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Labor Market Reforms and Their Impact over Formal Labor Demand and Job Market Turnover
The Case of Peru

Jaime Saavedra and Máximo Torero

2.1 Introduction

After the expansionary phase of the "heterodox" experiment (1986–1987) of the García government, the Peruvian economy fell into a very deep recession. Output fell between 1988 and 1990, in the midst of a hyperinflationary process. The Fujimori government implemented a harsh macroeconomic stabilization program in August 1991, and a few months later a comprehensive set of structural reforms was launched. Peru experienced one of the fastest trade liberalization processes and one of the deepest labor market reforms in Latin America. These reforms were accompanied by a downsizing of the public sector, the start of a privatization process, the abolition of all state-owned monopolies, and a tax reform. In addition, restrictions to capital account transactions were eliminated while the financial sector was deregulated.

The Peruvian Labor Code, developed during the import substitution period, had been termed one of the most restrictive, protectionist, and cumbersome of Latin America (International Labor Organization [ILO] 1994). The code was extremely complex and comprised a collection of overlap-

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ping decrees that had undergone many changes over time. The military government of 1969–1975 made firing extremely difficult by sanctioning job stability after a probationary period. In 1985, the García government reduced the probationary period to just three months, during what was the period of most rigid labor market legislation. In 1991, labor market regulations were relaxed through a succession of reforms. Firing costs diminished sharply through the progressive elimination of job stability regulations, the reduction in red tape for the use of temporary contracts, and changes in the severance payment structure. In addition, firms in the formal sector faced high nonwage costs: payroll taxes, social security and health contributions, a tenure bonus, training fund contributions, family allowances, and a long thirty-day vacation period. During the 1990s overall nonwage costs increased slightly.

One first adjustment mechanism to a restrictive labor legislation is the use of informal contracts. In this sense, changes in firing costs expected by the firm and in nonwage labor costs have an impact on the distribution of employment between the formal and informal sectors but not necessarily on overall employment. If firing costs are perceived by firms as a tax imposed on layoffs, a reduction, like the one observed in Peru given the fall in expected severance payments, and the abolition of job stability and the facilities given for the use of temporary contracts will increase the equilibrium employment level. Moreover, reductions in expected firing costs may have an effect on the response pattern of firms to changes in product demand, which may be reflected in larger employment-output elasticities. In this paper, we analyze the impact of changes in expected severance payments and labor costs by estimating labor demand functions for the formal sector. We use data from firm-level surveys for formal firms in Metropolitan Lima. With these data we construct a pseudo-panel data set of ten economic sectors observed bimonthly during the period 1987–1997 and three shorter panels of about 400 firms for the periods 1987–1990, 1991–1994, and 1995–1997, dates dictated by sample changes.

Also, reductions in labor legislation–related firings costs typically accelerate the process of job creation and job destruction, therefore increasing turnover and reducing job duration, particularly in the formal sector. We examine changes in job duration and labor market turnover using data from a series of annual household surveys, with which we analyze changes in mean tenure in both the formal and informal sectors. Informality is conceptualized here as a state chosen by firms and workers depending on a cost-benefit analysis. Many firms, typically smaller ones, operate totally underground, fire and hire at will, and do not pay any kind of socially mandated benefits. In most of the cases, their productivity is too low for them to afford to pay any kind of benefits. To operationalize this we define a worker as working in the formal sector if he or she receives social benefits or belongs to a union. In addition, using the Living Standards Measurement Survey, we construct complete and incomplete employment spells with which we calculate empirical hazards for different subsamples, and we estimate exponential hazard models.

The paper proceeds as follows. In section 2.2 we analyze the legal context regarding the probationary period, severance payments, nonwage costs, and temporary contracts, all factors that affect firm and worker behavior. We also describe changes in employment in Metropolitan Lima during the period of analysis and discuss how informality and temporary contracts have been mechanisms through which firms avoid paying mandated benefits and firing costs. In section 2.3 we present results of labor demand estimations at both the sectoral and firm levels. Finally, in section 2.4, we analyze basic patterns of employment duration. In order to assess possible impacts of labor laws, we compare patterns of the self-employed with those of wage earners in the formal and informal sectors. We present a comparison of job duration among different groups of workers using empirical hazards, and we show the results of exponential hazards functions.

2.2 Changes in the Regulatory Framework during the 1990s

Prior to the reforms, the Peruvian Labor Code was extremely complex and comprised a large collection of overlapping decrees. Formal workers enjoyed several employment stability provisions, payroll taxes and social security contributions were high, and collective bargaining and other regulations gave unions great power. Since 1991, labor market regulations were relaxed through a succession of reforms. In this section we describe the changes in firing costs determined by the severance payment and job stability regulations, the changes in regulations and in the use of temporary contracts, and the evolution of nonwage labor costs.

2.2.1 Severance Payments and Job Stability

The costs of firing in Peru comprised two main elements, mandated severance payments upon dismissal and the costs imposed by job stability regulations. The military government of General Velasco introduced severance payments in 1970, as a fixed value equivalent to three months' wages upon dismissal without "just cause." It was conceived as a compensation to the hardship of dismissal and simultaneously as an unemployment insurance device. In addition to severance payments, Peruvian labor laws had very rigid employment protection clauses, which increased firing costs dramatically. During the period 1971–1991, a worker who completed the probationary period—the length of which was changed a few times—was granted absolute job stability. That meant that if a firm dismissed a worker and could not prove just cause in labor courts, he or she could choose between being reinstated in the job and receiving the severance payment. This made the severance payment the lower bound of the firing cost, as workers

Table 2.1 Probationary Period and Job Stability Regulations

	Probationary Period Length	Job Stability Status	Temporary Contracts Availability
Before June 1986	3 years	Granted after 3 years	Low
June 1986–October 1991	3 months	Granted after 3 months	Low
November 1991–July 1995	3 months	In effect only for workers hired before November 1991	High
After July 1995	3 months	Abolished	High

had the incentive to ask to be reinstated in their jobs, and then settling out of court became a larger severance payment. This setting also implied high administrative and litigation costs. Just cause did not include economic reasons, and workers could be fired due only to serious misdemeanor or through complicated collective layoffs. From the employers' perspective, a worker was effectively "owner of his post."

In 1978, the length of the probationary period was increased to three years (see table 2.1). The severance payment schedule was raised, and workers with less than three years in a firm received the equivalent of three months' wages if fired without notice, while workers with longer tenures received twelve months' wages upon dismissal. During the probationary period, the employer had to inform the worker in advance if he wanted to fire him to avoid the severance payment.

Since June 1986, the probationary period was reduced again, to just three months, and a large portion of workers suddenly acquired total job stability. An interesting feature here is that the change was announced in June 1985, about a year before the law was effectively sanctioned. Casual evidence for that year shows that employers did not increase layoffs massively among workers with less than three years of tenure who had not concluded their probationary period. Given that the economy was starting an expansionary period, it is probable that business expectations regarding higher demand were on the rise, which reduced the incentive of employers to fire workers that could potentially receive job stability rights. Still, the announcement of the policy change, ceteris paribus, must have had a positive effect on turnover for these workers. The severance payment was set to the equivalent of three months' wages for those workers who had been employed between three months and one year, six months' wages for those with one to three years of tenure, and twelve months' wages for those with more than three years of tenure (see García schedule in figure 2.1 and table 2.2).

The June 1986 changes in labor laws by the García administration made the 1986–1991 period the one with the highest degree of rigidity, as severance payments were high, the probationary period was short, and job stability rights were still in place. Rigid job protection pushed firms to seek ways to get around these regulations. One way was to lobby for the gener-

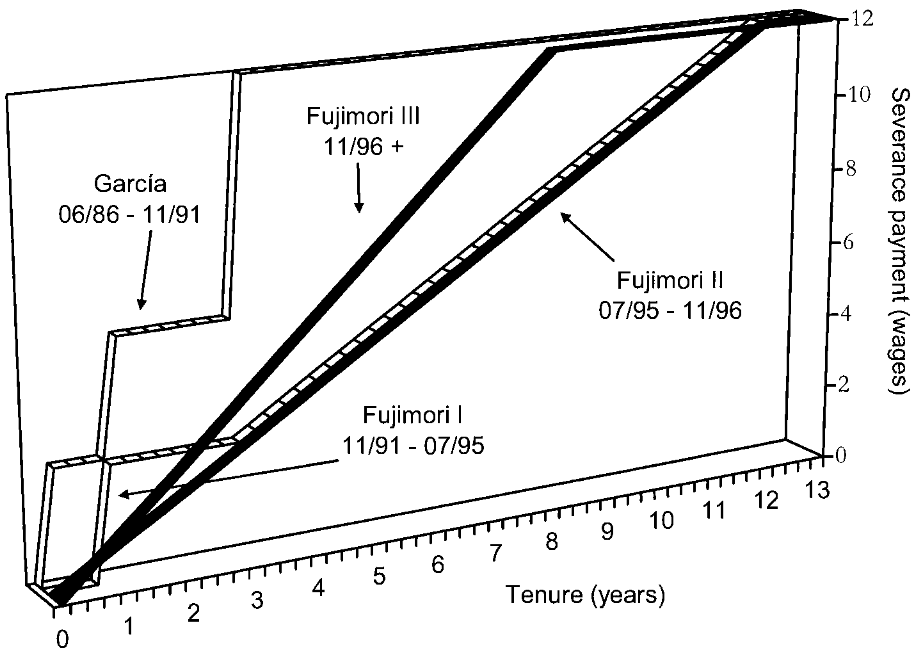


Fig. 2.1 Severance payment regimes

ation of the so-called Emergency Employment Program. The other form was to fire workers a few days before they completed the three-month probationary period and then rehire them. Another form of eluding these regulations was making workers sign an undated letter of resignation at the beginning of the contract period.

In 1991, the government introduced several changes aimed at reducing the extreme rigidity imposed by labor laws. The intention of the drafters of the Law Decree 726 of November 1991 was to abolish job stability. However, the right to job stability was written in the 1979 Constitution, so, in principle, only through a two-year process could the Congress pass a law approving a constitutional change. The outcome was the creation of a dual regime in which workers with contracts signed before November 1991 maintained their job security rights, while new workers would only have protection against unjustified dismissal. This meant that these workers could be dismissed at will upon payment of a severance benefit. In addition, just cause clauses were extended to include issues related to workers' productivity.[1] Also, the severance payment rule was modified in order to reduce firing costs. It was fixed at one months' wage for every year of tenure for workers with more than one year in the firm, with a minimum of three

1. In practice, it was very difficult for firms to use these clauses due to administrative problems in proving a reduction of productivity.

Table 2.2 **Severance Payment Legislation: Rules and Examples**

	Rule		Worker's Tenure		
	Tenure	Severance	2 Years	10 Years	20 Years
June 1986–November 1991	3 months–1 year =	3 months' wages			
	1–3 years =	6 months' wages	6 months' wages	12 months' wages	12 months' wages
	More than 3 years =	12 months' wages			
November 1991–July 1995	3 months–1 year =	0 months' wages			
	1–3 years =	3 months' wages	3 months' wages	10 months' wages	12 months' wages
	3–12 years =	1 months' wage per year			
	More than 12 years =	12 months' wages			
July 1995–November 1996	3 months–12 years =	1 months' wage per year	2 months' wages	10 months' wages	12 months' wages
	More than 12 years =	12 months' wages			
November 1996+	3 months–12 years =	1.5 months' wages per year	3 months' wages	12 months' wages	12 months' wages
	More than 8 years =	12 months' wages			

months' wages and a maximum of twelve months' wages, as shown in figure 2.1 (Fujimori I schedule).

In July 1995, with the second wave of labor reforms, the severance payment schedule was simplified to one month per year of work up to a maximum of twelve months (Fujimori II schedule in figure 2.1). As the 1993 Constitution replaced the right to job stability with the right to unjustified dismissal, the 1995 law eliminated job security rules and the two-tier regime. These changes, plus the reduction in severance payments, implied a sharp reduction in firing costs, which may be interpreted as a lower level of the tax on dismissals perceived by firms. This may have the effect of giving formal firms more flexibility to adapt to output changes, of increasing the formal employment level, and also of increasing the output elasticity in labor demand estimations for formal firms. In addition, reductions in firing costs typically accelerate the process of job creation and job destruction, therefore increasing turnover. Finally, in November 1996 the severance payments rule was again modified. Instead of receiving one months' wage for each year in the firm, the employee received one and a half months' wages, an important large increase in the firing costs of low-tenured workers. The maximum cap of twelve months' wages remained unaltered (Fujimori III schedule in figure 2.1).

Quantifying the Severance Payment

The severance payment rule has an effect on the amount of resources firms have to reserve to finance dismissals. Given that in Peru, as in many other Latin American countries, these payments are linked to tenure, these reserves will vary depending on the tenure structure of the workforce of the firm. In turn, the firm's tenure structure may be endogenous to the severance payment rule, as firms will try to avoid hiring workers who will later be relatively more expensive to dismiss. The tenure structure will also depend on technology and other characteristics of the firm and sector.[2]

We calculated the evolution of potential reserves for severance payment as a commodity contingent on a firing (F) or a hiring (H) state of the economy.[3] We may therefore think of a firm as choosing among probability distributions or "prospects" whose uncertain consequences are to be received with respective state probabilities $\pi = (\pi_H, \pi_F)$. Specifically, expected severance payment is calculated by state and sector using the evolution of the tenure structure, an estimate of the firing probability for each tenure group, and the corresponding mandated severance payment. The following formula describes how it is calculated (time subscripts have been eliminated):

2. For instance, the share of long-tenure workers will generally be larger in the manufacturing sector, where firm- and sector-specific knowledge is more important than in trade.

3. This is following the expected utility rule of John Von Neumann and Oskar Morgenstern.

$$E(\mathrm{sp})_i = \pi_F\left[\sum_X \lambda(X)_{i,F} \cdot N_X \cdot \mathrm{sp}(X)\right] + \pi_H\left[\sum_X \lambda(X)_{i,H} \cdot N_X \cdot \mathrm{sp}(X)\right]$$

$E(\mathrm{sp})_i$ is the expected severance payment, which is a probability-weighted average for the severance payments in each of the states, hiring and firing, and sector i. The first bracketed portion corresponds to the severance payment for the firing state and the second to that for the hiring states, which are weighted by π_H and π_F, the probabilities of being in a hiring (H) or in a firing (F) state of the economy, respectively. The severance payment in each of the bracketed sections for sector i is calculated by multiplying a time-invariant sector-specific and state-contingent firing probability, $\lambda_{i,\mathrm{state}}(X)$; by the number of workers in a specific tenure group (N_X); and by the mandated severance payment that will have to be paid to employees in that group if they are fired, $\mathrm{sp}(X)$. X denotes the tenure group. To calculate this firing probability we used the average employment reduction by tenure group in each possible state (hiring and firing), and when employment grew we assumed zero variation. Because of this, we obtained a constant probability across the whole period that was different across sectors, tenure groups, and states. Data on the structure of tenure groups and employment changes by sector come directly from the Quarterly Survey of Wages and Salaries (QSWS).[4]

Figure 2.2 shows the evolution of $E(\mathrm{sp})$ for the period 1986–1996 as a percentage of total wages. Note that we are fixing the sector-specific firing probability, so, in this aggregate, changes may only be attributed to changes in the employment share of different sectors and changes in legislation. The first large fall in the index is at the end of 1991, and it reflects the reduction in the mandated severance payment schedule. Further changes are related to increases in the share of short-tenure groups. Changes observed in June 1995 coincide with a further reduction in mandated severance payments, while the increase in August 1998 coincides with an increase in these payments. On average, reserves that firms had to maintain for severance payments were reduced from 16 percent of the wage bill to around 8 percent after the reforms.[5]

4. The survey includes a sample of workers per firm, from which we calculate the firm tenure structure. N_x is calculated with this structure and total firm employment. The characteristics of the QSWS will be described presently.

