Job Security *Does* Affect Economic Efficiency: Theory, A New Statistic, and Evidence from Chile *

Amil Petrin
Graduate School of Business, University of Chicago and NBER †

Jagadeesh Sivadasan
Stephen M Ross School of Business, University of Michigan ‡

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Abstract

In theory job security provisions can be completely undone by efficient contracts. The extensive empirical macro- and micro-level evidence on the impact of these provisions is largely inconclusive. We argue that the weak evidence is a consequence of the weak power of statistics used, which are often motivated by a 2-period model. We show this motivation does not necessarily extend to settings with imperfect competition/strategic interactions, non-convex adjustment costs, or an infinite horizon. All models do share one feature: firing costs drive a wedge between the marginal revenue product of labor and its marginal cost (wage). We examine changes in this gap as our test statistic. It is easy to compute and has a welfare interpretation. We use census data of Chilean manufacturing firms for the years 1979-1996 to look for real effects induced by two significant increases in the costs of dismissing employees. The traditional statistics provide weak evidence in our data (too). However, while we find little positive effect on the materials gap, we do find large and statistically significant increases in the gaps between the marginal product of labor and the wage for both white and blue collar workers.

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†email: amil.petrin@chicagogsb.edu

‡email: jagadees@umich.edu

1 Introduction

Job security regulations that increase the costs of dismissing employees are common in both the developing and developed world, and are a matter of intense policy debate, as the 2006 riots in France recently illustrated.[1] The wide body of theoretical work and of both macro- and micro-level evidence on the impact of job security is largely inconclusive, as the literature reviews in both Heckman and Pages (2004) and Addison and Teixeira (2001) illustrate.[2] The issue is further complicated by the insight of Lazear (1990), who shows that the distortion introduced by these provisions can be completely undone by efficient contracts, where the mandated firing costs are passed on to workers who willingly accept a lower wage. Overall, there is no clear evidence that job security negatively impacts economic efficiency.[3]

We argue that the mixed evidence is a consequence of the weak power of statistics being used. These statistics are based fundamentally on an aggregate or micro-level labor demand equation. Researchers have examined how changes in job security change aggregate employment and unemployment, micro- and aggregate- estimates of labor demand, the probability of adjustment, and the size of adjustment conditional on adjusting. The statistics are often motivated by a simple and intuitive 2-period model of plant-level employment decisions. We show these results do not necessarily extend to a world with imperfect competition, strategic interactions, non-convex adjustment costs, or an infinite horizon.[4] Also, while unobserved variables and aggregation bias are always potential problems in their own right, the theory we develop points directly to how - in these more general settings - unobserved variables and aggregation bias can weaken the test, if not destroy its econometric validity. We argue that job security provisions do have real effects and low power is the culprit in the reported ambiguous findings.[5]

One fact is certain from economic theory: firing costs drive a wedge between the marginal revenue product of labor and its marginal cost (wage). This idea has been widely explored in the literature analyzing the nature of labor adjustment costs

[1] The riots were spurred by an attempt to reduce firing costs for young workers.

[2] Depending upon the assumptions, some theory papers find a positive effect of increasing firing costs on employment (Bentolila and Bertola (1990), Alvarez and Veracierto (2001)), while others find a negative effect (Risager and Sorensen (1997), Bertola (1990), and Hopenhayn and Rogerson (1993).

[3] We agree with the sentiment from Heckman and Pages (2004) and Hammermesh (2004), that job security provisions have negative impacts, but remain with Topel (1998), skeptical about what the data to date have told us about the specifics of the labor demand equation.

[4] There is substantial evidence of non-convex adjustment costs for both capital (e.g. Cooper and Haltiwanger (1993), Cooper and Haltiwanger (2001)) and labor (e.g. Hamermesh (1989), Pfann and Palm (1993), and Cooper, Haltiwanger, and Willis (2004)).

[5] Thus, our argument is entirely consistent with Heckman and Pages (2004) claim that the inconclusive nature of results based on more aggregate cross-country time series data can in part be surmounted by the use of large micro data sets. We suggest a new statistic for plant-level data which may be more powerful.

(see the review in Bond and Van Reenen (2003)).[6] We develop a new test statistic, showing how to directly estimate this "gap" using plant-level production data. Our approach is robust to settings outside the 2-period model, and to many of the aggregation and estimation problems that we show can afflict the more traditional statistics. It is easy to compute, in our data has significant power, and, from the social planner's perspective, has a clear link to welfare, as increases in the absolute value of the gap are associated with greater economic inefficiency.

Inputs that do not have these costs associated with them and are presumably easy to adjust, like materials and electricity, should have gaps that are unaffected by the job security provisions. We use these gaps as "control" gaps. They tell us whether increases in the labor gaps - if found - might be related to some other unobserved phenomenon that affects the gaps of all inputs.

Using plant-level census data for Chilean manufacturing firms for the years 1979-1996, we look for real effects induced by two significant increases in job security. The first change, in late 1984, no longer exempted from severance pay firms that could show "economic cause" for dismissal. Severance was set equal to no less than a month's wages per year of tenure, with a five month ceiling. The second change, in 1991, increased the ceiling to 11 months, and added a 20% surcharge if the employer could not prove "economic cause." Both Edwards and Edwards (2000), who use aggregate time-series data, and Pages and Montenegro (1999), who use individual-level survey data, find *no* effect of these increases on aggregate employment levels, although the latter study does report that there are compositional effects.[7]

We find significant increases in the gaps between the marginal product of labor and the wage for both white and blue collar workers. Relative to the initial gap, the gap increases an average of 25% in the 1985-1990 period relative to pre-1985. When the maximum severance roughly doubles in 1990, the average gap becomes 50% greater than the initial gap. The timing of the increases are also linked to the changes. The gap is relatively unchanged from 1979-1985 (except for the deep recession in 1982-83) and then increases sharply after 1985 until it levels off at the larger gap in 1988. Then, in 1990 the gap begins steadily increasing again. Finally, we find little evidence of non-labor input gaps increasing; materials, which is on average half of revenue share, is virtually flat from 1981 onward.[8]

[6] Shapiro (1986) uses the Euler equations to recover dynamic demand for labor and capital, which is fundamentally identified using "gaps." See also Caballero and Engel (1993), Caballero and Engel (1999), Cooper and Willis (2001), Cooper and Willis (2003), Caballero, Cowan, Engel and Micco (2004), and Eslava, Haltiwanger, Kugler, and Kugler (2006).

[7] The young workers and the women tend to lose to the older men.

[8] Electricity exhibits a time-series pattern similar to materials. It suffers from less measurement error bias than materials because it is not an aggregate (it has a small share though). Like materials, fuels is also an aggregate,

We then turn to the question of what we would find if we used the usual statistics. As reported above, two studies have found no effects on aggregate employment or unemployment. We use the *plant-level* variation to estimate micro-level labor demand equations, looking for systematic changes, and find little conclusive evidence. We also look at the coefficient on lagged labor, which is often expected to move towards "1" as firing costs increase. The evidence is also inconclusive from the OLS and fixed effects estimators, although the Bond and Blundell (1998) IV estimates are consistent with this prediction when *all* parameters of the demand function are allowed to freely vary in each period.[9]

We estimate changes in the probability of plant-level adjustment in response to job security. The two period model's prediction is that it falls. In our data, unconditionally it increases from 13% to 17%, and it continues to be increasing until we condition on plant-level productivity, wage, lagged labor, and aggregate control factors, in which case it declines. We also look at whether hiring and firing conditional on adjustment are attenuated towards zero, as the 2-period model predicts. Unconditional hiring is attenuated to zero, but once all micro-level conditioning variables are used, it then is increasing in firing costs. Unconditional firing is attenuated to zero, and becomes more so when the micro-level conditioning variables are used. Overall, without the gap analysis, we view the evidence on whether firing costs have real effects in Chile - of which there is quite a bit - as being less than convincing.

We also develop one additional statistic that is implied by the Lazear (1990) insight. We calculate the implied wage change necessary to undo the distortion of the job security legislation under several different scenarios. The actual change in the average wage after the first dose of job security is several magnitudes larger than what is necessary to offset the firing cost. There is no change in average wage after the second dose. This new statistic also appears to be inconclusive in this Chilean data.

The paper is organized as follows. Section 2 develops the theory. Section 3 introduces the new Gap statistic. Section 4 summarizes the key economic reforms in Chile over the period that we examine (1979-1996). Section 5 describes the plant-level data. Section 6 reports the evidence from the Gap, Section 7 presents the traditional statistics, and Section 8 concludes.

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which does not change in the first treatment and then increases significantly in the second treatment.

[9] Heckman and Pages (2004) raise an econometric problem with this specification in aggregate cross-country studies.

2 Theory

We start with the impact of job security on labor demand in a 2-period model. These effects can be undone by efficient contracts, as in Lazear (1990). Also, we show that the 2-period case does not extend to a world with imperfect competition/strategic interactions, fixed costs, or an infinite horizon.

One implication is that the estimated labor demand "function" may be a correspondence if the necessary aggregation conditions fail. For example, under imperfect competition a firm's labor demand can be a function of its own productivity and the productivity of all of its competitors. If only own-firm productivity is conditioned on (usually done by conditioning on output), then multiple levels of labor will be associated with it, a different one (possibly) for every different set of competitors' productivity levels that could occur. Simple aggregation problems of this type present potentially serious problems for estimation of statistics related directly to labor demand and its derivatives.

In the cases when estimated labor demand is a function, it is an implicit one, and is not (generally) linear in its arguments. The appropriate estimator may require non-parametric techniques, which can lead to weak tests in finite data. Furthermore, the statistics currently used to look for real effects often rely on either the probability of adjusting being monotonically decreasing in firing costs, or that hiring and firing conditional on adjusting is attenuated towards zero as firing costs increase. When we add either strategic interactions, fixed costs, or an infinite horizon to the two-period model, these monotonicity results no longer necessarily hold.

While the "gap" between the current period marginal revenue product (MRP) and the wage is not necessarily monotonic in firing costs, looking to changes in the gap for evidence of firing cost effects is likely to be more powerful because the gap is *linear* in the MRP and wage, so standard root-n convergence results may be available. In addition, the gap is based directly on the first-order condition, which is the fundamental primitive in this exercise.

Readers not interested in the technical details underlying these results can skip to section 3 for a discussion of our "Gap" test and Section 6 for the results.

2.1 Two-Period Setup with no Fixed Costs

The two-period model, $t = 1, 2$, is similar to that presented in Heckman and Pages (2004). We assume the production function

$$Q(l, \theta) = \theta f(l)$$

that has a single input (labor), and $f(l)$ is increasing and concave function in labor, with productivity θ. Wages are exogenously set at w per unit of labor. The firing cost function is given as

$$C(l_t - l_{t-1}; c) = \begin{cases} c * |l_t - l_{t-1}| & \text{if } l_{t-1} < l_t \\ 0 & \text{if } l_{t-1} \geq l_t \end{cases} \tag{1}$$

where c denotes the per-unit labor cost of firing.[10] We write it as linear in workers because this is way the cost enters the firm problem in Chile and many other countries (see Heckman and Pages (2004)).[11]

The firm operates for two periods. In each period a demand (or production) shock θ_t $t = 1, 2$, is realized before the labor decision for that period is made. θ_2 is unknown when labor in period one is set, but θ_1 is known. With no firing costs, the profit function for the firm in either period is:

$$\pi_t = \theta_t f(l_t) - w l_t, \tag{2}$$

where w is the wage. The optimal choice of labor in each period t equates marginal revenue with marginal cost:

$$f'(l_t^*) = \left(\frac{w}{\theta_t}\right). \tag{3}$$

With no firing costs, the firm hires in period 2 if θ_2 is greater than θ_1, and fires if θ_2 is less than θ_1.

After l_1 has been chosen and θ_2 is known. Then, l_2^* solves

$$l_2^* = argmax_l \;\; \theta_2 f'(l) - wl - C(l - l_1; c). \tag{4}$$

In the face of adjustment costs a firm that has productivity shock θ_2 will adjust labor if

$$\theta_2 f'(l) - wl - C(l - l_1; c) > \theta_2 f'(l_1) - w l_1 \tag{5}$$

for some attainable l. When θ_2 is greater than θ_1, the firm will hire according to the same decision rule in (3), that is, as if there were no firing costs, because there is no future consequence of hiring more labor in the final (second) period. When θ_2 is less than θ_1, the firm will want to adjust down, but will only do so if the savings in costs $w - c$ exceeds the marginal revenue $\theta_2 f'(l)$ evaluated at $l = l_1$, that is, when $\theta_2 < \frac{w-c}{f'(l_1)}$. In this case, the firm will fire workers up to the point that

$$\theta_2 f'(l_2) = w - c.$$

[10] Thus we abstract from differences in severance payments that arise because of differences in tenure of employees.
[11] It is straightforward to add hiring costs to this setup, which serve to further exacerbate the estimation issues.

