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ABSTRACT

Economists have examined the impact of labor market regulations on the level of employment. However, there are many reasons to suspect that the impact of regulations differs across types of workers. In this paper we take advantage of the unusual large variance in labor policy in Chile to exa mine how different labor market regulations affect the distribution of employment and the employment rates across age, gender and skill levels. To this effect, we use a sample of repeated cross-section household surveys spanning the period 1960-1998 and measures of the evolution of job security provisions and minimum wages across time. Our results suggest large distribution effects. We find that employment security provisions and minimum wages reduce the share of youth and unskilled employment as well as their employment rates. We also find large effects on the distribution of employment between women and men.

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1. 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 sub-groups of workers bear the brunt or enjoy the benefits of such regulations has been much less studied.[2] 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.[3] Less attention has been paid to the issue of whether minimum wages particularly affect women or men or unskilled versus skilled workers. Similarly, very little attention has been paid to the effect that job security provisions may have on particular sub-groups of the labor force.

In this paper, 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. 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 the particular sub-groups employment rates. 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, in 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 the 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 these differential effects are driven by differences in labor supply elasticities or wage adjustments across sub-groups. Instead, our findings suggest that labor market regulations produce unequal shifts in labor demand across groups of workers.

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[2] 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.

[3] 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 and Krueger (1994, 2000) and Card, Katz and Krueger (1993) find no evidence of such an effect.

The rest of the paper is organized as follows. Section 2 reviews the arguments that predict non-neutral effects of regulations. Section 3 describes the evolution of job security and minimum wage regulations in Chile. Section 4 describes the data used in our empirical section. Section 5 describes the methodology implemented to estimate the effects of regulations on the distribution of employment. Section 6 describes our results for both the distribution of employment and the overall effect on employment rates. Finally, Section 7 concludes.

2. 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 sub-groups. 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.

2.1 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.[4]

Job security provisions will have differential effects across sub-groups of workers if changes in legislation bring changes in hiring and layoff rates that have a larger impact on some sub-populations 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 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

A 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.[5] 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 it can be defended that 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 sub-populations.[6] 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, who either are unemployed or hold temporary, part-time or fixed terms jobs without job security.[7] 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.

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[4] 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.

[5] See Pagés and Montenegro (1999) for a more formal development of this argument in the context of a partial equilibrium model.

[6] See for instance, Lindbeck and Snower (1988).

[7] 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.

Finally, differences in labor supply elasticity may contribute to differential effects across sub-populations 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.[8]

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

2.2 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 based on some form of imperfect competition predict wages lower than the marginal product, and thus, an increase in minimum wages can increase wages without reducing employment rates.[9]

On average, youth, women, and the unskilled tend to have lower wages than older, male or skilled workers. Therefore, since 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 sub-population. If that gap is larger in some groups than in others, an increase in minimum wages may have "competitive" effects on some groups and "non-competitive" effects on others. Given this ambiguity, the sign and magnitude of the effects become an empirical question.

3. 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.[10]

[8] See Hamermesh (1993).

[9] There are many situations that give raise to imperfect competition in the labor market, such like monopolistic power by part of employees, incomplete information or imperfectly mobile workers.

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

3.1 Job Security Provisions

Among the most interesting aspects of the Chilean experience is that, in the 39 years covered by our sample, Chile has gone from a situation of dismissal at will to a rigid labor market by OECD standards.[11] Since their inception in 1966, job security provisions have favored full-time indefinite employment over part-time, fixed-term of temporary contractual relationships. To this end, in case of a firm-initiated separation, labor codes regulate (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 noticed 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 widely varied over the years. This is the variance that we exploit in our empirical work.

Table 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.[12] In that event, the employer could choose between paying the mandatory compensation–plus wages foregone during trial—or reinstate the worker in his/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 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 a 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

[11] Heckman and Pagés (2000).
[12] Romaguera et al. (1995).

liable to pay compensation equal to one month's pay per year of work with a maximum amount of 11 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).[13] This measure is computed as follows:

$$JS_t = \sum_{i=1}^{T} \beta^i \delta^{i-1}(1-\delta)(b_{t+i} + a_t SP_{t+i}^{jc} + (1-a_t)SP_{t+i}^{uc})$$

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, SP_{t+1}^{jc} is the mandated severance pay in such event to a worker that has been i years at the firm, and finally, SP_{t+1}^{uc} denotes the payment to be awarded to a worker with tenure i in case of unjustified dismissal.[14]

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.

Based on the legal information summarized in Table 1 and assumptions regarding β, δ, a, and T, we obtain a measure of JS. We take β to be a constant value such that the average real interest is equal to 8.4%, 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 the ones observed in the US.[15] Davis and Haltiwanger (1992) report an average annual turnover rate of 12%. 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 courts. For the period 1966-1984, although despite that economic needs of the firm were considered just cause in the law, we assume a to be larger than zero and depending on labor courts stand.

[13] 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.

Finally, we assume T = 25. See Table 2 for a complete description of the parameters used in the computation of the JS measure.

The evolution of this variable over time is depicted in Graph 1. After some years of relatively low employment protection, JS increases eight-fold 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.