5. Figure 2.2 also shows an "adjusted" $E(\mathrm{sp})$ for the period 1992–1995. The increase in the calculated $E(\mathrm{sp})$ between 1992 and 1995 is related to an undersampling of newer firms. During those years the sample was not renewed, so only "deaths" were registered. As no new firms entered the sample, older firms, which tend to have older workers, are overrepresented. This implies a tenure structure biased toward older workers, therefore increasing the $E(\mathrm{sp})$. In the calculation of the employment series this problem was tackled through expansion factors that weighted the original data in order to take into account sample changes in the structure of firms by size.

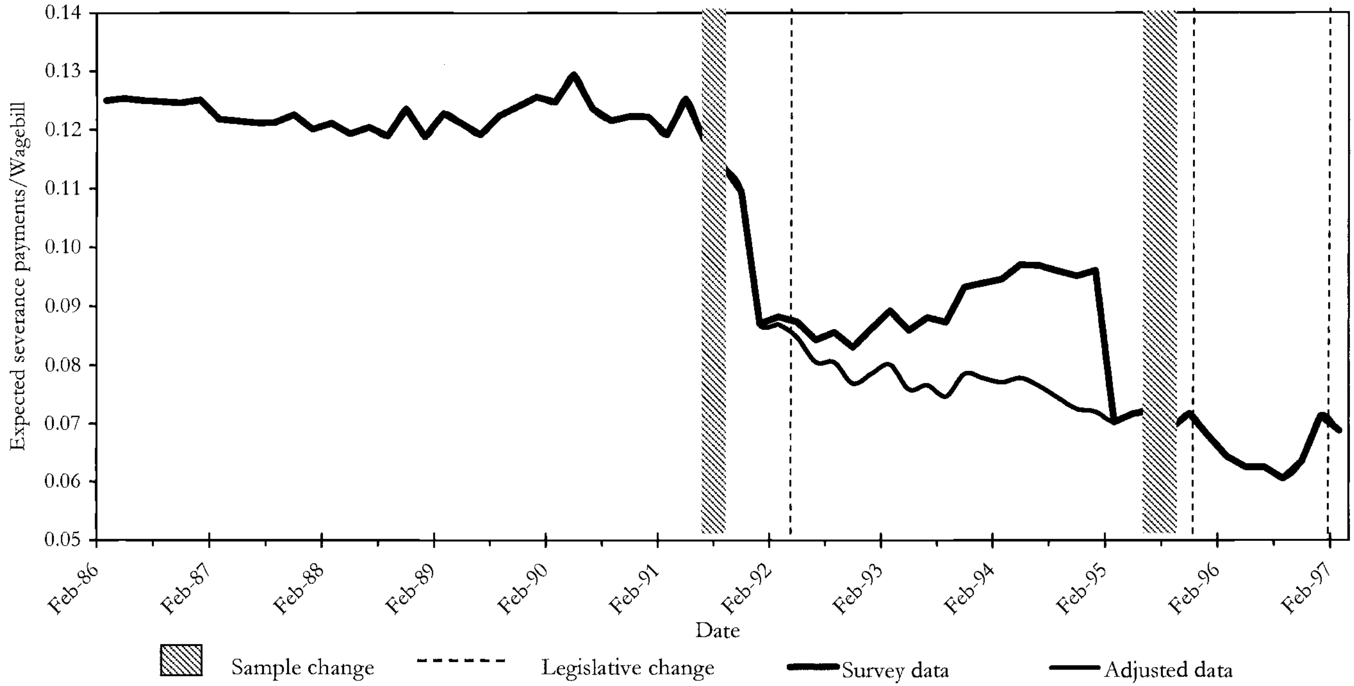


Fig. 2.2 **Expected severance payments as a percentage of total wages, 1986–1996**

2.2.2 Reducing Rigidities: Temporary Contracts

One possible way of bypassing the large adjustment costs imposed by employment protection policies is through lobbying the government to introduce short-term or temporary contracts. Temporary contracts were introduced in 1970. Firms required prior authorization from the Ministry of Labor in order to use them, and contracts were allowed under very specific circumstances. In practice, the high administrative costs this process implied restricted their use heavily. As shown in table 2.3, between 1986 and 1990, around 20 percent of workers in formal firms were under temporary contracts. Most of them carried full social benefits but had no employment protection clauses (*contratos sujetos a modalidad*), and important proportions of temporary workers were probationary-period workers. During the short-lived populist boom of 1987, in the midst of a period of extreme job protection, firms were allowed to hire using short-term temporary contracts through an emergency employment program (*Programa Ocupacional de Emergencia,* or PROEM). These contracts, which could last up to a year, were used mainly by large formal firms.

In August 1991, with the first wave of labor reforms, red tape for the use of fixed-term contracts was significantly reduced, and the reasons that could be used to justify hiring a worker under this type of contract were increased. The Ministry of Labor confined its role to record keeping and charging a fee for each contract. In general, in contexts of restrictive job protection regulations the output elasticity of temporary contracts is larger than that of permanent contracts, given that usually they do not carry firing costs (Bentolila and Saint-Paul 1992). In Peru, despite the reduction in firing costs for new workers under permanent contract in 1991, firms still preferred the now easier-to-use temporary contracts. The share of workers under these contracts increased from 20 percent in 1991 to 31 percent in 1992, and most of formal private employment growth observed

Table 2.3 Metropolitan Lima: Structure of Total Private Formal Salaried Employment, 1986–1997 (%)

	1986	1987	1989	1990	1991	1992	1993	1994	1995	1997
Total	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0	100.0
Permanent	80.7	82.1	82.9	80.8	80.1	68.6	67.9	64.8	59.8	56.0
Temporary	19.3	17.9	17.1	19.2	19.9	31.4	32.1	35.2	40.2	44.0
Fixed-term contract	19.3	17.7	14.3	19.2	19.6	30.0	29.8	33.3	39.4	39.9
Youth contracts	0.0	0.1	0.0	0.0	0.3	0.0	0.4	0.0	0.3	2.5
Probationary period	—	—	2.7	—	—	1.4	1.9	1.9	0.5	1.6

Source: Encuesta de Hogares del MTPS 1986–1995, Encuesta Nacional de Hogares del INEI 1997.

Note: Not all the surveys between 1986 and 1997 allow the separation between workers under fixed-term contracts and those under probationary periods. Dashes indicate that there was no probationary period in those years.

during the 1990s was explained by temporary contracts. Moreover, even after the elimination of the two-tier system in 1995 with the elimination of job stability for all workers, as well as an additional reduction in severance payments, temporary contracts continued growing, covering 44 percent of private formal wage employment in 1997.[6] This could be explained by the fact that firing costs for permanent workers, even if smaller than before, are still high or that firms may be reluctant to hire workers as permanent employees because they fear a setback in the progress toward flexibility. In fact, a change in the severance payments schedule in 1997 implied an increase in firing costs.[7] In our estimations we cannot distinguish permanent from temporary contracts; however, the lower administrative costs of using temporary contracts should imply a larger output elasticity after the reforms.

2.2.3 Nonwage Costs

In Peru an important source of public finance is payroll taxation. This burden has been heavily criticized, mainly along the lines that these contributions increase labor costs, reduce competitiveness, and have possible negative effects over employment. Peru has a complicated and unstable structure of nonwage labor costs, a description of which follows.

- *Public and private retirement plan payments.* Between 1986 and 1993, the employer had to pay to the public pension agency, the Instituto Peruano de Seguridad Social (IPSS), a contribution of 6 percent of the employee's wage, while the employee had to pay 3 percent. Poor and corrupt management, increasing numbers of retirees, and inflation led to the near collapse of the pay-as-you-go public system. In 1993, a private pension system was created, with individually held accounts managed by institutions called the Administradoras de Fondos de Pensiones (AFPs). Currently, both pension systems coexist. In 1995, after a few changes, the rate was set at a total of 11 percent in both systems, and the entire contribution had to be paid by the employee.[8]
- *Health plan payments.* The public health plan offered by IPSS is still the only option for workers. The total contribution rate has been fixed at 9 percent during the last few years. However, its composition with respect to employers and employees has changed: Before 1995 the em-

6. By 1997, according to Household Survey data, 316,000 private salaried workers in Lima had signed temporary contracts. According to the administrative records of the Ministry of Labor, 434,000 new contracts were signed that year. As a percentage of the total employment in Lima (i.e., including public workers and the informal sector), the share of workers under this type of contract reached 24 percent.

7. A surprisingly large output elasticity of temporary contracts was also observed in Spain in 1986, when the economy picked up and restrictions for the use of temporary contracts had been lifted, and almost all job creation was explained by this type of contract. Between 1987 and 1990, the share of temporary contracts increased from 15 percent to 32 percent.

8. See details in table 2A.1.

ployer had to pay 6 percent and the employee had to pay 3 percent. Currently, the employer must pay the entire contribution fee.

- *Accident insurance.* The employer is required to pay a accident insurance for his blue-collar workers. The amount is calculated as a rate of the employee's salary. This rate varies depending on the level of risk involved in the job and averages around 2 percent.
- *Manufacturing training fund (SENATI).* This is paid by the employers of firms in manufacturing industries. Initially it was set at 1.5 percent of the worker's income. In 1995, it was reduced to 1.25 percent, in 1996 to 1 percent, and in 1997 to 0.75 percent.
- *National Housing Fund (FONAVI).* Originally created as a contribution to workers' housing needs in the late 1970s, the National Housing Fund (FONAVI) rapidly resulted in a costly payroll tax, mainly due to inefficient and faulty management of collected funds.[9] Up to 1988, the FONAVI contribution paid by the employer was 4 percent of the employee's wage, while the employee's rate was 0.5 percent, and the maximum taxable wage was set at eight tax units (UITs). In November of that year, the employer's contribution rate was increased to 5 percent and the employee's rate to 1 percent. In May of 1991 the employer's rate was set at 8 percent, while the employee's rate remained unchanged, raising the total contribution to 9 percent and further widening the gap between the amount paid by the employer and the amount received by the employee. In January 1993 the employer's contribution responsibilities were abolished altogether, and the employee's rate was set at 9 percent. Even though the total contribution rate remained constant (at 9 percent), the maximum effective taxable wage was abolished, which might have increased the effective rate. Ten months later, due to harsh political pressures, the employee's contribution rate was diminished to 3 percent and the employer's rate was increased to 6 percent. In August of 1995 the employee's contribution was abolished and the employer's contribution rate was set at 9 percent. Finally, in January of 1997, the total contribution was reduced to 7 percent (paid completely by the employer), but the Christmas and holiday bonuses of a monthly salary were included in the taxable base.
- *Individual savings account (Compensación por Tiempo de Servicios, or CTS).* This is additional wage paid by the employer to the employee for every year of worker tenure. Prior to January 1991, the employer paid a maximum bonus of ten minimum wages if the employee's wage

9. As a result of this, FONAVI became an important issue in political discussion, as opposition parties used it as justification to attack the government, while the latter constantly shifted the FONAVI rate back and forth between employers or employees and altered its total level, to satisfy political and financing needs. Throughout the document, when talking about the payroll tax, we refer to this contribution.

was higher than that amount. Employers were allowed to keep those funds until an employee left the firm (the only obligation being to register it in the firm's balance sheet as a liability). The system failed due to employers' lack of compliance in actually keeping these bonuses for workers. In actuality, when a worker was fired, the payment of this bonus worked as an additional firing cost. Since January 1991 the employer has had to deposit 50 percent of an employee's monthly salary in an individual account in the worker's name in a commercial bank on May and November of each year.

- *Christmas and national holiday bonuses.* On December 1989, it became obligatory for the employer to pay two additional months' wages to his employees (on July and December of each year). However, this was already a common practice before the law was established, especially in medium-sized and large firms. In the public sector, these bonuses had been paid regularly to employees for several years, since the mid-1980s, but the amount varied.

Figure 2.3 shows the evolution of the effective rate paid by a firm in the case of a blue-collar worker who is affiliated to a public pension plan. To calculate the nonwage costs' effective rate it was necessary to estimate each of the nonwage costs just listed. The main difficulty in the estimation was to combine the effect of the different rates with the maximum and minimum taxable bases, so we calculated each of the nonwage costs separately and then summed them together. Most of the sources of change are related to cap changes in the tenure bonus and changes in the payroll tax rate. In

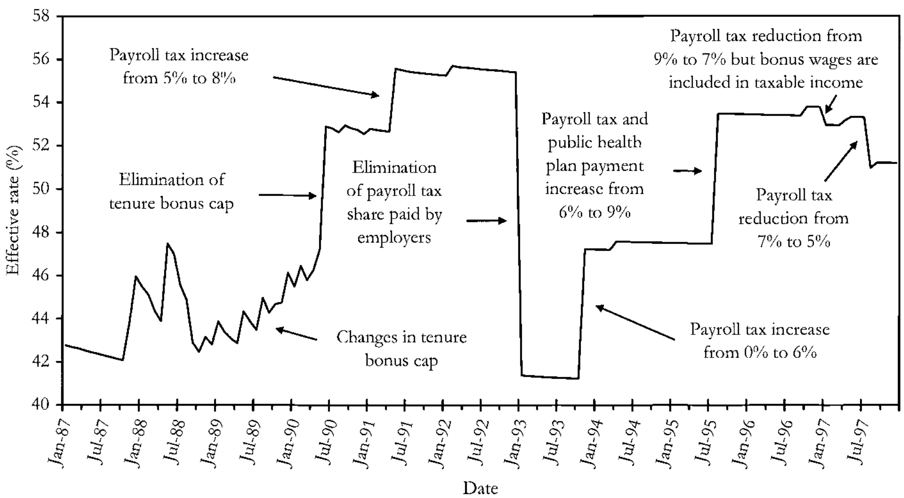


Fig. 2.3 Evolution of nonwage costs paid by employers

addition, on several occasions different rates were changed in such a way that the total employer contribution remained unaltered. This is the variable used later in the labor demand estimations.

