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Labor Market Adjustment in Chile

As a result of the Asian crisis, Chile experienced a substantial economic slowdown in 1998. Growth of the gross domestic product (GDP) among Chile's trading partners in 1998 was a full 2 percent below the previous five-year average, while terms of trade fell by over 5 percent.[1] In addition, the Central Bank of Chile applied a contractionary monetary policy aimed at minimizing the nominal devaluation and reining in the current account deficit. Interest rates on loans rose substantially in 1998—from preshock levels of 8 percent to above 18 percent in September 1998. Annual output growth fell to 3.2 percent in 1998 and then to −1.0 percent in 1999, a full 8.0 percent below the average growth rate of the previous ten years.

This collapse in output growth had an immediate effect on unemployment, which rose to an average of 8.3 percent in 1999, up from an average of 6.1 percent in the previous year and an average of 6.9 percent over the previous five years. More surprising than then initial upward jump in unemployment, however, was the slow recovery. The Chilean economy grew at an average rate of 2.9 percent between the first quarter of 2001 and the last quarter of 2002. Over this period unemployment remained relatively high, falling only 0.3 percent and averaging 9.2 percent.

The sluggish response of unemployment to improved economic conditions after 2000 led to a prolonged policy debate in Chile. One of the main

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1. External demand is measured as the growth rate of gross domestic product (GDP) among the country's trading partners.

concerns was that Chile experienced a labor-saving structural change in the late 1990s that aggravated the consequences of the aggregate demand shock mentioned above.[2] Explanations for this change included the increasing use of labor-saving technologies, the additional nonwage costs of realized (and expected) labor reforms, and the differential impact of the external demand shock and subsequent monetary policy contraction on the relatively labor-intensive small and medium-sized enterprises (SMEs).[3]

The distinction between a temporal (cyclical) demand shock brought about by deteriorating economic conditions and a permanent technological shock to demand has important policy implications. The former is temporary, and both employment and real wages should return to precrisis levels once external demand recovers. The latter shock is permanent (or at least strongly persistent in the case of the SME hypothesis), so employment and real wages will remain low even after external demand returns to precrisis levels.

These competing hypotheses, however, have nothing to say directly about the level and persistence of unemployment rates. Indeed, falling labor demand will only translate into higher unemployment rates if real wages are rigid. In the standard classical model, wages do all the adjusting to changes in labor demand, keeping employment stable at full employment levels.

If the quantity of labor supplied is procyclical, then changes in unemployment will mask much larger changes in labor demand. This is the case in the Chilean labor market, where labor participation has moved in step with employment over the last decade and a half. Over the period 1988–2003, changes in unemployment were generally much smaller than changes in employment growth, and the rise in unemployment following the Asian crisis was no exception.[4] Because of this procyclical labor supply, the drastic fall in employment following the Asian crisis did not translate into even higher unemployment rates. Without this extra margin, the aggregate

2. For example, senators used this argument to push for more active employment policies during the period; see Senado de la República de Chile, "Propuesta para una política activa de empleo," 22 July 2000.

3. Bergoeing and Morandé (2002), for example, argue that labor code reforms may explain labor market performance during this period. See also Pagés and Montenegro (1999) for a discussion of this issue. For a summary of the two sides of the SME debate, see Cabrera and others (2002).

4. See Bellani and Restrepo (2002); Duryea and Pagés (2001).

unemployment rate would have risen by 4.1 percentage points between 1997 and 2000 instead of the 3.3 percent actual increase.

In a related paper, we show that the response of labor supply to labor market conditions not only is procyclical, but also varies substantially across employment categories.[5] Consequently, the rise in the unemployment rate over the period 1997–2000 was highest among highly educated young workers, but the slowdown in employment growth was largest among young workers with a low level of education. This apparent contradiction is explained by the much larger procyclical response of the labor supply among young uneducated workers.

Because of this distinction between the path of the unemployment rate and the level of employment, and because the underlying initial shock to the Chilean labor market in 1998 was by all accounts a labor demand shock, in what follows we put unemployment rates aside and concentrate our discussion on employment and employment demand. We find little evidence for the hypothesis of structural change in employment demand. We show that the slowdown in employment in the late 1990s is in line with what would be predicted by a stable demand function given the path of wages and aggregate demand. We argue, therefore, that the slowdown in employment growth after 1997 is the result of a negative aggregate demand shock combined with downward rigidity of real wages. Two institutional characteristics of the Chilean labor market contributed to this wage rigidity: a statutory minimum wage that became increasingly binding in the years following the Asian crisis and the prevalence of long-term price-indexed wage contracts. In particular, we use microeconomic data from the Chilean employment survey to show that close to 6 percent of workers where affected by the rise in the minimum wage that took place between 1997 and 2000. We also show that the evolution of the aggregate wage is well described by inflation-adjusted two-year contracts, suggesting that the practice of subscribing indexed long-term contracts goes beyond the relatively small fraction of workers involved in collective wage bargaining. These rigidities observed at the microeconomic level seem to lie behind the slow recovery of aggregate employment during the period of mild output growth experienced after 1999. High wage rigidity also explains why Chile experienced a relatively larger fall in employment than other developing countries that faced substantial output reductions in the 1990s.

5. See Cowan and others (2004).

Fall in Quantity Demanded, Not a Shift in Demand

Following the Asian crisis, Chile experienced a fall in private employment growth across all sectors of economic activity. Indeed, the only sector in which employment was anticyclical was the social and communal services sector, which is dominated by public employment. Over the three-year period from 1997 to 2000, total employment grew by a meager 0.5 percent a year, while private employment (measured as total employment net of employment in the social and communal services sector) fell by close to 1.0 percent, well below the 2.6 percent annual growth of the previous seven years. The largest drop in employment growth took place in the construction sector, followed by the utilities sector (electricity, gas, and water). The largest drop in the level of employment took place in the mining sector (figure 1).

This slowdown is not at all surprising given the behavior of output and wages. To demonstrate this, we estimate a simple model of labor demand and then conduct stability tests and out-of-sample forecasts based on the model. We find no evidence of structural change in labor demand, suggesting that concerns about the impact of rising nonwage costs or dwindling investment opportunities may have been largely unfounded.[6]

We start by constructing estimates of employment growth based on a simple production function, such that labor demand is given by

$$L_d = \frac{\alpha}{\gamma}\frac{PY}{W}, \tag{1}$$

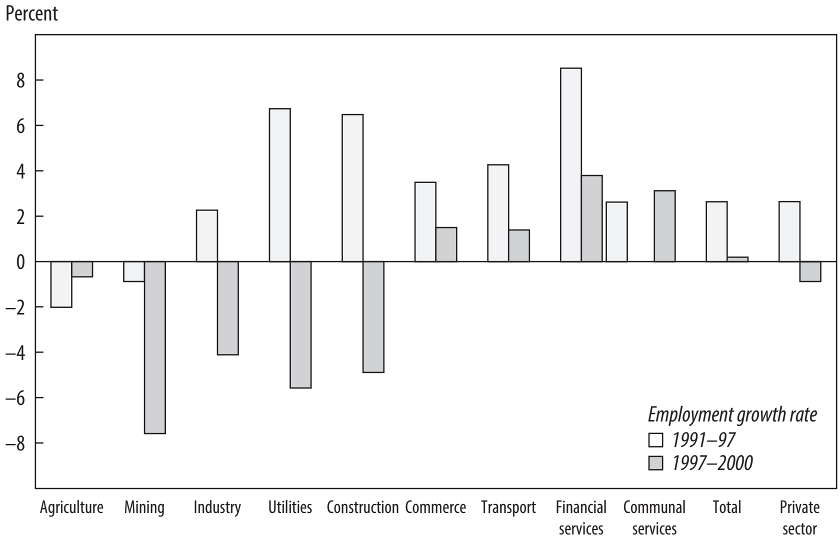
where α is the employment-output elasticity, γ is the markup, PY is nominal output, and W represents nominal wages.