3.2 Minimum Wages

Columns two and three in Table 3 present the hourly real minimum wage in pesos of 1998.[16] These indices were constructed using Chile's Central Bank Bulletins. It is interesting to note that since 1989 there has been a lower minimum wage for workers 18 years old or younger. This wage has been fixed at a level between 15 and 20 percent of the adult wage. Graph 2 summarizes the evolution of the minimum wage in relation to the average wage for teen and adult workers. The graph shows that minimum wages are much higher, relative to each group average rate, for teen 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% and teen minimum wages by 104%. However, because average ages rose by more, 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 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 suggests that the minimum wage is binding, others like Bravo and Vial (1997) suggest that it is not.[17]

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[15] 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, since it is well established that dismissal costs in the U.S. are very small.

[16] Per hour minimum wages are constructed as monthly minimum wages divided by 4.2*40 hours.

[17] 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

4. 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 de 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 information contained in them is comparable across years.[18] During the period 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 de Chile represented about one third of Chile's total population, and a higher proportion of GDP.[19] 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 15 years old and no older than 65.

We merge labor policy and macro variables 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 3. 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 influence of unions in 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. Since union membership is not available either 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 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 18 or younger and for adult workers. The next two columns summarize the index of bargaining (column four presents the original index, and column five presents the smoothed index). The evolution of these variables over time is depicted in Graph 5. Higher values of this measure, like those registered from 1960 to 1970, reflect

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(1996). A more recent survey on the international evidence of minimum wages can be found in Dowrick and Quiggin (2003).

[18] In this study we use data from 1960 on, because the previous years (1957-1959) do not have reliable data.

[19] According to the 1992 Census, the metropolitan area accounted for 39 percent of the total population.

periods of higher union centralization.[20] The next seven columns summarize the average hourly wage broken down by sex (columns six and seven); skill level (columns eight and nine) and age group (columns ten, eleven and twelve). Column thirteen summarizes the deviation of the GDP from its potential or trend value. Finally, columns fourteen, fifteen and sixteen 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 15 and 65 years old. These three rates are also depicted in Graph 3, which jointly with Graph 4 (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.[21] Some additional indicators describing the performance of the Chilean economy are summarized in Table 4.

5. Methodology

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

$$y^*_{ijt} = X_{it}*\beta_1 + X'_{it}*Z_t*\beta_2 + \gamma_t + \varepsilon_{ijt} \qquad (1)$$

where

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

and y^*_{ijt} is an unobservable variable that determines whether an individual i, in sub-population j, at time t will be employed or not, and y_{ijt} is the observable employment status of this individual. In addition, X_{it} is a vector of variables that summarizes the personal characteristics of the individual i at time t, 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 (15-24, 25-50, and 51-65) and two skill levels (9 years of education or less, and more than 9 years). Adding the skill and the age groups to the gender division, we have 12 different sub-populations, $j=1,\ldots 12$

[20] 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.

In the vector of aggregate variables 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 18 and younger. By construction, the vector of coefficients on the interaction of X_{it} and Z_t, β_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 $\phi(.)X_{it}\beta_2$, where $\phi(.)$ is the normal density function.

Our original intention was to estimate

$$y^*_{ijt} = X_{it}*\beta_1 + X'_{it}*Z_t*\beta_2 + Z_t*\beta_3 + \varepsilon_{ijt} \quad (1')$$

With such a specification we could recover the *total* marginal effect of a labor policy on sub-population j as $\phi(.)(X_{it}\beta_2 + \beta_3)$. 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 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.

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 $i \in j$, $j=1,..,12$, at period t, the average wage of all employed workers in group j at period t.

We minimize the risk of omitted variable biases and spurious correlations in five 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 introducing time dummies, we control for macroeconomic trends and cycles as well as policy changes that affect the overall population. Third, by controlling for effect of changes in the business cycle (using GDP deviations from its trend) across individuals (that is, including $X'_{it} * Z_t$, where 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. Fourth, 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. Lastly, 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

[21] The Chilean economic performance has been extensively documented by Edwards and Cox-Edwards (1987, 2000), de la Cuadra and Hachette (1991), Wisecarver (1992), Bosworth, Dornbusch and Laban (1994), Hudson (1994), Soto (1995), and Cortazar and Vial (1998).

covered), we assess whether we are capturing the effect of policy, or instead, the effect of some unobservable correlate.

6. Empirical Results

6.1 The Effect of Job Security on the Distribution of Employment

Our results indicate that job security provisions have a differential impact across demographic sub-groups. In Table 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 sub-populations of workers are reported in Table 6. The t-tests, reported next to the coefficients are robust to the presence of heteroskedasticity of unknown kind 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 (26-50) 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 JS measure with dummies for age (young and older), gender (women) and skill level. A negative (positive) sign indicates that periods of more stringent JS provisions are associated with a decline (increase) in the probability of employment of a particular sub-population relative to the omitted category. We find strong age effects. The coefficient on the young-JS interaction is negative and statistically significant while the coefficient on the older-JS 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-JS interaction is negative and statistically significant, suggesting that JS provisions reduce the probability of employment of unskilled workers relative to skilled ones. Lastly, the coefficient on the female-JS interaction suggests a negative effect of JS 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% level, suggesting that job security provisions benefit 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 since 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 suggest that more stringent JS regulations increase the probability that women and the unskilled are employment 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 JS on the distribution of employment, but also help to infer the channels through which JS 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 Column (1)-(4) We find that an increase in JS 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, JS 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 JS 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 than in column (6) but controlling in addition by the average wage of each sub-population 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 sub-populations. 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 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.

