Minimum Wage and Productivity: Evidence from Chilean Manufacturing Plants

ROBERTO ÁLVAREZ
University of Chile

RODRIGO FUENTES
Pontificia Universidad Católica de Chile

I. Introduction

Economic policies and institutions—such as market regulations, trade and financial openness, taxes, and other distortions—have been found to be relevant for explaining differences in total factor productivity (TFP) across countries (Parente and Prescott 2000; Nicoletti and Scarpetta 2003). More specifically, the literature indicates that labor market regulations that increase firing and hiring costs at the firm level could be an important determinant of productivity at the macroeconomic level (Hopenhayn and Rogerson 1993; Lagos 2006; Petrin and Sivadasan 2006).

In this paper, we empirically analyze the effects of a combination of minimum wage increases and firing costs on firm-level TFP. Our definition of TFP goes beyond the parameter of the production function. By TFP, we mean the ratio of output to a combination of inputs used given the technology available and the policy constraints to use that combination. In other words, there is a crucial distinction between TFP as a theoretical concept (like a parameter of the production function) and TFP as an empirical/residual concept (i.e., measured TFP). A change in the legal minimum wage affects the relative price of low skilled labor, but in the absence of adjustment costs, this would be absorbed by an optimal variation in input demands (substitution and scale effect). However, in the presence of high severance payments, it could be costly for firms to adjust their workforce to the new optimal level. Then even mini-

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mum wage changes do not affect the technological TFP; the actual TFP will be lower during the adjustment because the firm would be using a larger amount of labor compared to the case without adjustment costs. We expect that this effect would be stronger in sectors with higher exposure to minimum wage, specifically unskilled-labor-intensive industries.

The theoretical framework for expecting a negative relationship between labor regulations and productivity comes from a pioneer study by Hopenhayn and Rogerson (1993) that analyzed the effect of policies that create a tax on job destruction. The calibration of the model shows that this tax generates important negative effects on aggregate productivity and welfare. In the same vein, Lagos (2006) provides a model where observed TFP depends on the underlying distribution of shocks and labor market characteristics. Specifically, he shows that taxes affecting job destruction (such as, employment subsidies and firing costs) reduce observed TFP due to changes in the composition of labor employed. In contrast, Autor, Kerr, and Klugeret (2007) argue that firms try to offset the effects of labor regulations by screening new hires more stringently, leading to a favorable compositional shift in workforce productivity. In such a case, an increase in minimum wage would increase productivity. In this paper, we follow the approach presented by Wickens (2008), which allows us to show how—in the presence of adjustment costs—an increase in minimum wages may reduce observed firm productivity.

On the empirical side, our framework is related to the microeconomic misallocation mechanism studied by Hsieh and Klenow (2009). Using firm-level data for China and India, they find a large effect of resource misallocation on firms' TFP and emphasize labor-market regulation as one of the possible causes of this misallocation. In our empirical testing, we use plant-level data for the Chilean manufacturing industry during the period 1992–2005 and a difference-in-differences approach, exploiting differences in minimum wages over time and in industry exposure to changes in these labor costs. Given its relevance, we calculate measured productivity using three methods: (i) TFP, measured as in Olley and Pakes (1996) and Levinsohn and Petrin (2003); (ii) estimation of TFP using a nonparametric method, with labor share calculated from the data, which assumes constant return to scale (van Biesebroeck 2007); and (iii) a measure of labor productivity calculated as real value added per worker. For all of our estimations, we check the robustness of our results using different productivity measures and specifications.

Chile is an appropriate setting for conducting this analysis for two reasons. First, the minimum wage was greatly increased at the end of the 1990s. Second, after being a successful example of growth during the 1986–97 period (TFP grew at 3%), the aggregate productivity growth rate dramatically de-

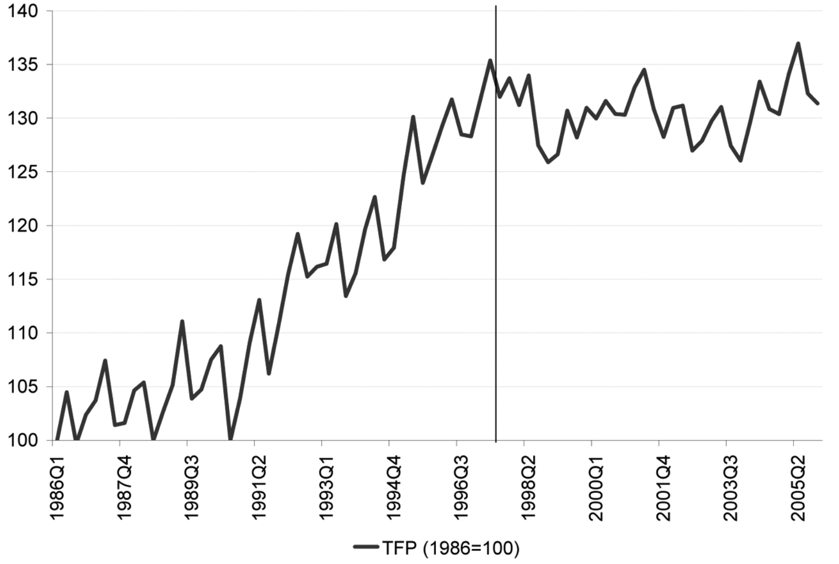

Figure 1. Total factor productivity index, estimated as the residual of an aggregate production function using capital and labor. Source: Fuentes, Gredig, and Larraín (2008). A color version of this figure is available online.

clined over the following 10 years (fig. 1). We analyze whether the increase in the minimum wage can explain this productivity slowdown. The firing costs existed before the start of the aggregate TFP growth slowdown but may not have become binding until the economy experienced a negative shock. A combination of the 1998 international financial crisis plus the significant increase in the minimum wage created the needed negative shock.[1] This idea is consistent with the evidence presented by Blanchard and Wolfers (2000) for member countries of the Organization for Economic Cooperation and Development, suggesting that observed changes in labor outcomes—for example, increased unemployment—are the result of interactions between economic shocks and labor institutions.

Most of the literature on the minimum wage has focused on its impact on wages and employment, but much less is known about how productivity can be affected by increased labor costs. Most closely related to our empirical results is the work by Autor et al. (2007), who find a negative effect of employment protection on productivity for firms in the United States. They argue that this negative impact occurs because the adoption of dismissal protection

[1] De Gregorio (2007) discusses some alternative hypotheses for the productivity slowdown of the Chilean economy, including potential problems of TFP measurement and resource reallocation from high- to low-productivity growth sectors.

alters production choices and causes employers to retain unproductive workers. Similarly, Petrin and Sivadasan (2006) construct a dynamic model to illustrate how job security affects economic efficiency. Using Chilean data, they find that increases in firing costs reduce productivity by increasing the gap between workers' marginal revenue product and wages. We contribute to this literature by directly estimating the impact of changes in minimum wage on plants' productivity and by using a particular episode of national policy changes that established in-advance increases in the minimum wage for a 3-year period.

In the case of Chile, one of the challenges for conducting this type of empirical study is that the minimum wage is established at the national level. Unlike with other countries, we cannot exploit regional variations to identify the causal effect of policy changes. These differences in the timing of regulation changes across states have been used previously to identify the effects of policy changes in countries like the United States and India.[2] Our identification strategy is thus based on the idea that the effect of regulations depends on the exogenous exposure to regulatory changes. In the specific case of minimum wage changes, we identify the differential effect on plants depending on the industry's exposure to increased labor costs. To measure this exposure, we use the ratio of unskilled to skilled labor for each industry in the first year of the sample. The reason for using the first year of the sample is to avoid potential endogenous response of this ratio to later changes in the minimum wage over time.

Our results show that the minimum wage increase at the end of the 1990s is partially accountable for the slowdown in firm productivity, reflected in the aggregate TFP evolution. Specifically, our estimates indicate that a real increase of about 22% in the minimum wage during the period 1998–2000 reduced TFP by 2% in industries with fewer unskilled workers and 3.7% in industries with more unskilled workers. These results are robust to different measures of productivity, to the inclusion of other sector- and firm-specific variables, and to sample selection and endogeneity issues.

The paper is structured as follows: Section II discusses the relationship between minimum wages and productivity in the presence of firing costs. Section III describes the main policy changes in the Chilean economy during the period under study. Section IV explains the methodology used to identify the effect of minimum wages on productivity. Section V describes the data set, while Section VI presents our econometric results. Section VII shows several robustness checks, and Section VIII concludes.

