EFFECTIVE LABOR REGULATION AND MICROECONOMIC FLEXIBILITY*

Ricardo J. Caballero
Kevin N. Cowan
Eduardo M.R.A. Engel
Alejandro Micco

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Abstract

Microeconomic flexibility is at the core of economic growth in modern market economies because it facilitates the process of creative-destruction, The main reason why this process is not infinitely fast, is the presence of adjustment costs, some of them technological, others institutional. Chief among the latter is labor market regulation. While few economists object to the hypothesis that labor market regulation hinders the process of creative-destruction, its empirical support is limited. In this paper we revisit this hypothesis, using a new sectoral panel for 60 countries and a methodology suitable for such a panel. We find that job security regulation clearly hampers the creative-destruction process, especially in countries where regulations are likely to be enforced. Moving from the 20th to the 80th percentile in job security, in countries with strong rule of law, cuts the annual speed of adjustment to shocks by a third while shaving off about one percent from annual productivity growth. The same movement has negligible effects in countries with weak rule of law.

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1 Introduction

Microeconomic flexibility, by facilitating the ongoing process of creative-destruction, is at the core of economic growth in modern market economies. This basic idea has been with economists for centuries, was brought to the fore by Schumpeter fifty years ago, and has recently been quantified in a wide variety of contexts.[1] In US Manufacturing, for example, more than half of aggregate productivity growth can be directly linked to this process.[2]

The main obstacle faced by microeconomic flexibility is adjustment costs. Some of these costs are purely technological, others are institutional. Chief among the latter is labor market regulation, in particular job security provisions. The literature on the impact of labor market regulation on the many different economic, political and sociological variables associated to labor markets and their participants is extensive and contentious. However, the proposition that job security provisions reduce restructuring is a point of agreement.

Despite this consensus, the empirical evidence supporting the negative impact of labor market regulation on microeconomic flexibility has been scant at best. This is not too surprising, as the obstacles to empirical success are legions, including poor measurement of restructuring activity and labor market institutions variables, both within a country and more so across countries.[3] In this paper we make a new attempt. We develop a methodology that allows us to bring together the extensive new data set on labor market regulation constructed by Botero et al. (2004) with comparable cross-country cross-sectoral data on employment and output from the UNIDO (2002) data-set. We also emphasize the key distinction between *effective* and official labor market regulation.

The methodology builds on the simple partial-adjustment idea that larger adjustment costs are reflected in slower employment adjustment to shocks.[4] The accumulation of limited adjustment to these shocks builds a wedge between frictionless and actual employment, which is the main right hand side variable in this approach. We propose a new way of estimating this wedge, which allows us to pool data on labor market legislation with comparable employment and output data for a broad range of countries. As a result, we are able to enlarge the effective sample to 60 economies, more than double the country coverage of previous studies in this literature.[5] Our attempt to measure *effective* labor regulation interacts existing measures of job security provision with measures of rule of law and government efficiency.[6]

[1] See, e.g., the review in Caballero and Hammour (2000).

[2] See, e.g., Foster, Haltiwanger and Krizan (1998).

[3] On a closely related literature, there is an extensive body of empirical work, pioneered by Lazear (1990), that has put together data on job security provisions across countries and over time, and measured the effect of these provisions on aggregate employment. A recent survey of this literature can be found in Heckman and Pages (2003). Results are mixed. On the one hand, Lazear (1990), Grubb and Wells (1993), Nickell (1997) and Heckman and Pages (2000) find a negative relationship between job security and employment levels. On the other hand Garibaldi and Mauro (1999), OECD (1999), Addison, Texeira and Grosso (2000), and Freeman (2001) fail to find evidence of such a relationship.

[4] For surveys of the empirical literature on partial-adjustment see Nickell (1986) and Hammermesh (1993).

[5] To our knowledge, the broadest cross-country study to date – Nickell and Nuziata (2000) – included 20 high income OECD countries. Other recent studies, such as Burgess and Knetter (1998) and Burgess et al. (2000), pool industry-level data from 7 OECD economies.

[6] See Loboguerrero and Panizza (2003) for a similar interaction term in their study of the relation between labor market institu-

Our results are clear and robust: countries with less effective job security legislation adjust more quickly to imbalances between frictionless and actual employment. In countries with strong rule of law, moving from the 20th to the 80th percentile of job security lowers the speed of adjustment to shocks by 35 percent and cuts annual productivity growth by 0.85 percent. The same movement for countries with low rule of law only reduces the speed of adjustment by approximately 1 percent and productivity growth by 0.02 percent.

The paper proceeds as follows. Section 2 presents the methodology and describes the new data set. Section 3 discusses the main results and explores their robustness. Section 4 gauges the impact of effective labor protection on productivity growth. Section 5 concludes.

2 Methodology and Data

2.1 Methodology

2.1.1 Overview

The starting point for our methodology is a simple adjustment hazard model, where the change in the number of (filled) jobs in sector j in country c between time $t-1$ and t is a probabilistic (at least to the econometrician) function of the gap between desired and actual employment:

$$\Delta e_{jct} = \psi_{jct}\mathrm{Gap}_{jct} \qquad \mathrm{Gap}_{jct} \equiv e^{*}_{jct} - e_{jc,t-1}, \tag{1}$$

where e_{jct} and e^{*}_{jct} denote the logarithm of employment and desired employment, respectively. The random variable ψ_{jct}, which is assumed i.i.d. both across sectors and over time, takes values in the interval $[0,1]$ and has country-specific mean λ_c and variance $\zeta_c\lambda_c(1-\lambda_c)$, with $0 \leq \zeta_c \leq 1$. This model can be obtained from a generalization of Sargent (1978) and Calvo (1983) (see below). The case $\zeta_c = 0$ corresponds to the standard quadratic adjustment model as in Sargent (1978), the case $\zeta_c = 1$ to the Calvo (1983) model. The parameter λ_c captures microeconomic flexibility. As λ_c goes to one, all gaps are closed quickly and microeconomic flexibility is maximum. As λ_c decreases, microeconomic flexibility declines.

Equation (1) hints at two important components of our methodology: We need to find a measure of the employment gap and a strategy to estimate the average (over j and t) speeds of adjustment (the λ_c). We describe both ingredients in detail in what follows. In a nutshell, we construct estimates of e^{*}_{jct}, the only unobserved element of the gap, by solving the optimization problem of a sector's representative firm, as a function of observables such as labor productivity and a suitable proxy for the average market wage. We estimate λ_c from (1), based upon the large cross-sectional size of our sample and the well documented heterogeneity in the realizations of the gaps and the ψ_{jct}'s (see, e.g., Caballero, Engel and Haltiwanger (1997) for US evidence).

tions and inflation.

2.1.2 Details

A sector's representative firm has output and demand:

$$y = a + \alpha e + \beta h, \tag{2}$$
$$p = d - \frac{1}{\eta} y, \tag{3}$$

where y, p, e, a, h, d denote output, price, employment, productivity, hours worked and demand shocks, and η is the price elasticity of demand. We let $\gamma \equiv (\eta - 1)/\eta$, with $\eta > 1$, $0 < \alpha < 1$ and $0 < \beta < 1$. All variables are in logs.

