# Minimum Wage and Productivity: Evidence from Chilean Manufacturing Plants

**Article** *in* Economic Development and Cultural Change · March 2018

DOI: 10.1086/697557



2 authors:

Roberto Alvarez
University of Chile


Rodrigo Fuentes
Pontifical Catholic University of Chile

# Minimum Wage and Productivity: Evidence from Chilean Manufacturing Plants[+]

**Roberto Álvarez**
*University of Chile*
robalvar@fen.uchile.cl

**Rodrigo Fuentes**
*Pontificia Universidad Católica de Chile*
rodrigo.fuentes@uc.cl

[+] We wish to thank José de Gregorio, Luis Opazo, Carmen Pagés, Klaus Schmidt-Hebbel, Chad Syverson, and an anonymous referee, along with seminar attendants at Alberto Hurtado University, Pontificia Universidad Católica de Chile, The Chilean Economic Society Congress 2009, and the DIME-ISGEP 2010 Workshop on Firm Selection and Country Competitiveness for their useful comments and suggestions. We also thank Harald Beyer for providing us the minimum wage data and Damian Vergara for valuable research assistance. Roberto Alvarez also thanks the Millennium Science Initiative (Project NS 100017 'Centro Intelis') for their financial support. Corresponding author postal address: Diagonal Paraguay 257, Of.1504, Santiago, Chile. Phone: (56-2)9783411, Fax: (56-2) 9783413

**Abstract**

This paper analyzes the effects of the minimum wage on a firm's productivity. The main hypothesis is that an increase in the minimum wage has a negative effect on total factor productivity (TFP) due to the existence of labor adjustment costs. Using data from Chilean manufacturing plants for the period 1992-2005, and a differences-in-differences methodology, we find that an increase in minimum wage had a negative effect on TFP. Our estimates indicate that real increase of about 22% in the minimum wage during the period 1998-2000 reduced TFP by 5.8% in low unskilled-intensive industries and 9.7% in high unskilled-intensive-industries. 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 minimum wage changes.

Keywords: productivity, regulations, minimum wage, firing costs

JEL Codes: O17, J30

## 1. 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, this 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; Petrin and Sivadasan, 2006; Lagos, 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 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 minimum wage changes do not affect the technological TFP, the actual TFP will be lower during the adjustment because firm will 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), which 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 (like employment subsidies and firing costs) reduces observed TFP due to changes in the composition of labor employed. In contrast, Autor et al. (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 differences-in-differences approach, exploiting differences in minimum wages over time and in industry exposure to changes in these labor costs. Given the relevance of measured productivity, we calculate it using three methods: i) TFP measured as in Olley and Pakes, (1996) and Levinsohn and Petrin (2003); ii) estimation of TFP using a non-parametric 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 1986-1997 (TFP grew at 3%), the aggregate productivity growth rate dramatically declined over the following 10 years (Figure 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 OECD countries, 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 effects of 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 U.S. They argue that this negative impact occurs because the adoption of dismissal protection 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 worker´s marginal revenue product and wages. We contribute to this literature by directly estimating the impact of minimum wage changes on plants' productivity and by using a particular

[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.

episode of national policy changes that established in-advance increases in the minimum wage for a three-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. In contrast to 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 U.S. 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% 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 2 discusses the relationship between minimum wages and productivity in the presence of firing costs. Section 3 describes the main policy changes

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

in the Chilean economy during the period under study. Section 4 explains the methodology used to identify the effect of minimum wages on productivity. Section 5 describes the dataset while Section 6 presents our econometric results. Section 7 shows several robustness checks, and Section 8 concludes.

## 2. Conceptual Framework

This section conceptually addresses why an increase in legal minimum wage, under the presence of relevant costs of workers dismissal, 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:

$\Pi_{t+s} = A_{t+s} f(k_{t+s}, l_{t+s}) - r_k k - w_{t+s} l_{t+s} - C(l_{t+s} - l_{t+s-1})$,

$A_{t+s} = \frac{y}{f(k_{t+s}, l_{t+s})}$ 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,

$r$ 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:

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

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

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

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

Where the parameter $\theta$ captures the relevance of adjustment costs. When $\theta$ 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:

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

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 non-optimal capital/labor ratio. Note that this conclusion still holds in the presence of adjustment costs for capital.[3]

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

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 $L_{t+s} = L_{t+s-1}$.

On the other hand, for some firms the mandatory minimum wage will be binding. This is the case of 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) — "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's call $l^*$ the optimal amount of labor when minimum wage increases and there is no adjustment cost, and $l'$ the amount of labor when the adjustment cost exists Then, we have that, given that $l^* < l'$, the firm's TFP will be lower than its optimal level according to:

$$A_t = TFP_t^* = \frac{y_t}{f(k_t, l_t^*)} > \frac{y_t}{f(k_t, l_t')} = TFP_t'$$

This difference between the two TFP levels will be zero when the amount of labor adjusts to its optimal level, i.e. when θ=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 before the minimum wage increases (t-1), firms were optimally choosing labor, 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 that its optimal level and TFP lower that the maximum. This is $TFP_t < TFP_{t-1}$. This is 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, therefore mandatory dismissal costs will drive the marginal productivity of labor below its wage. If the wage rate does not adjust because the presence of downward rigidity (like the minimum wage in our case) 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 equation 4). If the worker values the dismissal protection benefit more than the marginal cost of provision, a firm and a 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

[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).