[2] Autor et al. (2007) use changes in employment protection across American states, and Aghion et al. (2008) use changes in entry regulation and labor market policies across states in India.

II. Conceptual Framework

This section conceptually addresses why an increase in the legal minimum wage, under the presence of relevant costs of worker dismissals, may reduce firm productivity. There are several arguments in the literature indicating that higher minimum wages may be associated with a reduction or increase in productivity (TFP). In this section, we present first a simplified version of the labor adjustment cost model developed in Wickens (2008). We use this framework—with minor variations—with the purpose of showing how an increase in minimum wage, in the presence of adjustment costs, reduces measured TFP. Second, we add additional arguments on this issue linking minimum wage and productivity.

We consider that each firm maximizes the present value of profits, taking the prices of output and factors as a given. The objective function is

$$V_t = \sum_{s=0}^{\infty} \frac{\Pi_{t+s}}{(1+r)^s}.$$

Where $\Pi_{t+s} = A_{t+s} f(k_{t+s}, l_{t+s}) - r_k k - w_{t+s} l_{t+s} - C(l_{t+s} - l_{t+s-1}), A_{t+s} = y/[f(k_{t+s}, l_{t+s})]$ is total factor productivity, $C(l_{t+s} - l_{t+s-1})$ is the labor adjustment cost, $f(k_{t+s}, l_{t+s})$ represents a production function with positive and decreasing marginal productivity of factors and positive cross derivative between capital and labor, w_{t+s} is the wage rate, r_k is the rent of capital, and r is the constant discount rate.

The problem can be written as

$$\max V_t = \sum_{s=0}^{\infty} \frac{A_{t+s} f(k_{t+s}, l_{t+s}) - r_k k_{t+s} - w_{t+s} l_{t+s} - C(l_{t+s} - l_{t+s-1})}{(1+r)^s}.$$

The first-order conditions of this problem are

$$\frac{\partial V_t}{\partial k_{t+s}} = A_{t+s} f_k - r_k = 0 \tag{1}$$

and

$$\frac{\partial V_t}{\partial l_{t+s}} = \frac{A_{t+s} f_l - w_{t+s} - C'_t}{(1+r)^s} + \frac{C'_{t+1}}{(1+r)^{s+1}} = 0. \tag{2}$$

Assuming a quadratic cost of adjustment $C = (\theta/2)(l_t - l_{t-1})^2$, equation (2) becomes

$$\theta(l_{t+s} - l_{t+s-1}) = \frac{\theta}{1+r}(l_{t+s+1} - l_{t+s}) + A_{t+s} f_l - w_{t+s}, \tag{3}$$

where the parameter θ captures the relevance of adjustment costs. When θ is zero, firms hire the optimal level of labor, which is obtained equating the wage rate with the marginal productivity of labor.

For each level of capital stock, k, equation (3) is a difference equation in l, thus iterating forward on equation (3), we obtain an expression for the variation of the labor amount hired by the firm:

$$\theta(l_{t+s} - l_{t+s-1}) = \sum_{s=0}^{\infty} \frac{A_{t+s} f_l - w_{t+s}}{(1 + r)^s}. \tag{4}$$

Equation (4) states that an actual change in the amount of labor depends on the future differences between the marginal productivity of labor and the wage rate. The adjustment will be slower the larger parameter θ is. For instance, an increase in the binding minimum wage will reduce the level of employment, but while it adjusts to the optimal level of labor, the firm will be using a nonoptimal capital/labor ratio. Note that this conclusion still holds in the presence of adjustment costs for capital.[3]

This simple framework also illustrates the heterogeneous impact of minimum wages across firms, heterogeneity that we explore in our empirical approach. Firms with high productivity (and paying high wages) will not be affected by the increases in compulsory minimum wages. In such a case, w_{t+s} is higher than $w_{\min}$ and—according to equation (4)—$A_{t+s} f_l = w_{t+s}$ and $L_{t+s} = L_{t+s-1}$.

But for some firms, the mandatory minimum wage will be binding. This is the case with low-wage firms (with low-skilled workers), where an unexpected increase in minimum wages, such that $w_{t+s+1} = w_{\min} > w_{t+s}$, will generate an employment reduction. The higher θ is, the lower this will be. In other words, even when the parameter of the production function A (TFP) does not change, the actual TFP level will be lower during the adjustment period since firms will be using a larger amount of labor compared to the case without adjustment costs and binding minimum wages.

Taking the definition of TFP made by Parente and Prescott (2000, 59)—"the maximum output that can be produced given not only the technology constraints, but also the constraints on the use of technologies arising from policies"—we compare the level of TFP without and with adjustment costs. Let l^* be the optimal amount of labor when minimum wage increases and there is no adjustment cost and l' be the amount of labor when the adjustment cost exists. Then, we have that, given $l^* < l'$, the firm's TFP will be lower than its optimal level according to

[3] A similar argument is made by Bond and Söderbom (2005) using a related framework.

$$A_t = \mathrm{TFP}_t^* = \frac{y_t}{f\left(k_t, l_t^*\right)} > \frac{y_t}{f\left(k_t, l_t'\right)} = \mathrm{TFP}_t'.$$

This difference between the two TFP levels will be 0 when the amount of labor adjusts to its optimal level, that is, when $\theta = 0$. In general, θ may be associated with several labor market regulations such as mandatory benefits and job security rules that introduce constraints at the plant level.[4] All these regulations generate adjustment costs when firms want to change employment levels, and it may reduce their productivity.

Thus, assuming that firms were optimally choosing labor before the minimum wage increases ($t - 1$), TFP was the maximum to be reached by the firm. If in time t minimum wages are increased, we expect that for affected firms employment will be higher than its optimal level and TFP lower than the maximum. This is $\mathrm{TFP}_t < \mathrm{TFP}_{t-1}$ and the main implication that we test in the empirical section of this paper.

Using a similar argument, Autor et al. (2007) consider two classes of models to analyze the relationship between mandatory dismissal protection and productivity. In these models, a legal minimum wage may play an additional role. First, they consider a competitive model where workers may value the dismissal protection less than the marginal cost of provision, resulting in mandatory dismissal costs driving the marginal productivity of labor below its wage. If the wage rate does not adjust because of the presence of downward rigidity (like, in our case, the minimum wage), the difference between the private and the social marginal productivity will generate a deadweight loss. Since the adjustment cost is paid when a worker is dismissed, the firm will compare the present value of the deadweight loss and the dismissal cost (this is similar to the argument given in eq. [4]). If the worker values the dismissal protection benefit more than the marginal cost of provision, a firm and worker will continue the contractual relationship as long as the deadweight loss is lower than the present value of the productivity shortfalls. Therefore, a firm will keep unproductive workers, reducing labor productivity. But the authors also argue that firms—anticipating the dismissal costs—might be more rigorous in the screening process in future hiring, improving average labor productivity. As a result, the net effect of increasing labor cost on productivity is ambiguous.

The same authors consider the equilibrium unemployment model based on Mortensen and Pissarides (1994). In this context, the dismissal protection has

[4] These regulations are aimed to protect workers in the case of accidents or health problems and to diminish the costs of being laid off, as well as balancing workers' bargaining power when negotiating with firm owners. However, benefits for employed workers may also have negative effects, such as lower protection for the unemployed (Freeman 1993).

two effects. On one hand, it reduces the threshold productivity at which firms are willing to dismiss a worker (it increases the cost of having a vacant position), thus reducing labor productivity. On the other hand, it increases the threshold productivity at which firms are willing to hire a worker, increasing productivity at the firm level in the future.[5] Similar to the predictions of the competitive model, the overall effect of higher labor costs on productivity is ambiguous, and an empirical analysis is needed to settle the question.

Hamermesh and Pfann (1996) analyzed different types of adjustment-cost functions and reviewed empirical studies on the topic. Their main result is that hiring and firing costs prevent the firm from fully adjusting employment to exogenous profitability shocks, which in our case would be associated with a minimum wage increase. Similarly to Parente and Prescott (2000), this creates a gap between the optimal level of employment and the actual level, which in turn generates a lower level of TFP.

The models analyzed in this section assume that a worker's productivity is independent of the wage rate. In a context where the worker's effort depends on the wage rate, as in the efficiency wage model (see Yellen 1984), an increase in the minimum wage may increase effort and thus productivity. Once again, the answer to how all these labor regulations affect productivity should be found on the empirical side.