2.3 Evidence of the Effect of Labor Laws on Labor Demand

We can identify at least three main changes in labor legislation during the period 1986–1996 that had an effect on labor demand: changes in severance payments and job stability, changes in nonwage costs, and changes in the use of temporary contracts. The difficulty lies in isolating these changes from the effect of the cycle over labor demand. In the Peruvian case in this particular period, even if it is very probable that the legislative changes had a large impact on the level and structure of demand, the economy underwent a very drastic process of structural adjustment (see Saavedra 1996a,b). The purpose of this section is to estimate labor demand functions and assess the effect of changes in two specific regulations in Peru: firing costs and nonwage costs. In 1991, absolute job stability was eliminated for new hires, and in 1995, after the constitutional change of 1993, job stability was totally abolished. Severance payments rules were simplified, and the severance profile was made less steep. This, together with the reduction in red tape for the use of temporary contracts, implied a drastic reduction in firing costs in two steps, one in 1991 and the second in 1995. On the other hand, nonwage costs were increased in 1987 and in 1990, first due to changes in caps and minimums in several contributions, and later through the increase in FONAVI, a plain payroll tax, and the pension contribution. We limit the analysis to labor demand for the formal sector, which is precisely the one affected by regulations. However, being formal (i.e., being in the universe of this study) is endogenous. One of the first consequences of high firing and nonwage costs in a low-productivity economy is informality, so we start the analysis by looking at how informal and formal salaried employment adjusted between 1986 and 1996.

2.3.1 Informality, the First Way to Avoid Regulations

Firms and workers adjust to the labor market regulatory framework through multiple mechanisms. Job protection legislation and severance payments constitute firing costs that increase uncertainty about the actual costs of labor and render labor a quasi-fixed factor. Given the regulatory framework that prevailed until 1991, Peruvian firms devised ways to reduce the costs of adjusting labor to their desired levels. The first adjustment mechanism was—and for many firms still is—the informal sector. Informality is conceptualized here as a state chosen by firms and workers depending on a cost-benefit analysis. Many firms, typically small ones, operate totally underground, fire and hire at will, and do not pay any kind of socially mandated benefits. In most of these cases, their productivity is too

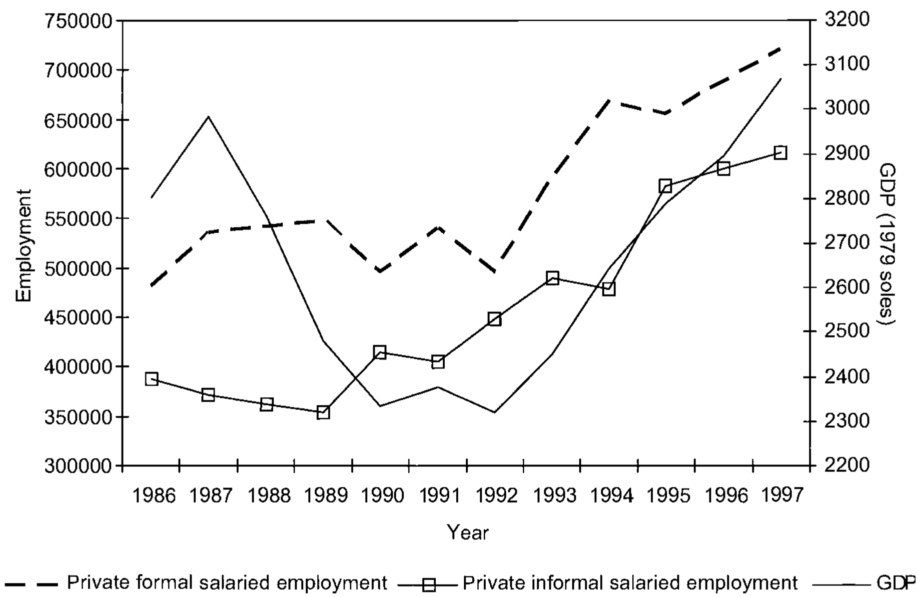


Fig. 2.4 Metropolitan Lima: Private formal and informal salaried employment and GDP, 1986–1997

Source: INEI, Encuesta de Hogares del MTPS 1986–1995, Encuesta Nacional de Hogares del INEI 1997.

low for them to afford to pay any kind of benefits. Both for the firm and for the worker, any kind of mandated benefit is a luxury. However, many firms operate in the gray area. In fact, there is a continuum of firms with different levels of productivity, and there is a cutoff point at which the firm decides that it has to operate formally. The decision to become formal entails a cost-benefit analysis. Firms evaluate the costs and benefits of formality (mandated benefits compliance and a larger volume of business, respectively) against the costs and benefits of informality (fines adjusted by the probability of being caught and savings in mandated benefits and firing costs, respectively).

Given changes in the regulatory framework, the balance in this cost-benefit analysis determines the evolution of formal and informal salaried employment. We used data from household surveys and defined formal salaried workers as those who show signs of working in a firm that complies with regulations.[10] As shown in figure 2.4, salaried informal employment increased since 1987 throughout the period of analysis. However, employ-

10. Operationally, formal salaried workers were defined as those who had health insurance, had a retirement plan, or belonged to a union. An application of this definition is found in Saavedra and Chong (1999).

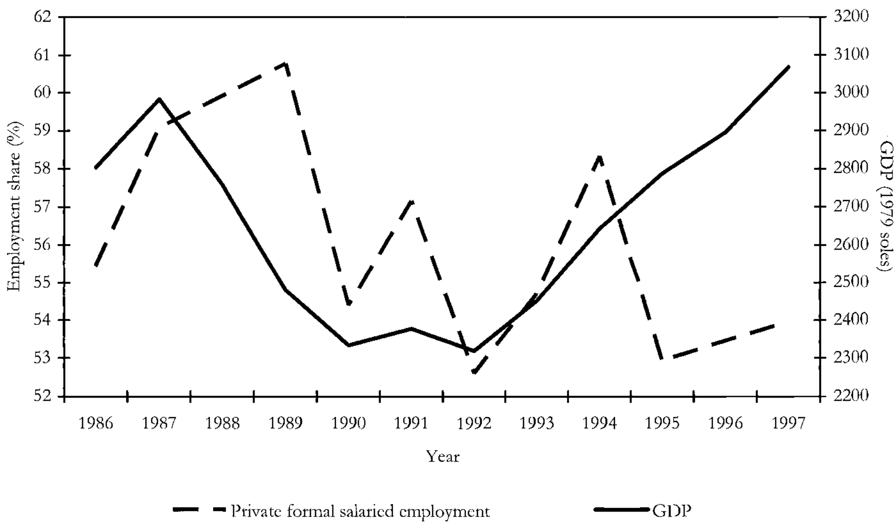


Fig. 2.5 Metropolitan Lima: Share of private formal salaried employment in total private salaried employment and GDP, 1986–1997 (percentage)

Source: INEI, Encuesta de Hogares del MTPS 1986–1995, Encuesta Nacional de Hogares del INEI 1997.

ment among formal salaried workers was more responsive to the business cycle. It fell slightly between 1987 and 1992 and then increased rapidly since 1993. It could be argued that the rigidities in labor legislation in the 1980s prevented formal employment from falling dramatically. Conversely, the more flexible environment of the 1990s allowed for a quick employment expansion. Looking at the shares of formal and informal salaried employment in total private salaried employment (figure 2.5), it is clear that the former fell sharply during the downturn and tended to increase timidly as output bounced back during the 1990s after the launching of the reforms.

2.3.2 Formal Labor Demand Estimations

Using household surveys, we only have annual data for ten years, so a formal analysis of the labor demand is not possible. Notwithstanding, it seems to be clear that, ceteris paribus, as the volume of business falls (as in 1988–1992) the costs of operating formally increase and outweigh the benefits, so more firms go underground, or more new firms decide to launch operations informally rather than formally. As of 1993, output rose again, and so did productivity; consequently, more firms should have found it profitable to operate formally. But to complicate matters, firms' decisions involve increasing or decreasing the share of their payroll that goes underground or not, and other developments also affect this decision. Reductions in firing costs could have had a positive effect on formal labor demand, but at the same time, nonwage labor costs increased, with the opposite effect over this demand.

In what follows, and with the purpose of analyzing formally the effects of these changes, using the quarterly data sets for the formal sector described next, we first perform static estimations of the labor demand at the sector level and at the firm level.[11] We show the results of different specifications, in which we analyze elasticity of wages, payroll contributions—taxes, health insurance, and pension and other contributions—and expected severance payments.

The Data

The main data source used to estimate static and dynamic labor demand functions for formal firms in Lima was the Quarterly Survey of Wages and Salaries (QSWS) conducted by the Ministry of Labor. The QSWS is a quarterly firm survey that collects pooled data on both the firm and individual worker levels. This survey collects approximately 600 private firms of ten or more workers in Metropolitan Lima (composed of the province of Lima and the constitutional province of Callao) and 8,000 workers from the same firms. The survey is divided into two sections. Part A provides firm-specific information that covers the gross wage bill divided into wage and nonwage costs, levels of employment, and presence of collective bargaining, each specified by category of employment (blue collar, white collar, and executive) and standardized international industrial code (SIIC). In Part B, five to twenty-five workers (according to the size of firm) from each firm are surveyed at random, thus providing individual-level information on age, gender, tenure, salary breakdown, and specific occupation, as well as employment category.

In 1986 the method of sample selection changed from a univariate distribution to one stratified across ten categories of economic sector and four categories of firm size.[12] This methodology ensures adequate representation of each cross section of firm sectors and sizes—totaling forty-eight groups of firms, among which a multivariate probability distribution is determined according to number of firms in each group, while minimizing total wage variance per group with standard optimal sampling methods.[13]

11. Using only formal firms—registered in the Ministry of Labor data sets—generates a selection bias for which we do not control.

12. The survey has been conducted since 1957, although at several points it has undergone important modifications. Due to the significance of modifications, data prior to 1986 are inappropriate for analytical comparisons with those of later periods. Furthermore, only hard copy tabulations of data from this period have been preserved.

13. Firms are divided into four size categories: 10–49 workers, 50–99, 100–499, and 500 or more. The economic sectors are agriculture, mining, manufacture of consumption goods, manufacture of intermediate and capital goods, utilities, construction, wholesale trade, retail trade, financial activities, insurance and real estate, transportation and communications, and services. Agricultural firms have been dropped from the sample.

Thus, the extent of survey information useful for analysis is restricted to the period 1986–1997, which comprises ten years of bimonthly data and quarterly data since 1996, representing a total of sixty-eight distinct points in time.[14]

During 1986 to 1997 there were three different samplings of firms, from the Ministry of Labor's "*Hoja de Resumen de Planillas*" (HRP) of 1986, 1990, and 1994. The HRPs are summary payroll forms that all private formal firms are legally required to present annually. The degree of compliance is high among large firms, and the probability of compliance increases with size. Total number of sampled firms per period remains around 500, but they were not replaced if the firm died or did not report during that period. Therefore, for the economic-sector estimations, we pool the data of all the firms in each sector and use expansion factors to calculate sector-level aggregates; we also use part B of the survey to calculate tenure structures by sector, which we then use for constructing the expected severance payment variable. With these, we build a pseudo panel at the sector level with fifty-six time points per sector. In addition to this firm database, we constructed time series of gross domestic product (GDP), which varied yearly by economic sector. To make the sector pseudo panel comparable to the firm-level panel described, we divide it into three pseudo subpanels according to the sampling dates, 1987–1990, 1991–1994, and 1995–1997. Although they roughly coincide with three distinct periods in terms of labor legislation (recall that the two main laws were enacted in November 1991 and July 1995) there is variability within periods, particularly regarding payroll contributions.

Figure 2.6 shows the evolution of employment of formal firms in Lima throughout the period. The gray bars show the periods in which the sample changed. Using the same data set, we constructed a sample of workers for each sector in each period. From that sample, we analyzed some basic worker characteristics. The results confirm the trends observed from household survey data. In particular, it is found that in the 1990s the proportion of younger workers increases, there is a slight increase in the share of female employment, and average tenure falls.

Finally, using this 1986–1997 QSWS survey data, we constructed three firm-level panels comprising all firms that remain in the sample set throughout the subperiods. The panels were constructed according to the

14. Data from all surveys prior to 1991 were stored only on eight-inch diskettes formatted with the antiquated XENIX system, which required the use of a Radio Shack TRS-16B computer and an eight-inch hard drive. None of those machines in Peru were in operating condition. The data were translated into a readable format by a software company based in Indianapolis, and the information was processed in order to recover the shape of the original databases. Only a few internal documents from the Ministry of Labor prior to 1990 describing the data existed. Fortunately, the survey did not undergo any methodological changes during that period, according to several current and former employers of the Direccion Nacional de Empleo y Formación Profesional (DNEFP) that were interviewed.

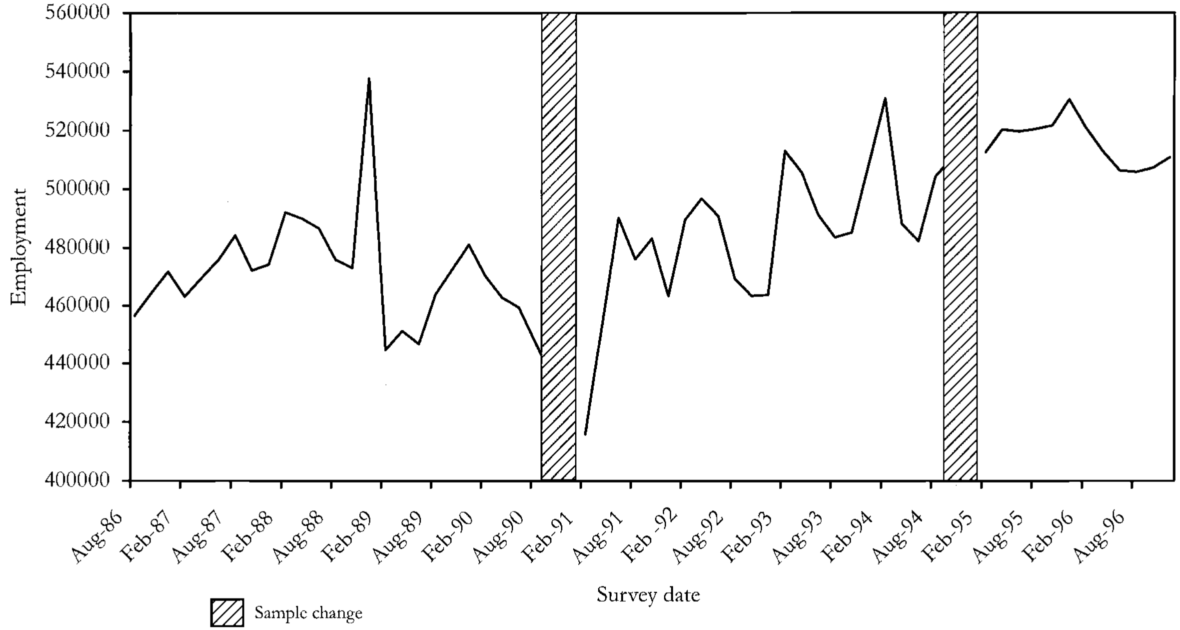


Fig. 2.6 Metropolitan Lima: Total employment in formal firms of more than ten workers
Source: Encuesta de Sueldos y Salarios del MTPS 1986–1996.

sampling periods and identify firms that remained in the survey throughout each subperiod. The first panel is 1987–1990, and all firms were drawn from the 1985 summary payroll forms of formal firms registered with the Ministry of Labor. The panel comprises 389 firms observed during twenty quarters. The second panel goes from 1991 through 1994, comprising 408 firms observed during twenty-four periods, drawn from the 1989 payroll forms. These two panels are the largest due to the fact that firms were never resampled from the total population of registered firms during this period. In other words, the list of panel observations is altered only by the death of firms that were originally sampled, and thus its size is determined solely by the mortality rate of those firms. In contrast, between 1995 and 1997, surveyed firms were resampled yearly from an updated sample set. Despite this greater variation in sampled observations, our third panel is only slightly smaller than its earlier counterpart (341 firms), largely because both the population (from payroll forms) and sample populations of surveyed firms have been considerably enlarged in recent years.[15]

Econometric Labor Demand Specifications

The objective here is to specify a static labor demand function from which the impact of different regulations may be inferred. We are mainly interested in analyzing the effect of payroll contributions—taxes on wages and social security payments—and severance payments on labor demand. With this objective, we will specify a static labor demand function following Hamermesh (1993).