Firms are competitive in the labor market but pay wages that increase with the number of hours worked, H:

$$w = k^o + \log(H^\mu + \Omega).$$

This can be approximated by:

$$w = w^o + \mu(h - \overline{h}), \tag{4}$$

with w^o determined by k^o and Ω, and $\overline{h}$ constant over time and interpreted below. In order to ensure interior solutions, we assume $\alpha\mu > \beta$ and $\mu > \beta\gamma$.

A key assumption is that the representative firm within each sector only faces adjustment costs when it changes employment levels, not when it changes the number of hours worked (beyond overtime payments).[7] It follows that the sector's choice of hours in every period can be expressed in terms of its current level of employment, by solving the corresponding first order condition for hours.

In a frictionless labor market the firm's employment level also satisfies a simple *static* first order condition for employment. Our functional forms then imply that the optimal choice of hours, $\overline{h}$, does not depend on the employment level. A patient calculation shows that

$$\overline{h} = \frac{1}{\mu} \log\left(\frac{\beta\Omega}{\alpha\mu - \beta} \right).$$

We denote the corresponding employment level by $\widehat{e}$ and refer to it as the *static employment target*:

$$\widehat{e} = C + \frac{1}{1 - \alpha\gamma}[d + \gamma a - w^o],$$

with C a constant that depends on μ, α, β and γ.

In the absence of adjustment costs the firm's cash flow, R, is maximized at $\widehat{e}$, taking the value $\widehat{R}$. A second order Taylor approximation of the firm's revenue function, net of adjustment costs, around $\widehat{e}$ then yields

$$R \cong \widehat{R} - C'(e - \widehat{e})^2, \tag{5}$$

[7] For evidence on this see Sargent (1978) and Shapiro (1986).

with $\widehat{R}$ unaffected by the firms' choice variables. Without loss of generality we set $C' = 1$ in what follows.

Firms' labor adjustment costs are assumed quadratic, with a stochastic proportionality factor k. The k's are independent (over time and across sectors within a country), identically distributed, and take both the value zero and infinity with positive probabilities, thereby allowing for both smooth and lumpy labor adjustments. More precisely:

$$k_t = \begin{cases} 0 & \text{with prob. } \pi_0 \\ K & \text{with prob. } \pi_k \\ \infty & \text{with prob. } \pi_\infty \end{cases} \tag{6}$$

where K a fixed number, $0 < K < \infty$, $\pi_i \geq 0$, $i = 0, k, \infty$, and $\pi_0 + \pi_k + \pi_\infty = 1$.[8]

The firm's profit maximization problem at time t then is equivalent to:

$$\min_{e_t} \mathrm{E}_t \left[\sum_{j\geq 0} \rho^j \left\{ (e_{t+j} - \widehat{e}_{t+j})^2 + k_{t+j}(e_{t+j} - e_{t+j-1})^2 \right\} \right], \tag{7}$$

with ρ denoting the firm's discount factor. In Appendix A we solve the corresponding Bellman equation and show that the firm's optimal employment choice satisfies

$$\Delta e_t = \psi_t (e_t^* - e_{t-1}), \tag{8}$$

with the *dynamic employment target*, e_t^*, defined via

$$e_t^* = (1-\tau) \sum_{j\geq 0} \tau^j \mathrm{E}_t[\widehat{e}_{t+j}],$$

for some constant $\tau \in (0,1)$, and

$$\psi_t \equiv \psi(k_t) = \begin{cases} 0 & \text{if } k_t = \infty \\ \nu & \text{if } k_t = K \\ 1 & \text{if } k_t = 0 \end{cases} \tag{9}$$

with $\nu \in (0,1)$ an explicit function of K, ρ, π_0, π_k and π_∞.

It follows that a fraction π_∞ of the time the representative firm does not adjust its employment, a fraction π_0 it adjusts fully to e^* and the remaining periods, it closes part of the gap between its dynamic employment target and actual employment.

Denoting $\lambda \equiv \pi_0 + \nu\pi_k$ we have that:

$$\begin{aligned} \mathrm{E}[\psi] &= \lambda, \\ \mathrm{Var}[\psi] &= \lambda(1-\lambda) - \pi_k \nu (1-\nu), \end{aligned}$$

with $\mathrm{Var}[\psi]$ taking values between 0 (quadratic adjustment: $\pi_k = 1$, $\nu = \lambda$) and $\lambda(1-\lambda)$ (Calvo model:

[8] The results that follow can be extended to the case where K is drawn from a distribution that takes positive values.

$\pi_k = 0$, $\pi_0 = \lambda$). Furthermore, as π_k decreases from one to zero, $\text{Var}[\psi]$ covers the full range of values between $\lambda(1-\lambda)$ and zero.

Having derived our estimating equation from first principles, we next turn to deriving a proxy for the dynamic employment target e^*. For this, note that the relation between the employment gap and the hours gap follows from the expressions obtained above for $\widehat{e}$, $\overline{h}$ and the first order condition satisfied by h:

$$\widehat{e} - e \;=\; \frac{\mu - \beta\gamma}{1-\alpha\gamma}(h - \overline{h}). \tag{10}$$

This is the expression used by Caballero and Engel (1993). It is not useful in our case, since we do not have information on worked hours. Yet the argument leading to (10) also can be used to express the employment gap in terms of the marginal labor productivity gap:

$$\widehat{e} - e \;=\; \frac{\phi}{1-\alpha\gamma}(v - w^o),$$

where v denotes marginal productivity, $\phi \equiv \mu/(\mu - \beta\gamma)$ is decreasing in the elasticity of the marginal wage schedule with respect to average hours worked, $\mu - 1$, and w^o was defined in (4). Note that $\widehat{e} - e$ is the difference between the static target $\widehat{e}$ and realized employment, not the dynamic employment gap $e^*_{jct} - e_{jct}$ related to the term on the right hand side of (1). However, if we assume that $d + \gamma a - w^o$ follows a random walk (possibly with an exogenously time varying drift) —an assumption consistent with the data[9] — we have that e^*_{jct} is equal to $\widehat{e}_{jct}$ plus a constant δ_{ct}. It follows that

$$e^*_{jct} - e_{jct-1} \;=\; \frac{\phi}{1-\alpha\gamma_j}\left(v_{jct} - w^o_{jct}\right) + \Delta e_{jct} + \delta_{ct}, \tag{11}$$

where we have allowed for sector-specific differences in $\alpha\gamma$. Note that both marginal product and wages are in nominal terms. However, since these expressions are in logs, their difference eliminates the aggregate price level component.

We estimate the marginal productivity of labor, v_{jct}, using output per worker multiplied by an industry-level labor share, assumed constant within income groups and over time.

Two natural candidates to proxy for w^o_{jct} are the average (across sectors within a country, at a given point in time) of either observed wages or observed marginal productivities. The former is consistent with a competitive labor market, the latter may be expected to be more robust in settings with long-term contracts and multiple forms of compensation, where the salary may not represent the actual marginal cost of labor.[10] We performed estimations using both alternatives and found no discernible differences (see below). This suggests that statistical power comes mainly from the cross-section dimension, that is, from the well docu-

[9] Pooling all countries and sectors together, the first order autocorrelation of the measure of Δe^*_{jct} constructed below is -0.018. Computing this correlation by country the mean value is 0.011 with a standard deviation of 0.179.