III. Economic Policy Changes

In this section, we discuss the main policies and institutional changes regarding Chile's labor markets during the period 1992–2005. Most Chilean economic policies are uniform across regions and industries, including labor market regulations. Nevertheless, policy changes may have different effects across sectors because of their different exposure to regulations and other idiosyncratic characteristics. There are several reforms in the Chilean labor market that could have impacted firms' inputs allocation, including the minimum wage.

The minimum wage has been one of the main tools for labor market regulation in Chile. Figure 2 shows the evolution of the real minimum wage (real Wmin), deflated by the consumer price index, and the minimum wage as a fraction of unskilled worker wages (Wmin/Wu) calculated by Beyer (2008). The real minimum wage increased significantly during the period 1992–2005. The total growth was 72%. It is interesting to note that compared to the average wage received by unskilled workers, this increase was actually only 27% over the same period. As shown in figure 2, this occurred mainly during a short period of time. These figures suggest that, especially since 1998, the increase in

[5] For more details, see Autor et al. (2007).

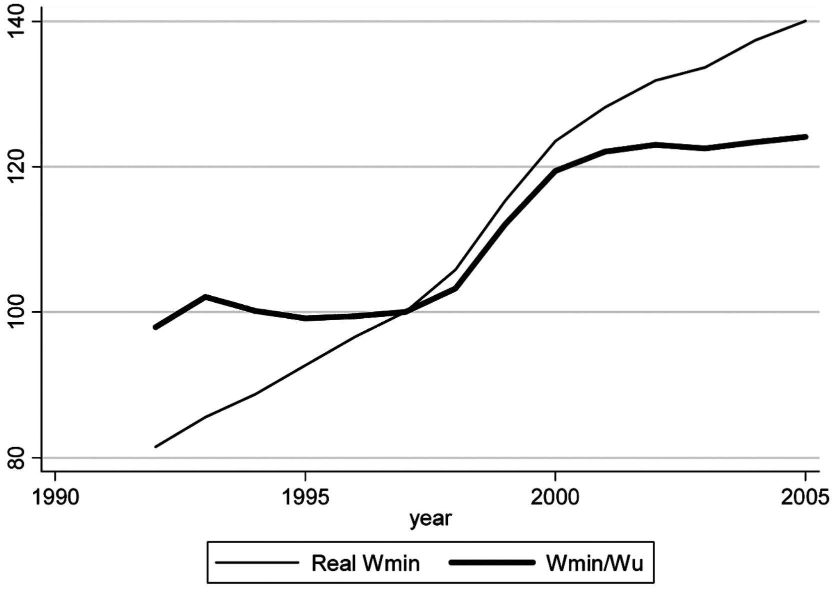

Figure 2. Evolution of real minimum wage and relative to unskilled wage. Real Wmin is an index of real minimum wage, that is, the nominal minimum wage deflated by the consumer price index; Wmin/Wu is the ratio between minimum wage and the average unskilled wage. Source: Beyer (2008). A color version of this figure is available online.

minimum wage could have been more binding for those plants that are more reliant on unskilled labor.

The sudden rise in the minimum wage took place between 1998 and 2000, when the minister of finance implemented a predetermined increase over 3 years.[6] This was in contrast to previous periods when minimum wage increases were planned on a yearly level. This change is important for our identification strategy because average annual increases may be affected by economic conditions, such as the evolution of TFP and other macroeconomic variables. We mainly exploit the exogenous increase in minimum wage implemented in 1998. As shown in figure 3, the real increase in minimum wage was, on average, 7.3% in those 3 years, well above the increase in preceding and succeeding years. During the beginning of the 1990s, the minimum wage increased, on average, 4.2% per year, and it increased 2.5% per year during the period 2001–5.

The authorities thought that establishing a 3-year schedule would generate more predictability in labor regulations. It also allowed them to avoid difficult negotiations with trade unions, the business sector, and congressional repre-

[6] There has been discussion in Chile on how these increments in minimum wage have affected unemployment, which remained very high until about 2005. However, there is not much empirical evidence on this issue, except in the work of Cowan et al. (2005).

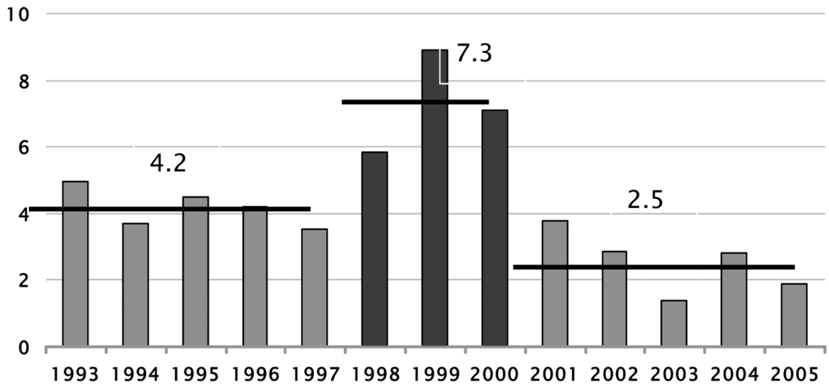

Figure 3. Real minimum wage growth rate for 1993–2005, with author elaboration based on Beyer (2008). A color version of this figure is available online.

sentatives each year. However, they did not predict the negative effects of the Asian crisis during 1998 and 1999 and that the minimum wage increased far over increases in labor productivity and unskilled wages. As discussed by Velásquez (2009), the gap between labor productivity and minimum wage growth was 19.4% by 2002. In the same year, the minimum wage was about 65.4% of the overall average wage and 90% of the average wage of unskilled workers in the construction sector, compared to 50% in 1995.

A second aspect of labor regulation is social security contributions. According to Lora (2001), social security contributions in Chile dropped from 25% to 21% in 1994. However, they increased by 3% in 2002 mainly due to a law that established an unemployment insurance mechanism. We complement social contributions with a job security index developed by Heckman and Pagés (2000) and updated for Chile by Pagés and Montenegro (2007) that combines information on notice periods, compensation for dismissal, the likelihood that firm difficulties would be considered a justified cause of dismissal, and how much the severance payment would be in such a case.

The following index, described in detail by Pagés and Montenegro (2007), measures the expected cost of dismissing a worker in terms of monthly wages:

$$\mathrm{Index}_t = \sum_{i=1}^{T} \beta^i \delta^{i-1}(1-\delta)\left(b + a\mathrm{SP}_{t+i}^{jc} + (1-a)\mathrm{SP}_{t+i}^{uc}\right),$$

where β is the discount factor, which is assumed to be equal to 0.92; δ is the probability of keeping the job equal to 0.88; $(1-\delta)$ is the probability of losing the job; $\delta^{i-1}(1-\delta)$ is the probability of a worker losing the job after i periods in the same job; b is the advance notice cost equal to 1; and a is the probability that a court will declare the dismissal justified, which is equal to 0 during the period 1992–98 and equal to 1 in the period 1999–2002.

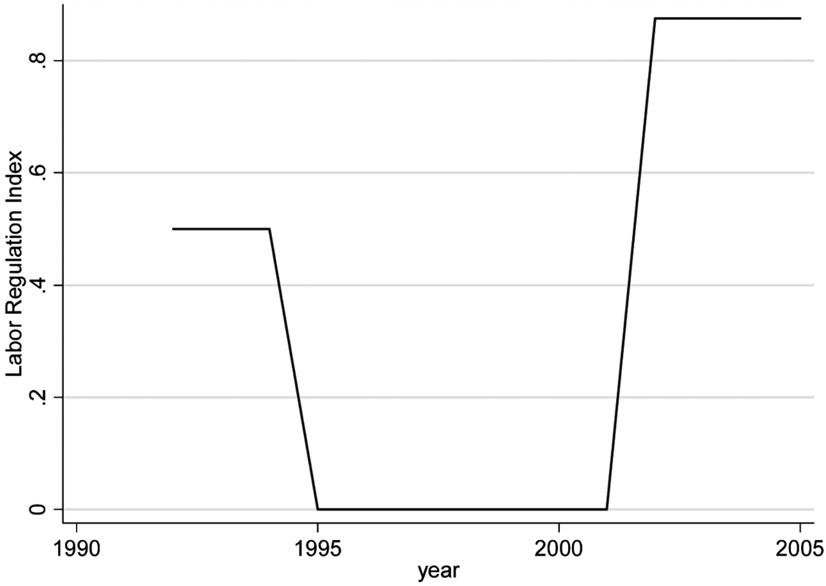

Figure 4. Labor regulation index, with own calculations based on Pagés and Montenegro (2007). A color version of this figure is available online.