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 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 costs 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 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.

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

## 3. Economic Policy Changes

In this section, we discuss the main policies and institutional changess regarding the 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 CPI, 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 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 three 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

[6] There has been a 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 the work by Cowan et al. (2005).

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 three 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 2.5% per year during 2001-2005.

The authorities thought that by establishing a three-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 representatives 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. At 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), which combines information on notice periods, compensation for dismissal, the likelihood that firm difficulties be considered as justified cause of dismissal, and how much the severance payment is in that 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:

$$Index_t = \sum_{i=1}^{T} \beta^i \delta^{i-1} (1-\delta)(b + aSP_{t+i}^{jc} + (1-a)SP_{t+i}^{uc})$$

Where $\beta$ is the discount factor, which is assumed to be equal to 0.92; $\delta$ 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; $a$ is the probability that a court will declare the dismissal justified, which is equal to zero during the period 1992-1998 and equal to one in the period 1999-2002.

$SP^{jc}$ is the tenure related to severance payments under justified cause for dismissal, and $SP^{uc}$ is 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 de facto a 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 et al. (2006), a reduction in trade barriers and financial development may have played a role in explaining the evolution of productivity in Chile. Since 1979, Chilean import

---

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

tariffs were uniform across sectors with few exceptions like 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 GDP (Beck, et al. 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%.

[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.

## 4. 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-industry or cross-regional variation. Thus, the identification strategy used in this paper is a variant of the differences-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:

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

Where $y_{it}$ is total factor productivity (in logs), ***i*** denotes a plant, ***j*** a sector, and ***t*** a year. $x_{it}$ is a vector of plant characteristics, WM is the real minimum wage, and $U_{jo}$ is a measure of industry's exposure to changes in minimum wages at the initial year of the sample period. Z is a vector of other time-varying variables than can affect TFP differentially depending on industry characteristics denoted by W. The vector Z contains three time-varying variables described in Section 3: 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, also due to the adjustment costs, it will affect TFP over time. In this dynamic specification, the short-run effect of changes in minimum wage are given by $\phi$ and in the long-run by $\phi / (1-\alpha)$.

Empirically, most of the microeconomic literature shows that productivity shocks tend to be persistent (Baily, et al. 1992; Foster, et al. 2002; and Foster et al. 2008). Moreover, the methodologies for estimating production functions assume that productivity follow a first-order Markovian process (Olley and Pakes, 1996; Levinsohn and Petrin, 2003; among others).

The omission of the lagged productivity generates serious problems in the estimation. 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 this 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 form and its first-difference form. They suggest applying a system 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 ($U_{jo}$).[9] Our identification strategy is based on the idea that a minimum wage increase

[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.

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-collars workers to white collars workers (measured also in logs) in the first year of the sample.[10] By measuring this variable at the beginning of the period, we can use it 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 3-digit ISIC industries in the U.S. — 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 3-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 3-digit ISIC industries. It is assumed that this indicator reflects technological reasons of why some industries depend more on external financing than others.

[10] We use the median across plants within each 3-digit industry to reduce the effects of potential outliers.

## 5. 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 plant characteristics for each plant that has at least 10 employees. All monetary variables were converted to constant pesos using 3-digit ISIC level price deflators. In addition, plants are classified according to the International Standard Industrial Classification (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 5,300 plants in 2005. Table 2 presents the distribution of plants by industry for 2005.[11] More than one third are in the food sector (311), followed by fabricated metals (381), and wood (331), 9.0% and 6.7% 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 Alvarez 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 present descriptive statistics of the variables used in the estimations. In panel A, we show variables corresponding to firm characteristics such as total factor productivity, the unskilled ratio, and size. In panel B, we summarize the variables relative to industry characteristics and economic regulations. We called then macroeconomic variables.

[11] Given that the number of plants in certain 4-digit industries is very low, we use 3-digit level industries.

## 6. 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 non-parametric 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 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 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 between 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

[12] All of these results for TFP estimations using system GMM for the production function are available upon request.
[13] Note that we do not include Log(MinWage) in the estimation because this variable only varies over time, and its effect is already captured by time fixed effects.

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 first column, 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 run, 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, manufacturing 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 the 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 1 to 4 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 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 general valid, as it can be appreciated in the rest of the results, for the TFP measures, in particular for the parametric estimation of Levinsonh and Petrin (2003).

Regarding the effect of other regulations, we find a mainly non-significant 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, 2004), access to a higher variety of inputs (Amiti and Konings, 2007) or trade liberalization increase 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 than larger firms could have lower labor adjustment costs[15] than smaller firms or might be in a better financial position for covering to costs of firing of low productivity workers. Second, we include a triple

[14] Bergoeing *et al* (2006) also using Chilean plants' information 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 than larger firms face smaller labor adjustment costs see Lapatinas (2009).

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 were 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 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 minimum wage and the unskilled ratio,

[16] Results are available upon request.

confirming that increases in the minimum wage reduces 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.