[10] While we have assumed a simple competitive market for the base salary (salary for normal hours) within each sector, our procedure could easily accommodate other, more rent-sharing like, wage setting mechanisms (with a suitable reinterpretation of some parameters, but not λ_c).

mented and large magnitude of sector-specific shocks. In what follows we report the more robust alternative and approximate w^o by the average marginal productivity, which leads to:

$$e^*_{jct} - e_{jct-1} \;=\; \frac{\phi}{1-\alpha\gamma_j}(v_{jct} - v_{\cdot ct}) + \Delta e_{jct} + \delta_{ct} \;\equiv\; \text{Gap}_{jct} + \delta_{ct}, \tag{12}$$

where $v_{\cdot ct}$ denotes the average, over j, of v_{jct}, and we use this convention for other variables as well. The expression above ignores systematic variations in labor productivity across sectors within a country, for example, because (unobserved) labor quality may differ systematically across sectors. The presence of such heterogeneity would tend to bias estimates of the speed of adjustment downward. To incorporate this possibility we subtract from $(v_{jct} - v_{\cdot ct})$ in (12) a moving average of relative sectoral productivity, $\widehat{\theta}_{jct}$, where

$$\widehat{\theta}_{jct} \equiv \frac{1}{2}[(v_{jct-1} - v_{\cdot ct-1}) \;+\; (v_{jct-2} - v_{\cdot ct-2})].$$

As a robustness check, for our main specifications we also computed $\widehat{\theta}_{jct}$ using a three and four periods moving average, without significant changes in our results (more on this when we check robustness in Section 3.2). The resulting expression for the estimated employment-gap is:

$$e^*_{jct} - e_{jct-1} \;=\; \frac{\phi}{1-\alpha\gamma_j}(v_{jct} - \widehat{\theta}_{jct} - v_{\cdot ct}) + \Delta e_{jct} + \delta_{ct} \;\equiv\; \text{Gap}_{jct} + \delta_{ct}, \tag{13}$$

where $\alpha\gamma_j$ is constructed using the sample median of the labor share for sector j across year and income groups.

Rearranging (13), we estimate ϕ from

$$\Delta e_{jct} \;=\; -\frac{\phi}{1-\alpha\gamma_j}(\Delta v_{jct} - \Delta v_{\cdot ct}) + \kappa_{ct} + \upsilon_{it} + \Delta e^*_{jct} \;\equiv\; -\phi z_{jct} + \kappa_{ct} + \varepsilon_{jct}, \tag{14}$$

where κ is a country-year dummy, Δe^*_{jct} is the change in the desired level of employment and $z_{jct} \equiv (\Delta v_{jct} - \Delta v_{\cdot ct})/(1-\alpha\gamma_j)$. We assume that changes in sectoral labor composition are negligible between two consecutive years. In order to avoid the simultaneity bias present in this equation (Δv and Δe^* are clearly correlated) we estimate (14) using $(\Delta w_{jct-1} - \Delta w_{\cdot ct-1})$ as an instrument for $(\Delta v_{jct} - \Delta v_{\cdot ct})$.[11]

Table 1 reports the estimation results of (14) for the full sample of countries and across income and job security groups. The first two columns use the full sample, with and without two percent of extreme values for the independent variable, respectively. The remaining columns report the estimation results for each of our three income groups and job security groups (more on both of these measures in Section 2.2). Based on our results for the baseline case, we set the value of ϕ at its full sample estimate of 0.4 for all countries in our sample.

It is important to point out that our methodology has some advantages over standard partial adjustment

[11] We lag the instrument to deal with the simultaneity problem and use the wage rather than productivity to reduce the (potential) impact of measurement error bias.

Table 1: ESTIMATING ϕ

Specification:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Change in Employment (ln)							
z_{jct}	-0.280	-0.394	-0.558	-0.355	-0.387	-0.363	-1.168	-0.352
	(0.044)	(0.068)	(0.135)	(0.119)	(0.116)	(0.091)	(.357)	(0.103)
Observations	22,810	22,008	8,311	6,378	7,319	7,730	6,883	7,036
Income Group	All	All	1	2	3	All	All	All
Job Sec. Group	All	All	All	All	All	1	2	3
Extreme obs. of instrument	Yes	No	No	No	No	No	No	No

Standard errors reported in parentheses. All estimates are significant at the 1% level. All regressions use lagged $\Delta w_{ict} - \Delta w_{\cdot ct}$ as instrumental variable. As described in the main text, z_{jct} represents the log-change of the nominal marginal productivity of labor in each sector, minus the country average, divided by one minus the estimated labor share. All regressions disregard the 2% observations with most extreme change in employment values and include a country-year fixed effect (κ_{ct} in (14)). Income groups are 1: High Income OECD, 2: High Income Non OECD and Upper Middle Income, and 3: Lower Middle Income and Low Income. Job Security Groups correspond to the highest, middle an lowest third of the measure in Botero et al. (2004).

estimations. First, it summarizes in a single variable all shocks faced by a sector. This feature allows us to increase precision and to study the determinants of the speed of adjustment using interaction terms. Second, and related, it only requires data on nominal output and employment, two standard and well-measured variables in most industrial surveys. Most previous studies on adjustment costs required measures of real output or an exogenous measure of sector demand.[12]

2.1.3 Regressions

The central empirical question of the present study is how cross-country differences in job security regulation affect the speed of adjustment. Accordingly, from (1) and (13) it follows that the basic equation we estimate is:

$$\Delta e_{jct} \;=\; \lambda_{ct}(\mathrm{Gap}_{jct} + \delta_{ct}), \tag{15}$$

where Δe_{jct} is the log change in employment and λ_{ct} denotes the speed of adjustment. We assume that the latter takes the form:

$$\lambda_{ct} \;=\; \tilde{\lambda}_1 + \tilde{\lambda}_2 \mathrm{JS}_{ct}^{\mathrm{eff}}, \tag{16}$$

where $\mathrm{JS}_{ct}^{\mathrm{eff}}$ is a measure of *effective* job security regulation. In practice we observe job security regulation (imperfectly), but not the rigor with which it is enforced. We proxy the latter with a "rule of law" variable,

[12] Abraham and Houseman (1994), Hammermesh (1993), and Nickel and Nunziata (2000)) evaluate the differential response of employment to observed real output. A second option is to construct exogenous demand shocks. Although this approach overcomes the real output concerns, it requires constructing an adequate sectorial demand shock for every country. A case in point are the papers by Burgess and Knetter (1998) and Burgess et al. (2000), which use the real exchange rate as their demand shock. The estimated effects of the real exchange rate on employment are usually marginally significant, and often of the opposite sign than expected.

so that

$$\mathrm{JS}_{ct}^{\mathrm{eff}} = a\mathrm{JS}_{ct} + b(\mathrm{JS}_{ct}\mathrm{RL}_{ct}), \tag{17}$$

where a and b are constants and RL_{ct} is a standard measure of rule of law (see below). When $b = 0$ there is no difference between *de jure* and *de facto* regulation. Substituting this expression in (16) and the resulting expression for λ_{ct} in (15), yields our main estimating equation:

$$\Delta e_{jct} \;=\; \lambda_1 \,\mathrm{Gap}_{jct} + \lambda_2 \,\left(\mathrm{Gap}_{jct} \times \mathrm{JS}_{ct}\right) + \lambda_3 \,\left(\mathrm{Gap}_{jct} \times \mathrm{JS}_{ct} \times \mathrm{RL}_{ct}\right) + \widetilde{\delta}_{ct} + \varepsilon_{jct}, \tag{18}$$

with $\lambda_1 = \tilde{\lambda}_1$, $\lambda_2 = a\tilde{\lambda}_2$, $\lambda_3 = b\tilde{\lambda}_3$, and $\widetilde{\delta}_{ct}$ denotes country$\times$time fixed effects (proportional to the δ_{ct} defined above).