The equations to be estimated will be derived first from a profit-maximizing framework. Consider the following firm's profit function:

$$\pi = F(K, L) - E(w)L - rK, \tag{1}$$

where K is capital, L is labor, and w and r are the cost of labor and capital, respectively. $E(w)$, which is the expected cost of labor, is used in order to account for the expected costs the firm would incur in the event of layoffs. This is important because w in the firm's maximization problem is not fully represented by the observed salaries, making it necessary to add other factors to appropriately represent the relevant cost per worker (following the distinction made in Hamermesh 1993).

The problem of the firm is to choose (K, L) such that it maximizes profit.

$$\max[F(K, L) - E(w)L - rK], \tag{2}$$

15. We attempted the construction of a panel of all firms that appeared continuously in the survey data between 1990 and 1997. This panel spanning both subperiods is by far the smallest, and, given its obvious biases, we will not include it in our estimations. On account of the fact that in 1995 a new sample of firms was selected (largely at random) from an updated payroll census for the first time since 1991, very few firms from the 1991–1994 period are resampled in 1995 and reappear continuously in the 1995–1997 sample populations.

where

$$E(w) = w + p + E(\text{sp}), \tag{3}$$

where w is the wage paid to the employee, p is all payroll contributions paid by the firm, and E(sp) is a measure of the expected severance payments as described in section 2.2.

A wide variety of functional forms have been developed in the past decade, although the derived factor demand functions are still analyzed under the same optimizing behavior (Merrilees 1982). The question remains as to which flexible production function will best suit our hypothesis testing. Here we use one of the approaches proposed in Hamermesh (1986) and estimate a simple and flexible functional form without any imposition of the restrictions that factor demand must be homogeneous of degree zero in all factor prices:

$$\ln L_i = a + \sum b_j \ln E(w_j)_i + c \ln Y_i + \beta \cdot \mathbf{Z}_i, \tag{4}$$

where j indicates the factor, i indicates the sector or the firm, w_j corresponds to two production factor prices, w and r, and $\mathbf{Z}_i$ is a vector of other explanatory variables at the sector or firm level. As mentioned by Hamermesh (1993), equation (4) should be viewed as part of a complete system of factor-demand equations, but given that we do not have data on all factors it is not possible to estimate a complete system.

Our initial objective is to see the effect of changes in labor regulations over labor demand in the formal sector of the economy. We analyze how labor cost elasticity changes as we add payroll contributions and the expected severance payment in a marginal productivity condition. We do not attempt to estimate labor supply relationships under the plausible assumption that the labor supply to the formal sector, in an economy with a very large informal sector, tends to be nearly horizontal. We estimate two variants of equation (4) that measure the effects of the different components of labor costs over employment:

$$\ln L_i = a + b_1 \ln w_i + b_2 \ln E(\text{sp})_i + c \ln Y_i + \beta \cdot \mathbf{Z}_i \tag{5}$$

$$\ln L_i = a + b_1 \ln(w_i + p_i) + b_2 \ln E(\text{sp})_i + c \ln Y_i + \beta \cdot \mathbf{Z}_i \tag{5'}$$

In equation (5) we include the average wage of the sector or firm and the expected severance payment as the two main labor costs. In equation (5′) we add to the average wage the average nonwage costs (public and private pension contributions, health contributions, accident insurance, etc.; see section 2.2.3) mandated by law that the employer had to pay in addition to the wage. These contributions are added to the salary because they are monthly charges paid by the employer, in contrast to the expected severance payment, which depends on the tenure structure of the employees.

Additionally, we estimate labor demand functions with sector-aggregated data and firm-level data with our three panels of the Peruvian firms (1987–1990, 1991–1994, and 1995–1997).[16] Following a modified version of Bentolila and Saint-Paul (1992), the econometric specification of labor demand is

$$(6)\quad \ln L_{i,t} = a + b_1 \ln[w_{i,t} + p_{i,t}] + b_2 \ln E(\mathrm{sp})_{i,t} + c \ln \hat{Y}_{t-L} + d \ln \hat{L}_{i,t-L} + e \ln \hat{L}_{i,t-L} \cdot \ln E(\mathrm{sp})_{i,t} + \delta t + \beta \mathbf{Z}_{i,t} + \varepsilon_{i,t},$$

where wages (w) and payroll taxes (p) represent the labor costs, $E(\mathrm{sp})$ represents the expected severance payments, $\hat{Y}$ is the quarterly output by economic sector as a proxy of firm output—instrumentalized with the lag—$\hat{L}_{t-L}$ is the number of workers in the previous period instrumentalized with the rolling regressions technique and using one- to four-period lagged employment, and t is a time trend.

Lagged employment is also included to measure the speed of adjustment to changes in output. The coefficient of this variable can lie between zero and 1; a large value is associated with a slower speed of adjustment, and a small value implies that the adjustment is instantaneous.

Finally, following Burgess and Dolado (1989), we try to measure the adjustment costs of changes in employment by including the interaction between lagged employment and expected severance payment as the main firing costs. The coefficient of this interaction measures whether there are increasing marginal costs of changing employment, and therefore a positive coefficient is expected.

Empirical Results

Using quarterly data for ten economic sectors observed between 1987 and 1997, we estimated the constant output labor demand wage elasticity for equations (5) and (5′).[17] As can be observed in table 2.4, all the components of $E(w)$ from equation (3) are significant and have the expected negative sign when included individually. The estimate of –0.19 for the labor demand wage elasticity (in the model in which labor costs included wages plus payroll contributions [b]) lies within the typical range for static labor demands using sector data (Hamermesh 1986, 1993).[18]

16. As mentioned before, the periods roughly coincide with three different legislation regimes.
17. This estimation is only done for the sector pseudo panel and not for the firms panel because we cannot generate a panel for the whole time period (1987–1997) given the structure of the survey.
18. As a sensitivity test, we also carry out a constant elasticity of substitution (CES) estimation, which included a proxy of the price of capital. The results of the CES specification were an elasticity of –0.13 for the wage and payroll cost variable and a positive elasticity for the price of capital. The latter reflects the positive cross-price elasticity of demand due to substitutability of labor for capital in production. Finally, the coefficient for the expected severance payment was –0.221.

Table 2.4 **Constant Output Labor Demand Estimation: Sector-Level Estimation (1987–1997)**

	Model 6 with Fixed Effects	Model 6′ with Fixed Effects
Constant	13.528***	13.701***
	(0.572)	(0.620)
$\ln(w)$	–0.174*	
	(0.096)	
$\ln(w + p)$		–0.191*
		(0.098)
$\ln[E(\mathrm{sp})]$	–0.406***	–0.401**
	(0.060)	(0.060)
$\ln(Y)$	0.047**	0.047**
	(0.022)	(0.022)
Log likelihood	–183.22	–182.97
$\chi^2(9)$	1,083.01***	1,084.59***
No. of observations	504	504

Note: Standard errors in parentheses.
***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

Moreover, as hypothesized, the coefficient of the average wage paid by the employer (b_1) is smaller by two points than the coefficient of the average wage plus all the payroll costs paid by the firm (b'_1). Therefore, as we include payroll taxes, the employment response to changes in labor costs increases. Additionally, we carried out an encompassing test on the model fit to select which specification should be used. We used a nonnested procedure and a Cox test for nonnested hypothesis (Greene 1997), and we were able to choose equation (5′) where $\ln(w + p)$ is used as the correct set of regressors. The Cox test, in which the null hypothesis was that equation (5) contained the correct set of regressors, was rejected with a p-value of 0.000 (Cox statistic = 5.27). On the other hand, when the null was that equation (5′) contained the correct set of regressors, we could not reject it at any significance level (Cox statistic = 3.56).

On the other hand, the coefficient of the expected severance payment, which varies across sectors and along time, also has the expected negative sign and is significant at the 99 percent level. This gives us evidence that the reduction in firing costs has a positive effect on employment level. Regarding the output elasticity, the coefficient for the whole period is around 0.05. This is a very small coefficient because in the models presented in table 2.4 we are including fixed effects by sector absorbing most of its variation—which is mainly across sectors rather than within. Specifically, when running the regressions without fixed effects the output elasticity is 0.17 and significant at the 99 percent level. We included the log of Y_i lagged six months, because the correlation between the errors and the actual output level that results from measurement error also biases ordinary least squares (OLS) output elasticity toward zero, and output measurement error can also bias the estimates of own-price elasticities. Griliches and Hausman (1986) demonstrate that when panel data are available, lead or lag of a variable subject to measurement error may be an appropriate instrumental variable. On the other hand, we are assuming that firms will not adjust immediately to changes in labor regulation, especially given the preexisting rigidities explained earlier.

Table 2.5 reports the results of equation (6), at both the sector and firm level, for the three subperiods determined by changes in the sample of firms: 1987–1990, 1991–1994, and 1995–1996. The first three columns are the results for the sectoral-level panels, and the last three columns show the results for the three firm-level panels. The variables used are the ones included in equation (5′) plus the instrumentalized lagged employment[19] as a measure of adjustment costs, its interaction with the expected severance payment, and a time trend. For the estimations we apply generalized least squares and correct for serial correlation with a correlation coefficient specific for each panel when needed. For the sector panel we include and test for sector fixed effects.[20]

In four out of six cases wage elasticities are negative and significant. Unfortunately, there are two exceptions: first at the sector level, for the first period in which the coefficient is positive and significant, and finally in the second subperiod on the firm-level data. It should be noted that variations in the measured price of labor may be the spurious result of shifts in the distribution of employment among subaggregates with different labor costs, as mentioned by Hamermesh (1986). It is difficult, however, to determine the extent of these potential problems. Regarding the expected severance payment, we found that in the first subperiod this variable had a negative and significant coefficient, –0.89 at the sector level and –0.31 at the firm level. In the last subperiod, the coefficient reduces to –0.31 at the sector level and to –0.14 at the firm level, losing its significance in both cases.[21] This result may be related to the fact that after 1995 there was not enough time variability in firing costs within the subperiod to establish an effect over the employment level, or that the variance of within-firm tenure structures had already fallen, reducing differences in expected severance payments across firms. In the firm panel data set, the interaction of the

19. This variable is instrumentalized using the rolling regressions technique with one- to four-period lagged employment.

20. We did not include fixed effects for the firm-level estimations because both the expected severance payment and the GDP were available only at the sectoral, and not at the firm, level.

21. We were not able to get evidence of statistically significant differences between these and other parameters when comparing different subperiods, using Wald tests. The limitation of these tests is that we assume that they are independent random samples, which is not true, given that large firms are always included in the samples.

Table 2.5 **Labor Demand Estimation Results for Panels at the Sector and Firm Level**

	Sector Level			Firm Level		
	1987–1990	1991–1994	1995–1997[a]	1987–1990[a]	1991–1994	1995–1997
Constant	8.262***	15.395***	13.657***	0.470***	0.032	1.678***
	(1.570)	(2.217)	(3.688)	(0.166)	(0.085)	(0.507)
$\ln(w + p)$	0.560***	–0.322***	–0.298**	–0.030***	0.028***	–0.053*
	(0.203)	(0.115)	(0.127)	(0.008)	(0.005)	(0.028)
$\ln[E(\text{sp})]$	–0.892**	–0.575	–0.315	–0.310***	–0.041**	–0.140
	(0.363)	(0.422)	(0.632)	(0.034)	(0.017)	(0.101)
$\ln(Y)$[b]	0.014	0.113	0.094*	0.249***	0.008	0.085**
	(0.067)	(0.101)	(0.053)	(0.008)	(0.007)	(0.033)
$\ln(L_{t-1})$[b]	0.070	–0.194	0.077	0.736***	0.942***	0.616***
	(0.147)	(0.215)	(0.310)	(0.027)	(0.016)	(0.088)
$\ln(L_{t-1}) \cdot \ln[E(\text{sp})]$	0.042	0.063	0.015	0.071***	0.006	0.040*
	(0.027)	(0.045)	(0.060)	(0.006)	(0.004)	(0.021)
Log likelihood	210.55	139.45	199.31	2,460.04	2,484.48	–1,389.821
χ^2	12,547.95***	4,537.29***	3,353.03***	230,609.34***	186,386.48***	2,728.07***
No. of observations	189	189	117.000	4,753	6,491	1,722

Notes: Standard errors in parentheses. Sector level includes fixed effects. For firm level, a time trend was included. It was significant for periods 1987–1990 and 1991–1994, but not for 1995–1997.

[a]Corrected for serial correlation when tests for autocorrelation were significant with a correlation coefficient specific for each of the panels because of the presence of lagged dependent variables.

[b]Instrumentalized with lagged values using rolling equations.

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

expected severance payment with the lag of employment, a measure of the marginal cost of changing employment, has a small but significant and positive coefficient that decreases over time.

In the sector-level estimations, the output elasticity increases from the first to the last subperiod, as shown in table 2.5. During the first subperiod it is 0.014 and not significant, while in the last subperiod it is 0.09 and significant at the 90 percent level.[22] This increase in output elasticity may be related to the fact that labor reforms made it easier for firms to adjust to the desired employment levels given changes in output. Given the lower level of the tax on dismissals generated by the reduction in severance payments and the abolition of job stability rights, and also given the lower administrative costs of using temporary contracts, formal firms enjoyed more flexibility in adapting to output changes. As shown in section 2.2, available evidence suggests that most of the increase in formal employment during the period seems to have been concentrated in temporary contracts. Nevertheless, this fact might also introduce a bias in the estimates, as our data aggregate employment and wages for both permanent and temporary contracts, and the true estimate for each of them might be different. This problem is dragged to the firm-level panel estimations also.[23] Output coefficients in this case are only significant for the first subperiod. It should be noted, however, that the output variable is defined at the sectoral level, so the coefficient cannot be interpreted as firm-level employment elasticity.