## 7. 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 of using only surviving plants. First, to rule out the possibility of over-controlling by including too many variables, we present estimation results using a simpler specification. This is shown in Table 6 and it 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 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 introduce also 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 and 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

[17] Unfortunately, we cannot compute other exposure measures, such as the percentage of workers receiving minimum wage because we do not have individual earnings in this dataset.

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 introduction of the inverse Mills ratio. The selection equation considers variable 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 squared of this variable in the selection equation, under the assumption that investment generates sunk costs that reduces plants closing. In general, we find that our main results remain similar to previous ones.

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

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 makes it more difficult to find a negative relationship between minimum wage changes and TFP as we do in most of our estimations.

## 8. Conclusions

There is a broad literature looking at the effects on 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 et al., 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 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 use them intensively. 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 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 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 the relevance of adjustment costs. 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 than 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 unskilled-labor intensive industries. 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.

## References


Aghion, P., Burgess, R., Redding, S. J and F. Zilibotti (2008). “The Unequal Effects of Liberalization: Evidence from Dismantling the License Raj in India.” *American Economic Review* 98(4): 1397–1412

Álvarez, R., and R. A. López (2005). “Exporting and Performance: Evidence from Chilean Plants.” *Canadian Journal of Economics* 38(4): 1384-1400.

Amiti, M. and J. Konings (2007). “Trade Liberalization, Intermediate Inputs, and Productivity: Evidence from Indonesia,” *American Economic Review*, 97(5): 1611-1638.

Autor, D. H., W. Kerr and A. D. Kluger (2007). “Does Employment Protection Reduce Productivity? Evidence from US States.” *The Economic Journal* 117(521): 189-217.

Baily, M. N., Ch. Hulten, and D. Campbell. (1992): “Productivity Dynamics in Manufacturing Establishments,” Brookings Papers on Economic Activity: Microeconomics, 187-249.

Bustos, P. (2011). “Trade Liberalization, Exports, and Technology Upgrading: Evidence on the Impact of MERCOSUR on Argentinian Firms,” *American Economic Review* 101(1): 304–340.

Becerra, G. (2006). “Arancel Efectivo de las Importaciones Chilenas: 2000-2005”, Economic Statistics Series 50, Central Bank of Chile.

Beck, T., A. Demigurc-Kunt and R. Levine (2000). “A New Database on Financial Development and Structure." *World Bank Economic Review* 14: 597-605.

Bergoeing, R., Hernando, A. and A. Repetto (2006). "Market Reforms and Efficiency Gains in Chile." *Working Paper* No. 372, Central Bank of Chile.

Beyer, H. (2008). "Mercado del Trabajo y Salario Mínimo." Puntos de Referencia No. 93, Centro de Estudios Públicos.

Blanchard, O., & Wolfers, J. (2000). "The Role of Shocks and Institutions in the Rise of European Unemployment: The Aggregate Evidence". *The Economic Journal*, 110(462), 1-33.

Bond, S. (2002). "Dynamic Panel Data Models: A Guide to Microdata Methods and Practice." *Portuguese Economic Journal* 1: 141-162.

Bond, S. and M. Söderbom. (2005). "Adjustment Costs and the Identification of Cobb Douglas Production Functions (No. 05/04). IFS Working Papers, Institute for Fiscal Studies (IFS).

Cowan, K., Micco, A., Mizala, A., Pagés-Serra, C. and P. Romaguera (2005). *Un Diagnóstico del Desempleo en Chile*, University of Chile and Inter-American Development Bank.

CORFO (2013): "Evolución de la Productividad Total de Factores en Chile,". http://repositoriodigital.corfo.cl/handle/11373/9104.

De Gregorio, J. (2007). "Algunas Reflexiones sobre el Crecimiento Económico en Chile." Documento de Política Económica No. 20, Central Bank of Chile.

Draca, M., Machin, S. and J. Van Reenen (2007). "Minimum Wages and Firm Profitability," *American Economic Journal: Applied Economics* 3(1): 129-151.

Foster, Lucia, John Haltiwanger, and C.J. Krizan. "The Link Between Aggregate and Micro Productivity Growth: Evidence from Retail Trade." NBER Working Paper 9120, 2002.

Foster, L., Haltiwanger, J., & Syverson, C. (2008). "Reallocation, Firm Turnover, and Efficiency: Selection on Productivity or Profitability?. American Economic Review, 98(1): 394-425.

Freeman, R. B. (1993). "Labor Market Institutions and Policies: Help or Hindrance to Economic Development?" Proceedings of the World Bank Annual Conference on Development Economics, Washington D.C. The World Bank.

Hamermesh, D. and G. Pfann (1996). "Adjustment Cost in Factor Demand" *Journal of Economic Literature*, 34(3): 1264-1292.

Hansen, L. P. (1982), "Large Sample Properties of Generalized Methods of Moments Estimators," Econometrica 50(4): 1029-1054.

Heckman, J. and C. Pagés (2000). "The Cost of Job Security Regulation: Evidence from Latin American Labor Markets." *Economia* 1(1): 109-144

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

Hsieh, C. and P. Klenow (2009). "Misallocation and Manufacturing TFP in China and India." *The Quarterly Journal of Economics* 124(4): 1403-1448..