In light of the different theories described in Section 2, how do we explain the results presented above? 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 above 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 that seems to be driven by a higher labor supply elasticity of this group.[22] Our results also suggest that these differential effects cannot be explained by insider-outsider theories, since in that case the effect would also be through wages. Instead, our results suggest that the differential 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 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 Higher rotation among women can explain why JS provisions affect young women less than young men. It can also explain why middle-aged and older women benefit less from JS 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. The model does not seem to able to explain why the effect on employment appears more negative on the unskilled than the on the skilled youth, but as we have seen, this effect seems to be driven by the more elastic labor supply of this group.

6.2 Distribution of the Effect of Minimum Wages

Table 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 similar qualitative effects on the distribution of employment across age and skill than stricter job security provisions. To account for contemporary employment policies and economic conditions we include measures of union activity, job

[22] 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.

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 shifts 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 wages hikes may increase the probability of employment for women relative to men.

Controlling for the sub-group effects of contemporary changes in policy and the business cycle does not alter the results reported in column (7) (See column (3)). The comparison between the results obtained from the wage employment and the self-employment specifications (column (3) and (4)) is also encouraging. As with the coefficients associated to 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.

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 rates of unskilled workers, particularly for middle-aged ones. 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 towards older workers. Lastly, 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 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 5.[23] The results reported in column (6) indicate that controlling for the time evolution of the average wage of sub-population $j = 1,\ldots,12$ does not alter the results reported in columns (3) to (5).

While most of our findings are consistent with the competitive model, some are difficult to explain with this paradigm. For instance, this model cannot explain why minimum wages tend to shift employment towards women. Assuming that women have higher marginal products than men and adding worker heterogeneity to the simple competitive model, this shift can be explained as a "flight to quality"

[23] Including such variables is tantamount to including a set of non-coverage 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.

effect. To see that, assume a population of heterogeneous workers that prior to the minimum wage increase were each paid their marginal value. After an increase in minimum wages, all workers with a marginal value below the new minimum wage cease to be employed. Assuming a perfectly elastic supply of all types of workers firms replace lower marginal value workers with higher value ones. This explanation, however, is at odds with the widespread observation that women's wages are lower than men's. Another 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 can expand both labor supply and employment rates.

6.3 Total Effects

In our previous results, all the estimated coefficients measured the effects of labor regulations on each particular sub-population *relative* to the omitted category, but they did not provide information on whether the employment probabilities of the different sub-groups 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-aged skilled men (the omitted category in the specifications reported in Table 5). To do so, we estimate the following error correction specification:

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

$$where \quad N_\tau^* = \gamma_0 + \gamma_1 Log(JS_t) + \gamma_2 Log(MW_t) + \gamma_3 Log(Union_t) \quad (2)$$

and where N_t denotes the employment rate–i.e. the employment to population ratio—of prime-aged 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 (1), employment changes in function of: previous period deviations from long-run equilibrium employment; GDP deviations from its trend; changes in wages and short run dynamics. Expression (2) 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 (1) and (2) is to test whether the variables are stationary. The first panel in Table 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 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 non-stationarity 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 non-stationarity of the employment rate, expression (1) 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 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 8 presents the results of estimating the error correction model (ECM) once expression (2) has been substituted into (1). We use the results of the AIC test to determine the optimal length of the lagged endogenous variable and determine that $L=1$. We estimate the ECM with and without wages to see whether introducing wages alters our results and find the results to be very similar in both cases. Essentially, we find that while job security provisions increase the long-run equilibrium rate of prime-aged skilled male employment. This is not totally surprising. As mentioned in Section 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. Since 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 JS 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-aged, 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 8, should be divided by (minus) the coefficient on the lagged employment variable, to obtain the coefficients in expression (2). They reflect the magnitude of the long-run effect of regulations on prime-age skilled male employment. The third and fourth columns of Table 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 6 to the long-run elasticities obtained from specification (1) in Table 8.[24]

The total effects reported in columns (3) and (4) suggest that job security provisions not only shift the distribution of employment towards 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 in about one third, reduced the employment rates of young unskilled male workers by 1.6 percentage points of the population.

We also find non-neutral effects of minimum wage spikes. Our estimates suggest that a 10% 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% increase in the minimum wage raises the employment rates of women by 0.46 percentage points.

7. Conclusions

The effect of regulations is far from neutral across demographic sub-groups. Paradoxically, job security and minimum wage regulations appear to be detrimental to those same 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, 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.

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

[24] 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.

8. References

Banco Central de Chile. 2001. "Indicadores Económicos y Sociales de Chile: 1960-2000." Banco Central de Chile, Santiago de Chile.

Bazen, S., and N. Skourias. 1997. "Is There a Negative Effect of Minimum Wages on Youth Unemployment in France?" *European Economic Review* 41(3-5): 723-32.

Bentolila, S. and G. Bertola. 1990. "Firing Costs and Labour Demand: How Bad is Eurosclerosis?" *Review of Economic Studies* 57: 381-402.