The main coefficients of interest are λ_2 and λ_3, which measure how the speed of adjustment varies across countries depending on their labor market regulation (both *de jure* and *de facto*).

2.2 The Data

This section describes our sample and main variables. Additional variables are defined as we introduce them later in the text.

2.2.1 Job Security and Rule of Law

We use two measures of job security, or legal protection against dismissal: the job security index constructed by Botero et al. (2004) for 60 countries world-wide (henceforth JS_c) and the job security index constructed by Heckman and Pages (2000) for 24 countries in OECD and Latin America (henceforth HP_{ct}). The JS_c measure is available for a larger sample of countries and includes a broader range of job security variables. The HP_{ct} measure has the advantage of having time variation.

Our main job security index, JS_c, is the sum of four variables, measured in 1997, each of which takes on values between 0 and 1: (i) grounds for dismissal protection PG_c, (ii) protection regarding dismissal procedures PP_c, (iii) notice and severance payments PS_c, and (iv) protection of employment in the constitution PC_c. The rules on grounds of dismissal range from allowing the employment relation to be terminated by either party at any time (employment at will) to allowing the termination of contracts only under a very narrow list of "fair" causes. Protective dismissal procedures require employers to obtain the authorization of third parties (such as unions and judges) before terminating the employment contract. The third variable, notice and severance payment, is the one closest to the HP_{ct} measure, and is the normalized sum of two components: mandatory severance payments after 20 years of employment (in months) and months of advance notice for dismissals after 20 years of employment $(NS_{tc} = b_{ct+20} + SP_{ct+20},\ t = 1997)$. The four components of JS_c described above increase with the level of job security.

The Heckman and Pages measure is narrower, including only those provisions that have a direct impact on the costs of dismissal. To quantify the effects of this legislation, they construct an index that computes the expected (at hiring) cost of a future dismissal. The index includes both the costs of advanced notice legislation and firing costs, and is measured in units of monthly wages.

Our estimations also adjust for the level of enforcement of labor legislation. We do this by including measures of rule of law RL_c and government efficiency GE_c from Kaufmann at al. (1999), and interact them with JS_c and HP_{ct}.[13] We expect labor market legislation to have a larger impact on adjustment costs in countries with a stronger rule of law (higher RL_c) and more efficient governments (higher GE_c).

The institutional variables as well as the countries in our sample and their corresponding income group are reported in Table 2. Table 3 reports the sample correlations between our main cross-country variables and summary statistics for each of these measures for three income groups (based on World Bank per capita income categories).[14] As expected, the correlation between the two measures of job security is positive and significant. Differences can be explained mainly by the broader scope of the JS_{ct} index. Also as expected, rule of law and government efficiency increase with income levels. Note, however, that neither measure of job security is positively correlated with income per capita, since both JS_{ct} and HP_c are highest for middle income countries.

2.2.2 Industrial Statistics

Our output, employment and wage data come from the 2002 3-digit UNIDO Industrial Statistics Database. The UNIDO database contains data for the period 1963-2000 for the 28 manufacturing sectors that correspond to the 3 digit ISIC code (revision 2). Because our measures of job security and rule of law are time invariant and measured in recent years, however, we restrict our sample to the period 1980-2000. Data on output and labor compensation are in current US dollars (inflation is removed through time effects in our regressions). Throughout the paper our main dependent variable is Δe_{jct}, the log change in total employment in sector j of country c in period t.

A large number of countries are included in the original dataset — however our sample is constrained by the cross-country availability of the independent variables measuring job security. In addition, we drop two percent of extreme employment changes in each of the three income groups. For our main specification the resulting sample includes 60 economies. Table 3 shows descriptive statistics for the dependent variable by income group.

3 Results

This section presents our main result, showing that effective job security has a significant negative effect on the speed of adjustment of employment to shocks in the employment-gap. It also presents several robustness exercises.

[13] For rule of law and government efficiency we use the earliest value available in the Kaufmann et al. (1999) database: 1996, since this is closest to the Botero et al. (2004) measure, which is for 1997.

[14] Income groups are: 1=High Income OECD, 2=High Income Non OECD and Upper Middle Income, 3=Lower Middle Income and Low Income.

Table 2: SAMPLE COVERAGE AND MAIN VARIABLES

		Job Security		Institutions			
WDI code	Inc Group	Botero et al	HP	Strong RL	Rule of Law	Gov. Eff.	High Gov. Eff
AUS	1	−0.19	−0.71	1	1.03	0.95	1
AUT	1	−0.15	−0.65	1	1.13	0.92	1
BEL	1	−0.11	−0.70	1	0.81	0.81	1
CAN	1	−0.16	−1.64	1	1.02	0.92	1
DEU	1	0.17	−1.56	1	1.04	0.92	1
DNK	1	−0.21		1	1.17	1.02	1
ESP	1	0.17	1.29	1	0.41	0.64	1
FIN	1	0.24	−0.82	1	1.22	0.89	1
FRA	1	−0.02	−1.09	1	0.81	0.78	1
GBR	1	−0.13	−1.00	1	1.09	1.05	1
GRC	1	−0.04	−1.05	1	−0.01	−0.06	1
IRL	1	−0.21	−1.40	1	0.92	0.82	1
ITA	1	−0.09	0.79	1	0.09	0.05	1
JPN	1	−0.14	−1.84	1	0.76	0.46	1
NLD	1	0.04	−1.53	1	1.09	1.25	1
NOR	1	−0.03	−1.55	1	1.23	1.13	1
NZL	1	−0.29	−2.21	1	1.22	1.25	1
PRT	1	0.37	2.05	1	0.53	0.24	1
SWE	1	0.06	−0.50	1	1.17	0.97	1
USA	1	−0.25	−2.43	1	0.95	1.01	1
ARG	2	0.11	0.56	0	−0.48	−0.37	0
BRA	2	0.36	0.61	0	−1.00	−0.82	0
CHL	2	−0.02	0.21	1	0.44	0.32	1
HKG	2	−0.32		1	0.86	0.81	1
ISR	2	−0.17		1	0.36	0.42	1
KOR	2	−0.07	1.14	1	0.02	−0.15	0
MEX	2	0.38	0.73	0	−0.86	−0.85	0
MYS	2	−0.24		1	0.05	0.18	1
PAN	2	0.34	1.37	0	−0.50	−1.19	0
SGP	2	−0.22		1	1.26	1.41	1
TUR	2	−0.13	1.54	0	−0.73	−0.69	0
TWN	2	0.01		1	0.21	0.49	1
URY	2	−0.30	−0.20	0	−0.26	−0.17	0
VEN	2	0.31	4.29	0	−1.38	−1.32	0
ZAF	2	−0.17		0	−0.42	−0.40	0
BFA	3	−0.10		0	−1.46	−1.38	0
BOL	3	0.24	2.32	0	−1.37	−1.12	0
COL	3	0.29	1.17	0	−1.19	−0.61	0
ECU	3	0.34	0.97	0	−1.13	−1.29	0
EGY	3	0.13		0	−0.53	−0.99	0
GHA	3	−0.17		0	−0.86	−0.78	0
IDN	3	0.10		0	−1.09	−0.55	0
IND	3	−0.14		0	−0.77	−0.79	0
JAM	3	−0.20	−0.44	0	−0.95	−1.06	0
JOR	3	0.22		0	−0.56	−0.54	0
KEN	3	−0.16		0	−1.48	−1.13	0
LKA	3	0.09		0	−0.48	−0.93	0
MAR	3	−0.22		0	−0.57	−0.73	0
MDG	3	0.23		0	−1.55	−1.39	0
MOZ	3	0.38		0	−1.92	−1.23	0
MWI	3	0.11		0	−0.94	−1.32	0
NGA	3	−0.07		0	−1.89	−1.68	0
PAK	3	−0.15		0	−1.16	−1.02	0
PER	3	0.37	2.25	0	−1.08	−0.87	0
PHL	3	0.24		0	−0.86	−0.54	0
SEN	3	−0.04		0	−0.92	−1.04	0
THA	3	0.10		0	−0.29	−0.32	0
TUN	3	0.05		0	−0.69	−0.24	0
ZMB	3	−0.33		0	−1.08	−1.44	0
ZWE	3	−0.13		0	−0.97	−0.86	0