The lagged employment was included to measure whether adjustment occurs instantaneously. As shown in table 2.5, the effect of this variable is only significant in the firm-level panels with coefficients ranging between 0.62 and 0.94. The magnitudes of these coefficients are within the range of the coefficients found by Abraham and Houseman (1994). Given that this is bimonthly data, a fall from 0.7 in the late 1980s to 0.6 in the mid-1990s would imply a reduction in the median adjustment—as, for example, from six to four quarters. The smaller coefficient in the last period could suggest an increase in the flexibility of the labor market that made it easier to reduce workforce levels during periods of slack demand as well as making employers more willing to hire during periods of rising demand. The speed of adjustment is, however, much lower than the one observed in the United States as reported by Abraham and Houseman (1994).[24]

22. It is important to mention that the coefficient is small because these models include sector fixed effects, but, despite this, in the last subperiod the coefficient is significant. If we exclude the fixed effects the coefficient is around 0.17.

23. Finally, appendix A tests for the implication that total labor demand should vary over the cycle due to employment composition changes (Bentolila and Saint-Paul 1992). When interacting the regressors with the cycle dummy to capture responses to the business cycles, the effects were not significant in practically all of our regressions, as shown in table 2A.1.

24. These authors report a speed of adjustment for the U.S. manufacturing sector of 0.383. On the other hand, the speeds of adjustment for West Germany, France, and Belgium were similar to our results: 0.837, 0.935, and 0.823, respectively.

2.4 Effects on Duration and Turnover of Changes in Labor Legislation

In this section, we analyze basic patterns of employment duration in Peru. We address the question of how long jobs last in Peru, if their duration is different in the formal and informal sectors and in different occupations, and if there are significant changes related to changes in labor legislation. Reductions in labor legislation–related firings costs—like the ones observed in Peru in the early 1990s through the reduction in severance payments and the abolition of job stability rights—typically accelerate the process of job creation and job destruction, therefore increasing turnover and reducing job duration, particularly in the formal sector. This is consistent with Lindbeck and Snower's (2002) insider-outsider theory, in which they maintain that labor turnover costs are important only in labor markets that are characterized by stringent job security legislation, such as Peru had. Moreover, the Peruvian reforms facilitated the use of temporary contracts. This had the effect of inducing firms to hire more during expansions and also to lay off more workers during downturns, which implies an increase in turnover. Using different data sets, we find a reduction in employment duration that cannot be explained only by cyclical movements of the economy. Using empirical hazards, we compare job duration and employment exit patterns of the self-employed with those of wage earners in the formal and informal sectors, and we also try to analyze the effects of certain regulations over duration patterns and their changes over time.

We first present trends in job duration using the series of ten annual household surveys from the Ministry of Labor. The main shortcoming of this source is that it only provides us with data on incomplete (elapsed) tenures. However, as long as we are precise about what we are measuring, we can exploit the fact that it allows us to analyze some time series and cross-sectional variations. Then we present empirical hazards and the results of exponential hazard models using data from the Living Standards Measurement Survey, which has the advantage of providing us with an (unfortunately) small sample of complete employment durations.

2.4.1 Analysis of Recent Trends Using Censored Data on Job Duration

We first analyze a repeated cross-sectional data set, the Annual Household Survey of the Ministry of Labor for all the years between 1986 and 1997, with the exception of 1988. This survey collects information regarding job characteristics and elapsed tenure in the case of the employed and time in unemployment for the unemployed. In the case of these surveys, the question is "How long have you been in your current job?" The data are recorded in years and months. The answer does not provide information on the length of a particular contract but only on a match between firm and employee. In the case of the self-employed, this question relates to the time performing the same occupation. All elapsed tenures refer to the main job.[25]

The data available from these surveys are reported as incomplete tenures. Following Lancaster (1990), we can assume that, given a probability density function (PDF) of complete tenures for a sample of the stock of employed workers, there is a related PDF for elapsed tenures. Moreover, it is possible to assume that for workers with some labor market history, the PDF of remaining duration is the same as for the elapsed duration. Therefore, the expected value of completed durations is double the expected value of incomplete (elapsed) durations. This will be true as long as the stationarity of the process is assured; that is, it may not be true for young workers starting their careers, women who enter and reenter the labor market, or older workers approaching retirement (Burgess and Rees 1996). Clearly, these data allow the analysis of the distribution of tenures among those employed at the time of the survey, but not the distributions of jobs.

Figures 2.7, 2.8, and 2.9 show mean elapsed tenures for several categories of prime-age workers (twenty-five to fifty-five years old). In general, it is clear that there is a downward trend in mean tenure. The trend is clear enough to dominate any possible cyclical fluctuations in tenure. During the sharp recession of 1988–1992, when an increase in mean tenure could be expected due to high separation rates and low hiring rates, mean tenure actually fell. Tenure rose only in 1991, when the Peruvian economy hit bottom.[26] In 1992–1993, right after the first changes in labor legislation, there was a sharp reduction in mean tenure. During the period 1994–1997 growth was fast, and hiring and separation rates increased, as usually happens in a booming economy, resulting in a further reduction in mean tenure. However, the 1997 figure was much lower than in 1986–1987, when the economy was also on an upswing. This gives an indication that the reduction in tenures may not be only a cyclical fluctuation but that it might be showing a secular trend.

The downward trend is clearer among prime-age males (figure 2.7). Given that the mean value of complete tenures should be about double the elapsed ones, in the mid-1990s mean completed tenure was about twelve years,[27] down from seventeen years in the mid-1980s. There is also a reduction in mean tenure among females (not shown), but it is harder to assume a stationary process in this case. First, because of maternity women enter and reenter the labor market, and second, during this period there is a rapid increase in labor force participation among women (Saavedra 1998).

25. In all surveys and years, the proportion of workers who declare having a second job fluctuates between 12 percent and 15 percent.

26. Tabulations not reported show that there is no clear trend in mean tenure among young workers.

27. Considering that the average schooling for males in Lima in this cohort is 8.5, and assuming retirement at 65, on average, each individual holds three jobs during his lifetime.

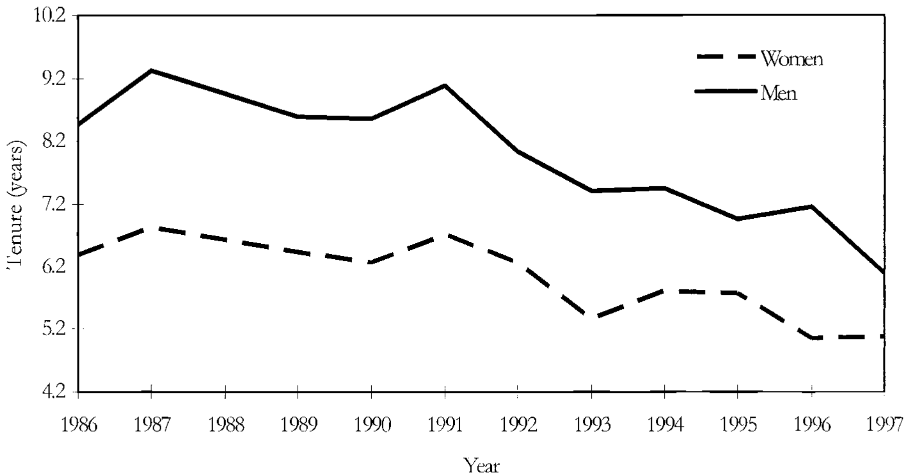


Fig. 2.7 Metropolitan Lima: Incomplete (elapsed) tenure of male and female workers aged twenty-five to fifty-five years, 1986–1997

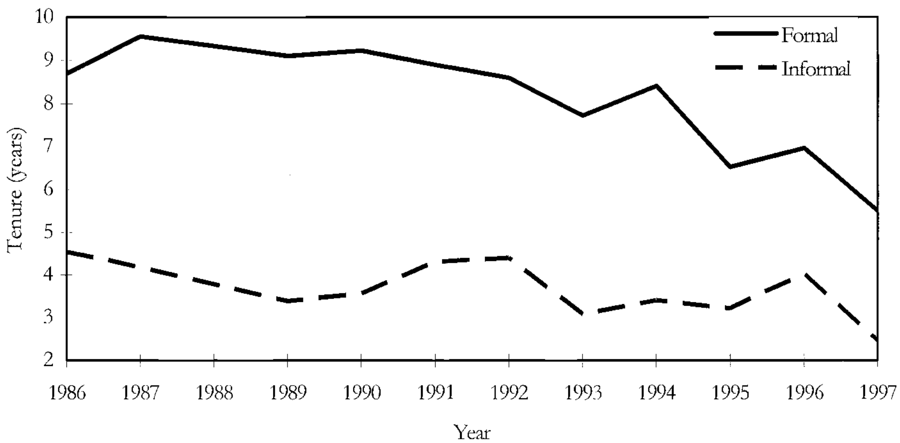


Fig. 2.8 Metropolitan Lima: Incomplete (elapsed) tenure of formal and informal male salaried private workers aged twenty-five to fifty-five years, 1986–1997

The differences in means between 1986 and 1991 and between 1991 and 1997 are statistically significant.

Figure 2.8 shows the evolution of mean elapsed tenures for prime-age male wage earners according to their formal or informal status. To define this status we use a legalistic definition: A worker works formally if he or she has health insurance or a pension plan or belongs to a union. The same definition is used in all the surveys. With this definition, the rate of formal

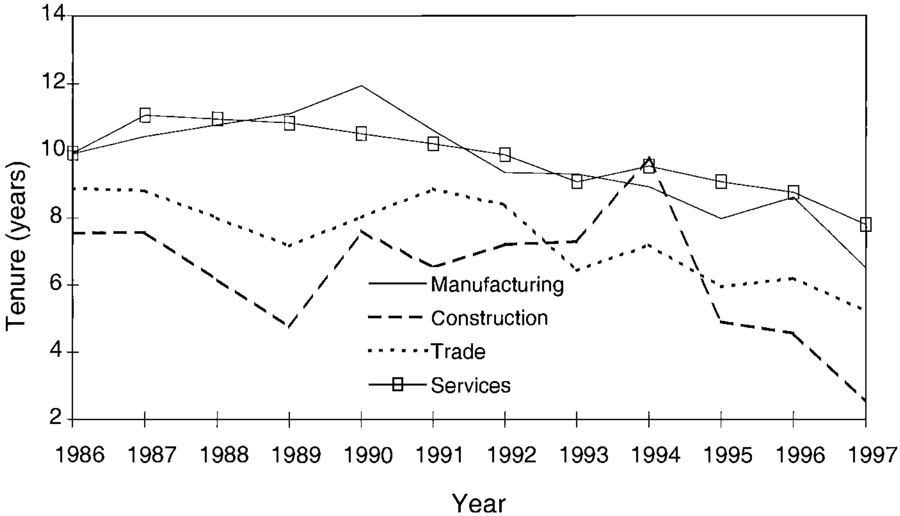


Fig. 2.9 Metropolitan Lima: Incomplete (elapsed) tenure of male formal workers in selected economic sectors aged twenty-five to fifty-five years, 1986–1997

employment fluctuated from 53 percent to 60 percent between 1986 and 1997. Several features are worth mentioning. Differences in mean elapsed tenures are large between formal and informal salaried workers. In fact, for formal salaried workers, mean tenure is between 9 and 6.8 years, while for informal workers, the mean fluctuates around 3. This difference is statistically significant in every year during the period 1986–1997, as shown in table 2.6.

The downward trend is more pronounced among formal workers,[28] in particular after 1991. Table 2.7 (panel A) shows tenure mean comparisons within formal and informal workers pairing different years. Within informal workers there is a significant reduction in mean tenure in the period 1986–1993 and a smaller and less significant reduction in the period 1993–1997. In the case of formal workers the fall is much larger and statistically significant in the postreforms period. From the results shown in panel B of table 2.7, it is clear that the differences in mean tenures between the formal and informal sectors have fallen during the 1990s. As discussed earlier, labor market reforms facilitated formal firms' adjustment to desired employment levels through temporary contracts and by reducing severance payments and eliminating job stability. In addition, unionization rates fell sharply, and union jobs have traditionally been held much longer than nonunion ones.

Figure 2.9 displays elapsed mean tenure calculations for prime-age for-

28. Tabulations not reported for self-employed workers show a downward trend among formal self-employed workers, but not among informal self-employed.

Table 2.6 **Tenure Mean Comparison Test**

Year	*t*-Test
1986	–7.377
1987	–8.400
1989	–10.678
1990	–10.291
1991	–7.715
1992	–7.676
1993	–9.492
1994	–9.416
1995	–7.444
1997	–6.285

Notes: H: (Informal worker tenure in period t – Formal worker tenure in period t) = 0. In all years the p-value was 0.000.

Table 2.7 **Mean Tenure Differences and Differences-in-Difference**

	1986–1993	1993–1997
Difference estimates		
Formal	–0.98	–2.23
	(0.55)	(0.57)
Informal	–1.45	–0.63
	(0.51)	(0.39)
Difference-in-difference estimates		
Formal – Informal	0.48	–1.60
	(0.75)	(0.67)

Notes: Differences of mean elapsed tenure for currently employed wage earners in Metropolitan Lima. Standard errors in parentheses.

mal male salaried workers in selected sectors. In the manufacturing sector, there is a smooth upward trend between 1988 and 1990, as the economy fell into a recession. Afterward, mean tenure falls as the economy picks up. We observe the same trend in services and, to a lesser extent, in trade. We also performed calculations controlling for age structure, and results were similar, which tends to suggest that these changes are not reflecting changes in the type of workers being fired but are an illustration of higher overall turnover.

Several factors may lie behind the reduction in tenure among formal prime-age workers. Before the reforms, high firing costs induced long employment spells among formal workers, but they also induced a lower rate of job creation in the formal sector, which also increased the relative size of the informal sector. The labor market reforms of 1991 facilitated an increase in hiring through temporary contracts and also reduced firing costs through a reduction in the severance payment and the elimination of job stability for new workers. The reforms were followed by an economic expansion that began in 1993 and increased employment, both formal and informal. The increase in net employment suggests that hirings were larger than layoffs. Layoffs in the private sector—also driven by trade liberalization and privatizations—were larger among older workers.[29] On one hand, the relative cost of firing a high-tenured worker fell tremendously with the reforms, in particular with the 1995 changes, when job stability was abolished for all workers. On the other hand, the increase in the demand for labor was larger for younger workers, who could more easily adapt to new technologies. Therefore, layoffs were biased toward older workers, while hirings were biased toward younger ones, with the effect of reducing mean tenures.

Table 2.8 shows mean job durations using elapsed-tenure data from several sources. The first two columns are from the same data sets discussed in the previous paragraphs, the third comes from the firm-level survey used in the labor demand analysis, and the rest are from the Living Standards Measurement Surveys (LSMS) described in the following section. All data sources confirm a reduction in mean tenure for formal workers during the 1990s.

2.4.2 The Duration of Employment Spells

The data used in this part of the analysis come from the LSMS.[30] The employment modules of the LSMS contain information about job characteristics like tenure in the current job, sector of activity, size of firm, whether contract was signed, union membership, type of employment (public/private/self-employed/wage earner), white- or blue-collar job, and so forth.[31] This information is collected regarding the job held in the previous seven days. In addition, individuals who are not working report whether they are looking for a job and number of weeks unemployed. The survey has another module that asks workers—either employed or unemployed—questions regarding their last job in the previous twelve months. If the worker has been unemployed during the last seven days, the survey asks for all the characteristics of the last job held during the previous year. If he or she has been working during the last seven days, the survey inquires if this job is the one held during the last seven days. If the job is different, the survey asks for the characteristics of that job. Two types of job spells are calculated with each survey. We use each survey separately and calculate right-censored spells for the sampled stock of employed workers and complete spells for the unemployed and for those who changed jobs during the last year. The detail of the duration data is as follows:

29. Saavedra (1998) shows that among workers older than 55, the employment-population ratio has not recovered with the employment growth observed in the 1990s and that unemployment has risen for this group of workers.