Bentolila, S. and G. Saint-Paul. 1994. "A Model of Labor Demand With Linear Adjustment Costs." *Labour Economics* (1): 303-26.

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

Bertola, G. 1991. "Labor Turnover Costs and Average Labor Demand." National Bureau of Economic Research Working Paper #3866. Cambridge, Massachusetts.

Bertola, G., F. Blau and L. Kahn. 2002. "Labor Market Institutions and Demographic Employment Patterns." National Bureau of Economic Research Working Paper #9043. Cambridge, Massachusetts.

Bosworth, B., R. Dornbusch and R. Laban, editors. 1994. *The Chilean Economy, Policy Lessons and Challenges*. The Brooking Institution, Washington D.C.

Bravo, D., and J. Vial. 1997. "La Fijación del Salario Mínimo en Chile: Elementos para una Discusión." *Colección de Estudios CIEPLAN* 43: 117-151.

Card, D., L.F. Katz and A.B. Krueger. 1993. "Minimum Wages and Employment: A Case Study of the Fast Food Industry in New Jersey and Pennsylvania." National Bureau of Economic Research Working Paper 4509. Cambridge, Mass.

Card, D., L.F. Katz and A.B. Krueger. 1994. "Employment Effects of Minimum Wages: Panel Data on State Minimum Wages Laws: Comment." *Industrial and Labor Relations Review* 47 (3): 487-97.

Card, D., and A. Krueger. 2000. "Minimum Wages and Employment: A Case Study of the Fast Food Industry in New Jersey and Pennsylvania: Reply." *American Economic Review* 90: 1397-1420.

Castañeda, T. 1983. "Salarios mínimos y empleo en el Gran Santiago: 1978 y 1981." *Cuadernos de Economía* 61.

Cortazar, R., and J. Vial. 1998. *Construyendo Opciones:* Propuestas Economicas y Sociales para el *Cambio de Siglo*. Santiago, Chile: Cieplan and DOMEN Editions.

Cowan, K., A. Micco, A. Mizala, C. Pagés, and P. Romaguera. 2003. "Un Diagnostico del Desempleo en Chile." Mimeo Inter-American Development Bank and Departamento Ingenieria Aplicada U. Chile.

Davis, S, and J. Haltiwanger. 1992. "Gross Job Creation, Gross Job Destruction, and Employment Reallocation." *Quarterly Journal of Economics* 107(3): 819-63.

De la Cuadra, S., and D. Hachette. 1991. "Lessons from the Chilean Stabilization and Recovery." In: B. Bosworth, R. Dornbusch and R. Laban, editors. 1994. "The Chilean Economy, Policy Lessons and Challenges." The Brooking Institution, Washington D. C.

Dowrick, S., and J. Quiggin. 1993. "A Survey of the Literature on Minimum Wages." Australian National University and University of Queensland. Mimeo.

Edwards, S., and A. Cox-Edwards. 1987. "Monetarism and Liberalization: The Chilean Experiment." The University of Chicago Press, Chicago, 2nd Edition, 1991.

Edwards, S. and Cox-Edwards, A. 2000. "Economic Reforms and Labour Markets: Policy Issues and Lessons from Chile." *Economic Policy* 15(30): 181-230.

Hamermesh, D. S. 1993. *Labor Demand.* Princeton, United States: Princeton University Press.

Heckman, J., and C. Pagés. 2000. "The Cost of Job Security Regulation: Evidence from Latin American Labor Markets." *Economía* 1(1): 147-151.

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

Hudson, R. 1994. "Chile: A Country Study." Library of Congress, Washington D. C.

Katz, L. and A. Krueger. 1992. "The Effect of the Minimum Wage on the Fast-Food Industry." *Industrial and Labor Relations Review* 46(1): 6-21.

Kosters, MH., editor. 1996. "Effects of the Minimum Wages on Employment." AEI Press for the American Enterprise Institute, Washington D.C.

Lazear, E. 1990. "Job Security Provisions and Employment." *Quarterly Journal of Economics* 105(3): 699-726.

Lindbeck A., and D.J. Snower. 1988. *The Insider Outsider Theory of Employment and Unemployment.* MIT Press, Cambridge and London.

Montenegro, Claudio E. 1998. "The Structure of Wages in Chile: 1960-1996: An Application of Quantile Regression." *Estudios de Economía* 25(1): 71-98.

Montenegro, Claudio E. 2002. "Unemployment, Job Security, and Minimum Wages in Chile: 1960-2001." Washington, DC. World Bank. Mimeo.

Pagés, C., and C.E. Montenegro. 1999. "Job Security and the Age Composition of Employment: Evidence from Chile." Inter-American Development Bank, Research Department, Working Paper # 398.

Paredes, Ricardo, and Luis Riveros. 1989. "Sesgo de Selección y el Efecto de los Salarios Mínimos." *Cuadernos de Economía* 79.

Partridge, M, and J. Partridge. 1998 "Are Teen Unemployment Rates Influenced by State Minimum Wage Laws?" *Growth and Change* 29(4):359-82.

