages.

The estimated effect of job security provisions and minimum wages on the employment rate can be used to infer the total effect of these regulations on the employment probabilities of other demographic groups. In order to do so, the coefficients on job security provisions and minimum wages, reported in table 7.8, should be divided by (minus) the coefficient on the lagged employment variable to obtain the coefficients in expression (3). They reflect the magnitude of the long-run effect of regulations on prime-age skilled male employment. The third and fourth columns of table 7.6 present our estimates for the total effects. They are obtained by adding the marginal effect reported in the first and second columns of table 7.6 to the long-run elasticities obtained from specification (1) in table 7.8.[21]

The total effects reported in columns (3) and (4) suggest that job security provisions not only shift the distribution of employment toward older and skilled workers, but also increase their employment rates. Instead, more stringent job security provisions reduce the employment rates of young workers. Moreover, job security provisions reduce employment opportunities for women while increasing those of men. The magnitudes of these estimated effects are substantial. According to them, the 1990 labor reform, which increased our measure of job security by about one-third, reduced the employment rates of young unskilled male workers by 1.6 percentage points of the population.

We also find nonneutral effects of minimum wage spikes. Our estimates suggest that a 10 percent increase in minimum wages reduces the probability of employment for young unskilled male workers by 0.51 percentage points. Lastly, we find that a 10 percent increase in the minimum wage raises the employment rates of women by 0.46 percentage points.

7.6 Conclusions

The effect of regulations on employment is far from neutral across demographic subgroups. Paradoxically, job security and minimum wage regulations appear to be detrimental to the very workers that they are supposed to help. Our results suggest that both minimum wages and job security regulations reduce the employment opportunities of the young and the unskilled—and particularly unskilled youth—while promoting the employment rates of skilled and older workers. We have also found indications that job security regulations may force some workers, particularly women and the unskilled, out of wage employment and into self-employment. This paper has only examined the effects on employment. A complete analysis of who benefits and who loses from regulations would require examining the effects of regulations on the distribution of wages and benefits as well.

There is an ongoing debate on whether raising minimum wages and job security provisions have any effects on aggregate employment rates. However, even if researchers concluded that job security provisions or minimum wages do not have an effect in the aggregate, it is important to care-

21. The long-run effect of job security on the employment rates of middle-age skilled workers is computed as 0.011 divided by 0.63, which is equal to 0.017.

fully consider these distributional effects when evaluating their desirability. At best, these policies will help some disadvantaged workers, although perhaps at the expense of other poor workers. At worse, they distribute jobs from less advantaged to better-off workers.

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