30. The LSMSs are a series of household surveys developed since 1985 under the technical and financial support of the World Bank and later implemented by Instituto Cuanto.

31. Sample sizes allow for the analysis of all these categories separately. As opposed to what is observed in developed economies, in Peru, as in other Latin American countries, self-employment rates reach 40 percent in urban areas.

Table 2.8 **Mean Job Tenure: Comparing Different Data Sources**

	Household Survey			LSMS								
				Self-Employed			Wage Earner			All		
	Formal Workers	All	Firm-Level Survey	Informal	Formal	All	Informal	Formal	All	Informal	Formal	All
1985				8.27	8.58	8.29	3.92	7.53	6.66	7.33	7.58	7.43
1986	8.87	6.87										
1987	8.97	7.28										
1989	9.41	7.00										
1990	9.70	6.97										
1991	9.45	7.45	10.08			7.64			7.07			7.34
1992	8.98	6.83	10.26									
1993	7.62	5.99	10.46									
1994	8.23	6.39	10.34	7.20	8.70	7.30	4.26	7.08	6.30	6.55	7.21	6.81
1995	7.48	5.85	7.44									
1996	6.11	5.74	6.93									
1997	6.63	5.11		7.14	5.89	7.14	3.6	5.89	5.22	6.3	5.89	6.15

Source: Encuesta de Hogares del MTPS 1986–1995, Encuesta Nacional de Hogares del INEI 1996–1997, Encuesta de Sueldos y Salarios del MTPS 1986–1996, Encuesta Nacional de Hogares sobre Nueles de Vida 1985, 1991, 1994 and 1997.

Notes: Household survey data for Metropolitan Lima, currently employed workers. Firm-level survey data for Metropolitan Lima, currently employed workers in firms of ten or more workers. LSMS data for Urban Peru, currently employed workers. Blank cells indicate that information is not available.

- We use right-censored spells for the stock of people currently working, using the question "How long have you been working as [occupation]?" (coded in weeks, months and years).[32]
- For those who declare that they have indeed changed jobs during the last twelve months, we construct two spells, a right-censored spell of less than twelve months and a complete previous spell. These data have two obvious biases. First, we have complete spells only for those who changed jobs during the last twelve months; if the current spell lasts more than that, we have no information about the previous spell. For these movers, we do not have information on possible unemployment periods between the two jobs. Second, for some workers who report a change in job, the change is within a firm. In those cases, we will not count that as a job change. We will isolate those cases by comparing all the job characteristics of the previous and current spells (occupation, sector, size of firm, public or private, etc.).
- We use complete job spell for those who are not currently employed and who answer positively to the question "Have you had a different job during the last twelve months?"[33]

The complete and incomplete employment spells that are constructed in our data sets are summarized in figure 2.10. According to the employment duration data for the years 1985 and 1994 from the LSMS, 78 percent of the job durations of 1985 are incomplete spells, while for the 1994 sample this figure is 86 percent.

We analyze the basic differences in job duration patterns using the LSMS employment duration data for the years 1985 and 1994, including both complete and incomplete employment spells. These spells are to be thought of as independent realizations of a random variable T with survivor function $\bar{F}(t)$. Using the complete and incomplete employment spells from the LSMS data, we use the Kaplan Meier estimator for the survivor function. Following Lancaster (1985), for homogeneous right-censored data the survivor function at t can be estimated by

$$\hat{\bar{F}}(t) = \prod_{t(j)<t} (1 - \hat{\theta}_j), \qquad t \geq 0, \tag{7}$$

32. The question, as written in the questionnaire, does not look very precise. However, two elements allow us to recognize them as job spells. First, personnel in charge of the fieldwork and of the interviewer's training process maintain that they insisted that the duration reported as an answer to that question should be the length of time working in a specific occupation and in a specific firm. Second, the survey allows for a second check mechanism from a separate question: "What was your main occupation during the last twelve months? Was this the same as your occupation during the last seven days?" In this case, the interviewer manual indicates that even a change in position within a firm should be considered a job change. If the respondent answers that the job was different, then he or she will answer for the characteristics of that previous job.

33. Note that we only have spells for those people—current unemployed or out of the labor force—that had a job during the last twelve months. For those unemployed or inactive for more than that period, we do not have any information.

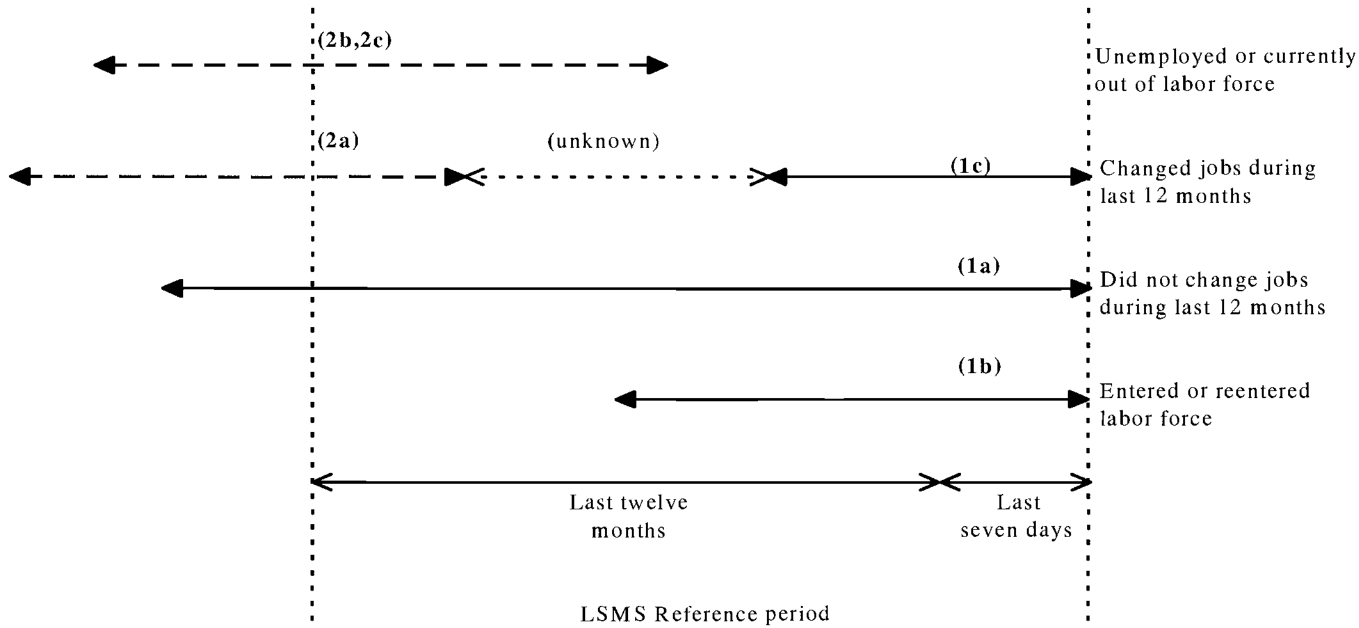


Fig. 2.10 Employment spells using the Peruvian LSMS: ***1*****, Right-censored job spells: for the currently employed, either (1a) people who didn't change jobs during the last year; (1b) newly employed entrants; (1c) people who changed jobs during the last year.** ***2*****, Complete job spells: (2a) currently employed that changed jobs during the last twelve months (this spell is the job held before the current one); (2b) unemployed workers whose unemployment spell is smaller than twelve months and held a job during the last twelve months; (2c) currently out of the labor force, that held a job during the last twelve months.**

for $\hat{\theta}_j = n_j/r_j$, where n_j is the number of employment spells—possible only one—observed to end at time t, and r_j is the risk set (spells that end at time t plus those censored at time t). θ_t is the probability of leaving the employment state (i.e., it is the hazard at time t). This estimator is a step function with steps at each observed (uncensored) exit time.

A shortcoming of this data set is that with the observational scheme of the survey complete spells are registered only for workers who are unemployed or out of the labor force and for workers who changed jobs during the year prior to the survey. Therefore, complete employment spell tenures are available only for a specific type of individual. However, as shown in appendix C, there is a similarity between the hazard function calculated using only the complete spells and the hazard function estimated using only the incomplete (censored) spells—as if they were completed despite the possible biases of the censored data.

In addition, the empirical analysis assumes a stationary economic environment. This assumption, which implies that the numbers of jobs created and destroyed are independent of time, allows to use each survey as a photograph of the distribution of their hazards assuming they will not be affected by the passage of time. It is difficult to assume stationarity in the Peruvian case, in particular, given the implementation of a set of structural reforms in the early 1990s. However, if we analyze each survey separately (1985 and 1994), despite the huge macro shocks observed in the Peruvian economy, no clear pattern of steady increase in the rate of job creation has been observed in the years previous to the surveys. In fact, a typical variable that could be used to condition the hazard function to the different environments confronted by different cohorts at their entry to or exit from employment is the rate of unemployment. That variable has fluctuated around a steady mean of 8.5 percent since 1974. Still, it is difficult to assure that a stationarity assumption can hold in volatile economies like Peru, in particular in the case of employment spells when we would need the same data generation process for a relatively long time.

Monthly hazards for a sample of censored and complete spells allow us to investigate duration patterns at the early stages of a job. In most of the cases there are spikes at months three, six, and twelve, which (at least in part) may be a heaping effect. In this sense, it will be important to compare changes through time and between categories. At the time of the fieldwork of the 1985 survey, the probationary period lasted three years, after which workers acquired total job stability. However, the authorities had already announced their intention of giving workers job stability rights after the third month.[34] In fact, the hazard function calculated with 1985 data for spells that started after 1983 and before June 1986 (left panel of figure 2.11) shows a spike at the third month. It is possible that employers in the formal

34. The change was actually put into effect in June 1986.

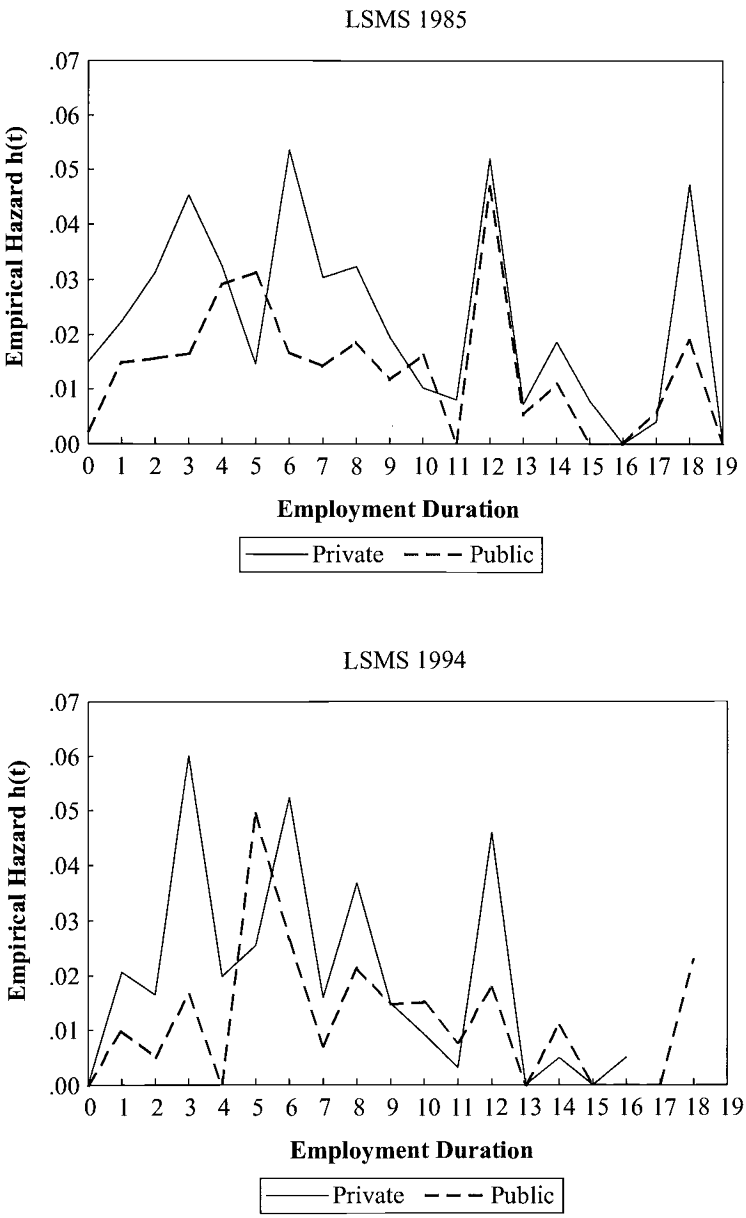


Fig. 2.11 Empirical hazards for formal public and private workers with less than three years of tenure (employment duration in months)

sector had already reacted to the announcement by dismissing workers right before they reached that tenure length. However, this spike is even larger among informal wage earners, who were not affected by regulations.

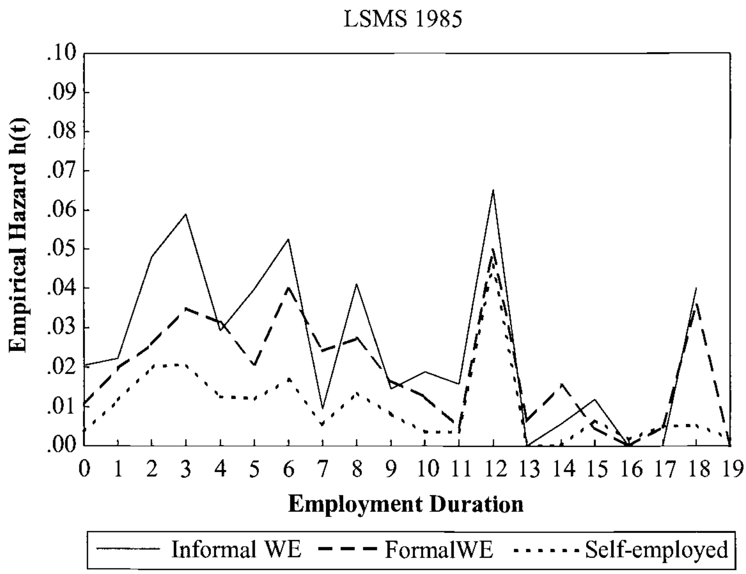
In 1994, labor legislation was more flexible, although only a few years had passed since the first wave of labor reforms in 1991. The probationary period was still three months, after which workers hired after 1991 obtained not job stability but only the right to a severance payment upon unjust dismissal. Therefore, firing costs were obviously lower than those perceived by firms in 1985. As shown in the right panel of figure 2.12, even if there is still a spike at the third month in 1994, the difference in the hazard functions between formal and informal workers is much smaller until the fourth month. Moreover, for tenures between five and eleven months the probability of leaving the state of employment is actually larger for formal workers than for informal ones. The hazard function for formal wage earners in 1994 is slightly above that for 1985. These higher hazards for formal workers in the postreform year may be related to the lower firing costs. They could also be related to an increased inflow of employment, but, as shown in section 2.2, inflows to informal employment were at least as great as those in the formal sector.