We define SP^{jc} as the tenure related to severance payments under justified cause for dismissal and SP^{uc} as the tenure related to severance payments without justified cause for dismissal. Pagés and Montenegro (2007) calculate this index for Chile until 1998. The severance payments law was changed on December 1, 2001, which increased the penalties paid by firms in case the court decided that the cause for dismissing a worker was unjustified; this was a de facto raise in the cost of dismissing a worker. That fine increased from 20% to a range between 30% and 100% of severance payments, depending on the fault. Additionally, the courts made it more difficult to prove justification for worker dismissal. To include this in the index, we use the same parameters as Pagés and Montenegro (2007) but use the maximum value of the fine incorporated in the 2001 legislation.

We compute an overall labor regulation index using both indicators: social contributions and the labor protection index. We standardize these indicators, taking the value 0 when regulation is less severe (the minimum value) and 1 when it is more severe (the maximum value).[7] Thus, the labor regulation index is the simple average of both standardized indicators, as shown in figure 4.

During this period, other policy reforms may have affected the TFP's evolution. As previously suggested by Bergoeing, Hernando, and Repetto (2006), a reduction in trade barriers and financial development may have played a

[7] We use the standardization $y^s = (y - y^{\min})/(y^{\max} - y^{\min})$.

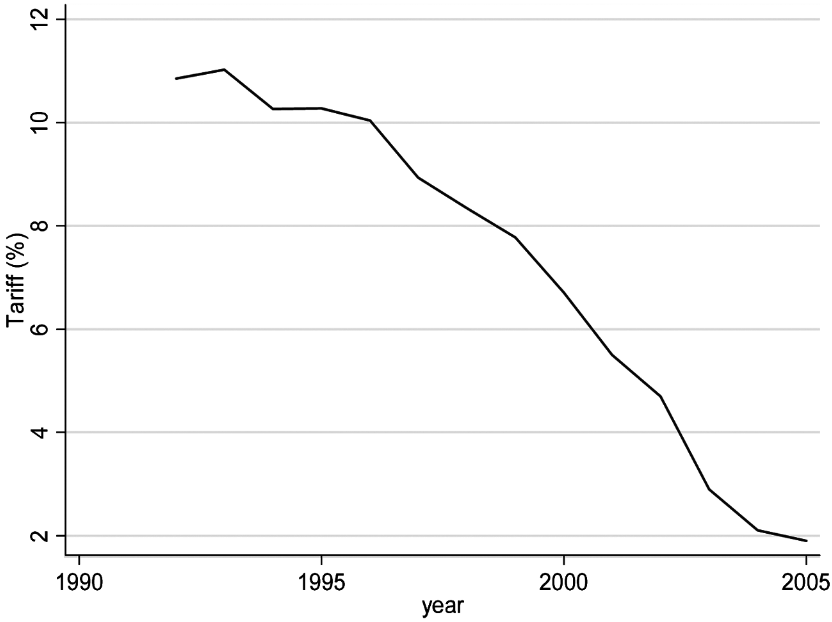

Figure 5. Effective average tariff rate, with own calculations based on Becerra (2006) and Bergoeing et al. (2006). A color version of this figure is available online.

role in explaining the evolution of productivity in Chile. Since 1979, Chilean import tariffs were uniform across sectors with few exceptions, such as price bands for specific crops and additional taxes on some luxury goods and alcoholic beverages. However, uniform tariffs changed during the 1990s due to free trade agreements with other countries. Therefore, the effective average import tariff decreased sharply from 10.9% to 1.9% during this period, as shown in figure 5.[8]

To analyze the potential effects of financial development on productivity, we consider private credit by tracking deposits in banks compared to gross domestic product (GDP; Beck, Demigurc-Kunt, and Levine 2000). Figure 6 shows the evolution of this variable over the sample period. Note that credit to the private sector expanded continuously since 1992, until contracting at the end of the sample period. Between 1992 and 2005, private credit over GDP increased from 46% to 67%.

IV. Empirical Methodology

The uniformity of most policies across regions and industries in Chile allows for a study of the effects of time-series variation in policies rather than cross-

[8] Becerra (2006) estimated the effective average tariff for the period 2000–2006. For the period 1992–2000, the estimation is from Bergoeing et al. (2006). Figure 5 merges both time series.

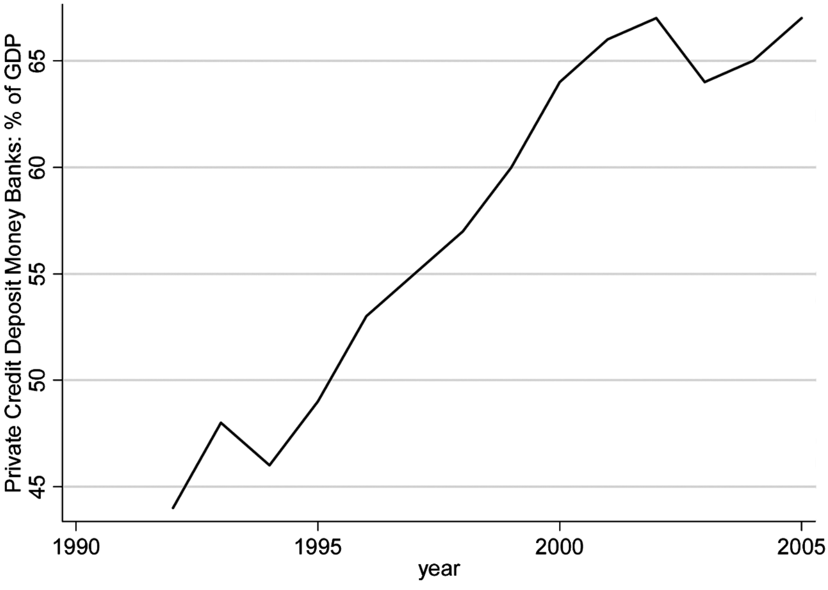

Figure 6. Financial development, with credit to private sector over gross domestic product taken from Beck et al. (2000). A color version of this figure is available online.

industry or cross-regional variation. Thus, the identification strategy used in this paper is a variant of the difference-in-differences approach developed by Rajan and Zingales (1998) and adapted to plant-level data. The effects of specific regulations are identified using plausible exogenous industry and plant characteristics to measure their ex ante exposure to these regulations.

The general specification for the plant-level estimation based on the differential effect of regulations on productivity is given by

$$y_{ijt} = \alpha_i + \eta_t + \phi y_{it-1} + \beta x_{it-1} + \lambda \log(\mathrm{WM})_{t-1} * U_{j0} + \delta Z_{t-1} * W_j + \varepsilon_{ijt},$$

where y_{it} is total factor productivity (in logs), i denotes a plant, j denotes a sector, and t denotes a year. The x_{it} is a vector of plant characteristics, WM is the real minimum wage, and U_{j0} is a measure of an industry's exposure to changes in minimum wages at the initial year of the sample period. The Z is a vector of other time-varying variables that can affect TFP differentially depending on industry characteristics denoted by W. The vector Z contains three time-varying variables described in Section III: the labor regulation index, tariffs, and financial development. All of these three variables vary overtime but not across industries.

The equation includes the lagged value of productivity as explanatory variable. There are theoretical and empirical justifications for this choice. Theoretically, based on the conceptual framework described in the previous section, it can be argued that changes in minimum wages affect TFP not only contemporaneously but, due to the adjustment costs, also over time. In this dynamic specification, the short-run effect of changes in minimum wage are given by ϕ and in the long-run by $\phi/(1-\alpha)$. With instantaneous adjustment ($\alpha = 0$), both effects should be equal.

Empirically, most of the microeconomic literature shows that productivity shocks tend to be persistent (Baily, Hulten, and Campbell 1992; Foster, Haltiwanger, and Krizan 2002; Foster, Haltiwanger, and Syverson 2008). Moreover, the methodologies for estimating production functions assume that productivity follows a first-order Markovian process (e.g., Olley and Pakes 1996; Levinsohn and Petrin 2003).