Lagos, R. (2006). "A Model of TFP." *Review of Economic Studies* 73(4): 983-1007.

Lapatinas, A. (2009). "Labour Adjustment Costs: Estimation of a Dynamic Discrete Choice Model using Panel Data for Greek Manufacturing Firms." *Labour Economics* 16(5): 521-533.

Levinsohn, J. and A. Petrin, (2003). "Estimating Production Functions Using Inputs to Control for Unobservables." *Review of Economic Studies* 70(2), 317-41.

Lileeva, A. and D. Trefler. (2010). "Improved Access to Foreign Markets Raises Plant-level Productivity…For Some Plants," *Quarterly Journal of Economics*, 125 (3): 1051-1099.

Lora, E. A. (2001). "Structural Reforms in Latin America: What Has Been Reformed and How to Measure It." Inter-American Development Bank (IADB), Working Paper No. 466.

MacDonald, J. M. (1994). "Does Import Competition Force Efficient Production?. *Review of Economics and Statistics*, 76(4): 721-727.

Micco, A. and C. Pagés (2006). "The Economic Effects of Employment Protection: Evidence from International Industry-Level Data." *IZA Discussion Paper* N° 2433.

Montenegro, C. and C. Pagés (2005). "Who Benefits from Labor Market Regulations? Chile 1960-1998." in Restrepo J. and A. Tokman (eds.). *Labor Market and Institutions*. Central Bank of Chile, Santiago, Chile.

Mortensen, D. T., and C. A. Pissarides. (1994). "Job Creation and Job Destruction in the Theory of Unemployment". *Review of Economic Studies*, 61(3): 397-415.

Nickell, S. (1981). "Biases in Dynamic Models with Fixed Effects". Econometrica, 49(6): 1399-1416.

Olley, S. and A. Pakes (1996). "The Dynamics of Productivity in the Telecommunication Equipment Industry." *Econometrica* 64: 1263-98.

Pagés, C. and C. Montenegro (2007). "Job Security and the Age-Composition of Employment: Evidence from Chile." *Estudios de Economía*. 34 (2): 109-139.

Parente, S. and E. Prescott (2002). *Barriers to Riches*. The MIT Press.

Petrin A. and J. Sivadasan (2006). "Job Security Does Affect Economic Efficiency: Theory, A new Statistic, and Evidence from Chile." *NBER Working Papers* 12757.

Rajan, R. and L. Zingales (1998). "Financial Dependence and Growth." *American Economic Review* 88(3): 559-586.

Restuccia, D. and R. Rogerson (2008). "Policy Distortions and Aggregate Productivity with Heterogeneous Establishments." *Review of Economic Dynamics* 11(4): 707-720.

Van Biesebroeck, J. (2007). "Robustness of Productividy Estimates." *The Journal of Industrial Economics* 55(3): 529-569.

Velásquez, M. (2009). "Flexibilidad, Protección y Políticas Activas en Chile," Serie Macroeconomía del Desarrollo, 78, CEPAL.

Wickens, M. (2008). *Macroeconomic Theory: A Dynamic General Equilibrium Approach*. Princeton University Press, Princeton, New Jersey

Wooldridge, J. (1995). “Selection Corrections for Panel Data Models under Conditional Mean Independence Assumptions.” *Journal of Econometrics*, 68: 115–32.

Yellen, J. L. (1984). “Efficiency Wage Models of Unemployment,” *American Economic Review*, 74(2): 200-205.

Table 1

**ENIA and Number of Plants**

| *Year* | *Plants* |
|---|---|
| 1992 | 5937 |
| 1993 | 5935 |
| 1994 | 6256 |
| 1995 | 5111 |
| 1996 | 5465 |
| 1997 | 5317 |
| 1998 | 4862 |
| 1999 | 4800 |
| 2000 | 4632 |
| 2001 | 4790 |
| 2002 | 5171 |
| 2003 | 5155 |
| 2004 | 5447 |
| 2005 | 5326 |

Source : ENIA

Table 2

**Distribution of Plants by Industries, 2005**

| *Industry* | *Description* | *Plants* | *Percentage* |
|---|---|---|---|
| 311 | Food | 1,499 | 28.0 |
| 313 | Beverages | 196 | 3.7 |
| 314 | Tobacco | 5 | 0.1 |
| 321 | Textiles | 275 | 5.2 |
| 322 | Wearing | 198 | 3.7 |
| 323 | Leather | 36 | 0.7 |
| 324 | Footwear | 78 | 1.5 |
| 331 | Wood | 356 | 6.7 |
| 332 | Furniture | 122 | 2.3 |
| 341 | Paper | 127 | 2.4 |
| 342 | Printing & Pub. | 284 | 5.3 |
| 351 | Industrial chemicals | 93 | 1.8 |
| 352 | Other chemicals | 208 | 3.9 |
| 353 | Petroleum refineries | 9 | 0.2 |
| 354 | Petroleum & coal | 14 | 0.3 |
| 355 | Rubber | 53 | 1.0 |
| 356 | Plastic | 292 | 5.5 |
| 361 | Pottery | 7 | 0.1 |
| 362 | Glass | 30 | 0.6 |
| 369 | Other non-metallic | 214 | 4.0 |
| 371 | Iron & steel | 70 | 1.3 |
| 372 | Non-ferrous | 98 | 1.8 |
| 381 | Fabricated metal | 477 | 9.0 |
| 382 | Machinery | 307 | 5.8 |
| 383 | Machinery elec. | 89 | 1.7 |
| 384 | Transport equ. | 88 | 1.7 |
| 385 | Prof. & scientific eq | 33 | 0.6 |
| 390 | Other manuf. | 68 | 1.3 |