Otherwise, for $\theta_2 \in \left(\frac{w-c}{f'(l_1)}, \frac{w}{f'(l_1)}\right)$, there is no attainable l such that a profit-maximizing firm adjusts and $l_2 = l_1$.

Assuming $w > c$, we can summarize the decision rule for second period labor l_2^* as:[12]

$$\text{For } \theta_2 \geq \theta^* = \frac{w}{f'(l_1)} \quad : \quad l_2^* = [f']^{-1}\left(\frac{w}{\theta_2}\right) \tag{6}$$

$$\text{For } \frac{w-c}{f'(l_1)} = \theta_* < \theta_2 < \theta^* \quad : \quad l_2^* = l_1 \tag{7}$$

$$\text{For } \theta_2 \leq \theta_* \quad : \quad l_2^* = [f']^{-1}\left(\frac{w-c}{\theta_2}\right) \tag{8}$$

Equations 6-8 characterize the effects of job security as c increases from zero. Firms choose levels of labor that are inefficient, as a "gap" between the marginal revenue product (MRP) and the wage is now present in (7) and (8). In this 2-period case, $\theta_*(c)$ is decreasing in c, which increases the size of the inaction zone, so the probability of adjusting labor falls in c. Finally, the concavity of $f(\cdot)$ means that, conditional on firing (demand shock $\theta \leq \theta_*$), the magnitude of the decrease in labor levels, $l_1 - l_2$, is decreasing in firing cost c.[13] The evidence is rather mixed on these last two predictions.

2.2 Undoing the Distortion with Contracts

If there is efficient bargaining between the worker and the employer, a contract can be written specifying a side payment from the worker to firm that fully offsets the firing cost. One such scheme, proposed in Lazear (1990), has the firm paying w in period 1 to the worker, with the worker agreeing to set aside c until period 2. In period 2, if $\theta_2 < \theta_1$, each worker who is fired receives c. All retained workers receive $w + c$. If $\theta_2 \geq \theta_1$, then retained workers receive $w + c$ and new hires get w.

This contract allows the firm to pay firing costs out of the worker's salary from the previous period. The optimal choices of labor and the hiring and firing rule remain unchanged from the non-distorted setting. The marginal cost faced by the firm is w in each period regardless of whether the firm hires or fires. The workers' labor force participation choice (not explicitly modeled here) is also unaffected, as they receive the same wage ($= w$) in present value terms, as in the zero firing cost

[12] We must assume that $w > c$ or there will never be adjustment in period 2. In a multi-period version, the present value of wages saved over the future time horizon needs to exceed the one time firing cost in order to observe any firing at all.

[13] Conditional on $\theta \leq \theta_*$, the magnitude of decrease in labor levels is $\left(l_1 - l_2 = l_1 - f'^{-1}\left(\frac{w-c}{\theta}\right)\right)$, which is decreasing in c because f'^{-1} is decreasing in its argument.

regime. This contract can be written for the infinite-horizon case, with the firm and the worker agreeing to a similar arrangement period-by-period. Since no distortions are introduced into the market, efficiency means welfare continues to be maximized.

Lazear (1990) argues that the inefficiency may be difficult to undo using side payments for many practical reasons. In particular, workers must be willing to make the side payments to the employer or into an insurance fund; apprehension on the part of workers regarding the future severance payment could prevent the distortion's undoing.[14] Also, from an efficiency standpoint, firing probabilities are dependent on worker characteristics and firm layoff experience, so any unemployment insurance plan that does not condition on these factors is not going to maximize welfare.

2.3 Imperfect Competition and Strategic Interactions

With imperfect competition or when firms' strategically interact with one another, current choice variables are generally a function of all relevant state variables. We consider the case without firing costs first, so labor is not a state variable. As in Section 2, firms differ in their productivity level, which is either an efficiency difference, a demand difference (differentiated product), or both.[15] We assume J single-product firms compete against one another in a differentiated product setting.[16] Labor demand for firm k solves

$$P_k(1+\epsilon_{P_kQ_k}) * \frac{\partial Q_k}{\partial l} = w,$$

where P_k and Q_k are price and quantity sold of k, and $\epsilon_{P_kQ_k}$ is the elasticity of the price of k with respect to quantity sold. Nash behavior requires that labor choices are best responses. When goods are substitutes, the price and quantity demanded of good j will be a function of the entire vector of plant-level productivities (and the wage), so the functions $P_k(\cdot)$ and $\epsilon_{P_kQ_k}(\cdot)$ are written with arguments $P_k(\theta_1, \ldots, \theta_J, w)$ and $\epsilon_{P_kQ_k}(\theta_1, \ldots, \theta_J, w)$. Labor demand is thus also a function of the same arguments:

$$l_t^{*k} = l_t^k(\theta_t^1, \ldots, \theta_t^J, w_t).$$

With firing costs or more complicated forms of strategic interactions, labor demand for firm k is determined by the wage, firm k's productivity, all competing firms' productivities, and similarly own- and competitor- stocks of labor:

$$l_t^{*k} = l_t^k(l_{t-1}^1, \ldots, l_{t-1}^J, \theta_t^1, \ldots, \theta_t^J, w_t). \tag{9}$$

[14]Even if the workers are willing to make side payments, other problems exist, including potential moral hazard problems like workers attempting to obtain the severance package early, or agency problems like managers colluding with workers to extract excess severance payouts in the face of full insurance.

[15]These are not separately identified in our setting.

[16]With multi-product firms, the analysis is more complicated, but the result is similar.

In either case, estimation of labor demand using only $(l^k_{t-1}, \theta^k_t, w_t)$ as regressors is misspecified, as

$$l^{*k}_t = l^k_t(l^k_{t-1}, \theta^k_t, w_t)$$

is a correspondence.

An example illustrates the latter case. Consider a market with two competitors. Given firm 1's labor level, productivity level, and the wage, firm 1's labor demand will also depend on firm 2's level of labor and productivity. When firm 2's labor and productivity are not conditioned upon, firm 1's labor demand relates its own arguments to different levels of labor, depending upon whether their competitor has low or high productivity and a low or high stock of current labor.

2.4 The Infinite Horizon Model with Fixed Costs

We add the infinite horizon back to the 2-period model. Non-convex adjustment costs are important in rationalizing many plant-level and aggregate data, so we also treat fixed costs here.[17]

Per-period profits in period t are given by current profits π_t minus the costs associated with adjusting labor:

$$\begin{array}{rlcl} & \pi(\theta_t, w_t, l_t) & - & C(l_t - l_{t-1}; c) \\ = & \theta_t f(l_t) - w_t l_t & - & C(l_t - l_{t-1}; c), \end{array} \tag{10}$$

where wages are now indexed by time. The adjustment cost function is redefined to allow for fixed costs F of adjustment, so

$$C(l_t - l_{t-1}; c) = \begin{cases} F + c * |l_t - l_{t-1}| & \text{if } l_{t-1} < l_t \\ 0 & \text{if } l_{t-1} \geq l_t \end{cases} \tag{11}$$

At time t, given l_{t-1}, the expected discounted profits for any given sequence of future labor levels $\{l_s\}_{s=t}^{\infty}$ is

$$\pi(\theta_t, w_t, l_t) - C(l_t - l_{t-1}; c) + E_t[\sum_{s=t+1}^{\infty} \beta^{s-t}(\pi(\theta_s, w_s, l_s) - C(l_s - l_{s-1}; c))], \tag{12}$$

with $\beta < 1$ denoting the discount rate and E_t denoting the expectation conditional on the information available at time t. The uncertainty about the future arises because (θ_t, w_t) evolve probabilistically. We assume they follow a first order Markov process, and we write the transition probability function as $dP(\theta', w'|\theta, w)$.[18] The firm observes (θ_t, w_t) before choosing l_t.

[17] We abstract from the entry/exit decision because the main issues can be illustrated without adding this complication.

[18] It is straightforward to allow the conditional distribution of θ to be higher order Markov.

The firm's problem is to choose $\{l_s\}_{s=t}^{\infty}$ to maximize (12). Assuming standard regularity conditions hold,[19] we can reexpress the problem of the choice of optimal labor at time t in the (recursive) value-function representation:

$$V(\theta_t, w_t, l_{t-1}) = max_l \; \theta_t f(l) - w_t l - C(l - l_{t-1}; c) + \beta \overline{V}(\theta_t, w_t, l), \qquad (13)$$

where $V(\theta_t, w_t, l_{t-1})$ is the expected discounted profits at time t, and the continuation value of being at (θ_t, w_t, l) is

$$\overline{V}(\theta_t, w_t, l) = \int_{\theta', w'} V(\theta', w', l) dP(\theta', w' | \theta_t, w_t).$$

With (θ_t, w_t) known, beliefs about probabilities associated with any sequence of outcomes $\{\theta_s, w_s\}_{s=t+1}^{\infty}$ determine the value of any contingency plan $\{l_s\}_{s=t+1}^{\infty}$. Labor demand at time t is then given by

$$l_t^*(\theta_t, w_t, l_{t-1}) = argmax_l \; \left(\theta_t f(l) - w_t l - C(l - l_{t-1}; c) + \beta \overline{V}(\theta_t, w_t, l) \right), \qquad (14)$$

We divide firms into two groups: those that adjust their labor, and those that do not.

2.4.1 Adjusting Firms

Lemma 1 summarizes the first-order conditions that hold for adjusting firms.[20]

Lemma 1. *Assume $V(\cdot)$ is differentiable in l. When $l_t^* > l_{t-1}^*$, l_t^* satisfies*

$$\theta_t f'(l_t^*) - w_t + \beta \frac{\partial \overline{V}(\theta_t, w_t, l_t^*; c)}{\partial l} = 0. \qquad (15)$$

When $l_t^* < l_{t-1}^*$, l_t^* satisfies:

$$\theta f'(l_t^*) - (w_t - c) + \beta \frac{\partial \overline{V}(\theta_t, w_t, l_t^*; c)}{\partial l} = 0. \qquad (16)$$

Proof
The result follows directly from optimization when an interior solution exists.#

In the special case when $c = 0$, the continuation value is unaffected by the choice of labor in the current period, so both (15) and (16) reduce to the "myopic" per period maximization solution that we reported earlier in the 2-period setting with no firing costs:

$$\theta_t f'(l_t^*) = w_t. \qquad (17)$$

[19] See Stokey and Lucas (1993) and Pakes (1996)

[20] A fixed cost that is paid to adjust and that is not changing in the size of adjustment does not affect the first order condition, conditional on adjustment.

If $Pr(\theta_t = \theta_{t-1}) = 0$, firms will adjust their labor every period.

Because Euler equations only hold for feasible alternatives, we must amend a standard two-period Euler equation setup to accommodate the possibility of the firm not adjusting. The program must be pushed forward until the firm adjusts again.[21] Let the variable τ take on positive integer values, and define the next adjustment time as $t + \tau^*$, so τ^* is given as

$$\tau^* = min_{\tau \geq 1}\{l^*(\theta_{t+\tau}, w_{t+\tau}, l_t) \neq l_t\}. \tag{18}$$

In general, τ^* is a random variable because the firm does not know with certainty the next period in which it will adjust.

Corollary 1 allows τ^* to be random.

Corollary 1. *When $l_t^* > l_{t-1}^*$, l_t^* satisfies*

$$\begin{array}{rcl} \theta_t f'(l_t^*) & - & w_t \\ & + & E_t[\sum_{\tau=1}^{\tau^*-1} \beta^\tau(\theta_{t+\tau} f'(l_t^*) - w_{t+\tau})] \\ & - & E_t[\beta^{\tau^*} 1\{l_{t+\tau^*}^* < l_t^*\}\, c] = 0, \end{array} \tag{19}$$

When $l_t^* < l_{t-1}^*$, l_t^* satisfies:

$$\begin{array}{rcl} \theta_t f'(l_t^*) & - & (w_t - c) \\ & + & E_t[\sum_{\tau=1}^{\tau^*-1} \beta^\tau(\theta_{t+\tau} f'(l_t^*) - w_{t+\tau})] \\ & - & E_t[\beta^{\tau^*} 1\{l_{t+\tau^*}^* < l_t^*\}\, c] = 0, \end{array} \tag{20}$$

Proof

These are the Euler equations. #

l_t^* is a fixed point to (19) or (20).[22] In theory there could be multiple solutions, and proving uniqueness without stronger assumptions is difficult (see e.g. Caballero and Engel (1999)).[23] If we assume it is a function, it is an implicit one, and is not (generally) linear in its arguments, as (19) and (20) illustrate. This is the reason that, without more prior information, a consistent estimator may require non-parametric techniques.