Note that in 1994 there still are large spikes in months 3 and 6. The spike in the third month may be explained by the fact that at that point workers acquired the right to a severance payment upon dismissal.[35] In addition, during this period employers still feared a possible reversal of the legislative changes and a return to a restrictive legislation, so many of them were still reluctant to hire workers under permanent contracts. They relied heavily on temporary contracts for short-term periods, usually three or six months, which in some cases were continually renewed.[36] There is a large spike at the twelfth month that may be related to the increase in the severance payment from zero to three months' wages after completing a year in the firm, so right before finishing that year firms had their last chance to dismiss the worker at zero cost. To summarize, there is an increase in the hazard function for formal wage earners between 1985 and 1994 and an increase in the hazard relative to that in the informal sector for workers with short durations.

An additional piece of evidence comes from the comparison between public and private formal wage earners. As shown in figure 2.11, there is a

35. The severance payment rule in 1994 stated that workers should get the equivalent of one month's salary per year worked if they had more than one year in the firm—with a minimum of three months' wages and a maximum of twelve months' wages. They acquired that right after the three-month probationary period, but the severance payments between the third and twelfth month were zero.

36. The spike in the third month observed in the informal sector may be a rounding effect or may also be some sort of "lighthouse effect."



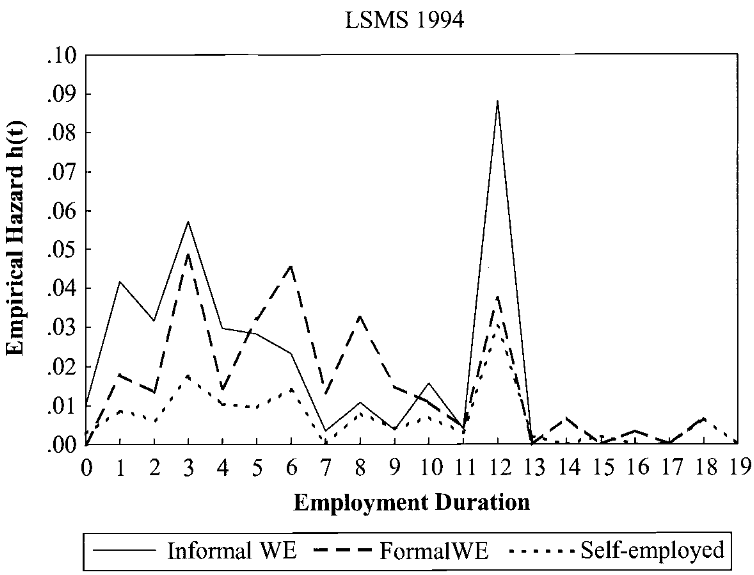


Fig. 2.12 Empirical hazards for wage earners and self-employed workers with less than three years of tenure (employment duration in months)

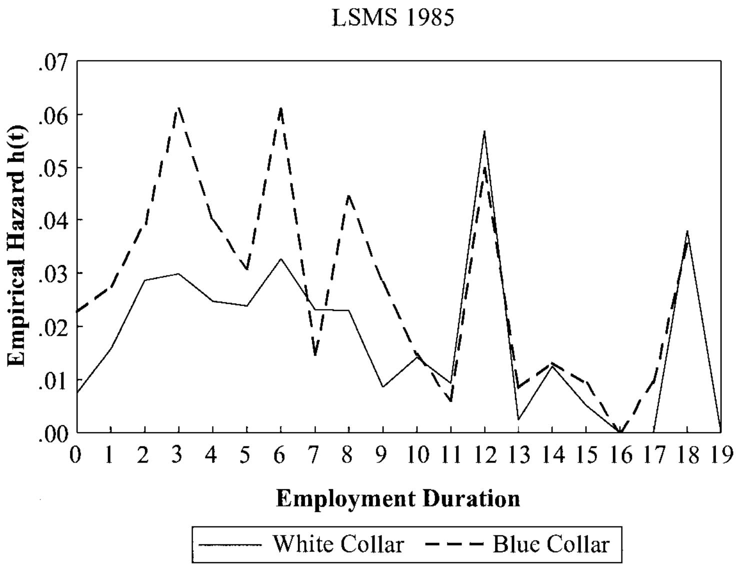
large spike in the third and sixth month for private formal workers, which is not observed for public workers. This could be consistent with firms' rehiring workers for two consecutive probationary periods. In general, the probability of exiting the employment state is much higher during the first months in the private sector, something that is not observed in the public sector. Spikes are also observed at one year of tenure, which is consistent with the increase in severance payment at that point in accordance with legislation. In 1994, however, the spike is smaller in the private sector and much lower for the public sector. This is probably related to the reduction in severance payment in the case of the private sector and to the public-sector downsizing that started in 1992.

An interesting change is observed when we compare hazards of blue-collar and white-collar workers. Clearly, during the first ten months of employment, hazards are higher for blue-collar workers, a result consistent with the common view that turnover is higher among those workers (see figure 2.13). In 1985, spikes at the third, sixth, and eighth months are very pronounced for blue-collar workers and are not observed among white-collars. However, after 1991, the spikes are observed in both groups, and, in general, differences in the hazard functions are much smaller.

Parametric Estimation of Hazard Functions

The sample employment spells just analyzed are not drawn from a homogeneous population. In order to adjust for the heterogeneity of observations and analyze patterns for different groups of workers, we estimate exponential hazard models using complete and incomplete spells. Table 2.9 shows the result of the estimation for three different years using employment spells of self-employed and salaried workers. Age shows the usual negative effect over the hazard, suggesting a lower turnover for older workers. The negative effect of age over the hazard is larger in 1991 and 1994, consistent with an increase in turnover among older workers. Education has a significant negative coefficient, suggesting lower hazards for the more educated, particularly after the reforms launched in 1991. Surprisingly, occupational training increases hazards in 1991. The results also confirm that the self-employed have lower hazards and much longer employment spells than formal wage earners, and that these in turn have longer spells than informal wage earners, the category of control. The negative coefficient for formal salaried workers is larger after the reforms, suggesting a relative increase in turnover for this group. However, the standard error is also larger, so the change may not be statistically significant.

Table 2.10 presents an extended model that limits the sample to wage earners. The 1985 estimates show that having a temporary contract increases the hazard, suggesting higher turnover among these workers. This effect disappears by 1994, although temporary contracts were intensively used, which may be related to a smaller difference in status within a firm



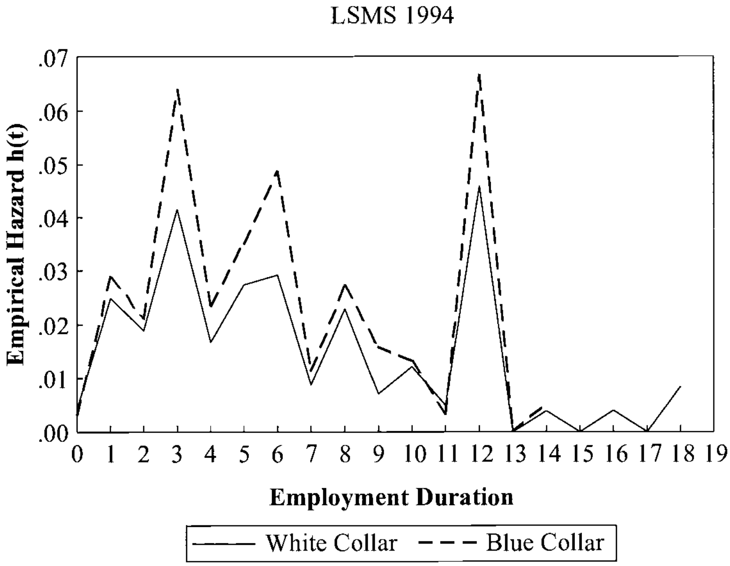


Fig. 2.13 Empirical hazards for wage earners (blue-collar and white-collar) with less than three years of tenure (employment duration in months)

Table 2.9 Exponential Hazard Model: Self-Employed and Wage Earners Sample

	1985	1991	1994
Male	–0.462***	–0.212**	–0.293***
	(0.064)	(0.103)	(0.092)
Age	–0.154***	–0.203***	–0.183***
	(0.012)	(0.019)	(0.017)
$Age^2 \cdot 10^{-2}$	0.111***	0.176***	0.143***
	(0.014)	(0.024)	(0.020)
Married	–0.348***	–0.351***	–0.048
	(0.074)	(0.124)	(0.107)
Years of schooling	–0.005	–0.054***	–0.023*
	(0.008)	(0.014)	(0.013)
Occupational training	0.073	0.480***	0.075
	(0.069)	(0.105)	(0.101)
Formal wage earner	–0.360***		–0.433***
	(0.094)		(0.138)
Self-employed	–0.979***		–0.976***
	(0.086)		(0.125)
Wage earner		0.704***	
		(0.114)	
No. of observations	6,144	3,570	4,561
Log likelihood	–4,461.59	–1,788.78	–2,656.25

Note: Standard errors in parentheses.
***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

between temporary and permanent positions.[37] Having social security coverage, a clear indication of formality, reduces the hazard rate, a result consistent with the higher empirical hazards found before for informal workers. Surprisingly, belonging to a private-sector union increases the hazard; however, as the influence of unions vanishes through time, the estimate for this variable is not significant during the 1990s. We also find that married workers tend to have longer employment spells, and hazards are larger for blue-collar workers, as was suggested in the empirical hazard analysis. Limiting the sample only to private workers does not modify the result significantly.

2.5 Concluding Remarks

Peru is one of the countries that made more progress in terms of labor market deregulation in Latin America as part of a package of structural re-

37. Saavedra and Maruyama (1999) show that before the reforms temporary workers tended to be younger, less experienced, and less educated than permanent ones. These differences diminished sharply after the reforms. Also, there was a significant reduction in the earning premia of permanent workers.

Table 2.10 Exponential Hazard Model: Wage Earners Sample

	1985	1991	1994
Male	–0.702***	–0.293**	–0.517***
	(0.097)	(0.139)	(0.134)
Age	–0.175***	–0.222***	–0.176***
	(0.019)	(0.030)	(0.032)
$Age^2 \cdot 10^{-2}$	0.146***	0.224***	0.123***
	(0.024)	(0.037)	(0.041)
Married	–0.355***	–0.463***	–0.010
	(0.096)	(0.164)	(0.139)
Years of schooling	0.050***	0.010	0.029
	(0.012)	(0.022)	(0.019)
Occupational training	–0.049	0.544***	–0.068
	(0.088)	(0.139)	(0.145)
Union	0.350**	0.128	–0.303
	(0.137)	(0.197)	(0.272)
Social security	–1.180***	–1.212***	–1.219***
	(0.117)	(0.171)	(0.170)
Temporary contract	0.182*		0.157
	(0.104)		(0.143)
Public worker	–0.362**	–0.188	–0.484**
	(0.157)	(0.274)	(0.221)
Blue collar worker	0.393***	0.269*	0.288**
	(0.103)	(0.156)	(0.146)
Union · public worker	0.019	–0.107	0.535
	(0.200)	(0.338)	(0.376)
No. of observations	3,344	1,945	2,330
Log likelihood	–2,557.92	–1,039.92	–1,481.19
χ^2(df)	1,171.71	517.49	592.49
Prob > χ^2	0.00	0.00	0.00

Note: Standard errors in parentheses.
***Significant at the 1 percent level.
**Significant at the 5 percent level.
*Significant at the 10 percent level.

forms that took place in the 1990s. One of the most important changes in labor legislation was the large reduction in firing costs, through the reduction in the steepness of the tenure-related severance payment profile since 1991, the progressive abolition of job stability, and the facilities given to the use of temporary contracts. To analyze the effect of changes in firing costs we constructed an expected severance payment indicator as a proxy of the monetary resources firms have to reserve in order to cover firing costs. We broke down the data into state-contingent components of firing and hiring states of the economy. Within each state, the severance payment was calculated by sector using the evolution of the tenure structure of workers, an estimate of the firing probability for each tenure group, and the corresponding mandated severance payment structure. These probabilities were allowed to vary only across sectors and were kept constant through time in order to reduce endogeneity. A series of nonwage costs was calculated by simulating the total labor costs paid by the firm as a proportion of the wage for different wage levels. This was necessary because several mandated benefits and the payroll tax had absolute lower and upper bounds that were continuously changed. In many cases, most of the changes in the effective rate paid were due to changes in these limits.

To analyze the effects of changes in labor costs and firing over labor demand, we used a pseudo-panel data set of ten economic sectors observed bimonthly during the period 1987–1997 and three shorter panels of about 400 firms for the periods 1987–1990, 1991–1994, and 1995–1997. There are four main empirical findings. The wage plus payroll elasticity is –0.19 for the whole period of study when using the sectoral-level panel. This price elasticity is larger when the payroll taxes are added as part of the labor costs than in an estimation in which only wages are included, and we were able to test that the latter was the model that should be used. In most of the subperiods, at both the sector and firm levels, labor costs have a negative and significant effect over labor demand. Labor demand elasticities may not be stable as the economy opens up, as happened in Peru with the trade liberalization process that started in 1991. However, Saavedra and Torero (2001) do not find significant changes in elasticities when interacted with proxies for changes in the trade regime.

The second main finding is that the coefficient of our measure of firing costs, the expected severance payment, is negative and significant, showing that job security provisions have a negative effect on employment. We also found that its magnitude decreases after 1995. This result may be related to the fact that after that year there was not enough time variability in firing costs within the subperiod to establish an effect over the employment level, or to the fact that the variance of within-firm tenure structures had already fallen, reducing differences in expected severance payments across firms.

Third, the output elasticity increases in the last subperiod. This may be related to the fact that labor legislation reforms made it easier for firms to adjust to the desired employment levels given changes in output. The reduction in severance payments and the abolition of job stability rights may be interpreted as a lower level of the tax on dismissals. In addition, the lower administrative costs of using temporary contracts made it easier for formal firms to adapt to output changes. Finally, and in line with the previous result, we also find a speedier employment adjustment during the postreform period.

As discussed previously, labor market reforms facilitated formal firms' adjustment to desired employment levels, through temporary contracts and by reducing severance payments and eliminating job stability. This reduction in firing costs may have the effect of increasing turnover, as firms will tend to increase hirings during expansions and firings during contractions. Using censored employment spells from different data sets that span the period 1985–1997, we find evidence that mean tenure fell since 1992, roughly coinciding with the beginning of labor market legislation changes, suggesting an increase in turnover in the Peruvian labor market. The reduction in mean tenure may also be related to the recovery initiated in 1993, when salaried employment was created, both in the formal and informal sector. However, even if mean tenure among informal workers fell, among formal workers the fall is much larger and statistically significant in the post–labor reform period. This is showing, therefore, as mentioned by Lindbeck and Snower (2002), that the smaller a firm's labor turnover costs, the more profitable it is for the firm to stop bargaining with its current employees (insiders) and start bargaining with the new potential hires (outsiders) instead. The differences in mean tenures between the formal and informal sectors also fell significantly during the 1990s.