Table 3: BASELINE SAMPLE STATISTICS*

Employment Growth (Yearly Avge.): 1980-2000					
Inc. Group	Obs.	Mean	SD	Min	Max
1	8,607	−0.01	0.06	−0.24	0.26
2	6,063	0.00	0.11	−0.43	0.42
3	7,063	0.02	0.16	−0.78	0.96
Total	21,733	0.00	0.11	−0.78	0.96

Job Security from Botero et al. (2004): JS					
Inc. Group	Countries	Mean	SD	Min	Max
1	20	−0.05	0.18	−0.29	0.37
2	15	−0.01	0.25	−0.32	0.38
3	25	0.05	0.21	−0.33	0.38
Total	60	0.00	0.21	−0.33	0.38

Job Security from Heckman and Pages (2001): HP					
Inc. Group	Countries	Mean	SD	Min	Max
1	19	−0.87	1.15	−2.43	2.05
2	9	1.14	1.30	−0.20	4.29
3	5	1.26	1.13	−0.44	2.32
Total	33	0.00	1.54	−2.43	4.29

Rule of Law from Kaufmann et al. (1999): RL					
Inc. Group	Countries	Mean	SD	Min	Max
1	20	0.88	0.37	−0.01	1.23
2	15	−0.16	0.72	−1.38	1.26
3	25	−1.03	0.42	−1.92	−0.29
Total	60	−0.18	0.96	−1.92	1.26

Government Effectiveness from Kaufmann et al. (1999): GE					
Inc. Group	Countries	Mean	SD	Min	Max
1	20	0.80	0.37	−0.06	1.25
2	15	−0.16	0.76	−1.32	1.41
3	25	−0.95	0.36	−1.68	−0.24
Total	60	−0.17	0.90	−1.68	1.41

Correlation Country Means					
		JS	HP	RL	GE
	JS	1.00			
	HP	0.66	1.00		
	RL	−0.36	−0.77	1.00	
	GE	−0.35	−0.77	0.97	1.00

*Income groups are: 1=High Income OECD, 2=High Income Non OECD and Upper Middle Income, 3=Lower Middle Income and Low Income.

3.1 Main results

Recall that our main estimating equation is:

$$\Delta e_{jct} = \lambda_1 \text{ Gap}_{jct} + \lambda_2 \left(\text{Gap}_{jct} \times \text{JS}_c\right) + \lambda_3 \ \left(\text{Gap}_{jct} \times \text{JS}_c \times \text{RL}_c\right) + \widetilde{\delta}_{ct} + \varepsilon_{jct}. \tag{19}$$

Note that we have dropped time subscripts from JS_c and RL_c as we only use time invariant measures of rule of law and job security in our baseline estimation. Note also that in all specifications that include the $(\text{Gap}_{ict} \times \text{JS}_{ct} \times \text{RL}_c)$ interaction we also include the respective $\text{Gap}_{ict} \times \text{RL}_c$ as a control variable.

Table 4: ESTIMATION RESULTS

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Change in Log-Employment							
Gap (λ_1)	0.600 (0.009)***	0.603 (0.008)***	0.607 (0.012)***	0.611 (0.012)***				
Gap×JS (λ_2):		−0.080 (0.037)**	−0.015 (0.051)	−0.025 (0.051)	−0.126 (0.041)***	−0.027 (0.052)	−0.038 (0.051)	
Gap×JS×DSRL (λ_3)			−0.514 (0.068)***			−0.314 (0.070)***		
Gap×JS×DHGE (λ_3)				−0.515 (0.068)***			−0.326 (0.071)***	
Gap×HP (λ_2)								−0.022 (0.007)***
Controls								
Gap×DSRL			−0.076 (0.015)***			0.086 (0.023)***		
Gap×DHGE				−0.091 (0.015)***			0.045 (0.023)*	
Observations	21,733	21,733	21,733	21,733	21,733	21,733	21,733	12,012
R-squared	0.60	0.60	0.60	0.60	0.61	0.61	0.61	0.62
Gap-Income Interaction	No	No	No	No	Yes	Yes	Yes	Yes
Gap-Sector Interaction	No	No	No	No	Yes	Yes	Yes	Yes

* significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parentheses. JS and HP stand for the Botero et al. (2004) and Heckman and Pages (2000) job security measures, respectively. DSRL and DHGE stand for strong Rule of Law and high Government Efficiency dummies (in both cases the threshold is given by Greece, see the main text), respectively, using the Kaufmann et al. (1999) indices. Each regression has country-year fixed effects. Gaps are estimated using a constant $\phi = 0.40$. Sample excludes the upper and lower 1% of Δe and of the estimated values of Gap.

We start by ignoring the effect of job security on the speed of adjustment, and set λ_2 and λ_3 equal to zero. This gives us an estimate of the average speed of adjustment and is reported in column 1 of Table 4. On average (across countries and periods) we find that 60% of the employment-gap is closed in each period. Furthermore, our measure of the employment-gap and country$\times$year fixed effects explain 60% of the variance in log-employment growth.

The next three columns present our main results, which are repeated in columns 5 to 7 allowing for different λ_1 by sectors and country income level.[15] Column 2 (and 5) presents our estimate of λ_2. This coefficient has the right sign and is significant at conventional confidence levels. Employment adjusts more slowly to shocks in the employment-gap in countries with higher levels of official job security.