Figure 2 plots predicted employment growth rates built from equation 1 against realized private employment growth over successive three-year periods in the 1990s. The value for 1997 thus corresponds to annualized private employment growth between 1994 and 1997, and so on.[7] In theory, if both the employment-output elasticity and the markup are constant, all observations should fall along the forty-five degree line. The figure shows, however, that actual employment growth is systematically higher than the

6. This result is in line with work by Martínez, Morales, and Valdés (2001).

7. We use three-year periods to allow for short-run deviations of employment from its optimal level as a result of adjustment costs.

FIGURE 1. Employment Demand: Sector Growth



Source: National Institute of Statistics (INE).

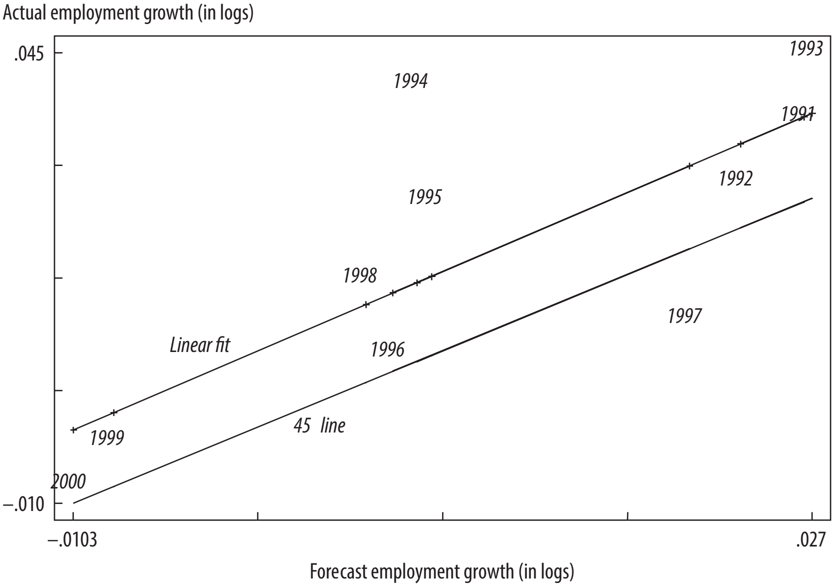
estimated values. To take this into consideration, we run a simple linear regression between predicted and actual employment growth. Our estimated coefficient shows that on average, actual employment growth is 1.0 percent higher than the predicted values.

All crisis years shown in figure 2 are above the 45 degree line, and no year is significantly below the predicted values when the average 1.0 percent annual change in α/γ is included. We obtain very similar results when we use total employment instead of private employment as our dependent variable. This suggests that the relation between employment, output, and wages did not change significantly after 1998.

To see how this relation plays out in more detail over time, we use equation 1 and sectoral data to obtain yearly out-of-sample predictions for annual employment growth over the period 1999–2001.[8] Figure 3 shows our results for two exercises: our employment growth predictions using actual

8. For further details on the sector results and the methodology, see Cowan and others (2004).

FIGURE 2. How Much Can We Explain with a Simple CD Framework?



Source: Authors' calculations, based on data from INE and Central Bank of Chile.

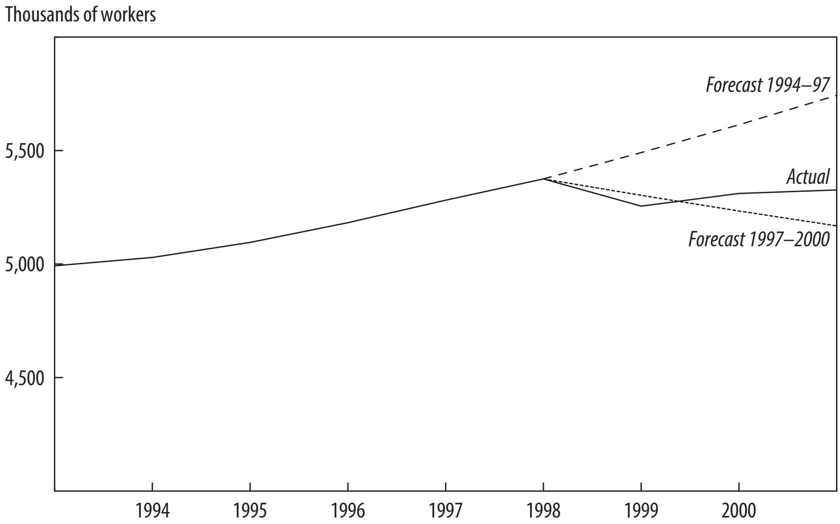
values of output and wages for each year and our employment predictions using average output and wage growth for the precrisis period 1994–97. The impact of the economic slowdown is clear from the gap between the two predictions. What is also apparent from both figure 2 and figure 3 is that employment was not significantly lower in 1999–2001 than could have been expected given the path of output and wages and the average relation between these three variables in previous periods.

With these results in mind, we turn now to a more careful statistical analysis of structural change. Assume that production takes place using the following standard production technology:

$$Y = A(\alpha L^{(\sigma-1)/\sigma} + (1-\alpha)K^{(\sigma-1)/\sigma})^{[\sigma/(\sigma-1)]\gamma}, \tag{2}$$

where Y is output, A is total factor productivity, L is labor, K is capital, α is the labor share, σ is the elasticity of substitution between labor and capital, and γ is a measure of returns to scale. In the absence of adjustment

FIGURE 3. A Second Look at the Employment Data



Source: Authors' calculations, based on data from INE and Central Bank of Chile.

costs, labor demand for a profit-maximizing firm operating with this technology in competitive product and factor markets is given by

$$l^* = \sigma \ln(\alpha\gamma) - \frac{(1-\sigma)}{\gamma} a - \sigma(w-p) + \left[1 + \frac{(1-\gamma)(1-\sigma)}{\gamma}\right] y, \tag{3}$$

where all lower case variables denote logs.

If the firm also faces quadratic adjustment costs, then actual employment demand will be a convex combination of the desired level given by equation 3 and lagged employment:

$$l_t = \lambda l_t^* + (1-\lambda) l_{t-1}. \tag{4}$$

When we substitute equation 3 into equation 4, and assuming that the log of total factor productivity, *a,* follows a linear trend, we obtain our empirical specification:[9]

$$l_t = \beta_1 + \beta_2 t + \beta_3(w_t - p_t) + \beta_4 y_t + (1-\lambda) l_{t-1} + \mu_t. \tag{5}$$

9. Martínez, Morales, and Valdés (2001) find that employment, wages, and output are cointegrated; we therefore estimate this specification in levels.

Table 1 reports the results we obtain from estimating equation 5 by ordinary least squares using quarterly data for the whole economy over the period 1986 to 2003. The first column reports our results using the log of total employment as our dependent variable; the second column presents our preferred specification, in which we replace total employment with employment net of communal services (a measure of private employment). All estimated coefficients have the expected signs. Furthermore, the estimated value of $1 - \lambda$ implies that the half-life of a shock to labor demand is 5.4 quarters, which is broadly consistent with other empirical studies on adjustment costs in the labor market.[10] The estimated coefficients also indicate that the elasticity of substitution between labor and capital is very close to 1 (which is equivalent to a Cobb-Douglas production function), so that a 1.0 percent change in output leads to an identical percent change in employment.

To test the stability of our estimated coefficients, we carry out a cumulative sum (CUSUM) test of structural change. The CUSUM test evaluates the degree to which the model generates adequate out-of-sample predictions by accumulating one-period-ahead out-of-sample prediction errors. If no structural change has occurred, the prediction errors should compensate each other and the cumulative sum should be close to zero. If structural change is present, however, the errors will take a positive or negative sign and the sum tends to increase (in absolute value). Figure 4 shows the CUSUM tests for the results reported in the first and second columns of table 1. Neither case shows evidence of a structural change in our estimated model of labor demand, since the cumulative sum stays within the confidence interval of a zero sum. We extend this analysis using sectoral data for the three largest sectors in terms of private employment: manufacturing, commerce, and financial services (unreported). In all cases, we fail to find evidence of a structural change.