Source : ENIA

Table 3

Descriptive Statistics

| Variable | Mean | Std. Deviation |
|---|---|---|
| Firm Characteristics | | |
| Log TFP (L&P) | 5,27 | 1,09 |
| Log TFP (N-P) | 3,11 | 0,89 |
| Log TFP (VA/L) | 8,73 | 0,98 |
| 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 | 0,37 | 0,38 |
| Volatility | 0,20 | 0,03 |
| Tariff | 0,07 | 0,03 |
| Trade | 0,62 | 1,09 |
| Finance | 0,57 | 0,08 |
| Financial Dependence | 0,27 | 0,29 |

Source: ENIA. Log TFP (L&P) is log of TFP using the methodology of Levinsohn and Petrin (2003), Log TFP (N-P) is the log of non-parametric estimation of TFP; Log TFP (VA/L) is the log of labor productivity, Unskilled ratio is the industry median of the ratio of blue-collar to white collar workers, Log (size) is the log of employment, log (K/L) is log of capital per worker, and minimum wage is an index of real minimum wage (1997=100), Labor Regulation is an index of regulations in the labor market, and Volatility is a measure of industry excess job reallocation from Micco and Pagés (2006), Tariff is the average imports tariff, Finance is a measure of financial development (private credit by bank deposits to GDP), and Financial Dependence is a measure of industry external finance dependence and is taken from Rajan and Zingales (1998).

Table 4

**Productivity and Minimum Wage**

| | TFP, L&P | TFP, non-parametric | Labor Productivity |
|---|---|---|---|
| Log TFP(-1) | 0.359 | 0.478 | 0.804 |
| | (3.02)** | (3.12)** | (7.76)** |
| Log(MinWage)*Log(Unskilled Ratio) | -0.166 | -0.123 | -0.029 |
| | (3.09)** | (2.46)* | (1.06) |
| Tariff*Trade | -0.003 | -0.017 | 0.007 |
| | (0.56) | (1.95) | (1.35) |
| Labor Reg. *Volatility | -3.827 | -3.947 | -4.133 |
| | (1.52) | (1.10) | (1.73) |
| Finance*Financial Dependence | -0.629 | -0.212 | -0.034 |
| | (2.41)* | (0.53) | (0.16) |
| Constant | 4.568 | 2.909 | 2.104 |
| | (9.63)** | (4.81)** | (2.41)* |
| Observations | 38805 | 38734 | 38734 |
| Plants | 6775 | 6756 | 6756 |
| AR (1) p-value | 0.000 | 0.000 | 0.000 |
| AR (2) p-value | 0.158 | 0.463 | 0.000 |
| Hansen p-value | 0.797 | 0.003 | 0.572 |

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. It includes 3-digit industry dummy variables and year fixed effects. Log(MinWage) is the log of real minimum wage, Log(Unskilled Ratio) is the log of the industry median ratio of unskilled to skilled workers, Tariff is average tariff, Trade is exports plus imports over industry output in 1992, Labor Reg. is the labor regulation index, and Volatility is a measure of industry excess job reallocation from Micco and Pagés (2006), Finance is a measure of financial development (private credit in bank deposits to GDP), and Financial dependence is a measure of industry external finance dependence and is taken from Rajan and Zingales (1998). All explanatory variables are one-year lagged.

Table 5

**Productivity and Minimum Wage: Additional Regressors**

| | TFP, L&P | TFP, non.parametric | Labor Productivity |
|---|---|---|---|
| Log TFP(-1) | 0.346 | 0.548 | 0.739 |
| | (3.44)** | (5.05)** | (10.25)** |
| Log(MinWage)*Log(Unskilled Ratio) | -0.167 | -0.108 | -0.023 |
| | (3.92)** | (2.54)* | (0.82) |
| Tariff*Trade | -0.005 | -0.008 | 0.009 |
| | (0.95) | (1.31) | (2.09)* |
| Labor Reg. *Volatility | -2.988 | 3.179 | -2.011 |
| | (1.65) | (1.54) | (1.42) |
| Finance*Financial Dependence | -0.568 | -0.754 | -0.057 |
| | (2.42)* | (2.48)* | (0.32) |
| Log(MinWage)*Size | 0.035 | 0.041 | 0.014 |
| | (5.08)** | (3.36)** | (3.18)** |
| Log(MinWage)*Log(Unskilled Ratio)*Labor Reg. | 0.030 | -0.014 | -0.004 |
| | (3.02)** | (1.01) | (0.42) |
| 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 | 0.000 | 0.000 | 0.000 |
| AR(1) P-Value | 0.218 | 0.077 | 0.002 |
| Hansen P-value | 0.797 | 0.002 | 0.733 |