The omission of lagged productivity generates serious problems in estimations. In fact, omitting this variable in a fixed-effects OLS estimation makes our minimum wage variable not significant. This is explained for two reasons. First, both variables are correlated because both are measured at $t-1$ and related as our equation states. Second, and more importantly, the existence of adjustment costs indicates that the impact of minimum wages is not instantaneous. To capture these dynamics effects, the inclusion of the lagged dependent variable is required.

It is well known that dynamic panel estimation with fixed effects suffers from endogeneity, even in the presence of strictly exogenous covariates. In this case, the endogeneity is generated by the introduction of lagged productivity with firm fixed effects, which is known as the Nickell bias problem (Nickell 1981). In order to solve this problem in the estimation of an autoregressive model with fixed effects, Blundell and Bond (1998) propose combining the moment conditions of its level and first-difference forms. They suggest applying a system generalized method of moments (GMM) estimator, using the lagged values of the endogenous variables as instruments in the difference equation and the first difference of the endogenous variables in the level equation.

The key element of our empirical exercise is the identification of the effect of minimum wage changes. This is captured by the interaction between the minimum wage and exposure to labor cost increases (U_{j0}).[9] Our identification strategy is based on the idea that a minimum wage increase will have a larger negative effect on those industries that use a larger proportion of unskilled workers. This exposure is calculated as the industry median of the ratio of blue-collar workers to white-collar workers (measured also in logs) in the first year of the

[9] We also estimate the model using dummy variables for the years when the minimum wage increased greatly, but it fails to capture any negative impact on productivity. Our interpretation is that actual change in minimum wage is the relevant variable.

sample.[10] By measuring this variable at the beginning of the period, we can use it as a predetermined characteristic of the industry that is not affected by changes in the minimum wage over time.

For labor regulations, we exploit the identification strategy used by Micco and Pagés (2006). They use data of exogenous differences in intrinsic volatility across industries to test whether the negative effects of more stringent employment protections are more pronounced in more volatile industries. In this paper, we use the same data as Micco and Pagés (2006)—labor turnover rates for three-digit International Standard Industrial Classification (ISIC) industries in the United States—to measure intrinsic volatility and see whether Chilean labor regulations have had differential effects across industries.

In the case of trade reforms, we introduce an interaction term between the average tariff and a measure of trade liberalization exposure. We measure industry-specific trade exposure as the ratio of imports plus exports over output. As we do not have a completely exogenous measure of trade exposure, we use information for each three-digit industry at the beginning of the period.

For financial development, we identify its differential effect using an interaction term between our proxy variable (private credit over GDP) with a measure of external finance dependence developed by Rajan and Zingales (1998). This variable is computed as the fraction of capital expenditures not financed with cash flow operations, using information for US firms by three-digit ISIC industries. It is assumed that this indicator reflects technological reasons for why some industries depend more on external financing than others.

V. Data Description

Our analysis is mainly based on information for Chilean manufacturing plants covering the period 1992–2005. The National Annual Survey of Manufactures (ENIA) collects information for more than 5,000 plants and contains data on several variables such as sales, output, employment, wages, exports, foreign ownership, and other characteristics for each plant with at least 10 employees. All monetary variables were converted to constant pesos using three-digit ISIC-level price deflators. In addition, plants are classified according to ISIC Rev. 2.

Table 1 shows the number of plants by year. There are approximately 6,000 plants at the beginning of the period but only about 5,300 plants in 2005. Table 2 presents the distribution of plants by industry for 2005.[11] More than

[10] We use the median across plants within each three-digit industry to reduce the effects of potential outliers.

[11] Given that the number of plants in certain four-digit industries is very low, we use three-digit-level industries.

TABLE 1
NUMBER OF PLANTS

Year	Plants
1992	5,937
1993	5,935
1994	6,256
1995	5,111
1996	5,465
1997	5,317
1998	4,862
1999	4,800
2000	4,632
2001	4,790
2002	5,171
2003	5,155
2004	5,447
2005	5,326

Source. National Annual Survey of Manufactures.

one-third are in the food sector (311), followed by fabricated metals (381) and wood (331), 9.0% and 6.7% of total manufacturing plants, respectively.

The data provided by the ENIA allows us to estimate total factor productivity at the firm level using the methodology developed by Olley and Pakes (1996) and extended by Levinsohn and Petrin (2003). This procedure has been used previously for the Chilean manufacturing industry by Bergoeing et al. (2006) and Álvarez and López (2005). In our regressions, given the methodological problems for computing TFP, we use two additional measures of productivity at the plant level to check the robustness of our results.

Table 3 presents descriptive statistics of the variables used in the estimations. First, we show variables corresponding to firm characteristics such as total factor productivity, the unskilled ratio, and size. Then we summarize the variables relative to industry characteristics and economic regulations. We called them macroeconomic variables.

VI. Main Econometric Results

Table 4 presents the main set of regressions for three different measures of productivity: (i) TFP computed using a parametric method (Olley and Pakes 1996; Levinsohn and Petrin 2003); (ii) TFP using a nonparametric method, with labor share calculated from the data, which assumes constant return to scale (van Biesebroeck 2007); and (iii) a measure of labor productivity calculated as real value added per worker.

Given the methodological difficulties in calculating productivity, we use these three measures to check the robustness of our results. The appendix shows the coefficients of the production function estimated with the methodology

TABLE 2
2005 DISTRIBUTION OF PLANTS BY INDUSTRY

Industry	Description	Plants	% of Total
311	Food	1,499	28.0
313	Beverages	196	3.7
314	Tobacco	5	.1
321	Textiles	275	5.2
322	Wearing	198	3.7
323	Leather	36	.7
324	Footwear	78	1.5
331	Wood	356	6.7
332	Furniture	122	2.3
341	Paper	127	2.4
342	Printing and publishing	284	5.3
351	Industrial chemicals	93	1.8
352	Other chemicals	208	3.9
353	Petroleum refineries	9	.2
354	Petroleum and coal	14	.3
355	Rubber	53	1.0
356	Plastic	292	5.5
361	Pottery	7	.1
362	Glass	30	.6
369	Other nonmetallic	214	4.0
371	Iron and steel	70	1.3
372	Nonferrous	98	1.8
381	Fabricated metal	477	9.0
382	Machinery	307	5.8
383	Machinery electrical	89	1.7
384	Transport equipment	88	1.7
385	Professional and scientific equipment	33	.6
390	Other manufacturers	68	1.3

Source. National Annual Survey of Manufactures.

developed by Levinsohn and Petrin (2003). In addition, the appendix exhibits a comparison of the three alternative productivity measures. Even though we find differences across the indicators, the main results are similar across measures. For an additional robustness check, we estimate TFP using system GMM for the production function estimation across sectors, but the results show some undesirable features. In particular, we find negative elasticities for some industries. However, dropping sectors with negative elasticities, the results confirm the negative impact of minimum wages on productivity, although the coefficient is larger than estimated with the other three measures.[12]

Our results in table 4 display evidence of persistence in within-plant productivity series; the coefficient of the one-period lagged TFP is between 0.4 and 0.8. As observed in table 4, our variable of interest, the interaction be-

[12] All of these results for TFP estimations using system GMM for the production function are available on request.

TABLE 3
DESCRIPTIVE STATISTICS OF VARIABLES USED IN THE ESTIMATIONS

Variable	Mean	SD
Firm characteristics:		
LogTFP (L&P)	5.27	1.09
LogTFP (N-P)	3.11	.89
LogTFP (VA/L)	8.73	.98
Unskilled ratio	4.56	6.36
Log(size)	1.90	2.21
Log(K/L)	4.23	5.79
Macroeconomic variables:		
Minimum wage	110.50	20.48
Labor regulation	.37	.38
Volatility	.20	.03
Tariff	.07	.03
Trade	.62	1.09
Finance	.57	.08
Financial dependence	.27	.29

Source. National Annual Survey of Manufactures.
Note. Definitions are as follows: logTFP (L&P) = log of total factor productivity using the methodology of Levinsohn and Petrin (2003); logTFP (N-P) = log of nonparametric estimation of TFP; logTFP (VA/L) = log of labor productivity; unskilled ratio = industry median of the ratio of blue-collar to white-collar workers; log(size) = log of employment; log(K/L) = log of capital per worker; minimum wage = index of real minimum wage (1997 = 100); labor regulation = index of regulations in the labor market; volatility = measure of industry excess job reallocation from Micco and Pagés (2006); tariff = average imports tariff; finance = measure of financial development (private credit by bank deposits to GDP); financial dependence = measure of industry external finance dependence, taken from Rajan and Zingales (1998).

tween the minimum wage and the unskilled ratio, is negative and significant for both measures of TFP but not significant in the case of labor productivity.[13] Thus our results confirm the hypothesis that a minimum wage increase reduces TFP and that this effect is larger—in absolute value—for industries with more unskilled labor, in other words, industries with higher exposure to minimum wage.