In sum, an across-the-board fall in private employment growth followed the aggregate shocks that hit the Chilean economy in 1997 and 1998. This slowdown in employment growth can be fully explained by the behavior of aggregate wages and output, which suggests that labor demand did not undergo a structural change in the late 1990s in Chile. This rules out

10. Caballero, Engel, and Micco (2003) find a half life of 2.5 quarters based on annual plant-level data. Martínez, Morales, and Valdés (2001) also find a half life of 2.5 quarters, using quarterly GDP.

TABLE 1. Testing Structural Change[a]

Explanatory variable	*(1)*	*(2)*
Real GDP	0.094**	0.139**
	(0.044)	(0.056)
Real Wages	−0.126*	−0.155*
	(0.070)	(0.081)
Employment t-1	0.886***	0.939***
	(0.045)	(0.031)
trend	0.000	−0.001
	(0.000)	(0.000)
constant	0.234	−0.661
	(0.239)	(0.451)
Summary statistic		
R^2	0.99	0.99
Period	1986:1–2003:2	1986:1–2002:4
No. observations	69	67

Source: Authors' calculations, based on data from National Institute of Statistics (INE) and Central Bank of Chile.
* Statistically significant at 10 percent.
** Statistically significant at 5 percent.
*** Statistically significant at 1 percent.
a. The dependent variable in column 1 is log total employment; in column 2, it is log employment net of communal services. The estimation method is OLS, using quarterly data for the whole economy over the period 1986 to 2002.

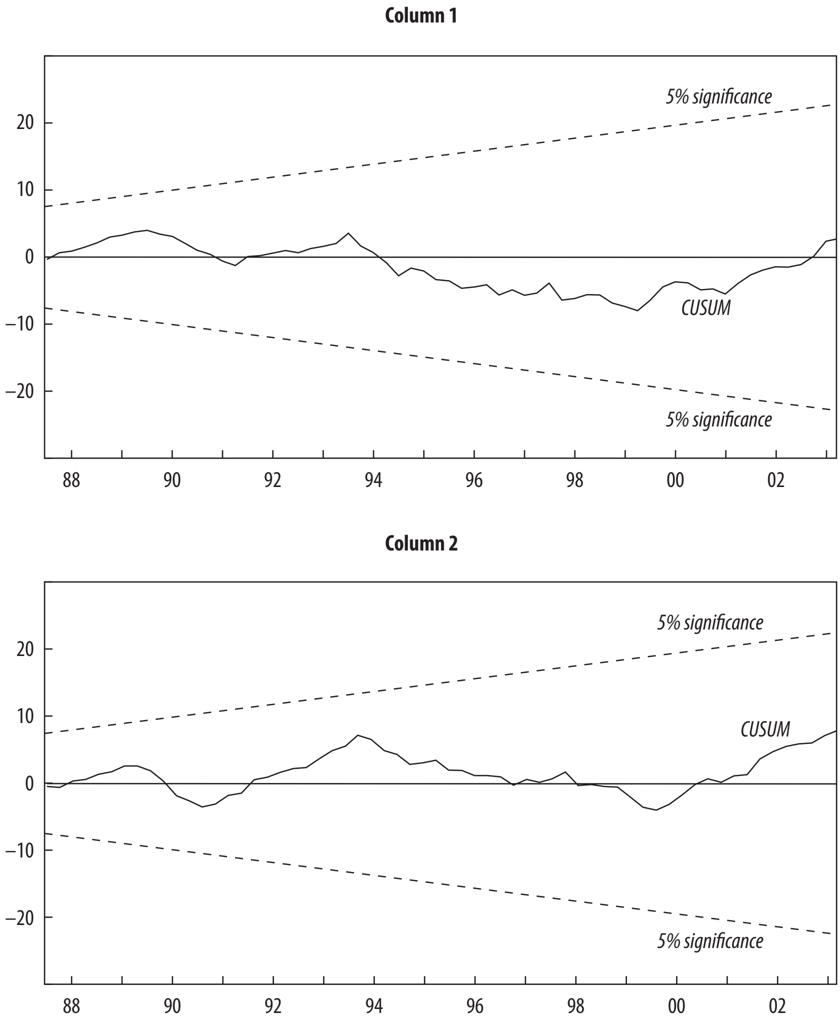
theories that explain the sluggish behavior of employment in the period 1998–2001 as resulting from the introduction of labor-saving technologies or low business confidence associated with expected labor reforms.[11]

What Is behind Chilean Wage Rigidity?

The previous section traced falling labor demand to a drop in output that was not compensated by a reduction in real wages. This section explores the reasons for this wage rigidity. We start by comparing Chile's wage flexibility with that of other economies that also experienced large external shocks in the 1990s, which shows that wages in Chile are relatively rigid by international standards. We then explore three institutional factors that may explain this downward rigidity: public sector wages, minimum wages,

11. If the slowdown in employment was caused by low hiring stemming from low business confidence, then observed employment would have been lower than the level predicted on the basis of wages and output. Business expectations could still play a role in explaining low employment if they caused low output levels.

F I G U R E 4 . CUSUM Tests of Structural Change[a]



Source: Authors' calculations.

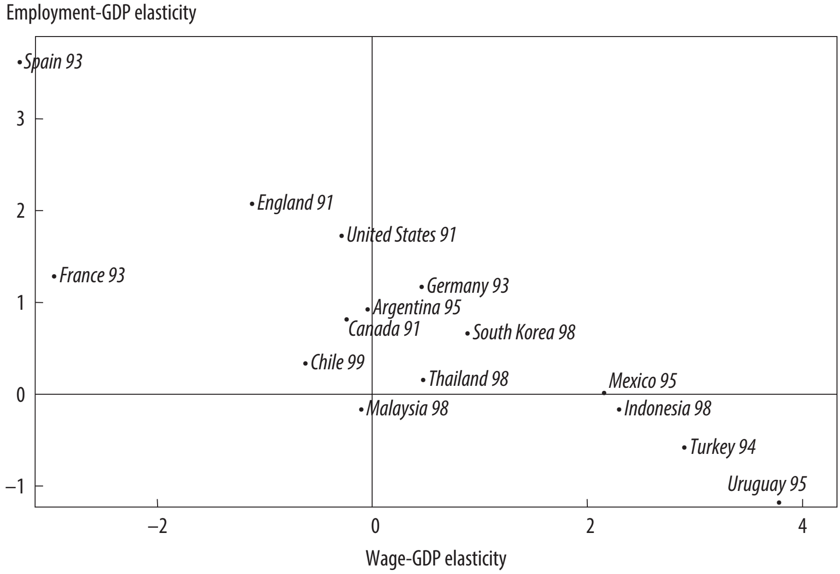
a. The figure shows the cumulative sum (CUSUM) tests for the results reported in the first and second columns of table 1.

and the type of private wage contract prevalent in Chile at the time of the slowdown. Our results indicate that upward pressure from minimum wages and the practice of using long, indexed wage contracts both played a part in the downward rigidity of Chilean wages.

Over the period 1991–97, real wages—which are constructed using the gross domestic product (GDP) deflator—grew at an average yearly rate of 6.2 percent. Wage growth slowed in 1997 and averaged 3.4 percent over the next three years. In no year, however, did real wages fall, despite the substantial slowdown in GDP growth. Real wages continued to grow at above 2.0 percent a year even when unemployment rose above 9.0 percent in 1999. This wage rigidity was observed for all occupational groups. The cost of labor grew fastest for professionals and technicians, but all groups experienced positive wage growth from 1998 onward.