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. It includes 3-digit industry dummy variables and year fixed effects. Log(MinWage) is the log of real minimum wage, Log(Unskilled Ratio) is the log of the industry median of unskilled to skilled workers ratio, Tariff is average tariff, Trade is exports plus imports over industry output in 1992, Labor Reg. is the labor regulation index, Volatility is a measure of industry excess job reallocation from Micco and Pagés (2006), Finance is a measure of financial development (private credit by deposit money banks to GDP), Financial dependence is a measure of industry external finance dependence taken from Rajan and Zingales (1998), and Size is log of employment. All explanatory variables are one-year lagged.

Table 6

## Productivity and Minimum Wage: Simple Specification

| | TFP, L&P | TFP, non.par. | Labor Productivity | TFP, L&P | TFP, non.par. | Labor Productivity |
|---|---|---|---|---|---|---|
| Log TFP(-1) | 0.304 | 0.516 | 0.735 | 0.449 | 0.510 | 0.651 |
| | (2.71)** | (4.66)** | (7.96)** | (2.17)* | (4.45)** | (2.65)** |
| Log(MinWage)*Unskilled Ratio | -0.125 | -0.084 | -0.015 | -0.212 | -0.125 | -0.021 |
| | (2.39)* | (2.40)* | (0.59) | (3.00)** | (2.00)* | (0.49) |
| 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) | (0.51) |
| Log(MinWage)*Size | 0.030 | 0.040 | 0.013 | | | |
| | (3.93)** | (2.39)* | (1.61) | | | |
| Log(MinWage)*Unsk. Ratio*Labor Reg. | | | | 0.008 | 0.005 | -0.019 |
| | | | | (0.30) | (0.17) | (0.77) |
| 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 | 38805 | 38734 | 38734 | 38805 | 38734 | 38734 |
| Plants | 6775 | 6756 | 6756 | 6775 | 6756 | 6756 |
| AR(1) p-value | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 | 0.000 |
| AR(2) p-value | 0.381 | 0.491 | 0.000 | 0.296 | 0.807 | 0.062 |
| Hansen p-value | 0.911 | 0.002 | 0.328 | 0.670 | 0.693 | 0.751 |

Robust sandard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. It includes 3-digit industry dummy variables and year fixed effects. Log(MinWage) is the log of real minimum wage, Log(Unskilled Ratio) is the log of the industry median of unskilled to skilled workers ratio, Labor Reg. is the labor regulation index, Volatility is a measure of industry excess job reallocation from Micco and Pagés (2006), and Size is log of employment. All explanatory variables are one-year lagged.

Table 7

**Productivity and Minimum Wage: Alternative Measure of Exposure**

| | TFP L&P | TFP Non-Parametric | Labor Productivity | TFP L&P | TFP Non-Parametric | Labor Productivity |
|---|---|---|---|---|---|---|
| Log TFP(-1) | 0.350 | 0.806 | 0.835 | 0.280 | 0.220 | 0.801 |
| | (3.43)** | (4.58)** | (10.48)** | (2.53)* | (1.39) | (8.71)** |
| Log(MinWage)*Log(BC Wage) | -0.322 | -0.746 | -0.037 | -0.451 | -0.616 | -0.103 |
| | (2.10)* | (2.85)** | (0.36) | (2.50)* | (2.05)* | (0.91) |
| Tariff*Trade | 0.008 | -0.026 | 0.009 | -0.010 | -0.009 | 0.002 |
| | (1.38) | (1.73) | (1.80) | (1.76) | (1.09) | (0.44) |
| Labor Reg. *Volatility | -7.368 | -12.512 | -4.737 | -5.416 | 1.493 | -4.429 |
| | (3.37)** | (2.40)* | (2.11)* | (2.56)* | (0.42) | (2.14)* |
| Finance*Financial Dependence | -0.596 | -0.065 | 0.034 | -0.302 | -0.472 | 0.136 |
| | (2.37)* | (0.15) | (0.16) | (1.15) | (1.15) | (0.58) |
| Log(MinWage)*Size | | | | 0.034 | 0.064 | 0.015 |
| | | | | (4.35)** | (2.46)* | (2.16)* |
| Log(MinWage)*Log(BC Wage)*Labor Reg. | | | | 0.062 | 0.064 | 0.033 |
| | | | | (1.55) | (0.81) | (0.78) |
| 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 | 0.000 | 0.000 | 0.000 | 0.001 | 0.000 | 0.000 |
| AR(2) p-value | 0.122 | 0.190 | 0.000 | 0.656 | 0.365 | 0.000 |
| Hansen p-value | 0.535 | 0.093 | 0.796 | 0.816 | 0.188 | 0.292 |

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. It includes 3-digit industry dummy variables and year fixed effects. Log(MinWage) is the log of real minimum wage, Log(BC Wage) is the log of industry average of blue-collar workers, Tariff is average tariff, Trade is exports plus imports over industry output in 1992, Labor Reg. is the labor regulation index, Volatility is a measure of industry excess job reallocation from Micco and Pagés (2006), Finance is a measure of financial development (private credit by deposit money banks to GDP), Financial dependence is a measure of industry external finance dependence taken from Rajan and Zingales (1998), Size is log of employment. All explanatory variables are one-year lagged.