Risager, O., and J.R. Sorensen. 1997. "On the Effects of Firing Costs When Investment Is Endogenous: An Extension of a Model by Bertola." *European Economic Review* 41(7): 1343-53.

Romaguera, P., C. Echevarria and P. Gonzalez. 1995. "Chile." In: G. Márquez, editor. *Reforming the Labor Market in a Liberalized Econom*". The Inter-American Development Bank, John Hopkins University Press, Baltimore, Maryland.

Soto, R. 1995. "Trade Liberalization in Chile: Lessons for Hemispheric Integration." In: E.L. Echeverri-Carroll, editor. *NAFTA and Trade Liberalization in the Americas*. Austin: University of Texas, Graduate School of Business, Bureau of Business Research in cooperation with the ICSS112 Institute.

White, H. 1980. "A Heterokedasticity-Consistent Covariance Matrix Estimator and a Direct Test for Heterokedasticity." *Econometrica* 48: 817-838.

Williams, N., and J. Mills. 2001. "The Minimum Wage and Teenage Employment: Evidence from Time Series." *Applied Economics* 33(3): 285-300.

Wisecarver, D. 1992. "El Modelo Económico Chileno." Centro Internacional para el Desarrollo Económico (CINDE). Instituto de Economía de la Pontificia Universidad Católica de Chile. Santiago de Chile.

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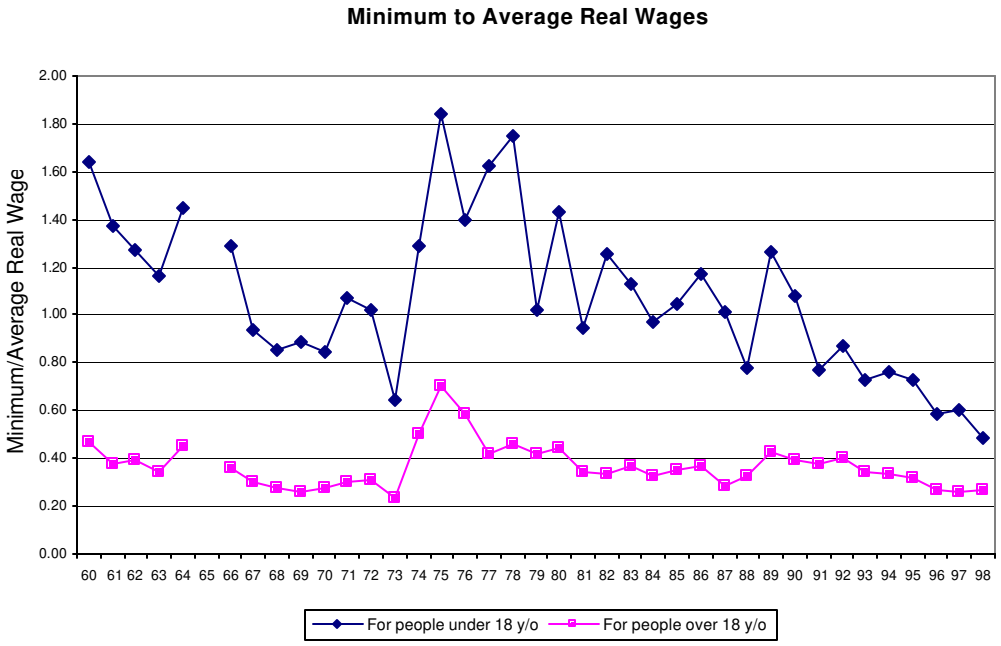
Graph 1

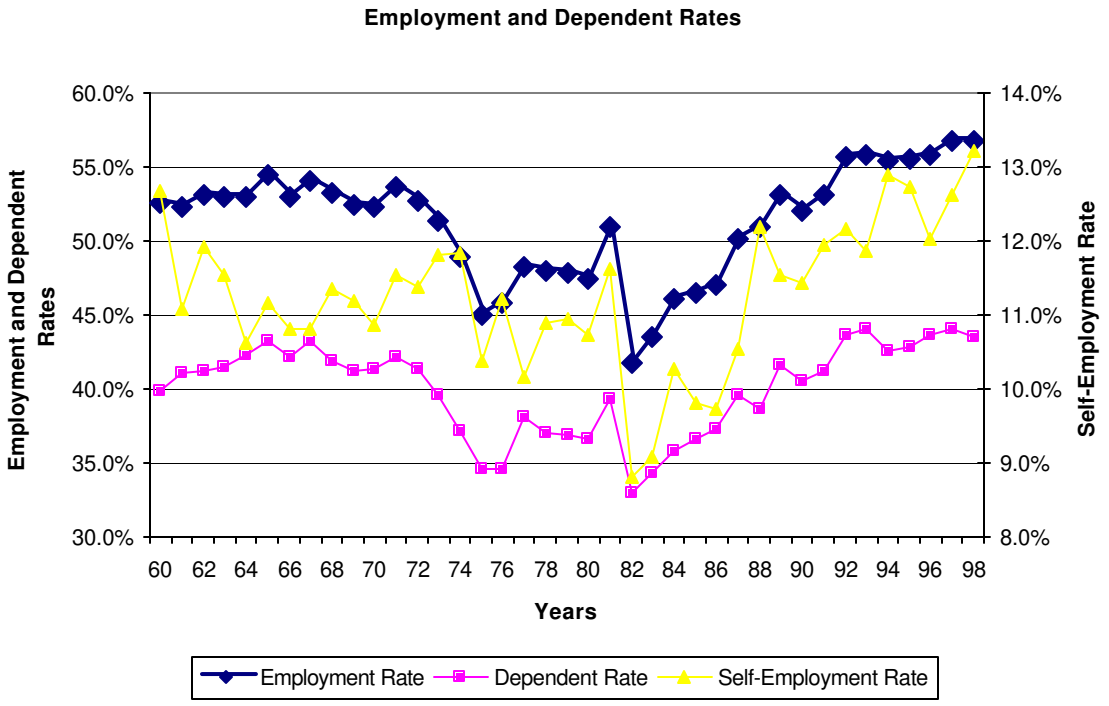
Source: Pagés and Montenegro (1999).