From estimates in the "TFP, L&P" column of table 4, the elasticity of TFP to minimum wage, evaluated at the sample unskilled ratio average, is -0.23 and -0.36 in the short and long runs, respectively. In quantitative terms, this implies that an accumulated increase of real minimum wages of about 22%, such as between 1998 and 2000, would have reduced TFP by about 8% in the long run for a plant with an average unskilled-to-skilled ratio. For industries with low (first decile) and high (ninth decile) intensity in unskilled labor, the estimated impact is a TFP reduction of 5.8% and 9.7%, respectively. The findings for the average firm in our sample can be compared to the results for the aggregated TFP in the manufacturing industry. Between 1998 and 2008, man-

[13] Note that we do not include log(MinWage) in the estimation because it only varies over time and its effect is already captured by time fixed effects.

TABLE 4
PRODUCTIVITY AND MINIMUM WAGE

	TFP, L&P	TFP, Nonparametric	Labor Productivity
LogTFP(−1)	.359	.478	.804
	(3.02)**	(3.12)**	(7.76)**
Log(MinWage)*log(unsk. ratio)	−.166	−.123	−.029
	(3.09)**	(2.46)*	(1.06)
Tariff*trade	−.003	−.017	.007
	(.56)	(1.95)	(1.35)
Labor reg.*volatility	−3.827	−3.947	−4.133
	(1.52)	(1.10)	(1.73)
Finance*financial dependence	−.629	−.212	−.034
	(2.41)*	(.53)	(.16)
Constant	4.568	2.909	2.104
	(9.63)**	(4.81)**	(2.41)*
Observations	38,805	38,734	38,734
Plants	6,775	6,756	6,756
AR (1) *p*-value	.000	.000	.000
AR (2) *p*-value	.158	.463	.000
Hansen *p*-value	.797	.003	.572

Note. Shown are the main sets of regressions for three different measures of productivity, including three-digit industry dummy variables and year fixed effects. Values in parentheses are robust standard errors. Definitions are as follows: log(MinWage) = log of real minimum wage; log(unsk. ratio) = log of industry median ratio of unskilled to skilled workers; tariff = average tariff; trade = exports plus imports over industry output in 1992; labor reg. = labor regulation index; volatility = measure of industry excess job reallocation from Micco and Pagés (2006); finance = measure of financial development (private credit in bank deposits to gross domestic product); financial dependence = measure of industry external finance dependence, taken from Rajan and Zingales (1998). All explanatory variables are 1-year lagged. See table 3 for other terms.
* $p < .1$.
** $p < .05$.

ufacturing TFP decreased by 0.55% (CORFO 2013). Our estimations indicate that the impact of minimum wage increases were larger than effective changes in the manufacturing industry.

One critical assumption for the validity of this GMM estimator is that the instruments must be exogenous. To test the validity of the instrument set used, we applied Hansen's (1982) test. We also report Arellano and Bond's (1991) test to detect first- and second-order autocorrelation of the error in the first-differences equation. The moments used in the estimation consider, in general, from one to four lags of all the explanatory variables as exogenous. However, in some cases, we vary the number of lags (and moment conditions) to pass the Hansen's and second-order correlation tests. As it can be appreciated in table 4, these tests do not reject either the null hypothesis of validity of the instruments (Hansen) or the null hypothesis of absence of second-order autocorrelation. This is generally valid—as it can be appreciated in the rest of the results—for the TFP measures and, in particular, for the parametric estimation of Levinsohn and Petrin (2003).

Regarding the effect of other regulations, we find a mainly nonsignificant effect for the interaction of the labor regulation index with industry volatility. Our results show that, in general, the labor regulation changes implemented in 2001 did not have any effect on productivity.

The coefficient for the interaction between tariffs and trade exposure is generally not significant, although it is negative and marginally significant in the second specification. This result is expected since tariff reductions have a positive effect on TFP, and this effect is larger for those sectors that are more exposed to international competition. This result is consistent with a wide variety of models in which higher international trade enhances productivity in domestic firms because they are exposed to fierce competition with foreign producers (MacDonald 1994), have access to a higher variety of inputs (Amiti and Konings 2007), or trade liberalization increased export opportunities (Lileeva and Trefler 2010) and incentives to innovate (Bustos 2011). In the case of financial development, we find a negative coefficient for the interaction between exposure to external financing and credit market deepening in one of our regressions, but this is not a robust result to alternative measures of productivity.[14]

As a robustness exercise, we include two additional covariates in the estimation. First, we include an interaction between minimum wage and size—measured as the lagged value of log of employment—for analyzing whether the increase in the minimum wage has a different effect on larger plants. We expect a positive coefficient for this interaction term considering that larger firms could have lower labor adjustment costs than smaller firms or might be in a better financial position for covering the costs of firing low-productivity workers.[15] Second, we include a triple interaction between minimum wage, unskilled intensity, and the index of labor regulations. By doing so, we test whether the negative impact of minimum wage on industries with more unskilled workers is higher when labor regulation changes increased the cost of worker dismissal.

Table 5 shows the results with these two additional explanatory variables for the three measures of productivity. We find no major changes in the rest of the parameters once these variables are introduced. Our results show that the interaction between size and minimum wage is positive, indicating that an increase in the minimum wage had a smaller effect on productivity for larger plants. Regarding the triple interaction, we do not find evidence of a greater

[14] Bergoeing et al. (2006), also using data on Chilean plants but for a different period, find a positive effect of both the measure of openness and for capital market development. However, they do not include all variables simultaneously; instead they use only one at a time in different regressions. In addition, they do not include lagged TFP as an explanatory variable.

[15] For evidence that larger firms face smaller labor adjustment costs, see Lapatinas (2009).

TABLE 5
PRODUCTIVITY AND MINIMUM WAGE: ADDITIONAL REGRESSORS

	TFP, L&P	TFP, Nonparametric	Labor Productivity
LogTFP(−1)	.346	.548	.739
	(3.44)**	(5.05)**	(10.25)**
Log(MinWage)*log(unsk. ratio)	−.167	−.108	−.023
	(3.92)**	(2.54)*	(.82)
Tariff*trade	−.005	−.008	.009
	(.95)	(1.31)	(2.09)*
Labor reg.*volatility	−2.988	3.179	−2.011
	(1.65)	(1.54)	(1.42)
Finance*financial dependence	−.568	−.754	−.057
	(2.42)*	(2.48)*	(.32)
Log(MinWage)*size	.035	.041	.014
	(5.08)**	(3.36)**	(3.18)**
Log(MinWage)*log(unsk. ratio)*labor reg.	.030	−.014	−.004
	(3.02)**	(1.01)	(.42)
Constant	3.733	1.353	2.305
	(10.47)**	(3.45)**	(3.54)**
Observations	38805	38734	38734
Plants	6775	6756	6756
AR(1) *p*-value	.000	.000	.000
AR(1) *p*-value	.218	.077	.002
Hansen *p*-value	.797	.002	.733

Note. See table 4 for definitions of terms.
* $p < .1$.
** $p < .05$.

negative effect of minimum wages after the changes in labor regulations in 2001. This last finding may be consistent with the fact that the changes in dismissal costs were minor compared to previous legislation. In fact, the change in labor laws in 1991—before the period under study—was more stringent since it increased the maximum severance payment from 5 to 11 months, and it greatly modified the causes that justified firing (Montenegro and Pagés 2005). Moreover, as can be appreciated in figure 4, the absence of a differential effect may be due to the large increase in minimum wage occurring at the same time as the reduction in the labor regulation index.

It can be argued that our unskilled ratio measure may be capturing other factor-intensity measures. For example, it can be the case that plants using more unskilled workers are also less capital intensive. To check that our results are also robust to this possibility, we estimate the regressions presented in table 4 again, including an interaction between the minimum wage and capital intensity. This variable is measured as the log of the stock of capital per worker. Our main results are robust to the inclusion of this additional interaction.[16] For all productivity measures, we find a negative parameter for the interaction between the

[16] Results are available on request.