The LSMSs for 1985 and 1994 allowed us to construct censored employment spells for currently employed workers and complete employment spells for the unemployed and for workers that changed a job during the twelve-month period before each survey. With this data we calculated empirical hazards for several groups of workers. We found spikes at three months of tenure, corresponding to the time at which the probationary period ended among formal workers. However, these spikes are also found in the informal sector. Also, spikes were found at the sixth and twelfth months, probably related to renewal of short-term contracts—as a way to avoid job stability measures—and to avoid discrete jumps in the severance payment. After the reforms, there is an increase in the hazard function for formal wage earners and an increase with respect to the hazard function of informal-sector wage earners. Large hazards in the third and sixth months are observed among private formal workers, and not among public ones, consistent with private firms' using short-term contracts in order to avoid job stability. Hazards are always higher for blue-collar workers, but the difference between blue- and white-collar workers diminishes after the reforms. Finally, we performed parametric estimations of hazard estimations in order to control for demographic characteristics of workers. These confirmed the results of higher hazards for informal, younger, private, and blue-collar workers. Education has a significant negative coefficient, suggesting lower hazards for the more educated, particularly after the reforms launched in 1991. There is evidence of a small relative increase in turnover for formal wage earners after the reforms. Having a temporary contract increases the hazard, suggesting higher turnover among these workers. This effect disappears by 1994, although temporary contracts were intensively used, which may be related to a smaller difference in status between temporary and permanent positions within a firm. Further work is needed, as 1994 is close to the beginning of the labor market reforms.

Appendix A

Table 2A.1 Evolution of Nonwage Costs Paid by Employer and Employee by Item, 1987–1997

							1993		1994		1995		1996		1997	
	1987	1988	1989	1990	1991	1992	IPSS	AFP	IPSS	AFP	IPSS	AFP	IPSS	AFP	IPSS	AFP
Nonwage Costs Paid by Employer																
Tenure bonus	8.33[a]	8.33	8.33	8.33[b]	8.33[c]	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33
National Housing Fund	4.00[d]	5.00[e]	5.00	5.00	8.00[f]	8.00	6.00[g]	6.00	6.00	6.00	9.00[h]	9.00	9.00	9.00	9.00[i]	9.00[i]
Holidays bonus	16.67	16.67	16.67[j]	16.67	16.67	16.67	16.67	16.67	16.67	16.67	16.67	16.67	16.67	16.67	16.67	16.67
IPSS payments	6.00[k]	6.00[l]	6.00	6.00[m]	6.00	6.00	6.00		6.00[n]							
Public health plan	6.00	6.00[o]	6.00	6.00	6.00	6.00	6.00	6.00	6.00	6.00	9.00[p]	9.00[p]	9.00	9.00	9.00	9.00
Accident insurance	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00	2.00
Industrial Training Fund	1.50	1.50	1.50	1.50	1.50	1.50	1.50	1.50	1.50	1.50	1.25[q]	1.25[q]	1.00	1.00	0.75	0.75
Vacations	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33	8.33
Total	52.83	53.83	53.83	53.83	56.83	56.83	54.83	48.83	54.83	48.83	54.58	54.58	54.33	54.33	54.08	54.08
Nonwage Costs Paid by Employee																
National Housing Fund	0.50[d]	1.00[e]	1.00	1.00	1.00	1.00	3.00[r]	3.00[r]	3.00	3.00	[h]	[h]				
IPSS payments	3.00[k]	3.00[s]	3.00	3.00	3.00	3.00	3.00		3.00		11.00[p]		11.00		13.00[t]	
AFP payments								10.0		10.0		8.00[p]		8.00		8.00
Solidarity payment								1.00		1.00	[p]	[p]				
Accident/burial expenses[21]								2.25		2.01		1.17		1.33		1.38
Percentual commission[21]								0.64		2.03		1.98		2.02		2.34
Public health plan	3.00	3.00[o]	3.00	3.00	3.00	3.00	3.00	3.00	[p]	[p]						
Total	6.50	7.00	7.00	7.00	7.00	7.00	9.00	19.89	9.00	21.04	11.00	11.15	11.00	11.35	13.00	11.72
Total nonwage costs	59.33	60.83	60.83	60.83	63.83	63.83	63.83	68.72	63.83	69.87	65.58	65.73	65.33	65.68	67.08	65.80

Source: Análisis Laboral (1987–1997).

Notes: IPSS = public retirement plan. AFP = private retirement plan. Blank cells indicate that information is not available.

[a]Last wage for every complete year of tenure. The maximum taxable wage equals ten Minimum Vital Wages.

[b]Since June 1990, the maximum taxable wage is the last wage, including holiday bonuses.

[c]Since January 1991 the employer must deposit the tenure bonus in an authorized financial institution on May and November.

[d]The maximum taxable wage is set at eight tax units.

[e]Changed in November 1988.

[f]Changed in May 1991.

[g]On January 1993 the employer's payment and the maximum taxable wage were abolished. On November 1993 it was set at 6 percent.

[h]Changed in August 1995.

[i]In January 1997 the employer's contribution was reduced to 7 percent, but holiday bonuses were included in the taxable wage. In August 1997 the contribution was reduced to 5 percent.

[j]Officially regulated since December 1989. However, this was already a usual practice long before this date.

[k]The minimum taxable wage equals one Minimum Vital Income, and the maximum equals ten Minimum Legal Incomes.

[l]Since January 1988, the maximum taxable wage was set at twenty Minimum Legal Incomes.

[m]In August 1990 the Minimum Legal Income is replaced by the Minimum Vital Remuneration. In October 1990, the maximum taxable wage is eliminated.

[n]Abolished in August 1995.

[o]The maximum taxable wage, set at ten Minimum Legal Incomes, was eliminated in January 1988.

[p]Changed in August 1995.

[q]The contribution decreases 0.25 percent every year until 1997 (0.75 percent).

[r]In January 1993, the employee's payment was set at 9 percent, and the maximum taxable wage was abolished. In November 1993, it was set at 6 percent.

[s]The maximum taxable wage for the employee was eliminated in January 1988.

[t]In January 1997, the employee's contribution increases to 13 percent.

[u]Market average.

Appendix B
Labor Demand Estimations

To test for cyclical variations of total labor demand due to employment composition changes (Bentolila and Saint-Paul 1992) we estimate the equation

$$\ln L_{i,t} = X_{i,t}\Omega' + d \ln \hat{L}_{i,t-L} + e \ln \hat{L}_{i,t-L} \cdot \ln E(\text{sp})_{i,t} + \delta t + \beta \mathbf{Z}_{i,t} + \text{cycle}_t(X_{i,t}\Omega') + \varepsilon_{i,t},$$

where

$$\Omega = (b_1, b_2, c, \delta)$$

$$X_{i,t} = [w + p,\ Y,\ E(\text{sp})],$$

where "cycle" is a dummy equal to zero in recessions and equal to 1 in expansions, and is interacted with all the regressors in the equation ($X_{i,t}$). This variable has a value of 1 when sectoral growth is 4 percent or more and zero otherwise. We used generalized least squares and correct for serial correlation with a correlation coefficient specific for each of the panels. The results of the estimations are shown in table 2B.1.

Appendix C
Equality of Empirical Hazard Functions

Graphical Analysis

In order to verify the equality of the hazard functions for complete and incomplete spells we assume that incomplete spells are completed ones and then compute the empirical hazard rates (Kaplan and Meier 1958) for both types of spells. These estimates are shown in the graph; note that the empirical hazard for incomplete spells has the same shape and spikes as the complete ones. Hazard functions for complete spells are above those using incomplete data, a fact that is consistent with lower mean tenures calculated using the former data set. Still, the pattern followed by the hazard function looks similar.

Kolmogorov-Smirnov Test

We use the Kolmogorov-Smirnov (K-S) statistic to formally test the equality of the empirical hazards functions between complete and incomplete spells (defined as uncensored spells). The test evaluates the closeness of the distributions λ^{is} and λ^{cs} (for incomplete and complete spells hazards)

Table 2B.1 Labor Demand Estimation Results for Panels at the Sector and Firm Level

	Sector Level			Establishment Level		
	1987–1990	1991–1994	1995–1997	1987–1990	1991–1994	1995–1997
Constant	8.189***	15.511***	14.900***	0.437***	0.027	1.554***
	(1.579)	(2.218)	(3.762)	(0.151)	(0.085)	(0.512)
$\ln(w + p)$	0.574***	–0.316***	–0.353***	–0.031***	0.026***	–0.056*
	(0.204)	(0.118)	(0.137)	(0.009)	(0.006)	(0.033)
$\ln[E(\text{sp})]$	–0.907**	–0.613	–0.443	–0.223***	–0.036*	–0.125
	(0.363)	(0.424)	(0.634)	(0.032)	(0.018)	(0.104)
$\ln(Y)$[a]	0.017	0.112	0.094*	0.206***	0.008	0.083**
	(0.067)	(0.102)	(0.054)	(0.009)	(0.008)	(0.035)
$\ln(L_{t-1})$[a]	0.074	–0.206	–0.005	0.787***	0.943***	0.613***
	(0.148)	(0.215)	(0.313)	(0.024)	(0.016)	(0.088)
$\ln(L_{t-1}) \cdot \ln[E(\text{sp})]$	0.041	0.066	0.028	0.053***	0.006	0.041*
	(0.027)	(0.045)	(0.060)	(0.005)	(0.004)	(0.021)
cycle dummy · $\ln(w + p)$	–0.035	–0.022	0.018	0.040***	0.005	0.006
	(0.052)	(0.049)	(0.018)	(0.010)	(0.008)	(0.043)
cycle dummy · $\ln(E(\text{sp})]$	0.046	0.030	–0.026	–0.009	–0.010	–0.023
	(0.063)	(0.066)	(0.024)	(0.013)	(0.011)	(0.059)
cycle dummy · ln(gdp)	0.017	0.020	–0.013	–0.078***	0.001	0.043
	(0.040)	(0.037)	(0.014)	(0.011)	(0.009)	(0.057)
Log likelihood	210.86	139.94	200.42	2,757.98	2,485.27	–1,388.166
χ^2	12,589.99***	4,561.53***	3,474.32***	100,574.42***	186,433.25***	2,736.63***
No. of observations	189	189	117	4,754	6,491	1,722

Notes: Standard errors in parentheses. Sector level includes fixed effects. For establishment level, a time trend was included. It was significant for the period 1991–1994, but not for 1987–1990 and 1995–1997.

[a]Instrumentalized with lagged values using rolling equations.

***Significant at the 1 percent level.

**Significant at the 5 percent level.

*Significant at the 10 percent level.

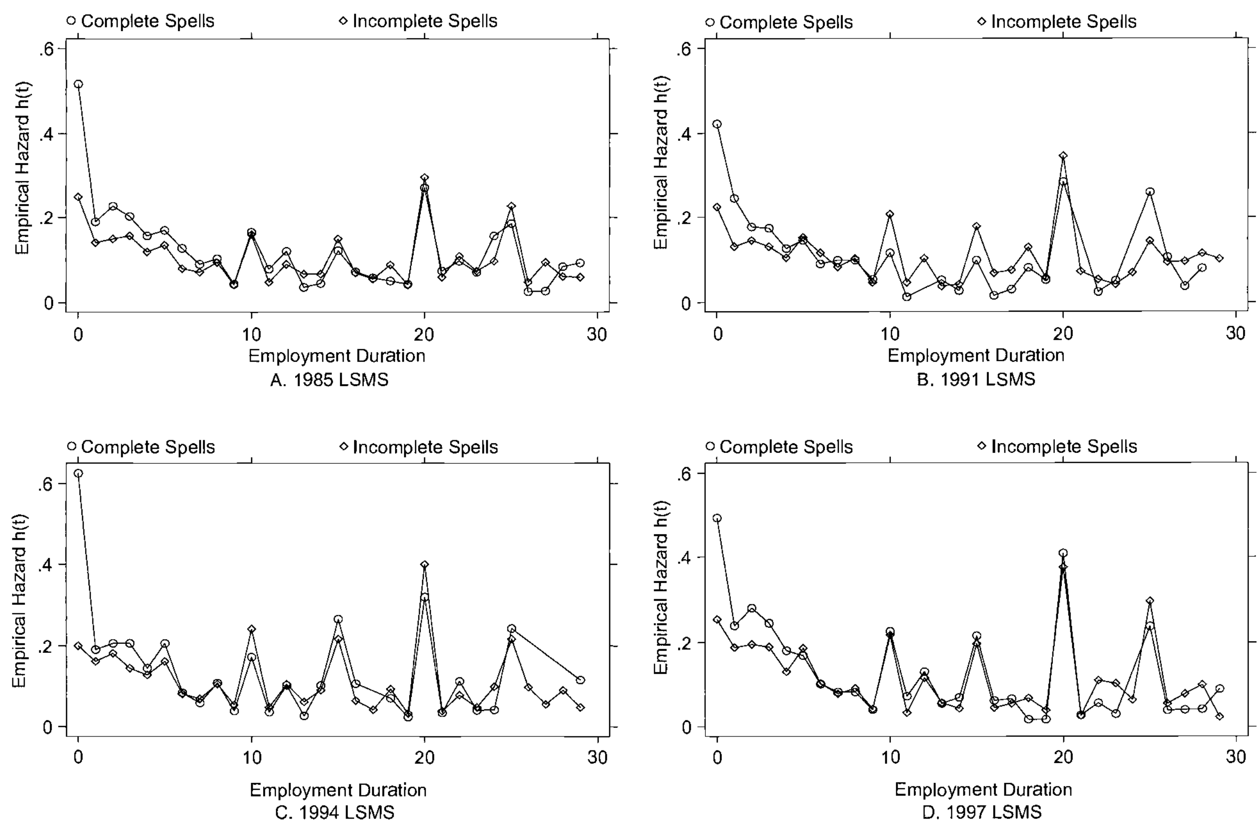


Fig. 2C.1 Empirical hazard functions of complete spells (assuming completeness of incomplete spells)

Table 2C.1 Kolmogorov-Smirnov *D* Statistic

LSMS	*D* statistic	*P* Value	Critical Value (95%)
1985	0.2143	0.228	0.22
1991	0.1721	0.661	0.23
1994	0.1884	0.470	0.23
1997	0.1779	0.553	0.22

by computing the least upper bound of all pointwise differences $|\hat{\lambda}^{is}(x) - \hat{\lambda}^{cs}(x)|$. We can write the K-S statistic D as

$$D = \sup_x[|\hat{\lambda}^{is}(x) - \hat{\lambda}^{cs}(x)|].$$

The null hypothesis ($H_0 : \lambda^{is} = \lambda^{cs}$) is accepted if λ^{is} is sufficiently close to λ^{cs}, in other words if the value of D is sufficiently small or smaller than the critical value at a certain significance level. The results are shown in table 2C.1. At the 95 percent significance level we cannot reject the null hypothesis that the two empirical hazard functions are equally distributed.

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