Graph 2

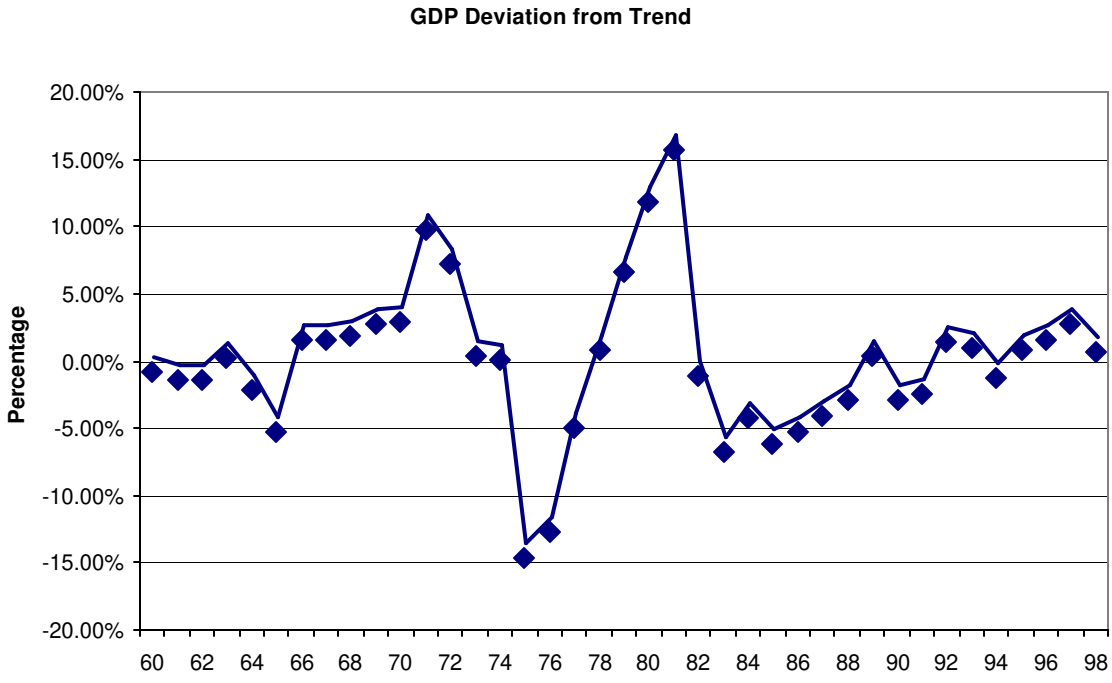


Source: Authors' calculations (see data section).

Graph 3

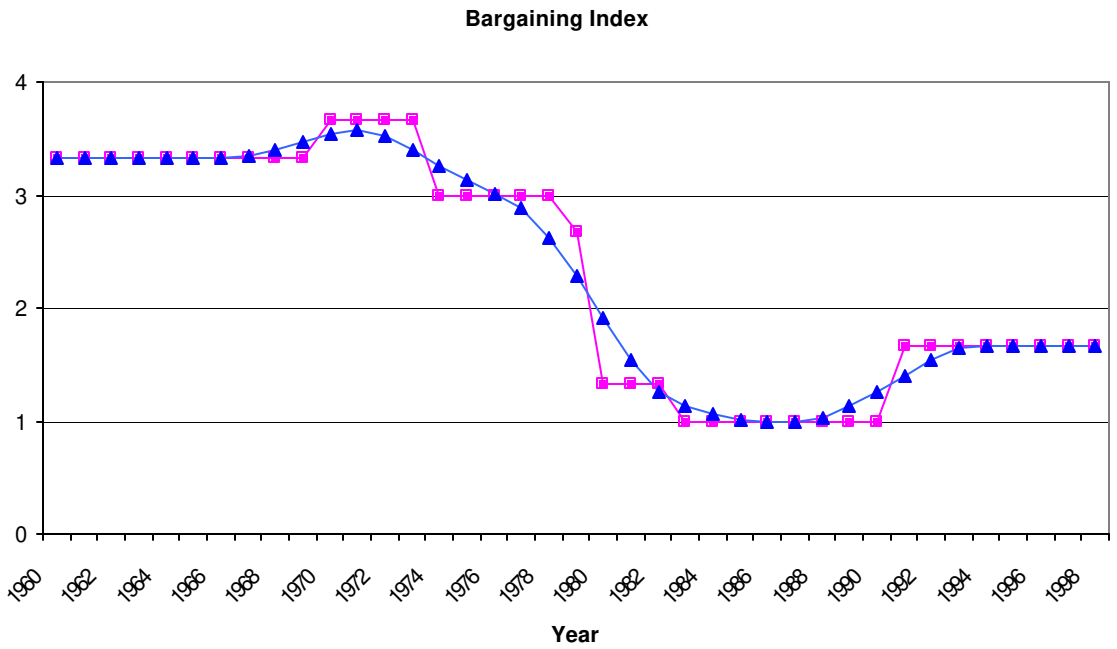


Graph 4



Source: Authors' calculations (see data section).