To put Chilean wage flexibility (or lack thereof) in context, we compare the relative response of wages, output, and employment in 1999 (the year in which output fell the most in Chile) with that of other economies that experienced substantial output drops in the 1990s. We include Argentina, Mexico, and Uruguay in 1995 (the year following the Tequila crisis); Indonesia, Korea, Malaysia, and Thailand in 1998 (the year following the Asian crisis); and a set of member countries of the Organization for Economic Cooperation and Development (OECD) that experienced substantial output drops in 1991 and 1993 (the year of the European Monetary Union's collapse). Figure 5 plots the elasticity of wages to GDP changes against the elasticity of employment to GDP. Since we show years of negative output growth, a positive wage or employment elasticity indicates a procyclical response in wages or employment (that is, a wage or employment decline), while a negative wage or employment elasticity indicates that employment or wages increased despite negative output growth. Countries that exhibit high wage elasticity appear at the right end of the figure, while countries such as Chile, which experienced a positive growth in wages in a period of GDP decline, are positioned to the left of zero, with a negative wage elasticity. Two things are apparent from this figure. First, the two variables display a strong negative correlation, as expected. Countries with high wage elasticities, such as Indonesia and Mexico, experienced relatively smaller employment elasticities than countries like Argentina or Chile, where real wages were more rigid. Second, real wages in Chile are among the most rigid in the group of emerging economies in our sample. In terms of wage and employment responses, Chile behaves more

FIGURE 5. Wage-GDP Elasticities versus Employment-GDP Elasticities



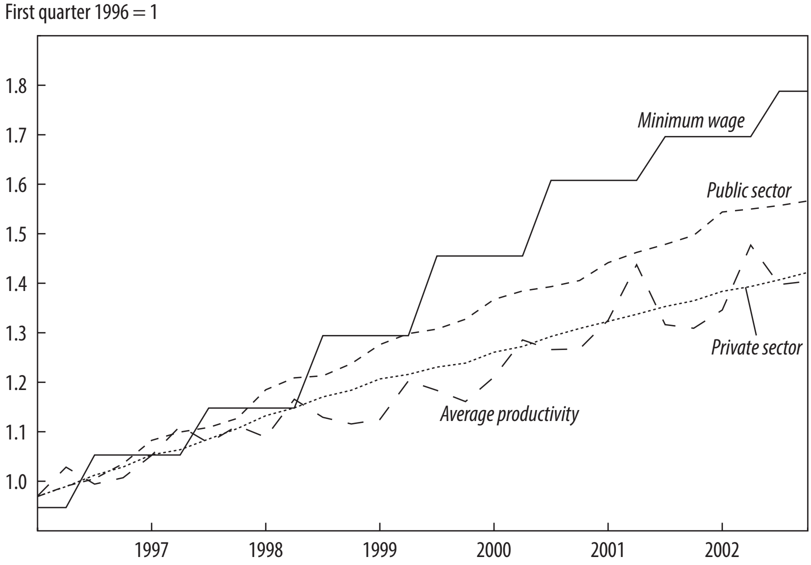
Source: Authors' calculations, based on data from the World Bank's *World Development Indicators*.

like the wage-rigid developed countries than the wage-flexible emerging economies.[12]

Figure 6 shows the behavior of average nominal private sector wages, wages in the social and communal services sector (which is largely made up of public sector employees), the mandated minimum wage, and nominal output per worker for the period 1996 to 2002. All four variables moved in step before the slowdown in 1998, with the three wage indices closely following average productivity. After 1997, however, public sector wages

12. Figure 5 also shows that relative to other countries, Chile exhibits a low employment-to-GDP flexibility given its level of wage-to-GDP elasticity, which implies that the employment decline in Chile could have been much higher if Chile had higher employment elasticities. Canada, the United Kingdom, and the United States experienced similarly high wage rigidity, but they had much higher responses in terms of employment. This suggests that employment regulations—which are more stringent in Chile than in these three developed countries (Heckman and Pagés, 2003)—reduced employment adjustment in Chile and induced adjustment in hours worked. Chile has the second highest hours-to-GDP elasticity in the sample of countries represented in figure 5.

FIGURE 6. Wage Rigidity: Minimum and Public Sector Wages[a]



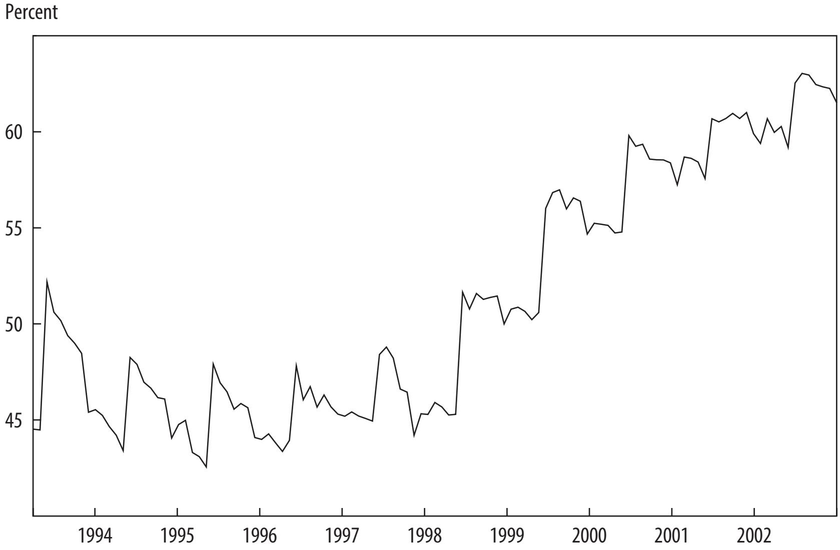
Source: INE and Central Bank of Chile.
a. Data are quarterly; years are labeled at first quarter.

and the minimum wage outpace both private wage growth and average productivity gains.[13]

The relative rise of minimum wages vis-à-vis private wages is even more apparent in figure 7, which plots the minimum wage as a percentage of the average nominal wages of unskilled workers. Minimum wages grew considerably faster than unskilled wages after 1998. As a result, the ratio of minimum wages to unskilled wages jumped from close to 45 percent at the end of 1997 to more that 60 percent by the end of 2002. A large part of this rise was due to a Chilean Congress decision in May 1998 to set minimum wage increases of 12.7 percent in June 1998, 12.4 percent in June 1999, and 10.5 percent in 2000 so as to achieve a target minimum monthly wage of 100,000 pesos by 2000. This policy entered into effect right before the

13. We obtain almost identical results if we use the public sector component of the social and communal services sector. This category was constructed by the National Institute of Statistics (INE) for this study.

FIGURE 7. Wage Rigidity: Minimum Wage versus Wages of Unskilled Workers[a]



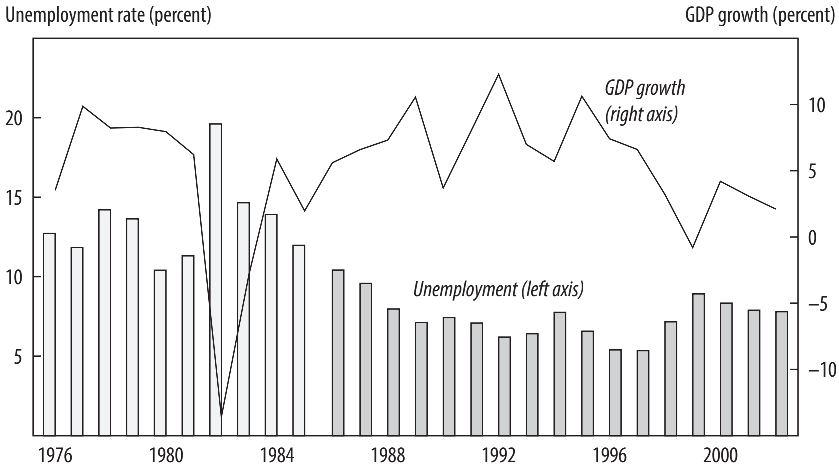
Source: INE.

a. Data are monthly; years are labeled at January.

deceleration of growth registered in the third quarter of 1998, when growth projections were still estimated to be above 6.0 percent a year. The relatively high growth rate of both public and minimum wages evident from the last two figures suggests that this is a plausible cause for the recent downward wage rigidity in Chile.

Unemployment rates above 8 percent are not a new phenomenon in Chile. During the debt crisis of the early 1980s, unemployment rates shot up to close to 20 percent. This is not surprising given the severity of the crisis in Chile—output fell by almost 15 percent in 1982 alone (see figure 8). What is surprising, however, is the length of time it took unemployment rates to drop to 5 percent (the precrisis level) despite prolonged periods of high growth: growth rates were above 5 percent in almost every year from 1984 to 1997, but the unemployment rate only fell to 5 percent of the workforce in 1996. Furthermore, it took five years of postcrisis recovery to bring unemployment rates below 10 percent in 1987.