Application of the implicit function theorem to (19) and (20) shows that the probability of adjusting is not necessarily monotonically falling as firing costs increase.

[21] For derivation of the Euler equations, we follow the standard approach, perturbing the firm off its optimal path - and then "as soon as possible" - putting the firm back on its optimal path. "As soon as possible" is the next period at which the firm adjusts its labor level. We follow the suggested approach outlined in Pakes (1996). Specifically, the idea is to increase l_t^* by ϵ in the current period and then at some future period decrease hired labor by ϵ. As long as the alternate plan is feasible, then the difference in profits between the unperturbed and the perturbed plan is non-negative for any $\epsilon \neq 0$, and is exactly zero at $\epsilon = 0$. Thus, optimization in the face of firing costs implies that the change in profits under the alternative program - when evaluated at $\epsilon = 0$ - is equal to zero, a condition which has implications for the adjustment behavior of labor.

[22] Hiring costs would add an additional term to each expression.

[23] In their flexible setting they are only able to prove that their investment policy function does not have an infinite number of solutions.

Additionally, hiring and firing conditional on adjusting is also no longer increasingly attenuated towards zero as firing costs increase.

Corollary 1 is easiest to understand if we write down the first order condition for a firm that knows τ^* and is hiring labor in period t. Let $F_{t+k} = E_t[1\{l^*_{t+k} < l^*_t\}]$ denote the probability of the event that a firm fires k periods from time t, given what is known at t. Conditional on the information set at time t and the choice of labor, this probability is given as

$$F_{t+k}(\theta_t, w_t, l_t, c),$$

which we write as F_{t+k}, where the probability is always calculated with respect to the information at time t. $l = l^*_t$ satisfies

$$\begin{aligned} \theta_t f'(l^*_t) &- w_t \\ &+ E_t[\textstyle\sum_{\tau=1}^{\tau^*-1} \beta^\tau (\theta_{t+\tau} f'(l^*_t) - w_{t+\tau})] \\ &- \beta^{\tau^*} F_{t+\tau^*} c = 0, \end{aligned} \tag{21}$$

where $F_{t+\tau^*}$ is known (either a zero or one), and we slightly abuse notation as E_t would be taken with τ^* known. For example, when $\tau^* = 2$, $l = l^*_t$ satisfies

$$\begin{aligned} \theta_t f'(l^*_t) &- w_t \\ &+ E_t[\beta(\theta_{t+1} f'(l^*_t) - w_{t+1})] \\ &- \beta^2 F_{t+2} c = 0. \end{aligned} \tag{22}$$

The firm accounts for the one period of inactivity by adding to the first-order condition expected discounted marginal revenue less expected discounted marginal costs for that inactive period. The choice of labor equates expected marginal revenue with expected marginal cost, but the expression is complicated by additional terms related to the periods of non-adjustment.

Corollary 2 is illustrative for a simple case of dynamics. Specifically, we consider a firm that adjusts this period and believes with probability 1 that it will adjust again in the next period. Let F_l be the derivative of the probability of firing workers in period 2 with respect to an increase in the number of workers in period 1. Define ε_{Fc} as the elasticity of the probability of firing tomorrow given an increase in firing costs today.

Corollary 2. *Assume a firm adjusting this period believes with probability 1 that it will also adjust next period. If $F_l > 0$, then hiring becomes increasingly attenuated as c increases if $\varepsilon_{Fc} < 1$, and firing becomes increasingly attenuated as c is increasing if $\varepsilon_{Fc} < \frac{1}{\beta F} - 1$.*

Proof

Using (19) and (20) with $\tau^* = 1$, when $l_t^* > l_{t-1}^*$, l_t^* satisfies

$$\theta_t f'(l_t^*) - w_t - \beta F_{t+1} c = 0, \tag{23}$$

and when $l_t^* < l_{t-1}^*$, l_t^* satisfies

$$\theta f'(l_t^*) - (w_t - c) - \beta F_{t+1} c = 0. \tag{24}$$

Using the implicit function theorem, for firms with $l_t^* > l_{t-1}^*$,

$$\frac{\partial l^*}{\partial c} = \frac{\beta(F_c c + F)}{\theta f'' - F_l \beta c}. \tag{25}$$

Similarly, for firms with $l_t^* < l_{t-1}^*$,

$$\frac{\partial l^*}{\partial c} = \frac{\beta(F_c c + F) - 1}{\theta f'' - F_l \beta c}. \tag{26}$$

The claims follow directly. #

In this case the change in the continuation value from Lemma 1 is given by the change in $-\beta F_{t+1} c$. If $F_c < 0$, increasing firing costs today decreases the probability of firing tomorrow. Many would consider this case to be the most relevant.

For hiring firms the requirement for monotonically declining upward adjustment is

$$\varepsilon_{Fc} = -\frac{F_c c}{F} < 1.$$

When $F_c < 0$, the elasticity of the fall in the probability of firing with respect to the firing cost must never exceed 1. If it does, then $\beta F_{t+1} c$ decreases towards zero as c increases, which works to increase the number of hires. The fall in the probability of firing is high enough such that the negative effect of the increase in the firing cost (to be paid in the future) is offset by the fall in the future probability that this worker is fired.[24]

For firing firms, the requirement for monotonic attenuation of number fired in c is

$$\varepsilon_{Fc} < \frac{1}{\beta F} - 1.$$

Since $\frac{1}{\beta F} - 1 > 0$, the base case of $F_c < 0$ yields monotonicity.[25] More general cases are usually more difficult to characterize, as the extra terms in Corollary 1 show.

[24] With $F_c \geq 0$ the monotonicity follows trivially.

[25] For $F_c > 0$, the increase has to be small. Also, all of this analysis has assumed $F_l > 0$. If the probability of firing is increasing the number of laborers that are employed, then these results must be revisited, as the number of possible outcomes increases significantly.

2.4.2 Non-adjusting Firms

For firms that do not adjust, expected discounted profits at l_{t-1} are higher than expected discounted profits for any other available choice of l. This inequality is given as

$$\begin{aligned} &\theta_t f(l_{t-1}) - wl_{t-1} + \beta \overline{V}(\theta_t, l_{t-1}; c) \\ &\geq \theta_t f(l) - wl - C(l - l_{t-1}; c) + \beta \overline{V}(\theta_t, l; c) \quad \forall l. \end{aligned} \tag{27}$$

Lemma 2 provides the necessary condition for the probability of not adjusting to be weakly increasing in firing costs.

Lemma 2. *Given l_{t-1} and firing costs $C(\cdot)$, let Θ_c be the set of θ such that the firm does not adjust labor:*

$$\begin{aligned} \Theta_c \;=\; &\{\theta : \;(\theta f(l) - wl) - (\theta f(l_{t-1}) - wl_{t-1}) \leq \\ &C(l - l_{t-1}; c) + \beta(\overline{V}(\theta, l_{t-1}; c) - \overline{V}(\theta, l; c)) \;\; \forall l\}. \end{aligned} \tag{28}$$

The probability that a firm does not adjust its labor is weakly increasing in the firing costs c if and only if $\theta \in \Theta_{c_1}$ implies $\theta \in \Theta_{c_2}$ for $c_2 \geq c_1$.

The proof is trivial. Unlike the 2-period case, the implications for how the probability of adjustment changes as firing costs increase is not clear.[26]

Corollary 3 provides a set of sufficient conditions for monotonicity to hold. The result relies on how, as firing costs increase, the difference in continuation values changes for firms with the same θ but different levels of labor. In general,

$$\frac{\partial \overline{V}(\theta, l; c)}{\partial c} \leq 0,$$

as the value of being in operation decreases as firing costs increase at any level of (θ, l).

Corollary 3. *Assume*

$$\frac{\partial^2 \overline{V}(\theta, l)}{\partial l \partial c} \leq 0.$$

Given l_{t-1}, assume

$$|l - l_{t-1}| \geq \beta \left| \frac{\partial(\overline{V}(\theta, l_{t-1}; c) - \overline{V}(\theta, l; c))}{\partial c} \right| \quad \forall l < l_{t-1}.$$

Then the probability that a firm does not adjust its labor is weakly increasing in the firing costs c.

[26] There is a large enough c_{MAX} such that there will never be adjustment.

Proof

Holding θ and l constant, the left hand side of the inequality (27) does not change. We show that the right hand side increases as c increases for all l holding θ constant. In the case of $l > l_{t-1}$, $\frac{\partial C(l-l_{t-1};c)}{c} = 0$. $\frac{\partial(\overline{V}(\theta,l_{t-1};c)-\overline{V}(\theta,l;c))}{\partial c} > 0$ by the Fundamental Theorem of Calculus because $\frac{\partial^2 \overline{V}(\theta,l)}{\partial l \partial c} < 0$. In the case of $l < l_{t-1}$, $\frac{\partial C(l-l_{t-1};c)}{c} = |l - l_{t-1}|$. By assumption this is greater than $\beta|\frac{\partial(\overline{V}(\theta,l_{t-1};c)-\overline{V}(\theta,l;c))}{\partial c}|$. Thus, holding θ constant, $\forall\, l$, the left hand side is constant and the right hand side is increasing.#

We interpret the conditions. Assume two firms with the same productivity level have the same beliefs about future productivity sequences. The first assumption requires firms with the same current productivity shock but more labor to experience greater decreases in value as firing costs increase. If one firm has more labor than the other, then the firm with more labor has its total potential liability in firing costs increase more than the firm with fewer laborers. Since both firms have the same beliefs about future demand/productivity shocks and there are no hiring costs, a firm with more labor could very well have a value that is decreasing in its labor level at a faster rate as firing costs increases.

The second assumption is harder to motivate. For a firm that currently has l_{t-1} employees and wants to fire $|l - l_{t-1}|$ employees, the change in current costs for a small increase in the per-unit firing cost is $|l - l_{t-1}|$. The assumption requires $l - l_{t-1}$ to be greater than the change in future discounted values between a firm with l_{t-1} and l employees that occurs due to the same small increase in firing costs. Further, this must be true for all values of labor less than l_{t-1}. If the condition is to hold in the aggregate, it must hold for all possible starting labor levels, i.e. for all l_{t-1}. Overall, the conditions are significantly stronger to achieve monotonicity than in the 2-period case.

3 The Gap Methodology

In models where job security policies are not undone through contracting, Section 2 tells us that firing costs drive a wedge between marginal revenue its marginal cost (the wage). From a social planners' perspective, firms should choose labor to equate the wage w_{it} with the marginal revenue product (MRP), given by

$$MRP_{it} = P_{it} * \frac{\partial Q_{it}}{\partial l},$$

where $\frac{\partial Q_{it}}{\partial l}$ gives the increase in output associated with a small increase in labor. We define the Gap, G_{it}, as the difference:

$$G_{it} = MRP_{it} - w_{it}.$$

Equation (17) shows that, in a setting with price-taking firms and no distortions, G_{it} is zero. With imperfect competition, strategic interactions or firing cost distortions, equations (9), (19), (20), and (27) illustrate the extra terms entering G_{it}. We start by focusing on the absolute values of the gaps, concluding as a social planner might that an increase in the average gap is suggestive of a negative impact on efficiency.[27]

In order to estimate the marginal product of labor at the firm level, we posit a Cobb-Douglas production function as below:

$$q_{it} = \beta_m m_{it} + \beta_s l_{it}^s + \beta_u l_{it}^u + \beta_k k_{it} + \beta_r r_{it} + \beta_f f_{it} + \varepsilon_{it}$$

where q_{it} is the log of the real output, m_{it} is log of real value of intermediate materials, l_{it}^s is the log of the number of skilled (white collar) employees, l_{it}^u is the log of the number of unskilled (blue collar) employees, k_{it} is the log of the real capital stock employed, r_{it} is log or real services used and f_{it} is log of the real value of fuels used by firm i in year t. The error, ε_{it}, is assumed equal to

$$\varepsilon_{it} = \omega_{it} + \eta_{it},$$

with ω_{it} is the transmitted (and predictable) component of firm specific productivity shock, and η_{it} is a firm specific iid productivity shock. We assume ω follows a first order Markov process (as in Olley and Pakes, 1996). The latter error is assumed to have no impact on the firm's decisions, while the former can be correlated with input choices.