Graph 5



Bargaining Index measures the degree of centralization of wage bargaining. It takes values from 1 to 4. Higher values of colective bargaining. Source Edwards and Edwards 2000

Table 1. Employment Protection Provisions in Chile: 1960 – 1998

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	All workers
1973-1978	1 month	Labor courts were much more pro-firms. Workers' claims were weaker.	Same as previous period	Same as previous period	All workers
1978-1980 June 15, 78 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 no longer considered just cause for dismissal.	zero	1 month' wage per year of work *with a maximum of 150 days*	All workers
1990- today (Nov. 1990) Firms need to justify dismissals	1 month	Firms have to justify dismissals but economic needs are considered just cause for dismissal.	Economic reasons: 1 month' wage per year of work with a maximum of 11 months' pay	1.2-1.5 months per year of work	All workers hired after August 1981

Table 2. Parameters used to compute *Index*

	β	δ	*b*	***a***	$\mathbf{SP^{fc}}$	$\mathbf{SP^{uc}}$
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-	.92	.88	1	.9	(4)	(5)

Notes: 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 Haltiwanger (1995) average turnover rates average 12% 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 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.

Table 3. Basic Statistics of the Sample

Year	Job Security Index	Minimum Wage		Bargaining Index		Average Wage							GDP deviation from Trend	Employ-ment Rate	Wage Employ ment Rate	Self Employ-ment Rate
		Under 18 y/o	Over 18 y/o	Original	Smoothed	By Sex		By Skill Level		By Age Group						
						Male	Female	Low	High	15-24	25-49	50-65				
	Col. (1)	Col. (2)	Col. (3)	Col. (4)	Col. (5)	Col. (6)	Col. (7)	Col. (8)	Col. (9)	Col. (10)	Col. (11)	Col. (12)	Col. (13)	Col. (14)	Col. (15)	Col. (16)
60	0.5199	119	119	3.33333	3.33333	302	152	157	475	133	283	306	-0.86%	52.5%	39.8%	12.7%
61	0.5199	114	114	3.33333	3.33333	370	179	171	554	164	331	435	-1.41%	52.2%	41.1%	11.1%
62	0.5199	126	126	3.33333	3.33333	373	203	181	615	162	361	418	-1.37%	53.2%	41.2%	11.9%
63	0.5199	109	109	3.33333	3.33333	376	206	n.a.	311	219	342	395	0.20%	53.0%	41.4%	11.5%
64	0.5199	107	107	3.33333	3.33333	268	160	n.a.	230	133	272	296	-2.15%	52.9%	42.3%	10.6%
65	0.5199	114	114	3.33333	3.33333	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	n.a.	-5.23%	54.4%	43.3%	11.2%
66	3.9090	118	118	3.33333	3.33333	380	211	187	591	179	376	434	1.50%	53.0%	42.2%	10.8%
67	3.9090	116	116	3.33333	3.34724	427	268	222	648	217	420	539	1.50%	54.0%	43.2%	10.8%
68	3.9090	111	111	3.33333	3.39543	466	278	224	699	251	450	502	1.79%	53.2%	41.9%	11.4%
69	3.9090	107	107	3.33333	3.46403	475	279	231	709	218	470	560	2.79%	52.4%	41.2%	11.2%
70	3.9090	133	133	3.66667	3.53596	549	351	256	804	248	536	693	2.97%	52.3%	41.4%	10.9%
71	3.9090	183	183	3.66667	3.57675	689	437	302	957	307	660	779	9.67%	53.7%	42.1%	11.5%
72	3.9090	195	195	3.66667	3.52856	712	457	342	929	359	698	729	7.28%	52.7%	41.3%	11.4%
73	3.9090	108	108	3.66667	3.40525	525	332	279	671	280	512	553	0.37%	51.4%	39.6%	11.8%
74	1.8642	204	204	3	3.26140	435	310	275	561	255	436	496	0.12%	49.0%	37.1%	11.8%
75	1.8642	245	245	3	3.12419	376	277	225	483	214	376	420	-14.58%	45.0%	34.7%	10.4%
76	1.8642	259	259	3	3.01390	486	352	249	635	280	474	542	-12.67%	45.8%	34.5%	11.2%
77	1.8642	269	269	3	2.88227	692	512	320	953	357	696	786	-5.01%	48.3%	38.1%	10.1%
78	1.0599	346	346	3	2.62090	868	517	360	1090	400	799	1072	0.87%	48.0%	37.1%	10.9%
79	1.0599	345	345	2.66667	2.27455	913	640	432	1150	496	904	1009	6.66%	47.8%	36.8%	10.9%
80	1.0599	354	354	1.33333	1.90434	890	611	424	1120	476	881	932	11.83%	47.4%	36.6%	10.7%
81	0.8772	334	334	1.33333	1.53353	1057	799	510	1338	590	1099	1016	15.64%	50.9%	39.3%	11.6%
82	0.8772	365	365	1.33333	1.25825	1235	852	508	1499	618	1206	1295	-1.15%	41.8%	33.0%	8.8%
83	0.8772	276	276	1	1.13070	842	622	345	1056	416	872	721	-6.79%	43.5%	34.4%	9.1%
84	0.8772	243	243	1	1.06209	843	573	355	1028	371	845	780	-4.19%	46.1%	35.8%	10.3%
85	2.2915	220	220	1	1.01390	699	480	312	808	323	683	725	-6.19%	46.4%	36.6%	9.8%
86	2.2915	215	215	1	1	653	471	301	742	314	634	731	-5.35%	47.0%	37.3%	9.7%
87	2.2915	199	199	1	1	796	539	288	932	355	764	907	-4.05%	50.1%	39.5%	10.5%
88	2.2915	222	222	1	1.02781	766	542	316	902	376	751	799	-2.93%	50.9%	38.6%	12.2%
89	2.2915	293	340	1	1.12419	869	679	376	981	434	868	973	0.41%	53.1%	41.6%	11.5%
90	2.2915	298	346	1	1.26140	1003	682	390	1074	462	960	1011	-2.83%	52.0%	40.5%	11.4%
91	3.0598	278	327	1.66667	1.40525	971	694	401	1046	470	951	949	-2.47%	53.2%	41.2%	11.9%
92	3.0598	293	340	1.66667	1.54247	904	726	455	998	503	914	900	1.47%	55.7%	43.6%	12.1%
93	3.0598	294	341	1.66667	1.63885	1072	832	496	1158	627	1054	1093	0.98%	55.9%	44.0%	11.9%
94	3.0598	294	342	1.66667	1.66667	1141	840	535	1194	624	1101	1163	-1.22%	55.4%	42.5%	12.9%
95	3.0598	302	351	1.66667	1.66667	1230	919	566	1310	657	1215	1199	0.81%	55.5%	42.8%	12.7%
96	3.0598	279	324	1.66667	1.66667	1329	1047	621	1412	725	1283	1465	1.59%	55.8%	43.7%	12.0%
97	3.0598	248	333	1.66667	1.66667	1392	1100	613	1505	775	1380	1335	2.79%	56.7%	44.1%	12.6%
98	3.0598	243	341	1.66667	1.66667	1356	1136	759	1427	792	1325	1500	0.70%	56.8%	43.6%	13.2%