FIGURE 8. Unemployment Is not a New Phenomenon[a]



Source: INE and Central Bank of Chile.

a. For the change in the GDP growth rate for the period 1976–85, the base is 1977; for 1986–96, the base is 1986; and for 1996–2 001, the base is 1996. Note that unemployment rates between 1976 and 1985 are not comparable to unemployment rates between 1986 and 2002, due to methodological changes.

The path of unemployment after the debt crisis suggests that downward wage rigidity is a persistent characteristic of the Chilean labor market. Thus, while minimum and public sector wages are probably among the causes of the recent downward rigidity, they are unlikely to constitute the full story. We therefore turn to another possible explanation for Chilean wage rigidity—namely, the length and indexation clauses of Chilean wage contracts.

Information on wage contracts in Chile is patchy at best.[14] Data are available, however, on the characteristics of contracts signed by workers and firms that participate in collective bargaining processes.[15] Three things stand out with respect to these private wage contracts: they are long, extending

14. One exception is Jadresic (1997), who finds, based on survey data and regression analysis, that aggregate private contracts are well described by two-year contracts that are revised every six months according to 100 percent of past inflation in the 1980s.

15. See Jadresic (1997) and statistics from the *Dirección del Trabajo*.

for two years, on average; they include clauses for full backward indexation every six months; and they have not changed significantly over the last fifteen years, despite falling inflation.[16]

Using an econometric approach, Jadresic finds evidence of important rigidities in Chilean wages in the 1980s.[17] He examines the behavior of aggregate wages in Chile in 1980–90 and shows that they are well described by two-year contracts revised every six-months based on 100 percent of inflation. This implies that the importance of long-term contracts in that period extended well beyond the share of employment under collective bargaining and affected the rest of the labor force. Since the above evidence suggests that collective agreements in the 1990s still follow this institutional model, we reexamine the behavior of Chilean aggregate wages for the 1990s to assess whether those institutional sources of aggregate wage rigidity are still at work.

The presence of long-term contracts dampens the response of wages to changes in aggregate economic conditions. For a start, they trigger a mechanical effect under which firms are only able to change wages once every two years. In addition, they set in motion a number of strategic interactions, which are emphasized extensively in the nominal rigidities literature. If a large share of firms cannot change their wages, then the incentives for those firms that can change their wages are reduced.[18]

To provide a quantitative idea of how large an effect this type of contract can have on employment, we reproduce a simulation exercise for a small open economy with floating exchange rates.[19] Figure 9 shows the response of employment to identical demand shocks under three types of contract. The first contract, *I* (our baseline simulation), corresponds to two-year contracts with adjustments for inflation every semester. The second type of contract, *P,* lasts two years but has no indexation clause. Finally, the flexible contract, *F,* lasts a year and has no indexation clause. Extending the length of the contract increases the persistence of the demand shock. The effects of the demand shock disappear after three semesters under flexible contracts, but they extend past the fifth semester for long-term contracts.

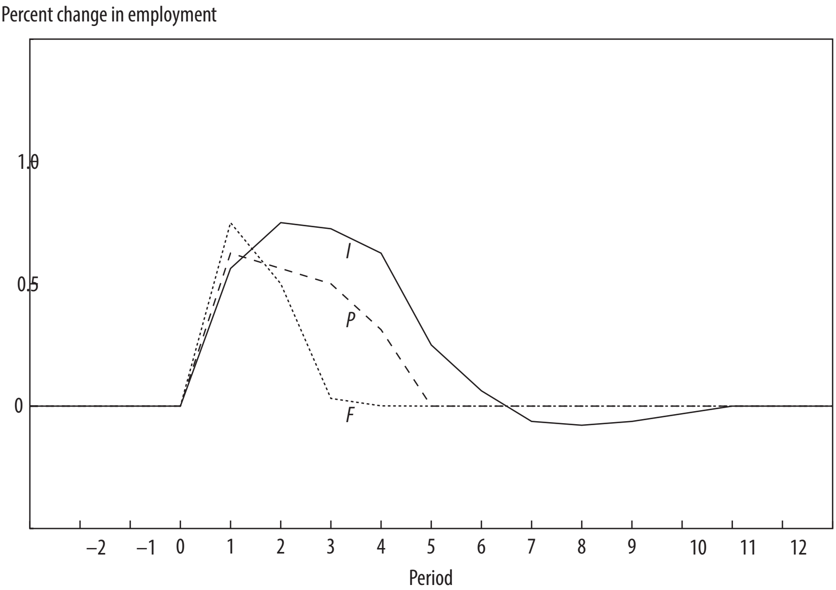
16. For a discussion of the consequences and causes of indexation, see Landerretche, Lefort, and Valdés (1998); Landerretche and Valdés (1997).

17. Jadresic (1997).

18. See Rotemberg and Woodford (1992).

19. The simulation is from Jadresic (1996).

FIGURE 9. Effect of a Demand Shock on Employment[a]



Source: Authors' calculations, based on Jadresic (1996).

a. The figure plots the economy's response to a demand shock under three types of contracts: *I*, which last two years and are adjusted for inflation every semester; *P*, which lasts two years but has no indexation clause; and *F*, which lasts a year and has no indexation clause.

While it is always the case that *P* and *I* are to the right of *F*, the relative ranking of *P* and *I* depends on the type of shock affecting the economy. In the particular case of a real demand shock and flexible exchange rates, indexation introduces additional persistence.

To piece together the impact of these factors on private sector wages in Chile and test their empirical relevance, we estimate a set of private wage equations based on the following specification:

$$\Delta \ln(W_t) = \alpha + \beta\pi_{6m,t} + \delta\Delta\ln(W_t^{\min}) + \varphi\mathbf{X}_t + \mu_t, \quad (6)$$

where $\Delta\ln(W)$ is the monthly change in nominal private wages, π_{6m} is the accumulated inflation over the last six months, $W^{\min}$ is the monthly change in minimum wages (which is only positive in June), and $\mathbf{X}$ is a vector of variables that capture desired changes in real wages.

The available information on wage contracts in Chile indicates that firms that participate in collective bargaining adjust their wages for past inflation every six months. Monthly nominal wages should therefore rise by the accumulated price change over the previous six months, multiplied by the share of firms that readjust in that specific month. If we assume that readjustments are evenly spaced over the year, then each month, one-sixth of the contracts should be adjusted for the accumulated inflation over the last six months, while the rest of the contracts remain unchanged.

We also know that minimum wages are adjusted once a year (in June), which should mechanically push up the wages of all workers earning the minimum. We expect this coefficient to be close to the share of workers directly affected by minimum wage changes.

Monthly changes in wages should also include the adjustments in the real wages of all firms that negotiated new wage contracts during that month. Our set of controls, **X,** thus includes the year-on-year growth rate of the monthly economic activity index (the *Indicador Mensual de Actividad Económica,* or IMACEC) and the average unemployment rate over the previous six months. We expect a positive coefficient estimate for output growth and a negative coefficient for unemployment. **X** also includes the year-on-year growth rates of nominal public wages and nominal minimum wages. We include this last variable in addition to $W^{\min}$ to capture the correlation between minimum wage growth and the growth rate in the wages of workers who are not directly affected by changes in the statutory minimum wage. We include the growth rate of public wages to capture the effect these may have on the opportunity cost of private workers.