Table 8

**Productivity and Minimum Wage: Interactions with Exposure**

| | TFP L&P | TFP Non-Parametric | Labor Productivity |
|---|---|---|---|
| Log TFP(-1) | 0.478 | 1.690 | -0.150 |
| | (5.19)** | (0.09) | (2.69)** |
| Log(MinWage)*Log(Unsk. Ratio) | -0.171 | -0.330 | 0.042 |
| | (1.97)* | (3.52)** | (0.68) |
| Tariff*Log(Unsk. Ratio) | 0.239 | 0.001 | -0.017 |
| | (0.49) | (8.10)** | (0.38) |
| Finance*Log(Unsk. Ratio) | -0.014 | 0.696 | 0.857 |
| | (1.84) | (2.89)** | (13.83)** |
| Tariff*Trade | -0.000 | -0.013 | 0.004 |
| | (0.07) | (2.18)* | (0.99) |
| Labor Reg. *Volatility | -2.735 | -2.612 | -2.591 |
| | (1.75) | (1.49) | (2.14)* |
| Finance*Financial Dependence | -0.748 | -0.405 | 0.187 |
| | (3.19)** | (1.17) | (0.98) |
| Constant | 3.851 | 2.318 | 1.473 |
| | (9.24)** | (5.49)** | (2.74)** |
| Observations | 38805 | 38734 | 38734 |
| Plants | 6775 | 6756 | 6756 |
| AR(1) p-value | 0.000 | 0.000 | 0.000 |
| AR(2) p-value | 0.004 | 0.002 | 0.000 |
| Hansen p-value | 0.520 | 0.016 | 0.808 |

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. It includes 3-digit industry dummy variables and year fixed effects. Log(MinWage) is the log of real minimum wage, Log(Unsk. Ratio) is the log of the industry median of unskilled to skilled workers ratio, Tariff is average tariff, Trade is exports plus imports over industry output in 1992, Labor Reg. is the labor regulation index, Volatility is a measure of industry excess job reallocation from Micco and Pagés (2006), Finance is a measure of financial development (private credit by deposit money banks to GDP), Financial dependence is a measure of industry external finance dependence taken from Rajan and Zingales (1998). All explanatory variables are one-year lagged.

Table 9

**Productivity and Minimum Wage: Sample Selection**

| | TFP, L&P | TFP, non-parametric | Labor Productivity |
|---|---|---|---|
| Log TFP(-1) | 0.251 | 0.421 | 0.721 |
| | (1.72) | (2.67)** | (5.67)** |
| Log(MinWage)*Log(Unskilled Ratio) | -0.135 | -0.122 | -0.018 |
| | (2.56)* | (2.26)* | (0.56) |
| Tariff*Trade | -0.003 | -0.009 | 0.008 |
| | (0.42) | (0.75) | (0.96) |
| Labor Reg. *Volatility | -2.652 | -4.789 | -1.671 |
| | (0.98) | (1.30) | (0.61) |
| Finance*Financial Dependence | -0.583 | -0.226 | -0.027 |
| | (2.15)* | (0.59) | (0.10) |
| Mills | -2.142 | -3.350 | -0.963 |
| | (1.76) | (1.87) | (1.02) |
| Constant | 4.723 | 3.488 | 2.362 |
| | (7.70)** | (4.71)** | (2.38)* |
| Observations | 34943 | 34889 | 34889 |
| Plants | 6325 | 6308 | 6308 |
| AR (1) p-value | 0.000 | 0.000 | 0.000 |
| AR (2) p-value | 0.654 | 0.527 | 0.000 |
| Hansen p-value | 0.914 | 0.000 | 0.680 |

Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1. It includes 3-digit industry dummy variables and year fixed effects. Log(MinWage) is the log of real minimum wage, Log(Unsk. Ratio) is the log of the industry median of unskilled to skilled workers ratio, Tariff is average tariff, Trade is exports plus imports over industry output in 1992, Labor Reg. is the labor regulation index, Volatility is a measure of industry excess job reallocation from Micco and Pagés (2006), Finance is a measure of financial development (private credit by deposit money banks to GDP), Financial dependence is a measure of industry external finance dependence taken from Rajan and Zingales (1998), and Mills is the inverse of the Mills ratio from a Probit model. All explanatory variables are one-year lagged.