TABLE 6
PRODUCTIVITY AND MINIMUM WAGE: SIMPLE SPECIFICATION

	TFP, L&P	TFP, Nonparametric	Labor Productivity	TFP, L&P	TFP, Nonparametric	Labor Productivity
LogTFP(−1)	.304	.516	.735	.449	.510	.651
	(2.71)**	(4.66)**	(7.96)**	(2.17)*	(4.45)**	(2.65)**
Log(MinWage)*unskilled ratio	−.125	−.084	−.015	−.212	−.125	−.021
	(2.39)*	(2.40)*	(.59)	(3.00)**	(2.00)*	(.49)
Labor reg.*volatility	−5.266	−2.378	−4.667	−5.479	−5.639	−1.719
	(2.71)**	(1.02)	(2.48)*	(1.49)	(1.41)	(.51)
Log(MinWage)*size	.030	.040	.013			
	(3.93)**	(2.39)*	(1.61)			
Log(MinWage)*unskilled ratio*labor reg.				.008	.005	−.019
				(.30)	(.17)	(.77)
Constant	4.067	1.491	2.587	4.793	3.042	3.333
	(9.29)**	(3.24)**	(3.26)**	(5.19)**	(3.98)**	(1.42)
Observations	38,805	38,734	38,734	38,805	38,734	38,734
Plants	6,775	6,756	6,756	6,775	6,756	6,756
AR(1) *p*-value	.000	.000	.000	.000	.000	.000
AR(2) *p*-value	.381	.491	.000	.296	.807	.062
Hansen *p*-value	.911	.002	.328	.670	.693	.751

Note. See tables 3 and 4 for details.
* $p < .1$.
** $p < .05$.

minimum wage and the unskilled ratio, confirming that increases in the minimum wage reduce productivity more in those industries that use unskilled workers more intensively. As in our basic regression, the negative effect of minimum wages is significant only for TFP measures and not for labor productivity.

VII. Robustness Check

We carry out several robustness checks. We analyze the case of controlling for too many variables, using alternative measures of minimum wage exposure, introducing interactions between macroeconomic variables and firms' exposure, and the potential impact of selection bias in using only surviving plants. First, to rule out the possibility of overcontrolling by including too many variables, we present estimation results using a simpler specification. As seen in table 6, this shows that the negative effect of a minimum wage increase remains statistically significant for TFP. Second, we use an alternative measure of exposure to minimum wage changes. This is computed as the average wage of blue-collar workers per industry, under the assumption that more exposed firms are in industries paying lower wages to unskilled workers.[17] Table 7

[17] Unfortunately, we cannot compute other exposure measures, such as the percentage of workers receiving minimum wage, because this data set does not include individual earnings.

TABLE 7

PRODUCTIVITY AND MINIMUM WAGE: ALTERNATIVE MEASURE OF EXPOSURE

	TFP, L&P	TFP, Nonparametric	Labor Productivity	TFP, L&P	TFP, Nonparametric	Labor Productivity
LogTFP(−1)	.350	.806	.835	.280	.220	.801
	(3.43)**	(4.58)**	(10.48)**	(2.53)*	(1.39)	(8.71)**
Log(MinWage)*log(BC Wage)	−.322	−.746	−.037	−.451	−.616	−.103
	(2.10)*	(2.85)**	(.36)	(2.50)*	(2.05)*	(.91)
Tariff*trade	.008	−.026	.009	−.010	−.009	.002
	(1.38)	(1.73)	(1.80)	(1.76)	(1.09)	(.44)
Labor reg.*volatility	−7.368	−12.512	−4.737	−5.416	1.493	−4.429
	(3.37)**	(2.40)*	(2.11)*	(2.56)*	(.42)	(2.14)*
Finance*financial dependence	−.596	−.065	.034	−.302	−.472	.136
	(2.37)*	(.15)	(.16)	(1.15)	(1.15)	(.58)
Log(MinWage)*size				.034	.064	.015
				(4.35)**	(2.46)*	(2.16)*
Log(MinWage)*log(BC Wage)*labor reg.				.062	.064	.033
				(1.55)	(.81)	(.78)
Constant	4.879	4.802	1.972	5.069	3.305	2.176
	(9.93)**	(4.28)**	(2.62)**	(8.81)**	(2.61)**	(2.48)*
Observations	38805	38734	38734	38805	38734	38734
Plants	6775	6756	6756	6775	6756	6756
AR(1) *p*-value	.000	.000	.000	.001	.000	.000
AR(2) *p*-value	.122	.190	.000	.656	.365	.000
Hansen *p*-value	.535	.093	.796	.816	.188	.292

Note. Log(BC Wage) = log of industry average of blue-collar workers. See table 4 for other terms.
* $p < .1$.
** $p < .05$.

shows that, for most of our specifications, the coefficient for the interaction between minimum wage and this measure of exposure is negative and significant. Again, this is particularly valid for TFP but not for labor productivity.

We also introduce interactions between other aggregate variables (international trade and financial development) and exposure to the minimum wage because these interactions may be driving our results. We show in table 8 that this is not the case. We find that most of these interactions for trade liberalization are not significant, nor are they positive and significant for financial development. However, for our measures of TFP, the negative impact of minimum wages survives the introduction of these additional covariates.

In table 9, we show that our results are robust to the correction for sample selection given that we employ an unbalanced panel. To address this problem, we follow the procedure developed by Wooldridge (1995) to deal with sample selection in panel data models. First, we estimate a probit model for the survival probability for each year and compute the inverse Mills ratio. In the second step, we estimate the fixed-effects panel data specification including the

TABLE 8
PRODUCTIVITY AND MINIMUM WAGE: INTERACTIONS WITH EXPOSURE

	TFP, L&P	TFP, Nonparametric	Labor Productivity
LogTFP(−1)	.478	1.690	−.150
	(5.19)**	(.09)	(2.69)**
Log(MinWage)*log(unsk. ratio)	−.171	−.330	.042
	(1.97)*	(3.52)**	(.68)
Tariff*log(unsk. ratio)	.239	.001	−.017
	(.49)	(8.10)**	(.38)
Finance*log(unsk. ratio)	−.014	.696	.857
	(1.84)	(2.89)**	(13.83)**
Tariff*trade	−.000	−.013	.004
	(.07)	(2.18)*	(.99)
Labor reg.*volatility	−2.735	−2.612	−2.591
	(1.75)	(1.49)	(2.14)*
Finance*financial dependence	−.748	−.405	.187
	(3.19)**	(1.17)	(.98)
Constant	3.851	2.318	1.473
	(9.24)**	(5.49)**	(2.74)**
Observations	38,805	38,734	38,734
Plants	6,775	6,756	6,756
AR(1) *p*-value	.000	.000	.000
AR(2) *p*-value	.004	.002	.000
Hansen *p*-value	.520	.016	.808

Note. Shown are three-digit industry dummy variables and year fixed effects, with robust standard errors in parentheses. Log(MinWage) = log of real minimum wage; log(unsk. ratio) = log of industry median of unskilled to skilled workers ratio; tariff = average tariff; trade = exports plus imports over industry output in 1992; labor reg. = labor regulation index; volatility = measure of industry excess job reallocation from Micco and Pagés (2006); finance = measure of financial development (private credit by deposit money banks to gross domestic product); financial dependence = measure of industry external finance dependence taken from Rajan and Zingales (1998). All explanatory variables are 1-year lagged.
* $p < .1$.
** $p < .05$.

introduction of the inverse Mills ratio. The selection equation considers variables such as productivity, size, and exports that have been traditionally associated with survival. For identification purposes, we introduce the investment-to-sales ratio and the square of this variable in the selection equation, under the assumption that investment generates sunk costs that reduce plants closings. In general, we find that our main results remain similar to previous ones.

There are potentially two problems with these estimations. First, we control only for time-specific shocks that are common to all industries and so there may be changes in the industry-specific real exchange rate (RER) that are unaccounted for. Unfortunately, we do not have RER data by industry. However, in our defense, the omission of this variable would only be a serious problem in our estimations should the RER changes during the period of minimum wage increases be negatively correlated with the degree of industry exposure to the minimum wage. We do not have a reason to think that this is the case.