Table 2 reports the results from estimating equation 6 using monthly data for the period January 1986 to March 2002. The first two columns include π_{6m} and either output growth or the unemployment rate over the last six months as regressors. The signs of all coefficients are as expected, although only π_{6m} is significant at conventional confidence levels. For the remaining specifications reported in table 2, we use unemployment as our control variable for labor market slackness; results are robust to exchanging unemployment for GDP growth. Columns 3 through 5 include, successively, the monthly change in the minimum wage, the year-on-year changes in the minimum wage, and the annual nominal growth rate of public wages. In each of these three specifications we find that the monthly change in minimum wages has a significant positive effect on private sector wages. The estimated coefficients on the twelve-month growth rate of public and minimum

TABLE 2. Aggregate Private Nominal Wages[a]

Explanatory variable	*(1)*	*(2)*	*(3)*	*(4)*	*(5)*	*(6)*	*(7)*
Inflation (last six months)	0.166***	0.160***	0.165***	0.165***	0.169***	0.166***	0.150***
	(0.015)	(0.016)	(0.015)	(0.015)	(0.015)	(0.017)	(0.023)
Unemployment (last six months)	−0.018		−0.033	−0.031	−0.024	−0.031	
	(0.030)		(0.028)	(0.028)	(0.028)	(0.030)	
GDP (year-on-year growth)		0.013					
		(0.012)					
Minimum wage (monthly growth)			0.015**	0.015**	0.015**	0.015**	0.011
			(0.007)	(0.007)	(0.006)	(0.007)	(0.007)
Minimum wage (year-on-year growth)				0.004	0.001	0.004	0.005
				(0.009)	(0.010)	(0.010)	(0.011)
Public wage (year-on-year growth)					0.010		0.002
					(0.013)		(0.016)
Alt. public wage (year-on-year growth)[b]						0.001	
						(0.015)	
Minimum wage (monthly)* POST97							0.033*
							(0.010)
Minimum wage (twelve-month)* POST97							−0.007
							(0.02)
Public wage (twelve-month)* POST97							0.029
Summary statistic							(0.037)
R^2	0.47	0.47	0.49	0.49	0.49	0.48	0.50
No. observations	188	180	183	183	183	180	183

Source: Authors' calculations, based on data from INE and Central Bank of Chile.

* Statistically significant at 10 percent.

** Statistically significant at 5 percent.

*** Statistically significant at 1 percent.

a. The dependent variable is the log change in the nominal private wage. The estimation method is OLS, using monthly data for the period January 1986 to March 2002. POST97 is a dummy variable equal to one after 1997; the main effect of the POST97 dummy is included in the regression but not reported.

b. Our alternative measure of the public wage is the growth of wages in the social and communal services sector.

wages are positive but not significant at conventional confidence levels. Column 6 includes an alternative measure of year-on-year public wage growth—namely, the growth of wages in the social and communal services sector—for robustness. Finally, we allow for differential effects of minimum and public wages before and after the economic slowdown by including a post-1997 dummy (POST97) interacted with our main wage variables (column 7). The estimated coefficient on the interaction between W^{min} and POST97 is positive and significant, suggesting that the effect of minimum wages on private wages was larger in the years following the Asian crisis.

What do these results imply for private wage setting in Chile? First, they indicate that the behavior of aggregate wages is consistent with all private wages being adjusted every six months for past inflation. Since the coefficient on past six-month inflation reflects the share of workers whose wages are adjusted according to past inflation, an estimated coefficient of 0.16 on π_{6m} suggests that each month, one in every six workers receives a wage adjustment for accumulated price changes over the previous semester, independent of whether that worker participates in collective agreements. Thus the high degree of indexation prevalent in collective agreements—which by 2000 only covered about 8 percent of all workers— extends to other private wage contracts, as was the case in the 1980s.

Second, the response of wages to changes in output or unemployment is sluggish. This result is robust to changes in the length of the moving average used to measure unemployment and to the use of lagged values of unemployment or year-on-year changes in unemployment as regressors.

Third, changes in private and public sector wages are positively correlated, although this correlation is not significant at conventional confidence levels. This result holds true for both of our measures of public wage growth, and it also holds when we allow the coefficients to vary before and after the crisis.

Fourth, we cannot reject the hypothesis that minimum wages had an effect on private wages. We find that minimum wage adjustments had a significant impact on private wages in June of each year. Year-on-year growth of the minimum wages, on the other hand, had a positive but insignificant effect on private wages. This result suggests that the minimum wage in Chile only has a direct impact on wages. In other words, minimum wage changes only affect workers who are close to the legal minimum, without spilling over to other wage categories. When we include the interactions with the post-1997 dummy, we find that $W^{\min}$ increases (and is significant) in the period after 1997. This suggests that the minimum wage became increasingly binding during the economic slowdown. Our point estimates suggest that prior to 1997, close to 1.2 percent of the workforce saw their wages go up in step with the minimum wage. This share rises to 4.4 percent after 1997.

Our results suggest an important role for institutional characteristics—namely, the length and indexation of contracts and changes in the minimum wage—and a relatively small role for labor market conditions in determining private wages in Chile. Because of a series of endogeneity issues, however, all of these results suggest correlations and not causal relationships, and they should be taken as suggestive evidence rather than hard econometric facts.

The Impact of Minimum Wages: The Microeconomic Evidence

In the previous section, we used time series data on nominal wages to provide evidence that the rise in the mandated minimum wage between 1997 and 2000 had a significant impact on average wages. To corroborate these results, this section draws on microeconomic data from the national employment survey conducted by the National Institute of Statistics (INE) to assess the effect of changes in the minimum wage on the wage distribution of employment.[20] We concentrate on the distribution of salaried workers between eighteen and sixty-five years of age, who declare to have worked more than forty hours in the reference week. This group represents 72 percent of workers in this age category and 69 percent of all workers fifteen years and older (averaged for 1996–2000).

Assume, following equation 1, that the demand for labor for a worker of type i is given by

$$L_i = \frac{\alpha_i}{\mu_i}\frac{P}{W_i}Y, \tag{7}$$

where Y denotes total value added, P is the product price, and α_i is output elasticity per worker of type i, and W_i and μ_i are, respectively, the wages and markup for workers of type i. If the elasticity and the markup by type of worker remain constant across time, changes in the distribution of employment will depend only on changes in the value added, Y, and real wages, W/P_i.

Equation 7 implies that the number of employees (in logs) that earn a real wage above $\overline{W}/\overline{P}$ is equal to

$$\mathrm{pcd}\left(\frac{\overline{W}}{P}\right) = \ln\left(\sum_{i\cdot\frac{W_i}{P}\geq\frac{\overline{W}}{P}} Li\right) = y + \ln\left(\sum_{i\cdot\frac{W_i}{P}\geq\frac{\overline{W}}{P}} \frac{\alpha_i}{\mu_i}\frac{P}{W_i}\right), \tag{8}$$

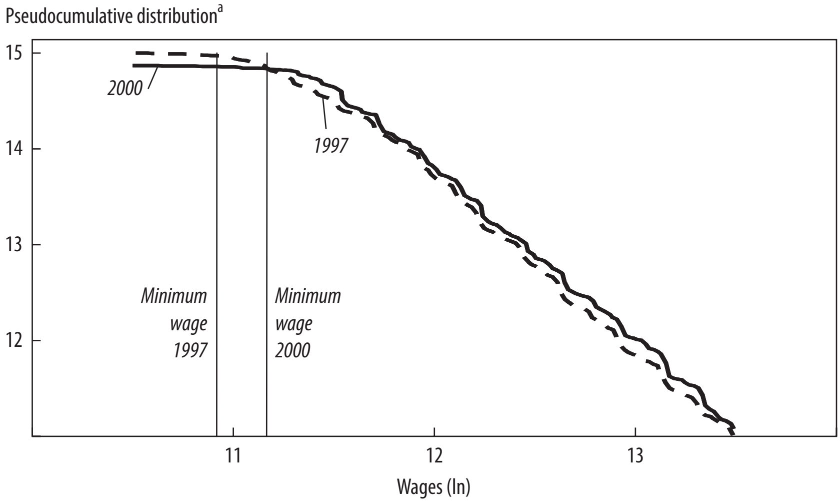
where pcd stands for pseudocumulative distribution.[21]

Figure 10 plots the number of total private sector workers (in logs) who in 1997 and 2000 received real wages (also in logs) above the level indicated in the x axis (that is, the pseudocumulative distribution). To isolate the effect of the minimum wage from the effect of changes in the value

20. The Spanish name of the survey is *Encuesta Nacional de Empleo*.

21. We refer to a pseudocumulative distribution instead of a cumulative distribution because in this case, the cumulative sum is not normalized between zero and one.