Given the production function specification and observed input levels, the marginal product is straightforward to calculate once one determines what "error" should be conditioned upon. For example, if one conditioned on both ω and η, then for skilled labor, firm i operating in year t has a marginal product $(\frac{\partial Q_{it}}{\partial l})$ given by:

$$\frac{\partial Q_{it}}{\partial l} = \;=\; \beta_s e^{\varepsilon_{it}} (l_{it}^s)^{\beta_s - 1} (l_{it}^u)^{\beta_u} (k_{it})^{\beta_k} (m_{it})^{\beta_m} (f_{it})^{\beta_f} (e_{it})^{\beta_e} = \beta_s * \frac{Q_{it}}{l_{it}^s}.$$

However, if η is measurement error, then we will only want to condition upon ω when the marginal revenue product is calculated.

Given the first-order Markov assumption on ω, we might try to predict ω_{it} given ω_{it-1} by projecting $\hat{\varepsilon}_{it}$ on $\hat{\varepsilon}_{it-1}$, and then calculating the marginal revenue product at

[27] For now we treat price as observable, addressing the case of measurement error later.

its predicted value. Assuming the firm does not know η when the choice of inputs are made, conditioning on this predicted value for ω is akin to calculating the marginal revenue product for the firm at its belief regarding this period's productivity.[28]

We look at the absolute value of the gap between marginal product and wage for skilled and unskilled labor, G^s_{it} and G^u_{it}, and for materials (typically 50% or more of the revenue share):

$$G^s_{it} = |MRP^s_{its} - w^s_{it}|$$

$$G^u_{it} = |MRP^u_{itu} - w^u_{it}|$$

$$G^m_{it} = |MRP^u_{itm} - P_{itm}|,$$

where P_{itm} is the price for materials. As noted earlier, the gap is linear in the marginal revenue product and the wage, so root-n consistency of estimators for each of these components leads to root-n consistent estimates of the gap.

In many cases, an aggregate price deflator is used to translate observed firm revenues $P_{it}Q_{it}$ into units of quantity. We denote this aggregate price index $\overline{P}_t$. When $P_{it} \neq \overline{P}_t$, a new term enters into the residual:

$$\varepsilon_{it} = \omega_{it} + \eta_{it} + lnP_{it} - ln\overline{P}_t.$$

If this new source of error - $lnP_{it} - ln\overline{P}_t$ - is uncorrelated with input choices, then no further estimation questions are raised (beyond simultaneity of inputs and ω).[29] In this case, projecting $\hat{\varepsilon}_{it}$ on $\hat{\varepsilon}_{it-1}$ gives us a forecast of the expected value of ω *plus* the difference between firm price and aggregate price index, both of which we will generally want to condition on. Examination of the change in the absolute gap remains a valid indicator for efficiency changes.

The analysis does change if we forsake recovery of the production function parameters. From this perspective, the marginal effects of varying input levels on revenue are recovered from regressing the log of total revenue on log-levels of inputs. Setting aside estimation issues, the recovered coefficients from the revenue production function reflect both the marginal product and the markup. This means estimated gaps will reflect both $P_i * \frac{\partial Q_i}{\partial l}$ and a markup term given by $\epsilon_{P_{it}Q_{it}}$, the elasticity of price with respect to quantity. For example, with a single-product firm, using estimates of the revenue production function produces an estimated gap:

$$G_{it} = P_{it}(1 + \epsilon_{P_{it}Q_{it}}) * \frac{\partial Q_{it}}{\partial l} - w_{it}.$$

[28]This assumes the firm constructs its beliefs in a similar manner.

[29]One alternative, proposed in Griliches and Klette (1996) and used in De Loecker (2005), is to derive this term based on assumptions made on the demand system.

The new term reflects lost revenue on infra-marginal demand due to the fall in price caused by increasing quantity (by increasing labor).

If information on demand elasticities are available, they can be combined with the plant-level information to hold this part of the gap constant. The analysis can proceed as before, looking at the absolute value of the gaps. When demand elasticities are not readily available, we must restrict ourselves to looking at plants that have positive gaps in juxtaposed periods. For these firms, efficiency falls if the magnitude of the gaps increase. For other firms, an increase in the absolute value of the gap could be associated with a decrease in the difference between the value of the marginal product and the wage. Specifically, the existence of the markup term makes the analysis of the gap ambiguous for observations where both gaps are not positive.[30]

4 The Chilean Job Security Reforms

Workers in Chile have traditionally been provided with job security through three means: advance notices for dismissal, limitations on the use of fixed-term labor contracts, and severance payments on dismissal.[31] Over the 1979-1996 sample period, advance notice was unchanged at one month, and limitations on the use of fixed-term contracts changed only slightly. However severance payments changed substantially on two occasions, particularly for workers that were fired for "economic" reasons. We look at these changes for evidence of an impact on economic efficiency.

In 1978, the Pinochet administration extended severance payments from white collar workers to all workers. He also required firms to pay one month's wages per year of service (subject to no upper limit) for any worker dismissed for "unjustified reasons". Economic and financial needs of the firm were not "unjustified", so firms letting workers go during downturns in demand were not required to pay severance. The *Labor Plan* of 1980 mandated that severance packages be part of the overall job contract negotiated between the employee and the employer. The law also restricted the minimum severance package for "unjustified reasons" to one month's wages per year of service, subject to a maximum of five months.

We first look at the impact of the enhancement in job security that occurred

[30]Consider a monopolist who chooses output such that marginal revenue equals marginal cost. Now suppose the monopolist sets marginal revenue equal to marginal cost plus an additional positive (cost) term. Efficiency unambiguously falls as we reduce output.

Now consider the monopolist that starts with a large negative gap (in absolute value) and then is found to have a small positive gap after treatment. We cannot conclude that efficiency increased because the large negative gap could have been closer to efficient than the small positive gap. At the social optimum, marginal revenue will be less than marginal cost.

[31]This section draws heavily from Edwards and Edwards (2000).

in June 1984, when economic and financial needs were reclassified to "unjustified". Firms firing workers in a downturn (e.g.) after 1984 had to pay up to five months of wages in severance. In December 1990, the new democratic regime strengthened the provision, raising the maximum severance package from five to eleven months' wages, one month per year employed. It also placed the burden of the proof of "economic cause" on the employer, and charged the employer a further penalty when the case could not be established to the satisfaction of the court.

Edwards and Edwards (2000), drawing on Pagés and Montenegro (1999), construct an index that reflects the expected present value of the firing costs associated with hiring a given laborer. Let β denote the discount factor, δ the probability of retention, b the cost of advance notice, a the probability that the courts admit economic distress as a "just cause", S^J_{t+s} the payment under justified cause, and S^U_{t+s} the payment under unjustified dismissal. $\delta^{(s-1)} * (1-\delta)$ is then the probability of firing after s years. The index is given as

$$C_t = \sum_{s=1}^{T} \beta^s \delta^{(s-1)} * (1-\delta) * \left(b + aS^J_{t+s} + (1-a)S^U_{t+s}\right).$$

Figure 1 is calculated using their estimates of the parameters for Chile, where a falls in the first treatment period and S^U increases in the second treatment period.

The index suggests that the effective firing costs in the pre-1984 period was quite low, close to 0.75 months of wages, driven by the cost of advance notice. Job security was then significantly enhanced by the reform in mid-1984 and then again by the reforms of late 1990. Over our panel time-frame (1979-1995), we analyze three distinct periods: Period 1 (1979 to 1984) has relatively low levels of job security, period 2 (1985 to 1990) has higher levels and period 3 (post-1990) has the highest levels of job security. For 41 OECD and Latin American countries together, over the period of our sample Chile went from having one of the smallest levels of firing costs to an average expected cost of firing at hiring of slightly more than 3 months wages. While above the sample median of 2 months wages, these firing costs were not nearly as high as countries at the top of the distribution, like Colombia, Brazil, Peru, and Ecuador, which ranged from 10-14 months wages.

There are a number of other political and economic changes taking place over the sample period, many of which have been analyzed in their own right. The Labor Plan reduced payroll taxes substantially in 1981. Gruber (1997) reports that these reductions were fully passed on to wages with no effect on unemployment. The bargaining power of unions was relatively low through the 1980s under the military government, but increased under reforms introduced by the democratic regime in 1991. Using aggregate data and time series analysis, Edwards and Edwards (2000) find that reduction of payroll taxes and decentralization of bargaining increased labor market flexibility and contributed to a reduction in employment. Finally, there was a severe recession in 1982 related to the Latin American debt crisis and the fall in copper prices, a major Chilean export. The recovery was also quite remarkable, with wages increasing at 5% a year and unemployment falling from 17% to 5.5% in the post-recession period.

Our empirical approach uses the 1979-1984 as the control period, and compares the two treatment periods to this initial period to look for evidence of an impact on efficiency. We condition on as many potential confounding factors that we observe, like unemployment, plant-level and aggregate output, and plant-level wages and productivity. We also look closely at the times when the reforms were put in place - before and after - in order to try to control for other unobserved confounding factors.

5 The Data

We use the annual Chilean Manufacturing Census (Encuesta Nacional Industrial Anual) conducted by the Chilean government statistical office (Instituto Nacional de Estadistica). The survey covers all manufacturing plants in Chile with more than 10 employees and has been conducted annually since 1979. Our data covers the seventeen year period from 1979 to 1996. There are about 5000 firms every year, with an entry rate and exit rate of about 5 percent over the panel period.

This survey has been used in a number of previous studies.[32] The survey provides an industry indicator, and measures of output, inputs, wages, employment and investment. A detailed description of how the longitudinal samples were combined into a panel from 1979-1986 can be found in Liu(1991). We extended this to 1996 following broadly the procedure used by Liu. Further, we supplemented the raw data files with additional data on price series for output, machinery and inputs from other sources including IMF's IFS database, data on price indices obtained from the Chilean government statistical office, and also with data from Edwards and Edwards (2000) and Edwards and Edwards (1991).

Plant-level real output is total revenue deflated with a 4-digit industry output deflator obtained from the Web site of the Chilean Government's statistical office. Industry real output is the sum of these plant-level real outputs by 4 digit (ISIC)

[32]See Levinsohn and Petrin (2003), Pavcik (2000), Roberts and Tybout (1996), Hsieh and Parker (2006)), for example.

industry. Real materials and fuels are both aggregates at the plant-level, and each have their own 3-digit price deflator. Electricity usage is separately reported from other energies.[33]

At each firm we calculate the nominal wage rate for each labor type as total employee costs divided by the number of those employees. The components of the wages are given as Wages, Bonus, Payroll Taxes, and Family Allowance Taxes, and do not appear to explicitly account for any firing costs. The real wage rate is obtained by deflating the wage rate using the output deflator.

The real capital series is constructed using the perpetual inventory method, following Liu (1991). Data on book value of capital is available for the years 1980-81 and 1992-96. We use the same methodology as Liu (1991) to construct the capital series for all firms for which we have data on book value for 1980-1991. For other firms, we build capital series backward and forward using the data on book value available for 1992-96. As in Liu, we assume 10 percent depreciation rate for machinery, 5 percent depreciation rate for buildings, 20 percent depreciation rate for vehicles and zero depreciation for buildings. We use a deflator for the construction sector to deflate investments in buildings and use a deflator for machinery to deflate investments in both machinery and vehicles. The capital series we use is defined as the capital series constructed using the 1980 base year, where missing values are replaced using the capital series constructed using 1981, 1992, 1993, 1994, 1995 and 1996 in that order.

We examine general trends in real wage rates in Figure A2 and unemployment and inflation rates in Figure A3. We find that both blue and white collar real wages dropped till about the mid 1980s and then grew through the late 1980s and early 1990s, starting after unemployment levels out around 5% in 1987. The spike in unemployment rates in the early 1980s seem to be driven by the deep recession around the same period. The overall trend in the real wage rate seems to be strongly influenced by macroeconomic trends.

6 The Gap Results for Chile, 1979-1996

Our goal is to estimate the marginal revenue product and the marginal input price. For the marginal revenue product, we estimate the production function separately by 3 digit ISIC code to allow for flexibility in the production function across industries. In total, there are 27 industries with more than 100 observations each over the sample period that we use. For our baseline analysis, we posit the gross output production

[33] We thank Andrés Hernando for providing us with these deflators.

function given above, estimating its parameters using OLS with fixed effects. This imposes that $\omega_{it} = \omega_i$ for all t.