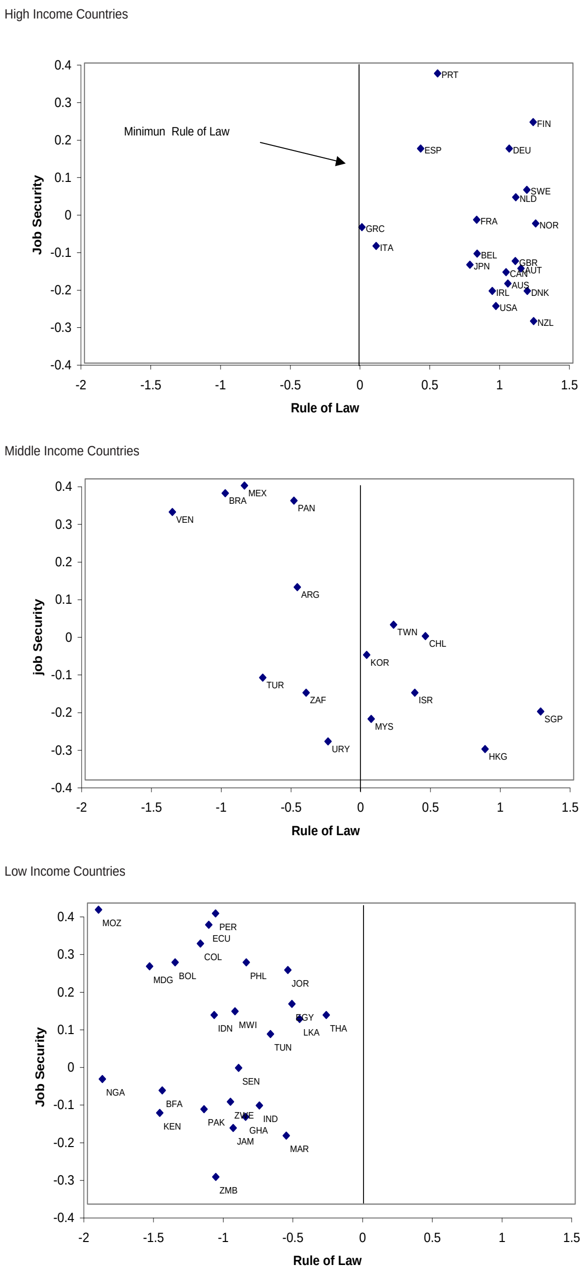
Next, we allow for a distinction between effective and official job security. Results are reported in columns 3 and 4 (and, correspondingly, 6 and 7) for different rules-enforcement criteria. In columns 3 and 6 the distinction between effective and official job security is captured by the product of JS_c and DSRL_c, where DSRL_c is a dummy variable for countries with strong rule of law ($\mathrm{RL}_c \geq \mathrm{RL}_{\mathrm{Greece}}$ — where Greece is the OECD country with the lowest RL score). The three panels in Figure 1 show the value of the job security index for countries in the high, medium and low income groups, respectively. Now λ_2 becomes insignificant, while λ_3 has the right sign and is highly significant. That is, the same change in JS_c will have a significantly larger (downward) effect on the speed of adjustment in countries with stricter enforcement of laws, as measured by our rule-of-law dummy. The effect of the estimated coefficients reported in column 3 is large. In countries with strong rule of law, moving from the 20th percentile of job security (-0.19) to the 80th percentile (0.23) reduces $\hat{\lambda}$ by 0.22. The same change in job security legislation has a considerable smaller effect, 0.006, on the speed of adjustment in the group of economies with weak rule of law. That is, employment adjusts more slowly to shocks in the employment-gap in countries with higher levels of effective job security.

Columns 4 and 7 address whether the negative coefficient on λ_3 is robust to other measures of legal enforcement. To do so we use an alternative variable from the Kaufmann et al. (1999) dataset – government effectiveness (GE) – and construct a dummy variable for high effectiveness countries ($\mathrm{GE}_c \geq \mathrm{GE}_{\mathrm{Greece}}$). Clearly, the results are very close to those reported in columns 3 and 7. Job security legislation has a significant negative effect on the estimated speed of adjustment when governments are effective – a proxy for enforcement of existing labor regulation.

Finally, the last column in Table 4 uses an alternative measure of job security. We repeat our specification from column 7 (including sector and income dummies) using the Heckman-Pages (2000) measure of job security. The HP_{ct} data are only available for countries in the OECD and Latin America so our sample size is reduced by half, and most low income countries are dropped. The flip side is that this measure is time varying which potentially allows us to capture the effects of changes in the job security regulation. As reported in column 8, we find a negative and significant effect of HP_{ct} on the speed of adjustment.

[15] We allow for an interaction between Gap_{jct} and 3 digit ISIC sector dummies (we also include sector fixed effects). We also control for the possibility that our results are driven by omitted variables, correlated with our measures of job security. For this, we include an additional interaction between Gap_{jct} and three income-group dummies.

Figure 1: Job Security and Rule of Law in Countries with High, Medium and Low Income



3.2 Further robustness

We continue our robustness exploration by assessing the impact of three broad econometric issues: alternative gap-measures, exclusion of potential (country) outliers, and misspecification due to endogeneity of the gap measure.

3.2.1 Alternative gap-measures

Table 4 suggests that conditional on our measure of the employment-gap, our main findings are robust: job security, when enforced, has a significant negative impact on the speed of adjustment to the employment-gap. Table 5 tests the robustness of this result to alternative measures of the employment-gap. Columns 1 and 2 relax the assumption of a ϕ common across all countries. They repeat our baseline specifications —columns 2 and 3 in Table 4— using the values of ϕ estimated per income-group reported in Table 1. In turn, columns 3 and 4 report the results of using values of ϕ estimated across countries grouped by level of job security. Countries are grouped into the upper, middle and lower thirds of job security. Next, columns 5 through 8 repeat our baseline specifications using a three and four period moving average to estimate $\widehat{\theta}_{jct}$. The final two columns (9 and 10) use an alternative specification for w^o_{jct} based on average wages instead of average productivity (see equation 13) to build Gap_{jct}. In all of the specifications reported in Table 5, our results remain qualitatively the same as in Table 4.

3.2.2 Exclusion of potential (country) outliers

Table 6 reports estimates of λ_2 and λ_3 using the specification from column 3 in Table 4 but dropping one country from our sample at a time. In all cases the estimated coefficient on λ_3 is negative and significant at conventional confidence intervals.

However, it is also apparent in this table that excluding either Hong Kong or Kenya makes a substantial difference in the point estimates. For this reason, we re-estimate our model from scratch (that is, from ϕ up) now excluding these two countries. In this case the value of ϕ rises from 0.40 to 0.42. Qualitatively, however, the main results remain unchanged. Table 7 reports these results.

3.2.3 Potential endogeneity of the gap measure

One concern with our procedure is that the construction of the gap measure includes the change in employment. While this does not represent a problem under the null hypothesis of the model, any measurement error in employment and ϕz_{jt} could introduce important biases. We address this issue with two procedures.