Source: Authors' calculations (see data section).

Table 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
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 Bank World Development Indicators Data Base and Gini coefficient from background data, Montenegro (1998).

Table 5. The Effect of Job Security and Minimum Wages
Probit Results

		(1)		(2)		(3) Wage		(4) Self		(5)		(6)		(7)	
Dependent variable:		Employed		Employed		Employment		Employment		Employed		Employed		Employed	
		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 mather	-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 Logarithm of Job Security with	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	-0.0546	-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 Logarithm of Minimum Wage with	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		
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		
	Number of Observations	303945		303945		303945		303945		303945		295318		303945	
	Pseudo R2	0.196		0.168		0.11		0.08		0.211		0.210		0.197	

Note: 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 0 otherwise (unemployed or inactive). The wage employment dummy variable is defined as 1 if the person is a dependent employee and 0 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 0 otherwise (dependent, unemployed or inactive).

Table 6. Marginal and Total Effects of Labor Market Regulations

	Marginal Effects		Total Effects	
	Job Security	Min. Wage	Job Security	Min. Wage
	(1)	(2)	(3)	(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 [0.000]	0.009	0.0596

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

Table 7. Unit Root and Cointegration Tests

Name 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	Δ *(logW)*	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	*Nt-1*	Constant	-1.6736	-2.9472
First diff. Lagged Emp. Rate	Δ *Nt-1*	Constant	-3.0433	-2.9499
Change in Log Minimum Wage	Δ *L(Minwage)*	Constant	-2.5591	-2.9499
Change in Log JS	Δ *L(Index)*	Constant	-2.655	-2.9499
Change in Log Union	Δ *L(Union)*	Constant	-2.3443	-2.9499

Panel 2: Johansen Cointegration Test

Series: *Nt-1 L(Minwage) L(JS) L(Union)*

Likelihood Ratio	5% critical Value	Hypothesized number of CE
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 5% (1%) significance level

Table 8. Level Effects on Male Prime-Age Employment

	(1)	(2)
Independent Variables:		
Nt-1	-0.63	-0.66
	(-3.05)	(-3.24)
Deviations GDPt	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)
Δ *Nt-1*	0.277	0.239
	(1.48)	(1.30)
N obs.	37	35
Adj. R squared	0.16	0.23
Long term Effect of *JS*	0.017	0.023
Long term Effect of *Minwage*	0	0

t-statistics shown in parenthesis.