We also use the value-added formulation for the Levinsohn-Petrin (2003) estimator (LP), which is easy to calculate, and addresses the simultaneity problem raised in Marchak and Andrews (1944).[34] The value-added formulation is also useful because it limits the amount of information necessary to calculate the derivative, which adds back more than 30,000 observations relative to the gross output production function setting.[35] The simultaneity correction allows for greater flexibility on the error specification relative to fixed effects, as ω_{it} is allowed to vary over time and be correlated with input choices.[36]

The second component of the Gap is the marginal cost for the input. For labor, it is the wage that must be paid for an additional year of labor. As described earlier, we observe total annual wage bill and divide by the number of laborers to get an average wage. The wage bill and number of workers is available for blue and white collar workers, allowing us to estimate a plant-level wage for each labor type. We assume this average wage is the marginal wage. That is, we assume that firms face a perfectly elastic supply of labor at observed average wage, ruling out well-known phenomena like overtime premiums, which add measurement error to our gap.[37]

Table 1 reports the Gap analysis for blue collar labor, white collar labor, and materials. In each column the absolute value of the gap for the input is the dependent variable. All regressions include plant-level fixed effects and two period-indicators for the different degrees of job security, one for 1985-1990, and one for 1991-1996. Columns 2, 4, and 6 also include the industry output growth rate as a control. The fixed effects allows for base-period plant-specific gaps, and so the magnitudes of the period dummies are identified by within-plant variation in the gap over time.

The regressions suggest that the base-period average gap for blue (white) collar labor was 111 (162) thousand pesos per year, approximately equal to the annual wage of one blue (white) collar worker. In the first period of treatment - with a maximum of five months severance pay - the blue collar gap increases by 24 thousand pesos/year, while the white collar gap shows no significant change. In the second period - with a maximum eleven months firing costs - the blue collar gap increases another 17

[34] Other proxy alternative methods are Olley and Pakes (1995) and the modified-LP estimator suggested by Ackerberg, Caves, and Fraser (2006).

[35] Many observations are missing for fuels, which makes estimation of the derivative unfeasible, and this severely attenuates the gross output production function sample size.

[36] Stata code is available for this setup, the gross output production function setup, and other formulations of the Levinsohn-Petrin estimator. See Petrin, Levinsohn, and Poi (2004).

[37] If fixed costs of employment are included in these numbers they too add measurement error to the calculation of the marginal wage.

thousand pesos/year to 41 thousand pesos/year. This second change in job security also increases the white gap by 16 thousand pesos/year.

The results in Table A1 are from the value-added Levinsohn-Petrin specification. The blue collar results are similar, except the intercept drops from 111 to 55 thousand pesos. The white collar results are starker, with the first period associated with an increase of 40 thousand pesos/year, and the second period jumping to 63 thousand pesos/year. Virtually all of the point estimates from both tables are statistically significant at standard testing levels.[38]

Figures 2 and A1 plot the coefficients on the year dummy variables that come from regressions similar to those reported in Table 1, but include the year-to-year dummy variables as opposed to just the two period controls. The 1980 coefficient is normalized to one. The timing of the results are consistent with the timing of the job security changes. In Figure 2, the labor gaps level off in post-recession 1982-83, are approximately flat until 1985, when they increase significantly in 1986-87 after the first application of job security. The gaps appear to level off in 1988 at the higher level of 132 and 202 thousand pesos respectively. Then, with the next application of job security in 1990, both labor gaps start to gradually increase, and do so until 1994, where they level off at 152 and 225 thousand pesos respectively. The results from value-added specification that controls for simultaneity are similar, although there are 30,000 additional observations (see Figure A1), with starker predictions for increases in gap magnitudes, although there are two dips, one prior to 1990, and one prior to 1992.[39]

We also look at the firms that have only positive gaps in juxtaposed years. For the OLS sample, for both blue and white collar labor approximately 3/4 of the observed gaps are positive. Of those firms that have positive gaps this period and are observed in the previous period, about half of the gaps get bigger. This is true for both blue and white collar workers. For these observations, the average increase in the gap across the three periods for blue collar labor goes from 68.73 to 84.21 thousand pesos. The white collar average increase in the gap goes from 85.95 to 110.73 thousand pesos.

As a control input we examine the gap for materials. Since materials are unaffected by the job security provisions, we should expect to find little change in the materials gap. Table 1 and Figure 2 show that there does appear to be a fall in the gaps early in the first period during the recession, but after 1982 the gaps are flat and constant

[38]The average (median) unconditional gaps for blue collar labor across periods are 111.22 (38.78), 140.70 (50.40), and 148.44 (45.01). The average (median) unconditional gaps for white collar labor across periods are 151.96 (87.18), 162.45 (87.58), and 187.98 (100.80). These results are robust to using a trans-log specification.

[39]We have also run the OLS with fixed effects specification dropping fuels as an input. The 30,000 extra observations do not change the flavor of the results.

over time. We conclude that there is no evidence that the materials gap increased concurrently with the white and blue collar gaps.[40]

Overall, both Tables and Figures are strongly suggestive of a negative impact on efficiency arising post-1985 and again in post-1990, with the most pronounced effects occurring right after the application of the job security provisions.

7 Traditional Labor Demand Statistics for Chile

We look for an impact of job security increases using traditional approaches. We start with statistics from the labor demand equation. We also ask whether the probability of adjustment and the size of adjustment change as firing costs increase. Finally, we look to see if wages change in a way consistent with the necessary offset of the job security provisions.

7.1 Labor Demand

We use the widely adopted log-linear regression specification (see, e.g. Blundell and Bond (1998)).[41]

$$l_{it}^u = \beta_0 + \beta_1 w_{i,t}^u + \beta_2 w_{i,t-1}^u + \beta_3 w_{i,t}^s + \beta_4 w_{i,t-1}^s + \beta_5 q_{i,t} + \beta_6 q_{i,t-1} + \beta_7 l_{i,t-1}^u + \epsilon_{it} \quad (29)$$

where i and t index firm and time, w^u is the log of unskilled (blue collar) wage rate, w^s is the log of skilled (white collar) wage rate, l^u is the log of unskilled (blue collar) employment and q is the log of value added.

We look at three statistics. Changes in the intercept and the slope coefficients suggest a change in demand, either through a change in the level or to the sensitivity in price (wage). Increases in the coefficient on lagged labor are also suggestive of a negative impact, as last period's labor becomes a better predictor of this period's level. As have others in the literature, we look for these effects.

We estimate the dynamic labor demand equations using three approaches: OLS, instrumental variables, and fixed effects. The instrumental variable approach, which many may regard as the preferred specification, follows Blundell and Bond (1998). We consider as endogenous blue collar wage, white collar wage, value added and lagged blue collar employment. The instruments we use are lagged industry output,

[40] We find a similar pattern for the electricity gaps. The fuels gap is unchanged in the first period, but then does increase in the second period.

[41] One motivation is given by considering a firm in a competitive environment with a Cobb-Douglas production function and an AR(1) disturbance. Similar dynamic labor demand equations have been estimated in a number of studies, including Hammermesh (1993), Sevestre and Trognon (1996), Fajnzylber and Maloney (2000), and the literature cited in Heckman and Pages (2004).

lagged industry average blue collar wage, lagged industry average white collar wage, two-period lagged industry output, two-period lagged industry average blue collar wage, two-period lagged industry average white collar wage, two-period lagged blue collar wage, two-period lagged white collar wage, two-period lagged materials, two-period lagged capital, two-period lagged value added and three-period lagged blue collar employment.

We recognize the potential problems with using lagged variables as instruments, but note that it is standard practice in the literature, in part no doubt because other instruments are not readily available. Indeed, weak instruments, or instruments which are correlated with the error, are two further reasons the labor demand equation may be problematic as a statistic of interest.

Table 2 presents the results for the instrumental variables specification and the OLS specification (fixed effects is in Table A2 in the appendix). In estimating the demand functions for each of the three relevant periods, we use two alternative approaches. First, we allow only the intercepts and the lagged blue collar employment term to vary by period. Second, we allow all the coefficients to vary by period. Both have been used in the literature so we look at the performance of both in the Chilean data.

Table 3 reports the IV estimates. We start with the restricted IV specification reported in column 1. The period intercepts increase after the first treatment and then decrease after the second. The lagged labor coefficient also first increases towards one, and then falls from 1. In the full specification, the intercept increases and then decreases. The coefficient on lagged labor monotonically increases towards 1, from 0.22 to 0.66. The price sensitivity parameter (on wage) first increases (in absolute) value, and then decreases. Overall, these results are hardly conclusive, although they may be suggestive.

Table 4 reports the OLS estimates. For the constrained model, OLS shows decreasing demand (intercept) over time, but an increasing then decreasing coefficient on lagged labor. The unconstrained model has a decreasing then increasing intercept, an increasing price sensitivity, and a very slightly increasing effect on the lagged labor coefficient (approximately a 2% increase in the coefficient). Fixed effects results are also mixed, and are reported in Table A2.[42] There seem to be few common themes across the different estimators.

[42]For the constrained model, the intercept dummy is first decreasing then increasing, and the lagged labor coefficient increases and then is flat. For the unconstrained model, the intercept decreases then increases, and the coefficient on lagged labor increases then decreases.

7.2 Probability of Adjustment

In Table 4, we examine whether firms less frequently change employment levels as firing costs increase. This statistic is motivated by Kugler (1999), for example, and others cited therein. The theory from Section 2 in this paper contrasts to the 2-period competitive case.[43] In this paper's more general setting, the probability of not adjusting is not necessarily increasing in firing costs.

With the plant-level data, one might view this demand side exercise as the supply-side "equivalent" of asking whether duration of (un)employment increases? The empirical evidence in that literature appears also to be mixed, although employment-level surveys seem to be more in agreement on increasing periods of duration being associated with increases in firing costs (see Heckman and Pages(2004)).

The dependent variable in our regressions equals one for firms that did *not* change net employment levels from the previous year.[44] Columns 1 and 2 indicate that the probability of not adjusting fell from 17% to 13% as firing costs increased, the opposite of what the two-period model predicts. With fixed effects there is no change in probabilities as job security increases. If we use the production function estimates to condition on productivity residuals, we find a small *increase* in the probability of not adjusting using the LP residual, although not statistically significant. Using the OLS residual, the probability of not adjusting increases 1.2% in the first treatment and then to 3.6% in the second treatment, although only the latter number is statistically significant.

7.3 Magnitudes Conditional on Adjusting

We look at whether fires conditional on firing decrease, and similarly, whether hires conditional on hiring decrease. The dependent variable is the year-to-year percentage change in the level of labor. Table 5 presents the results for the case of firing. Unconditionally, the magnitude of firing falls by 1.9% with the first dose, and then falls by 2.6% on the second dose (relative to the initial period). The results are similar with firm fixed effects. Once we condition on either productivity residual, the effects become more pronounced, increasing to -3.4% in the first period and up to -5.7% in the second period.

Table 6 present the hiring results. Unconditionally, in the first period hiring decreases by 1.2%, and the second period decreases further to 4.5%. With firm fixed

[43] Kugler (1999) uses differences in differences using covered vs. non-covered workers.

[44] We only observe net hires, so we (like much of the literature) can only talk about change in net employment levels. See Hammermesh and Pfann(1996).

effects, the numbers fall to 2.3% for the first period and 7.4% for the second period. However, conditional on either productivity residual, number of hires *increases*. For the OLS-residual, hiring increases 2.3% in the first period and 2.6% in the second period.

7.4 Wages

For an estimate of the per period reduction in wages required to offset the two job security changes introduced in Chile, we consider two "insurance" plans. Under the first, expected firing costs are recovered through premium payments over the lifetime of the worker in the firm. Under the second, the firm insures against the possibility of firing workers period by period.

In Appendix 1, we estimate the fair premium under these two insurance schemes for a change in job security equivalent to six months wages (comparable to the maximum increase in both of the Chilean changes). Our estimates suggest that a drop in wages in the range of 3% to 6% could provide the necessary offset.

To try to separate out the effect of job security changes on wages, we regress the estimated plant-level average real wage on period controls for the job security changes. The other controls include firm fixed effects, firm output growth rate, industry output and industry growth rate, and the unemployment rate. Unfortunately, we do not observe worker-specific covariates.

We report the estimates in Table 7. In all the specifications, there is a major decline in wages in period 2 (1985-1990). The extent of the decline, between 36% and 53%, is much larger than that required under our offset plans. In period 3 wages recover somewhat. Overall, there is no clear evidence that the job security changes were offset through lower wage rates.