FIGURE 10. Impact of Minimum Wage on All Employment



Source: Authors' calculations, based on data from INE.
a. The number of employees (in logs) who earn a real wage above a given wage.

added, we subtract the change in the private sector value added during the period 1997–2000 in the 2000 distribution (see equation 8). The vertical lines indicate the statutory minimum wage (in logs) in 1997 and 2000.

While the two curves overlap in the upper (high wage) section of the pseudocumulative real wage distribution, they diverge in the lower section, close to the minimum wage. The 2000 wage distribution is slightly above the 1997 distribution at real wage levels slightly above the 2000 minimum wage, and it is below the 1997 distribution (and flat) below the 2000 minimum wage. The flat segment for wages below the 2000 minimum wage, which we do not observe in the 1997 pseudocumulative distribution, suggests that the rising minimum wage pushed some of those workers close to the minimum wage out of employment. The fact that the distribution in 2000 is slightly above the 1997 distribution for wages just above the 2000 minimum wage, after being flat and below the 1997 distribution for wages below the 2000 minimum, suggests that the wages of some workers where pushed up by the rising minimum wage.

If minimum wages became binding over the period 1997–2000, then the effects should be most visible among workers with low levels of experience and education, since their wages are significantly lower than the wages of more experienced, better educated workers (see table 3). Figure 11 shows the 1997 and 2000 cumulative wage distributions for workers with fewer than ten years of experience and fewer than twelve years of schooling, while figure 12 focuses on the more experienced, better educated group. For the group with low levels of experience and education, the pseudocumulative distribution in 2000 is similar to or higher than the 1997 distribution for medium and high wages, but it becomes flat at a level slightly above the 2000 minimum wage. This was not the case in 1997. This shows that a large fraction of jobs with labor productivities between the 1997 and 2000 minimum wages were directly affected by the rise in the statutory minimum wages over this period. In contrast, the 1997 and the 2000 distributions are almost identical for the more experienced, better educated group, as expected.

To estimate the share of workers that where affected by the rise in minimum wages, we calculate the fraction of workers in the 1997 distribution that earned wages between the 1997 minimum and the 2000 minimum. We account for noncompliance by subtracting the fraction of workers in 2000 with wages between the minimum wages of 2000 and 1997 (which would be zero if all employers fully complied with the minimum wage). These calculations yield sizeable estimates. The estimated share of workers directly affected by the rise in the minimum wage ranges from 2 percent for workers with high levels of education and experience to 13 percent for workers with low levels of education and experience (see table 4). For the aggregate, this percentage is 6 percent, which is on the same order of magnitude as our previous calculations of the incidence of minimum wages using the evolution of aggregate wages (4.4 percent).

We can also use this methodology to assess the effects of the rise in the minimum wage by sector of economic activity. The share of workers affected by a minimum wage hike should be larger in sectors that traditionally pay low wages than in high-wage sectors. Agriculture, construction, commerce, and manufacturing have the highest shares of low-skill, low-education workers, while the mining, utilities, and financial sectors have the largest share of skilled and experienced workers.[22]

22. The shares of workers with zero to eleven years of schooling and zero to nine years of experience in each sector are as follows: agriculture, 10 percent; construction, 9 percent; commerce, 8 percent; industry, 7 percent; transportation, 5 percent; communal services, 4 percent; utilities, 4 percent; mining, 3 percent; and financial services, 3 percent. The employment sample is defined as wage employees aged eighteen to sixty-five.

TABLE 3. Employment and Wages by Education and Experience[a]

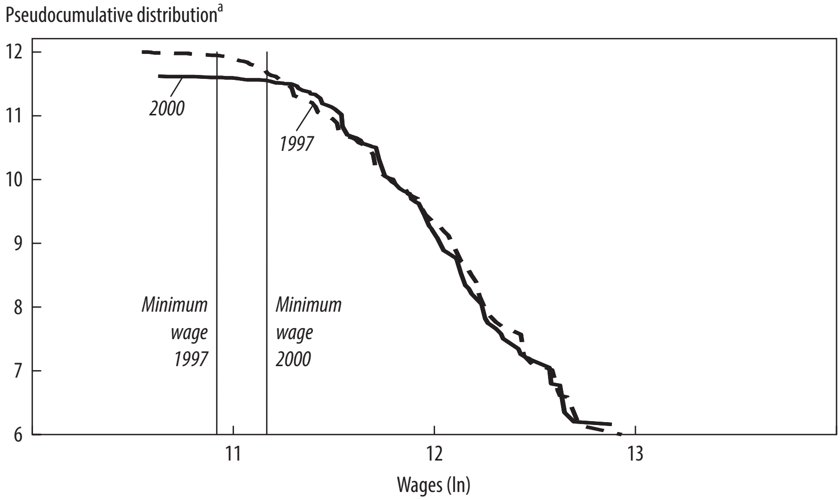
Year and indicator	*Group I*	*Group II*	*Group III*	*Group IV*
1996				
Employment	171,829	1,118,607	471,594	657,165
Share of population (%)	33	33	43	47
Real wage (pesos)	90,545	117,111	182,053	307,903
1997				
Employment	183,343	1,192,306	490,487	659,874
Share of population (%)	35	34	43	48
Real wage (pesos)	97,469	129,767	208,246	322,101
1998				
Employment	161,044	1,150,897	484,506	674,265
Share of population (%)	31	33	41	49
Real wage (pesos)	108,124	140,370	247,264	357,767
1999				
Employment	145,240	1,120,344	454,288	669,590
Share of population (%)	26	33	38	47
Real wage (pesos)	108,181	143,383	236,564	347,182
2000				
Employment	133,138	1,1419,62	458,907	663,558
Share of population (%)	26	33	36	47
Real wage (pesos)	107,624	146,111	208,515	372,155
Change, 1997–2000				
Employment (%)	–32	–4	–7	1
Share of population (%)	–30	–3	–17	–2
Real wage (%)	10	12	0	14
Average schooling	9.1	7.1	13.4	13.4
Average experience	6.2	27.8	5.1	18.8
Employment (% total employment)	7	47	19	27
Employment (% population)	37	50	43	52

Source: Author's construction, based on INE employment survey.

a. The groups are defined as follows: group I: 0–11 years of schooling and 0–9 years of experience; group II: 0–11 years of schooling and 10 years or more of experience; group III: 12 years or more of schooling and 0–9 years of experience; and group IV: 12 years or more of schooling and 10 years or more of experience.

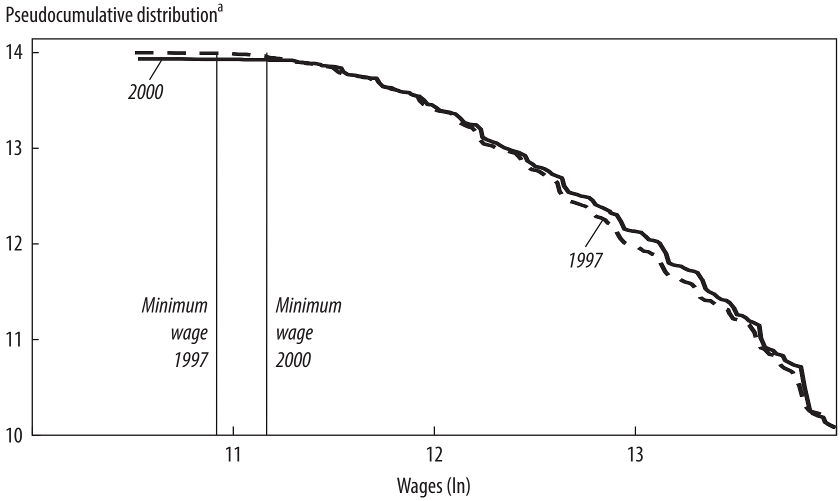
Table 5 shows the results of estimating the share of workers affected by changes in the minimum wages across industries and types of workers by applying the methodology described above. The table shows the share of workers—by economic sector and educational level—that in 1997 earned wages between the minimum wages of 1997 and 2000 (minus the share in 2000 that earned less than the 2000 minimum wage). The table confirms our previous results: we find that the estimated incidence of the minimum wage is higher for low-skilled workers than for high-skilled workers across all sectors. The table also shows that the sectors most affected by rising

FIGURE 11. Impact of Minimum Wage on Low-Skilled, Experienced Workers



Source: Authors' calculations, based on data from INE.
a. See figure 10.