The first procedure maintains our baseline specification, but instruments for the contemporaneous gap measure. Given that $\text{Gap}_{jct} = \phi z_{jt} + \Delta e_{jct}$ can be rewritten as $\phi z_{j,t-1} + \Delta e^*_{jct}$, a natural instrument is the lag of the ex-post gap, $\phi z_{jc,t-1}$. Unfortunately, the latter is not a valid instrument if it is computed with measurement error and this error is serially correlated. In our specification this could be the case because we use a moving average to construct the estimate of relative sectoral productivity, $\widehat{\theta}_{jct}$. To avoid this problem,

Table 5: Robustness of Main Results to Alternative Specifications

	ϕ varies across income groups		ϕ varies across job security groups		$\phi = 0.40$					
					$\theta = MA(3)$		$\theta = MA(4)$		dw	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)
	Change in Log-Employment									
Gap	0.568		0.574		0.564		0.529		0.590	
	(0.009)***		(0.008)***		(0.008)***		(0.008)***		(0.009)***	
Gap×JS	−0.094	−0.013	−0.027	0.046	−0.069	−0.008	−0.009	0.063	−0.108	−0.135
	(0.038)***	(0.051)	(0.037)	(0.051)	(0.037)*	(0.050)	(0.036)	(0.050)	(0.038)***	(0.050)***
Gap×DSRL		−0.051		−0.071		−0.085		−0.082		−0.106
		(0.015)***		(0.015)***		(0.015)***		(0.015)***		(0.016)***
Gap×JS×DSRL		−0.501		−0.532		−0.515		−0.538		−0.258
		(0.069)***		(0.069)***		(0.068)***		(0.068)***		(0.071)***
Observations	21,733		21,733		20,902		20,219		20,439	
R-squared	0.58	0.58	0.58	0.58	0.59	0.59	0.58	0.58	0.60	0.60
Gap-Sector Int.	No	Yes	No	Yes	No	Yes	No	Yes	No	Yes

* significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parentheses. JS stands for the Botero et al. (2004) job security measure. DSRL stands for high (above Greece, see main text) Rule of Law using the Kaufmann et al. (1999) measure. Columns (1), (2), (3) and (4) use values of ϕ estimated in Table 1. Samples exclude the upper and lower 1% of Δe and of the estimated values of Gap.

Table 6: Excluding One Country at a Time

	λ_2		λ_3			λ_2		λ_3	
Country	Coeff.	St. Dev.	Coeff.	St. Dev.	Country	Coeff.	St. Dev.	Coeff.	St. Dev.
ARG	−0.01	0.05	−0.51	0.07	KOR	−0.02	0.05	−0.52	0.07
AUS	−0.02	0.05	−0.52	0.07	LKA	−0.02	0.05	−0.51	0.07
AUT	−0.02	0.05	−0.52	0.07	MAR	−0.02	0.05	−0.51	0.07
BEL	−0.02	0.05	−0.52	0.07	MDG	−0.02	0.05	−0.51	0.07
BFA	−0.03	0.05	−0.50	0.07	MEX	0.00	0.05	−0.53	0.07
BOL	0.00	0.05	−0.52	0.07	MOZ	0.02	0.05	−0.55	0.07
BRA	−0.01	0.05	−0.52	0.07	MWI	−0.01	0.05	−0.52	0.07
CAN	−0.02	0.05	−0.52	0.07	NYS	−0.02	0.05	−0.46	0.07
CHL	−0.02	0.05	−0.53	0.07	NGA	0.00	0.05	−0.53	0.07
COL	−0.02	0.05	−0.51	0.07	NLD	−0.02	0.05	−0.51	0.07
DEU	−0.02	0.05	−0.52	0.07	NOR	−0.02	0.05	−0.51	0.07
DNK	−0.02	0.05	−0.52	0.07	NZL	−0.02	0.05	−0.53	0.07
ECU	−0.03	0.05	−0.50	0.07	PAK	0.02	0.05	−0.55	0.07
EGY	−0.02	0.05	−0.51	0.07	PAN	−0.01	0.05	−0.52	0.07
ESP	−0.02	0.05	−0.53	0.07	PER	0.06	0.05	−0.59	0.07
FIN	−0.02	0.05	−0.54	0.07	PHL	−0.03	0.05	−0.50	0.07
FRA	−0.02	0.05	−0.51	0.07	PRT	−0.02	0.05	−0.54	0.07
GBR	−0.02	0.05	−0.51	0.07	SEN	0.00	0.05	−0.53	0.07
GHA	−0.05	0.05	−0.48	0.07	SGP	−0.02	0.05	−0.52	0.07
GRC	−0.02	0.05	−0.51	0.07	SWE	−0.02	0.05	−0.53	0.07
HKG	−0.02	0.05	−0.37	0.07	THA	−0.01	0.05	−0.51	0.07
IDN	−0.02	0.05	−0.51	0.07	TUN	−0.02	0.05	−0.51	0.07
IND	0.01	0.05	−0.54	0.07	TUR	−0.03	0.05	−0.50	0.07
IRL	−0.02	0.05	−0.54	0.07	TWN	−0.02	0.05	−0.49	0.07
ISR	−0.02	0.05	−0.52	0.07	URY	−0.02	0.05	−0.50	0.07
ITA	−0.02	0.05	−0.51	0.07	USA	−0.02	0.05	−0.53	0.07
JAM	−0.02	0.05	−0.51	0.07	VEN	0.00	0.05	−0.53	0.07
JOR	−0.04	0.05	−0.49	0.07	ZAF	−0.02	0.05	−0.51	0.07
JPN	−0.02	0.05	−0.52	0.07	ZMB	−0.02	0.05	−0.51	0.07
KEN	−0.15	0.05	−0.38	0.07	ZWE	0.03	0.05	−0.55	0.07

This table reports the estimated coefficients for λ_2 and λ_3, for the specification in Column 3 of Table 4, leaving out one country (the one indicated for each set of coefficients) at at time.

Table 7: ESTIMATION RESULTS EXCLUDING HONG KONG AND KENYA

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Change in Log-Employment							
Gap (λ_1)	0.615 (0.009)***	0.620 (0.009)***	0.649 (0.012)***	0.652 (0.012)***				
Gap×JS (λ_2):		−0.105 (0.039)***	−0.156 (0.051)***	−0.163 (0.051)***	−0.204 (0.042)***	−0.171 (0.052)***	−0.183 (0.052)***	
Gap×JS×DSRL (λ_3)			−0.231 (0.062)***			−0.062 (0.072)		
Gap×JS×DHGE (λ_3)				−0.227 (0.070)***			−0.071 (0.072)	
Gap×HP (λ_2)								−0.021 (0.007)***
Controls								
Gap×DSRL			−0.121 (0.015)***			0.065 (0.023)***		
Gap×DHGE				−0.136 (0.015)***			0.023 (0.024)	
Observations	20,881	20,881	20,881	20,881	20,881	20,881	20,881	12,003
R-squared	0.61	0.61	0.61	0.61	0.62	0.62	0.62	0.62
Gap-Income Interaction	No	No	No	No	Yes	Yes	Yes	Yes
Gap-Sector Interaction	No	No	No	No	Yes	Yes	Yes	Yes

* significant at 10%; ** significant at 5%; *** significant at 1%. Robust standard errors in parentheses. JS and HP stand for the Botero et al. (2004) and Heckman and Pages (2000) job security measures, respectively. DSRL and DHGE stand for high (above Greece, see main text) Rule of Law and Government Efficiency dummies, respectively, using the Kaufmann et al. (1999) indices. Each regression has country-year fixed effects. Gaps are estimated using a constant $\phi = 0.42$. Sample excludes the upper and lower 1% of Δe and of the estimated values of Gap.

we construct an alternative measure of the ex-post gap letting wage data play the role of productivity data when calculating the v and θ terms on the right hand side of (13).