8 Conclusions and Extensions

Firing costs are commonplace in both the developing and developed world. In theory, the impact of firing costs can be completely undone by efficient contracts. To date the extensive empirical macro- and micro-level literature on the impact of these provisions is largely inconclusive, as are the results we and others have obtained for Chile using these same traditional statistics.

We argue that the mixed evidence is a consequence of the weak power of statistics being used. They are motivated by a simple 2-period model, the implications of which do not necessarily extend to a world with imperfect competition, strategic interactions, non-convex adjustment costs, or the infinite-horizon. The theory we develop also points to how - in these more general settings - unobserved variables and aggregation bias can weaken the test, if not destroy its econometric validity.

The theory is clear: firing costs drive a wedge between the marginal revenue product of labor and its marginal cost. We develop a new test statistic, showing how to directly estimate this "gap" using plant-level production data. Our statistic has a clear link to welfare, is easy to compute, and is robust to settings outside the 2-period model and to many of the aggregation and estimation problems that afflict the more traditional statistics. There also exist many "control" inputs whose gaps should not increase in response to firing costs. We find large and statistically significant increases in the gaps between the marginal product of labor and the wage for both white and blue collar workers, but little positive effect on the non-labor inputs.

We see three future directions. We have provided only one empirical example. It remains to be seen whether the gap finds evidence in other plant-level micro datasets with changes in firing costs. These exercises will be possible given the increasing availability of plant-level data and the widespread application of firing costs around the world. More generally, the gap can be used to examine other policies that introduce distortions. By construction, if the gap is increasing, then willingness to pay and cost of production are getting further apart.

A second direction is to look at data from an industry where markups are known to be important and where information on both demand and supply is available. For example, in the automobile industry, coupling demand and production information together, one could hold constant the markup terms when examining the gaps. These types of case-studies would inform us on the amount of noise added by the markup terms in cases where we suspect markups are important but we do not observe demand side data.

A final question is whether we can aggregate the gaps to a quantity that is tightly linked to changes in aggregate welfare. A natural setup within which to start is that outlined in Petrin and Levinsohn (2005), who show how to aggregate changes in plant-level productivity to recover the change in aggregate welfare. Alternatively, a more structural approach that estimates the complete dynamic program to recover labor demand, as in Aguirregabiria and Alonso-Borrego (1999) and Roata (2004), may prove amenable to calculating the change in welfare.

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A Side payment plans that offset firing costs

A.1 Plan 1: Insuring over the worker's life time

Under this plan, premia are collected over the worker's lifetime with the firm to offset the expected firing costs. The fair premia α_j for worker j, as a fraction of annual wages, is calculated by setting the expected present value of the dismissal costs equal to the present value of the premia collected:

$$\sum_{s=1}^{T} \beta^s \delta_j^s (1-\delta_j)(y_{j,t+s}) = \sum_{s=0}^{T-1} \beta^s \delta_j^s \alpha_j W_j$$

where β is the discount factor, δ_j is the probability of worker j being retained, $y_{j,t+s}$ is the severance cost in annual wages of firing worker j at end of s years, and T is the maximum tenure. Assuming that all workers in a firm have identical wages and dismissal probabilities, we can calculate the drop in wage levels (ie the premium payments) required to offset any increase in dismissal costs. We estimate how large the fall must be to offset the job security reforms introduced in Chile in 1984, assuming the interest rate (for discounting) is 5% and the maximum tenure is 20 years.

Current tenure	Dismissal rate	Implied wage change in year 1
all new	10%	**-3.09%**
all new	15%	**-4.25%**
all new	20%	**-5.19%**
all > 5 years	10%	**-4.17%**
all > 5 years	15%	**-6.25%**
all > 5 years	20%	**-8.33%**

A.2 Plan 2: Insuring period by period over the pool of workers

In this approach, the firm's expected firing cost for each period is insured by collecting a premium from all the workers of the firm. Assuming the same fraction of wages is collected from each worker, the fair premium in this case is obtained by setting:

$$\sum_{j=1}^{N_j} \delta_j y_j = \alpha \sum_{j=1}^{N_j} W_j$$

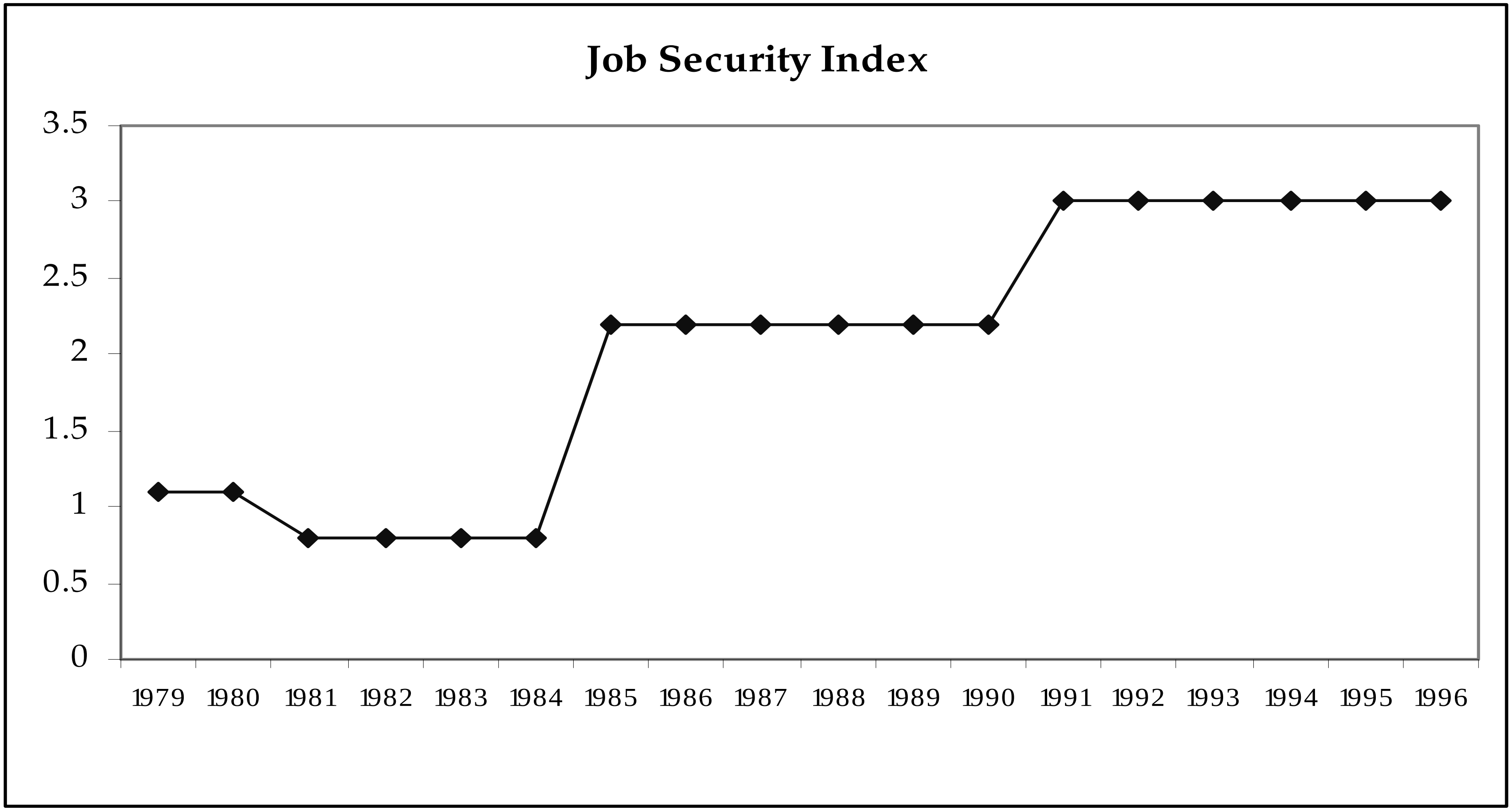
where δ_j is the probability of worker j being retained, y_{jt} is the severance cost in annual wages of firing worker j and N_j is the number of workers in firm j. Assuming that the workers in the firms are identical, we obtain the required drop in wage levels to pay for the insurance premia that offsets the 1984 increase as:

Current tenure	Dismissal rate	Implied wage change in year 1
all new	10%	**-0.83%**
all new	15%	**-1.25%**
all new	20%	**-1.67%**
all > 5 years	10%	**-4.17%**
all > 5 years	15%	**-6.25%**
all > 5 years	20%	**-8.33%**

If workers are identical and $y_{j,t+s}$ is constant over all j (as in the case where the tenure of all workers exceeds 5 years), the premium payments are the same for both plans and given by $(1-\delta)y$.

Since we expect the current average tenure of typical firm to be between the extremes considered in the tables above, we guess that the fall in wages required to neutralize the Chilean 1984 dismissal cost might lie in the range of 3 to 6 percent. Since the second job security change increases the maximum dismissal cost from 5 months to 11 months, we expect the additional drop required to be similar in magnitude.

Figure 1
The Change in Firing Costs in Chile
Expected discounted cost of dismissing a worker, in multiples of monthly wages



Source: Edwards and Edwards (2000).

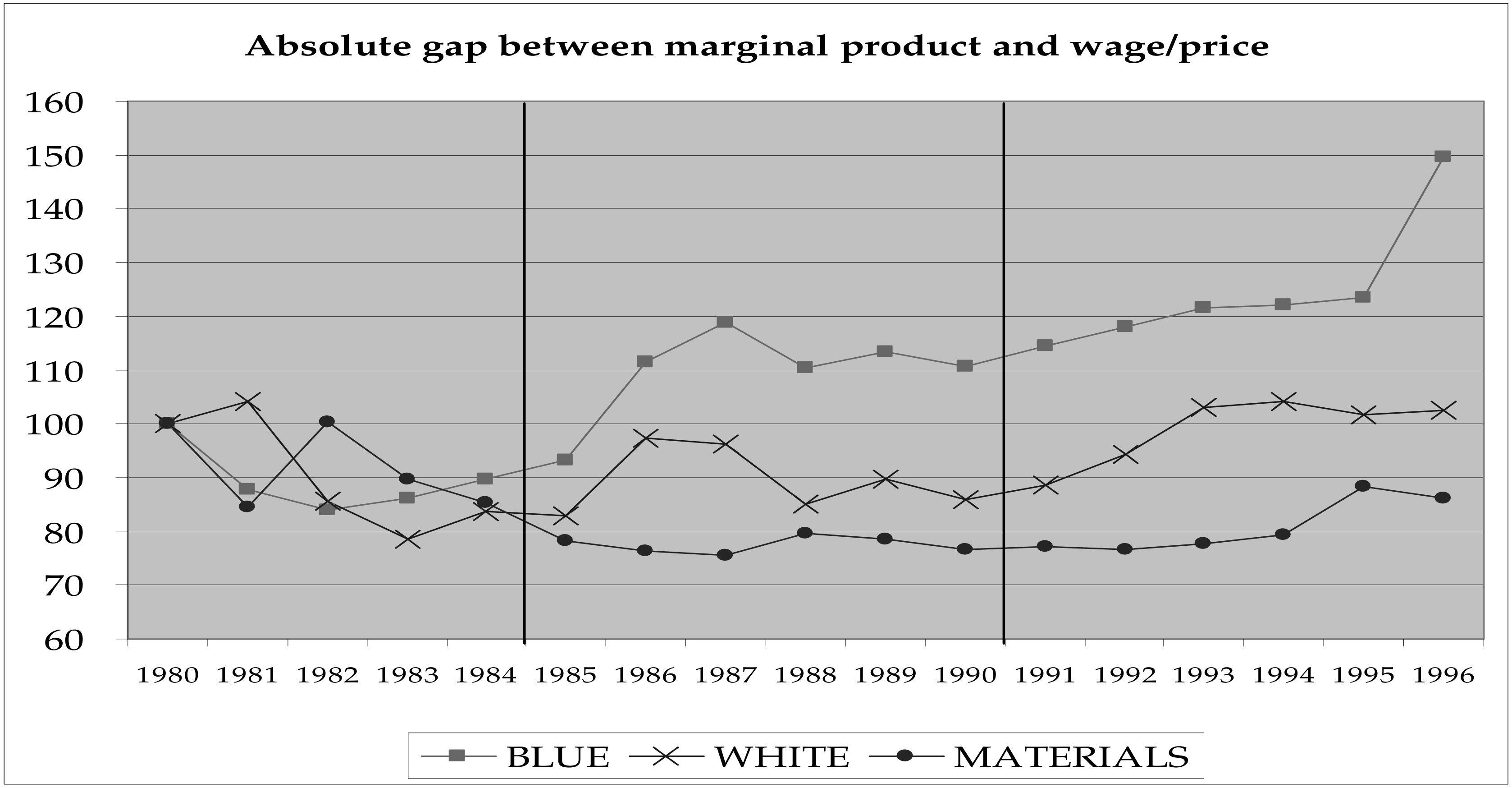
Table 1
Absolute Value of Plant-level Gap, OLS