TABLE 9
PRODUCTIVITY AND MINIMUM WAGE: SAMPLE SELECTION

	TFP, L&P	TFP, Nonparametric	Labor Productivity
LogTFP(−1)	.251	.421	.721
	(1.72)	(2.67)**	(5.67)**
Log(MinWage)*log(unsk. ratio)	−.135	−.122	−.018
	(2.56)*	(2.26)*	(.56)
Tariff*trade	−.003	−.009	.008
	(.42)	(.75)	(.96)
Labor reg.*volatility	−2.652	−4.789	−1.671
	(.98)	(1.30)	(.61)
Finance*financial dependence	−.583	−.226	−.027
	(2.15)*	(.59)	(.10)
Mills	−2.142	−3.350	−.963
	(1.76)	(1.87)	(1.02)
Constant	4.723	3.488	2.362
	(7.70)**	(4.71)**	(2.38)*
Observations	34,943	34,889	34,889
Plants	6,325	6,308	6,308
AR (1) *p*-value	.000	.000	.000
AR (2) *p*-value	.654	.527	.000
Hansen *p*-value	.914	.000	.680

Note. Terms are as in table 8.
* $p < .1$.
** $p < .05$.

Second, Chilean firms could have increased their subcontracting during this period. We do not have detailed information for dealing with this potential problem. However, we think that this phenomenon would work against our main hypothesis. If firms could subcontract some tasks and jobs in order to reduce the impact of minimum wage increases on labor costs, this would make it more difficult to find a negative relationship between minimum wage changes and TFP, as we do in most of our estimations.

VIII. Conclusions

There is a broad array of literature looking at the effects of employment regulations on economic performance and productivity. Related to our work, there is empirical evidence regarding the impact of changes in employment protection on productivity (Autor et al. 2007) and the effects of introducing minimum wages on firm profitability (Draca, Machin, and Van Reenen 2011).

The evidence presented in this paper suggests that a large increase in the minimum wage, such as during the period 1998–2000 in Chile, in the presence of relevant firing costs has a negative and significant effect on firms' total factor productivity. Our estimates indicate that a real increase of about 22% in the minimum wage during the period 1998–2000 reduced TFP by 5.8% in industries that do not use low-skilled workers intensively and 9.7% in ones that do. These results are robust to alternative measures of productivity and to the inclusion of several covariates to avoid confounding effects of other policy changes or firms' exposure to changes in the minimum wage.

Considering our results from a quantitative perspective, our estimations for the negative impact on the average firm in our sample are high compared to what happened with aggregate TFP in the manufacturing industry. Between 1998 and 2008, manufacturing TFP was reduced by 0.55%. Our findings indicate that the estimated impact of minimum wage increases was larger than the effective changes in the manufacturing industry. This indicates how relevant the impact of policy changes can be.

As expected, this negative effect on productivity is stronger in sectors with higher exposure to minimum wage, in particular, in more unskilled labor-intensive industries. This result is consistent with the idea that, in the presence of high severance payments, when the minimum wage increases, it may be costly for firms to adjust their workforce to the new optimal level. In this situation, firms are forced by regulations to keep an undesired amount of unskilled workers, affecting plant productivity. These findings are robust to several measures of productivity and the inclusion of other control variables. Notably, the conclusion remains after controlling for size, capital intensity, and a triple interaction of the minimum wage with labor regulations and minimum wage exposure.

Nevertheless, given that we do not have a direct measure of labor adjustment costs, we have not exactly identified their relevance. Our evidence suggests that adjustment costs may be important for understanding the relationship between wage increase and productivity.

Despite these results, the significant decline in TFP growth during the last decade in the Chilean economy does not yet have a clear explanation. Our evidence suggests that an increase in labor costs could be responsible for this productivity slowdown. This conclusion could be extended to the rest of the formal sector, especially to industries that rely heavily on unskilled labor. Nevertheless, more microeconomic work needs to be done to explain changes in firm productivity and how this is related to variations in economic policies in general and labor regulations in particular.

Appendix

Production Function Estimations and Productivity Comparisons

We show the elasticities calculated using the methodology developed for Levinsohn and Petrin (2003), which provides reasonable values for them. The median across three-digit industries for labor elasticity is 0.59 and for capital elasticity 0.35.

TABLE A1
LABOR AND CAPITAL ELASTICITIES OF THREE-DIGIT INDUSTRIES

Sector	Observations	Labor	Capital	Sum
311	18,567	.57	.38	.95
313	1,257	.15	.70	.85
321	4,248	.61	.33	.94
322	3,611	.52	.36	.88
323	555	.91	.05	.96
324	1,639	.70	.81	1.51
331	4,408	.45	.33	.78
332	1,906	.65	.16	.81
341	1,210	.58	1.27	1.85
342	3,065	.62	.71	1.33
351	832	.35	.66	1.01
352	2,438	.69	.37	1.06
353	61	.39	.20	.59
354	228	.23	.65	.88
355	779	.77	.05	.82
356	3,533	.51	.40	.91
361	201	.72	.81	1.53
362	297	.65	.59	1.24
369	2,164	.35	.32	.67
371	553	.30	.29	.59
372	562	.42	.21	.63
381	6,017	.66	.24	.9
382	2,964	.59	.21	.8
383	939	.50	.17	.67
384	1,339	.63	.41	1.04
385	336	.63	.26	.89
Median	1,298	.59	.35	.90

In table A2, we compare the different measures of productivity and find that they show similar broad patterns. We also show the correlations of the three measures. They are significant and between 0.4 and 0.7. The lowest correlation is between productivity measured as nonparametric total factor productivity (TFP) and labor productivity.

TABLE A2
YEARLY PRODUCTIVITY GROWTH RATES

Year	TFP, L&P (%)	TFP, Nonparametric (%)	Labor Productivity (%)
1993	7,90	5,28	9,14
1994	2,80	13,27	2,39
1995	3,93	3,92	4,20
1996	4,46	0,41	6,67
1997	4,57	−3,32	6,81
1998	0,17	3,90	0,92
1999	−0,62	−4,10	1,91
2000	4,20	−9,06	2,21
2001	−0,35	4,92	2,12
2002	−0,65	4,42	−0,02
2003	−1,02	0,59	−1,57
2004	3,49	7,66	2,63
2005	1,47	1,22	−0,13
Median	2,80	3,90	2,21
Average	2,33	2,24	2,87

Note. TFP = total factor productivity; L&P = methodology of Levinsohn and Petrin (2003).

Finally, we show the distribution of TFP—each measure is computed as the log difference with respect to the industry average—and find that there are not large differences.

TABLE A3
PRODUCTIVITY GROWTH CORRELATIONS

	TFP, L&P	TFP, Nonparametric	Labor Productivity
TFP, L&P	1		
TFP, nonparametric	.57	1	
Labor productivity	.67	.34	1

Note. See table A2 for terms.

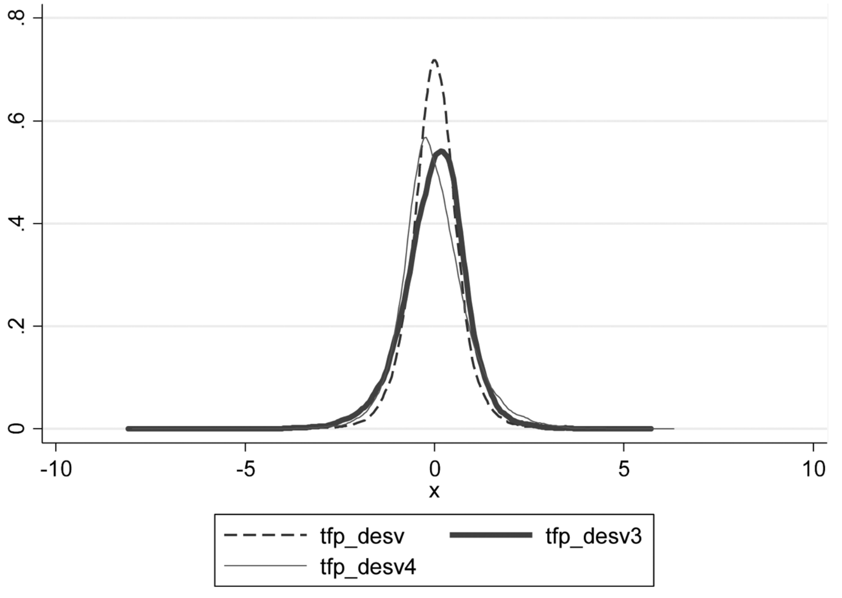

Figure A1. Here tfp_desv corresponds to total factor productivity (TFP) using the Levinsohn and Petrin (2003) procedure, tfp_desv3 corresponds to the nonparametric TFP, and tfp_desv4 corresponds to labor productivity.

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