## 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 these elasticities. The median across 3-digit industries for labor elasticity is 0.59 and for capital elasticity is 0.35.

| Sector | Obs | Labor | Capital | Sum |
|---|---|---|---|---|
| 311 | 18567 | 0,57 | 0,38 | 0,95 |
| 313 | 1257 | 0,15 | 0,70 | 0,85 |
| 321 | 4248 | 0,61 | 0,33 | 0,94 |
| 322 | 3611 | 0,52 | 0,36 | 0,88 |
| 323 | 555 | 0,91 | 0,05 | 0,96 |
| 324 | 1639 | 0,70 | 0,81 | 1,51 |
| 331 | 4408 | 0,45 | 0,33 | 0,78 |
| 332 | 1906 | 0,65 | 0,16 | 0,81 |
| 341 | 1210 | 0,58 | 1,27 | 1,85 |
| 342 | 3065 | 0,62 | 0,71 | 1,33 |
| 351 | 832 | 0,35 | 0,66 | 1,01 |
| 352 | 2438 | 0,69 | 0,37 | 1,06 |
| 353 | 61 | 0,39 | 0,20 | 0,59 |
| 354 | 228 | 0,23 | 0,65 | 0,88 |
| 355 | 779 | 0,77 | 0,05 | 0,82 |
| 356 | 3533 | 0,51 | 0,40 | 0,91 |
| 361 | 201 | 0,72 | 0,81 | 1,53 |
| 362 | 297 | 0,65 | 0,59 | 1,24 |
| 369 | 2164 | 0,35 | 0,32 | 0,67 |
| 371 | 553 | 0,30 | 0,29 | 0,59 |
| 372 | 562 | 0,42 | 0,21 | 0,63 |
| 381 | 6017 | 0,66 | 0,24 | 0,9 |
| 382 | 2964 | 0,59 | 0,21 | 0,8 |
| 383 | 939 | 0,50 | 0,17 | 0,67 |
| 384 | 1339 | 0,63 | 0,41 | 1,04 |
| 385 | 336 | 0,63 | 0,26 | 0,89 |
| Median | 1298 | 0,59 | 0,35 | 0,90 |

We compare the different measures of productivity and they show similar broad patterns. As we show below, the yearly growth rates are different across measures, but they show similar 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 non-parametric TFP and labor productivity.

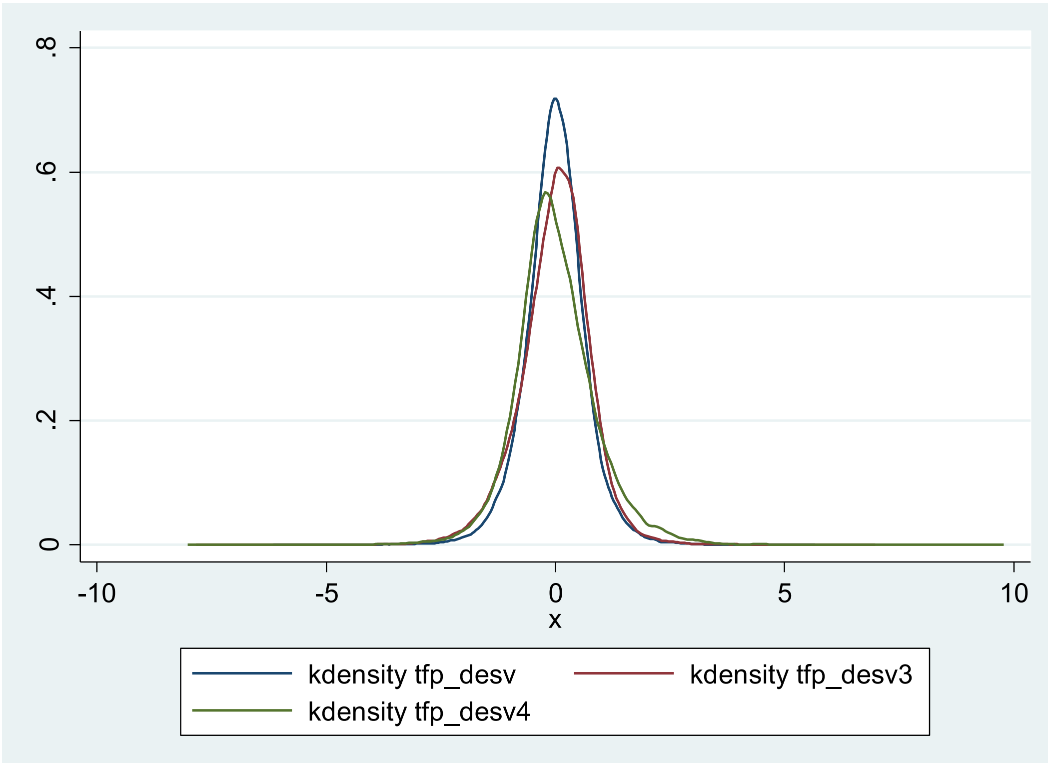
Yearly Productivity Growth Rates

| year | TFP L&P | TFP, non-parametric | 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% |

Productivity Growth Correlations

| | TFP L&P | TFP, non-parametric | Labor Productivity |
|---|---|---|---|
| TFP L&P | 1 | | |
| TFP, non-parametric | 0,57 | 1,00 | |
| Labor Productivity | 0,67 | 0,34 | 1 |

Finally, we show the distribution of TFP –each measured is computed as the log difference respect to the industry average – and it can be appreciated that there are not large differences.




Notes: tfp_desv corresponds to TFP using L&P procedure, tfp_desv3 corresponds to the non-parametric TFP, and tfp_desv4 to labor productivity.