Gap = Marginal Revenue Product - Marginal Cost

Dependent Variable	**Blue Gap**		**White Gap**		**Materials Gap**	
	(1)	(2)	(1)	(2)	(1)	(2)
Constant	111.467	111.678	162.815	163.149	0.283	0.283
	[9.551]**	[9.551]**	[3.637]**	[3.617]**	[0.010]**	[0.010]**
Period Dummy (1985-1990)	24.302	23.861	-0.164	-0.827	-0.044	-0.043
	[9.815]*	[9.811]*	[7.049]	[7.092]	[0.014]**	[0.014]**
Period Dummy (1991-1996)	41.459	41.005	16.857	16.203	-0.035	-0.034
	[16.615]*	[16.633]*	[6.581]*	[6.515]*	[0.012]**	[0.012]**
Industry Output Growth Rate		3.415		4.803		-0.01
		[2.507]		[3.070]		[0.008]
Firm Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	28,641	28,634	28,643	28,636	29,485	29,478
R-squared	0.79	0.79	0.64	0.64	0.54	0.54
Number of clusters	66	66	66	66	67	67

All variables are in thousands of 1979 pesos. Blue Gap is defined as the absolute value of the gap between the marginal product of blue collar labor and the blue collar wage rate, normalized by the 3-digit industry output price deflator. The marginal product of blue collar labor is estimated based on a gross output Cobb-Douglas production function, which is estimated using OLS with firm fixed effects. The blue collar wage rate is obtained by dividing the total blue collar wage bill by the number of blue collar employees. White Gap is defined similarly. Materials Gap is defined as the absolute value of the gap between the marginal product of materials and the normalized materials price index (also defined at the three digit industry level). The underlying production functions used for calculating the gaps are estimated separately for each 3 digit industry. Standard errors are clustered at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Figure 2
Differences Across Years in the Absolute Gap
Blue and White Collar Labor and Materials, OLS estimates



The graph plots the coefficient on year dummies in regression of absolute gap between marginal product of an input and its normalized price, including firm fixed effects and industry output growth rate. Year 1980 is normalized to 100.

Table 2
Labor Demand Regression
Blue Collar Employment, Instrumental Variables

	Blue Collar Labor			
	Period 1 (1979-84)	Period 2 (1985-90)	Period 3 (1991-96)	All Periods
Constant	-0.015	0.016	-0.044	-0.024
	[0.010]	[0.006]*	[0.006]**	[0.006]**
Blue wage	-0.123	-0.235	-0.104	-0.191
	[0.037]**	[0.025]**	[0.043]*	[0.017]**
Blue wage_t-1	-0.047	0.082	0.238	0.066
	[0.016]**	[0.019]**	[0.023]**	[0.011]**
White wage	0.117	0.023	0.029	0.054
	[0.023]**	[0.019]	[0.019]	[0.010]**
White wage_t-1	-0.026	-0.034	-0.023	-0.048
	[0.012]*	[0.008]**	[0.010]*	[0.005]**
Value added	0.104	-0.002	0.075	0.081
	[0.010]**	[0.024]	[0.017]**	[0.009]**
Value added_t-1	0.04	-0.006	-0.015	0.012
	[0.006]**	[0.006]	[0.008]+	[0.004]**
Blue Employment_t-1	0.12	0.435	0.664	0.457
	[0.055]*	[0.093]**	[0.071]**	[0.056]**
Period Dummy (1985-1990)				0.015
				[0.010]
Period Dummy (1991-1996)				0.005
				[0.009]
Period Dummy (1985-1990)*Blue Employment_t-1				0.35
				[0.106]**
Period Dummy (1991-1996)*Blue Employment_t-1				-0.283
				[0.092]**
Observations	8129	15272	18499	41900

Dependent variable is the log (Blue Collar Employment). All other variables are also in logs. The regressions are estimated in first differences. The endogenous variables are Blue Wage, White Wage, Value Added and Blue Employment_t-1. Instruments are lagged industry output, lagged industry average Blue Wage, Lagged industry average White Wage, two-period lagged industry output, double lagged industry average blue wage, double lagged industry average white wage, two-period lagged Blue Wage, two-period lagged White Wage, two-period lagged Materials, two-period lagged Capital, two-period lagged Value Added and three-period lagged Blue Collar employment. Standard errors are adjusted for clustering at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Table 3
Labor Demand Regression
Blue Collar Employment, OLS

	OLS			
	Period 1 (1979-84)	Period 2 (1985-90)	Period 3 (1991-96)	All Periods
Constant	0.174	0.082	0.095	0.141
	[0.038]**	[0.026]**	[0.030]**	[0.015]**
Blue wage	-0.194	-0.274	-0.296	-0.25
	[0.016]**	[0.018]**	[0.022]**	[0.007]**
Blue wage_t-1	0.095	0.177	0.202	0.154
	[0.017]**	[0.018]**	[0.017]**	[0.007]**
White wage	0.084	0.057	0.042	0.064
	[0.009]**	[0.011]**	[0.014]**	[0.004]**
White wage_t-1	-0.08	-0.048	-0.034	-0.057
	[0.008]**	[0.010]**	[0.012]**	[0.004]**
Value added	0.108	0.101	0.124	0.111
	[0.006]**	[0.007]**	[0.008]**	[0.003]**
Value added_t-1	-0.031	-0.02	-0.049	-0.033
	[0.004]**	[0.007]**	[0.007]**	[0.003]**
Blue Employment_t-1	0.859	0.879	0.88	0.858
	[0.007]**	[0.010]**	[0.008]**	[0.004]**
Period Dummy (1985-1990)				-0.047
				[0.011]**
Period Dummy (1991-1996)				-0.056
				[0.012]**
Period Dummy (1985-1990)*Blue Employment_t-1				0.024
				[0.003]**
Period Dummy (1991-1996)*Blue Employment_t-1				0.019
				[0.004]**
Observations	17650	20416	23155	61221
R-squared	0.91	0.92	0.9	0.91

Dependent variable is the log (Blue Collar Employment). All other variables are also in logs. Standard errors are adjusted for clustering at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Table 4
Probability of Not Adjusting Labor

	(1)	(2)	(3)	(4)	(5)	(6)
Constant	0.179	0.179	0.146	0.146	0.172	0.17
	[0.013]**	[0.013]**	[0.006]**	[0.007]**	[0.023]**	[0.032]**
Period dummy (1985-1990)	-0.044	-0.044	-0.003	-0.003	0.001	0.012
	[0.008]**	[0.008]**	[0.009]	[0.009]	[0.011]	[0.011]
Period dummy (1991-1996)	-0.044	-0.044	0.009	0.008	0.017	0.036
	[0.006]**	[0.006]**	[0.010]	[0.010]	[0.015]	[0.018]*
Industry output growth rate		-0.002		0.001		
		[0.003]		[0.003]		
Lagged labor					0	0
					[0.000]**	[0.000]**
Firm wage rate					0	0
					[0.000]	[0.000]+
Observed productivity (LP estimate)					-0.003	
					[0.005]	
Observed productivity (OLS estimate)						-0.008
						[0.013]
Fixed effects	None	None	Firm	Firm	Firm	Firm
Observations	73,705	73,689	73,705	73,689	55,204	31,891
R-squared	0	0	0.25	0.25	0.24	0.25
Number of clusters	89	89	89	89	60	84

Dependent variable is one if the firm does not adjust its employment level (compared to previous year). All other variables are in logs. Standard errors are adjusted for clustering at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Table 5
Firing Rates Conditional on Firing
Percentage change

	(1)	(2)	(3)	(4)	(5)	(6)
Constant	0.172	0.17	0.174	0.169	0.232	0.208
	[0.005]**	[0.005]**	[0.004]**	[0.003]**	[0.026]**	[0.022]**
Period dummy (1985-1990)	-0.019	-0.016	-0.025	-0.019	-0.034	-0.03
	[0.005]**	[0.005]**	[0.008]**	[0.007]**	[0.010]**	[0.013]*
Period dummy (1991-1996)	-0.026	-0.023	-0.026	-0.02	-0.057	-0.06
	[0.004]**	[0.004]**	[0.005]**	[0.004]**	[0.010]**	[0.014]**
Industry output growth rate		-0.011		-0.025		
		[0.005]*		[0.006]**		
Lagged labor					0.00	0.00
					[0.000]**	[0.000]**
Firm wage rate					0.00	0.00
					[0.000]**	[0.000]**
Observed productivity (LP estimate)					-0.02	
					[0.006]**	
Observed productivity (OLS estimate)						-0.028
						[0.007]**
Fixed effects	None	None	Firm	Firm	Firm	Firm
Observations	28,908	28,898	28,908	28,898	21,216	12,557
R-squared	0.01	0.01	0.37	0.37	0.39	0.410
Number of clusters	89	89	89	89	60	81

Firing rate is the absolute percentage decrease in employment, defined for firms that decrease employment. Standard errors are adjusted for clustering at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Table 6
Hiring Rates Conditional on Hiring
Percentage Changes

	(1)	(2)	(3)	(4)	(5)	(6)
Constant	0.244	0.244	0.258	0.258	0.297	0.269
	[0.011]**	[0.011]**	[0.007]**	[0.007]**	[0.024]**	[0.033]**
Period dummy (1985-1990)	-0.012	-0.013	-0.023	-0.023	0.014	0.023
	[0.006]+	[0.006]*	[0.007]**	[0.007]**	[0.006]*	[0.007]**
Period dummy (1991-1996)	-0.045	-0.046	-0.074	-0.074	0.004	0.026
	[0.010]**	[0.010]**	[0.013]**	[0.013]**	[0.013]	[0.009]**
Industry output growth rate		0.01		0.004		
		[0.007]		[0.005]		
Lagged labor					0.00	0.00
					[0.000]**	[0.000]**
Firm wage rate					0.00	0.00
					[0.000]**	[0.000]**
Observed productivity (LP estimate)					0.007	
					[0.005]	
Observed productivity (OLS estimate)						0.02
						[0.014]
Fixed effects	None	None	Firm	Firm	Firm	Firm
Observations	33,903	33,899	33,903	33,899	25,706	14,585
R-squared	0.01	0.01	0.34	0.34	0.39	0.410
Number of clusters	88	88	88	88	59	82

Hiring rate is the percentage change in total employment, defined for firms that increase employmen. Standard errors are adjusted for clustering at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Table 7
Explaining Movements in Real Wages

	(1)	(2)	(3)	(4)	(5)
Period Dummy (1985-1990)	-0.375	-0.378	-0.405	-0.363	-0.533
	[0.026]**	[0.024]**	[0.022]**	[0.027]**	[0.020]**
Period Dummy (1991-1996)	0.081	0.072	-0.007	0.092	-0.112
	[0.023]**	[0.022]**	[0.025]	[0.023]**	[0.028]**
Firm Output Growth Rate		0.008			
		[0.002]**			
Log(Industry Output)			0.127		
			[0.021]**		
Industry Output Growth Rate				0.001	
				[0.006]	
Unemployment Rate					-2.131
					[0.193]**
Constant	4.494	4.52	2.52	4.486	4.814
	[0.014]**	[0.014]**	[0.319]**	[0.016]**	[0.026]**
Observations	86176	73701	86160	80346	86176
R-squared	0.8	0.82	0.81	0.81	0.81
Number of clusters	89	89	89	89	89