FIGURE 12. Impact of Minimum Wage on High-Skilled, Experienced Workers



Source: Authors' calculations, based on data from INE.
a. See figure 10.

TABLE 4. Estimated Impact of the Minimum Wage on Employment[a]
Percent

Year	*Group I*	*Group II*	*Group III*	*Group IV*	*Total*
1997	20	14	6	3	9
2000	7	5	2	1	3
Difference	13	9	4	2	6

Source: Authors' calculations, based on INE employment survey.
a. Includes wage employees aged eighteen to sixty-five. The groups are defined as follows: group I: 0–11 years of schooling and 0–9 years of experience; group II: 0–11 years of schooling and 10 years or more of experience; group III: 12 years or more of schooling and 0–9 years of experience; and group IV: 12 years or more of schooling and 10 years or more of experience.

minimum wages are construction and agriculture, two sectors with a high incidence of unskilled workers. The magnitudes are large: in the construction sector, the rise in minimum wages affected 20 percent of the jobs of unskilled workers and 16 percent of total jobs. In agriculture, the estimated effect on total employment is 13 percent, which is lower than the effects in the construction sector because agriculture is characterized by low levels of compliance.

Taken together, the results strongly suggest that the large increase in the minimum wage between 1997 and 2000 directly affected a substantial number of workers, pushing up the wages of some at the expense of others who lost their jobs. These effects were stronger in agriculture and construction, a result of the high incidence of low-wage jobs in these sectors. As a final test, we assess whether the increase in minimum wages was more binding and therefore had larger effects in some regions than in others. Minimum wage increases are determined at the aggregate level, but wage distributions vary regionally. The decline in employment generated by a national minimum wage hike should be higher in regions with a high incidence of low-wage employment. To verify this, we estimate the following specification:

$$\Delta l_{rj} = \alpha + \beta\lambda_{rj} + D_r + \epsilon_{rj}, \tag{9}$$

where Δl_{rj} is the change in employment between 2000 and 1997 in region r and sector j, and λ_{rj} is the share of workers in region r and sector j whose wages in 1997 were above the 1997 minimum wage and up to the 2000 minimum wage, minus the share of workers whose wages in 2000 were below the 2000 minimum wage and above the 1997 minimum wage.

TABLE 5. Estimated Impact of Minimum Wages on Employment, by Sector and Skill Level[a]
Percent

	0–11 years of schooling		*12 years or more of schooling*		*Total*		
Sector	*Employment fraction*	*Employment growth*	*Employment fraction*	*Employment growth*	*Employment fraction*	*Employment growth*	*Evasion*
Agriculture	13.4	2.1	*10.2*	*18.2*	13.1	4.1	7.4
Mining	*3.9*	*–23.7*	*0.5*	*–19.5*	2.1	–21.4	0.2
Industry	4.1	–15.9	1.6	–15.4	2.9	–15.6	1.3
Electricity, gas, water	*5.0*	*–22.2*	*1.6*	*0.4*	*3.0*	*–7.5*	0.7
Construction	19.4	–38.4	8.2	–33.1	16.2	–36.9	2.2
Commerce	5.5	0.1	2.5	7.6	3.7	4.6	1.9
Transportation	4.2	4.6	3.6	15.0	3.9	9.9	2.5
Financial services	*5.6*	*36.5*	0.7	1.8	1.8	11.1	1.2

Source: Authors' construction, based on INE employment survey.

a. Employment fraction is our measure of potential minimum wage impact, constructed as the difference between the fraction of waged workers whose payment remains between the 1997 minimum wage and the 2000 minimum wage in 1997 and 2000. Employment growth was calculated between 1997 and 2000. Evasion was estimated as the fraction of waged workers whose wage in 2000 is less than the minimum wage in that year but above the 1997 minimum wage. Salaries was deflated using sectoral data. Numbers in italic are based on fewer than 500 observations.

Table 6 presents our results. A negative coefficient on λ indicates that those region-sectors where more people were affected by the minimum wage increase were the region-sectors where more jobs were lost due to the minimum wage policy. The number of individual observations for some region-sectors is too small for our exercise, so we aggregate some sectors to increase the number of observations, such that we end up with 102 region-sector pairs. The sign and statistical significance of the coefficient is maintained when we control for the 1997–2000 change in sector GDP, although this reduces the magnitude of the coefficient. Since our earlier results suggest that the minimum wage hikes only affected workers between the 1997 and 2000 minimum wage and not workers who were higher up in the wage distribution, the coefficient on the share has a very intuitive interpretation: it represents the share of workers who in 1997 received wages between the two minimum wages and who lost their jobs as a result of the minimum wage hike. The estimates in column 2 indicate that, on average, as many as 57 percent of the workers in this wage category could have lost their jobs in 2000.

These results confirm our previous findings. The large predetermined minimum wage increase led to a reduction of jobs, which was highest

TABLE 6. Estimates of the Effect of the Minimum Wage Hike on Employment, by Region and Sector[a]

Explanatory variable	*(1)*	*(2)*
Share (λ)	−0.951	−0.572
	(2.98)***	(2.06)**
GDP growth		0.026
		(2.06)**
Summary statistic		
R^2	0.18	0.26
Regional fixed effects	Yes	Yes
No. observations	102	102

Source: Authors' calculations, based on INE employment survey and data from the Central Bank of Chile.
* Statistically significant at 10 percent.
** Statistically significant at 5 percent.
*** Statistically significant at 1 percent.
a. The dependent variable is the change in employment between 2000 and 1997. The sample covers thirteen regions and nine sectors, resulting in over twenty observations per cell (in 1997 and 2000). Share (λ) denotes the share of workers that in 1997 had wages above that year's minimum wage and up to the 2000 minimum wage level, minus the share of workers that in the year 2000 had wages below the 2000 minimum wage and above the 1997 minimum wage. Absolute value of *t* statistics in parentheses; robust standard errors.

among unskilled workers, sectors intensive in unskilled labor, and region-sectors with a predominance of low-wage jobs.

Summing Up

In this paper we argue that the slowdown in employment growth in Chile after 1997 is the result of a negative aggregate demand shock in the presence of substantial downward rigidity of real wages. We find no evidence of a structural break in the demand for labor: falling employment and its slow recovery can be fully explained by the path of wages and output.

When viewed in relation to other economies that experienced large output reductions in the 1990s, Chilean wages appear to be relatively rigid. This wage rigidity is likely an important factor behind the slow recovery of employment during the mild expansionary period that Chile experienced after 1999.

We explore a series of possible explanations for this rigidity and find evidence supporting two of them: rising minimum wages and the prevalence of long-term price-indexed wage contracts. We show that the minimum wage became increasingly binding in the years following the Asian crisis. Indeed, our estimates indicate that close to 6 percent of workers where affected by the rise in the minimum wage over the period 1997–2000. This percentage doubles in the construction sector, which is a large employer of unskilled workers. We also find that regions and sectors with a predominance of low-wage jobs were more affected by the minimum wage hike than regions and sectors offering better pay.

We also find evidence that contracts in Chile are long and that indexation is prevalent. After a demand shock, the contract structure introduces real rigidities that produce an adjustment through quantity instead of price. Finally, we argue that the drastic drop in labor demand after the Russian and Brazilian crises did not translate one to one into higher unemployment because labor participation in Chile is cyclical. If labor participation had remained constant during this period, unemployment would have risen by 4.1 percentage points instead of the 3.3 percentage points observed between 1997 and 2000. This procyclical labor force is another factor behind the slow fall in unemployment in recent years: rising employment has been accompanied by rising labor force participation.