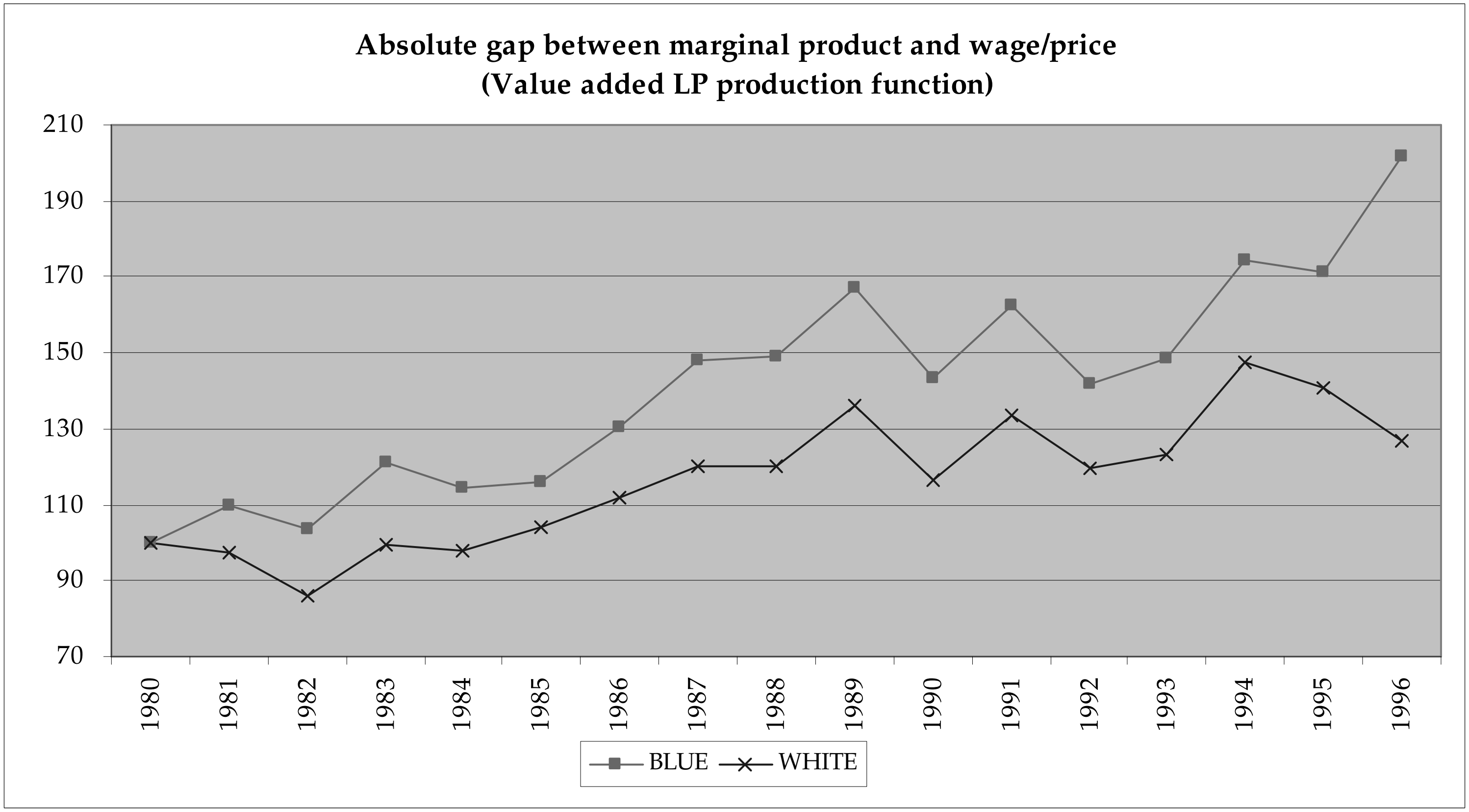
Dependent variable is Log (real wage rate). Real wage rate is the nominal wage rate deflated by the producer price index. Nominal wage rate is defined as the total wage bill/ number of employees. For each independent variable the first row gives the coefficient values and the second row gives the related t-values. All regressions include firm fixed effects. Standard errors are adjusted for clustering at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Table A1
Absolute Value of Gap Using Levinsohn-Petrin Value-Added Production Function Estimator

Dependent Variable	Blue Gap		White Gap	
	(1)	(2)	(1)	(2)
Constant	55.63	57.258	172.187	174.762
	[6.079]**	[6.274]**	[6.386]**	[6.604]**
Period Dummy (1985-1990)	17.766	15.959	40.423	37.414
	[8.786]*	[8.548]+	[7.600]**	[7.215]**
Period Dummy (1991-1996)	28.576	26.866	63.74	61.125
	[9.549]**	[9.313]**	[12.130]**	[11.929]**
Industry Output Growth Rate		1.685		18.537
		[2.351]		[5.287]**
Observations	63,824	60,021	63,868	60,064
R-squared	0.74	0.75	0.68	0.69
Number of clusters	60	60	60	59

All variables are in thousands of 1979 pesos. Blue Gap is defined as the absolute value of the gap between the marginal product of blue collar labor and the blue collar wage rate, normalized by the 3-digit industry output price deflator. The marginal product of blue collar labor is estimated using a value-added Levinsohn-Petrin estimator to address simultaneity. The blue collar wage rate is obtained by dividing the total blue collar wage bill by the number of blue collar employees. White Gap is defined similarly. The underlying production functions used for calculating the gaps are estimated separately for each 3 digit industry. Standard errors are clustered at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Figure A1
Differences Across Years in the Absolute Gap
Blue and White Collar Labor
Levinsohn-Petrin Estimator



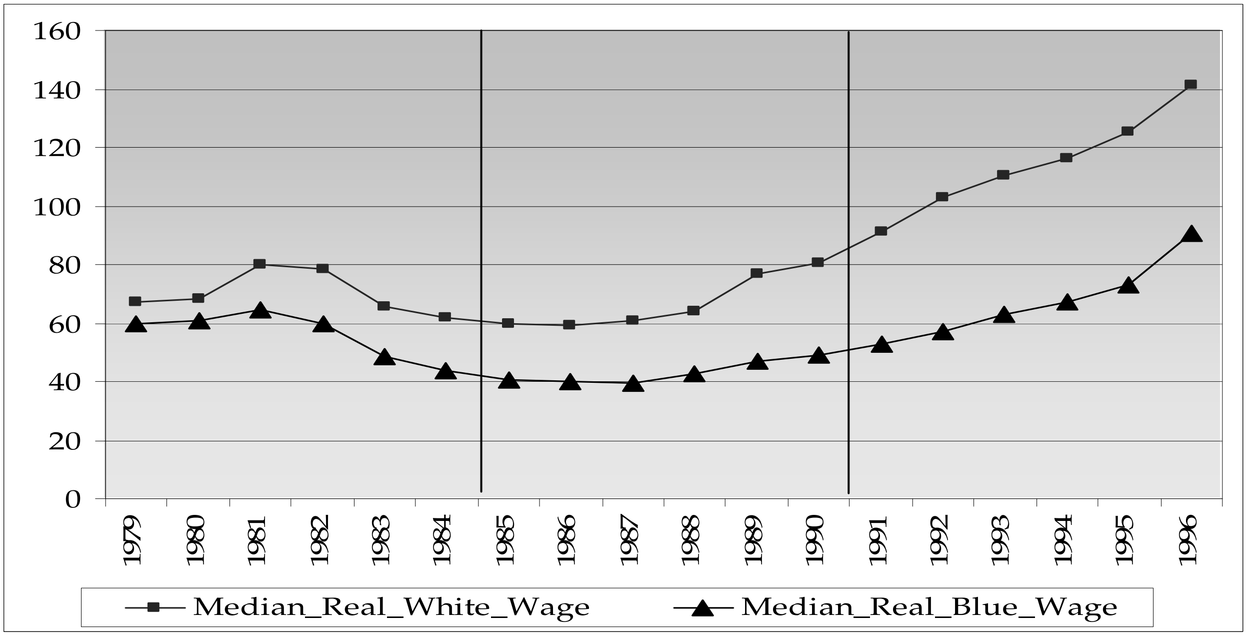
The graph plots the coefficient on year dummies in regression of absolute gap between marginal product of an input and its normalized price, including firm fixed effects and industry output growth rate. Year 1980 is normalized to 100. The underlying value added production function is estimated using the Levinsohn-Petrin procedure.

Table A2
Labor Demand Regression
Blue Collar Employment, Fixed Effects

	Blue Collar Labor			
	Period 1 (1979-84)	Period 2 (1985-90)	Period 3 (1991-96)	All Periods
Constant	1.593	1.554	2.081	0.892
	[0.094]**	[0.188]**	[0.236]**	[0.045]**
Blue wage	-0.151	-0.26	-0.282	-0.23
	[0.021]**	[0.023]**	[0.022]**	[0.008]**
Blue wage_t-1	-0.01	0.06	0.065	0.091
	[0.012]	[0.018]**	[0.017]**	[0.007]**
White wage	0.098	0.062	0.023	0.06
	[0.012]**	[0.013]**	[0.019]	[0.004]**
White wage_t-1	-0.027	-0.017	-0.017	-0.042
	[0.010]**	[0.008]*	[0.008]*	[0.004]**
Value added	0.11	0.101	0.117	0.115
	[0.008]**	[0.010]**	[0.010]**	[0.004]**
Value added_t-1	0.032	0.038	0.017	0.009
	[0.005]**	[0.004]**	[0.007]*	[0.003]**
Blue Employment_t-1	0.228	0.367	0.321	0.55
	[0.017]**	[0.025]**	[0.021]**	[0.010]**
Period Dummy (1985-1990)				-0.037
				[0.016]*
Period Dummy (1991-1996)				-0.029
				[0.019]
Period Dummy (1985-1990)*Blue Employment_t-1				0.023
				[0.005]**
Period Dummy (1991-1996)*Blue Employment_t-1				0.022
				[0.006]**
Observations	17650	20416	23155	61221
R-squared	0.95	0.95	0.94	0.93

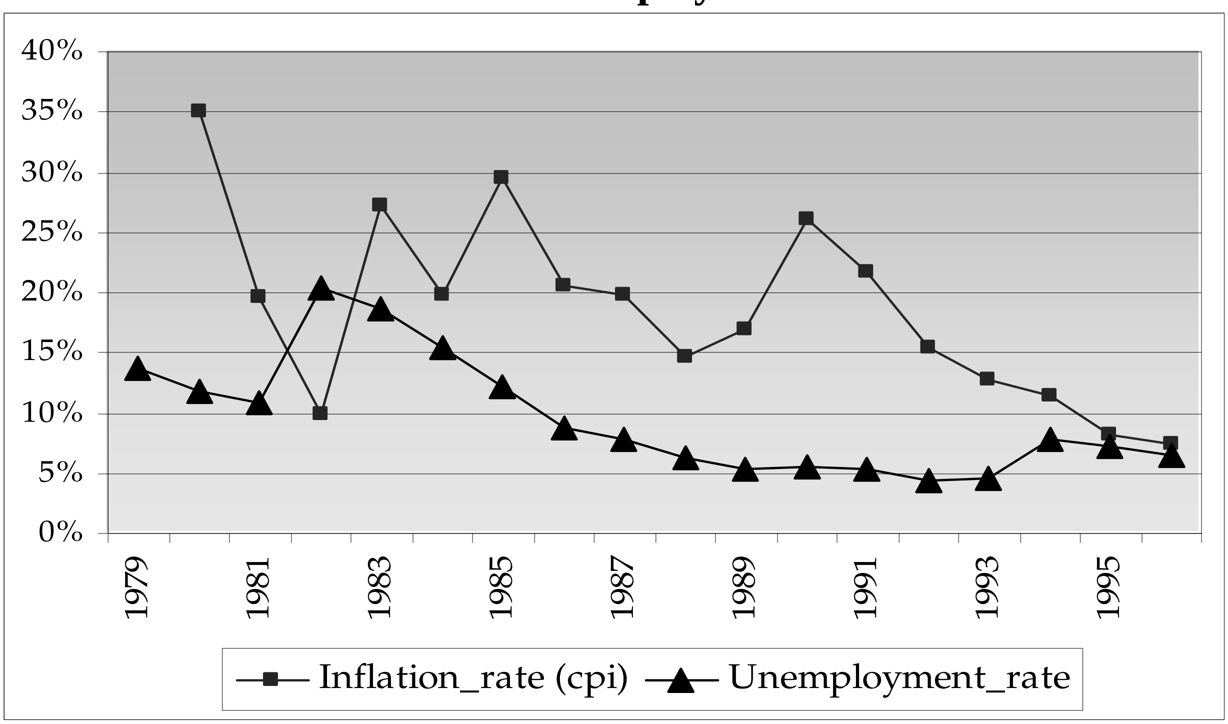
Dependent variable is the log (Blue Collar Employment). All other variables are also in logs. Standard errors are adjusted for clustering at the 4-digit industry level. + significant at 10%; * significant at 5%; ** significant at 1%.

Figure A2
Trends in Real Wages



Source: Authors' calculations

Figure A3
Inflation and Unemployment Rate



Source: Edwards & Edwards (2000), ILO statistics, authors' calculations.