The second procedure re-writes the model in a standard dynamic panel formulation that removes the contemporaneous employment change from the right hand side:[16]

$$\Delta \text{Gap}_{jct} = (1-\lambda_c)\Delta \text{Gap}_{jct-1} + \varepsilon_{jct}. \tag{20}$$

Table 8: IV ESTIMATION

	Average speed of adjustment	
Estimation Method	Point Estimate	Robust Standard Error
Baseline Model (Column 1 in Table 4)	0.600	0.009
Gap instrumented with wage data	0.570	0.065
Standard dynamic panel formulation	0.543	0.078

Table 8 reports the values of the average λ estimated with these two alternative procedures (note the significant decline in the precision of the estimates). For comparison purposes, the first row reproduces the first column in Table 4. The second row shows the result for the IV procedure based on using lagged changes in wages as instruments. Finally, Row 3 reports the estimate from the dynamic panel. It is apparent from the table that the estimates of average λ are in the right ballpark, and hence we conclude that the bias due to a potentially endogenous gap is not significant.

Finally, we note that the standard solution of passing the Δe-component of the gap defined in (13) to the left hand side of the estimating equation (15) does not work in our context. Passing Δe to the left suggests that the coefficient on the resulting gap will be equal to $\lambda/(1-\lambda)$. As shown in Appendix B, this holds only in the case of a partial adjustment model ($k = 0$ in the notation of Section 2.1). By contrast, in the case of a Calvo-type adjustment ($k = 1$), the corresponding coefficient will, on average, be negative.[17] More important, even small departures from a partial adjustment model (small values of k) introduce significant biases when estimating λ using this approach.

[16] To estimate this equation we follow Anderson and Hsiao (1982) and use twice and three-times lagged values of ΔGap_{jct} as instruments for the RHS variable. Similar results are obtained if we follow Arellano and Bond (1991).

[17] In the Calvo-case, for every observation either the (modified) gap or the change in employment is zero. The former happens when adjustment takes place, the latter when it does not. It follows that the covariance of Δe and the (modified) gap will be equal to minus the product of the mean of both variables. Since these means have the same sign, the estimated coefficient will be negative. See Appendix B for a formal derivation.

4 Gauging the Costs of Effective Labor Protection

By impairing worker movements from less to more productive units, effective labor protection reduces aggregate output and slows down economic growth. In this section we develop a simple framework to quantify this effect. Any such exercise requires strong assumptions and our approach is no exception. Nonetheless, our findings suggest that the costs of the microeconomic inflexibility caused by effective protection is large. In countries with strong rule of law, moving from the 20th to the 80th percentile of job security lowers annual productivity growth by close to one percentage point. The same movement for countries with weak rule of law has a negligible impact on TFP.[18]

Consider a continuum of establishments, indexed by i, that adjust labor in response to productivity shocks, while their share of the economy's capital remains fixed over time. Their production functions exhibit constant returns to (aggregate) capital, K_t, and decreasing returns to labor:

$$Y_{it} \;=\; B_{it} K_t L_{it}^{\alpha}, \tag{21}$$

where B_{it} denotes plant-level productivity and $0 < \alpha < 1$. The B_{it}'s follow geometric random walks, that can be decomposed into the product of a common and an idiosyncratic component:

$$\Delta \log B_{it} \;\equiv\; b_{it} = v_t + v_{it}^{I},$$

where the v_t are i.i.d. $\mathcal{N}(\mu_A, \sigma_A^2)$ and the v_{it}^{I}'s are i.i.d. (across productive units, over time and with respect to the aggregate shocks) $\mathcal{N}(0, \sigma_I^2)$. We set $\mu_A = 0$, since we are interested in the interaction between rigidities and idiosyncratic shocks, not in Jensen-inequality-type effects associated with aggregate shocks.

The price-elasticity of demand is $\eta > 1$. Aggregate labor is assumed constant and set equal to one. We define *aggregate productivity*, A_t, as:

$$A_t \;=\; \int B_{it} L_{it}^{\alpha} di, \tag{22}$$

so that aggregate output, $Y_t \equiv \int Y_{it} di$, satisfies

$$Y_t \;=\; A_t K_t.$$

Units adjust with probability λ_c in every period, independent of their history and of what other units do that period.[19] The parameter that captures microeconomic flexibility is λ_c. Higher values of λ_c are associated with a faster reallocation of workers in response to productivity shocks.

Standard calculations show that the growth rate of output, g_Y, satisfies:

$$g_Y \;=\; sA - \delta, \tag{23}$$

[18] Of course, a weak rule of law has an adverse impact on productivity through various channels not considered in this paper.

[19] More precisely, whether unit i adjusts at time t is determined by a Bernoulli random variable ξ_{it} with probability of success λ_c, where the ξ_{it}'s are independent across units and over time. This corresponds to the case $\zeta = 1$ in Section 2.1.

where s denotes the savings rate (assumed exogenous) and δ the depreciation rate for capital.

Now compare two economies that differ only in their degree of microeconomic flexibility, $\lambda_{c,1} < \lambda_{c,2}$. Tedious but straightforward calculations relegated to Appendix C show that:

$$g_{Y,2} - g_{Y,1} \cong (g_{Y,1} + \delta)\left[\frac{1}{\lambda_{c,1}} - \frac{1}{\lambda_{c,2}}\right]\xi, \tag{24}$$

with

$$\xi = \frac{\alpha\gamma(2-\alpha\gamma)}{2(1-\alpha\gamma)^2}\sigma^2,$$

where $\gamma = (\eta-1)/\eta$ and $\sigma^2 = \sigma_I^2 + \sigma_A^2$.[20]

We choose parameters to apply (24) as follows: The mark-up is set at 20% (so that $\gamma = 5/6$), $g_{Y,1}$ to the average rate of growth per worker in our sample for the 1980-1990 period, 0.7%, $\sigma = 27\%$,[21] $\alpha = 2/3$, and $\delta = 6\%$.

Table 9 reports the annual productivity costs of 20 percentile changes in job security regulation. These numbers are large. They imply that moving from the 20th to the 80th percentile in job security, in countries with strong rule of law, reduces annual productivity growth by 0.85%. The same change in job security legislation has a much smaller effect on TFP growth, 0.02%, in the group of economies with weak rule of law.

Table 9: PRODUCTIVITY GROWTH AND JOB SECURITY

Change in Job Security Index	Cost in Annual Growth Rate	
	Weak Rule of Law	Strong Rule of Law
20th to 40th percentile	0.002%	0.083%
40th to 60th percentile	0.007%	0.292%
60th to 80th percentile	0.008%	0.478%

Reported: change in annual productivity growth rates associated with moving across percentiles in the distribution of country job security measures computed in Botero et al. (2004). Lower values of job security index correspond to less job security. Values of speed of adjustment calculated using Column 3 in Table (4). The threshold for weak and strong rule of law is given by the OECD country with the lowest Rule of Law score (Greece). Changes in annual productivity growth calculated based on (24)). Parameter values used: $\gamma = 5/6$, $g_{Y,1} = 0.007$, $\sigma = 0.27$ $\alpha = 2/3$, and $\delta = 0.06$.

We are fully aware of the many caveats that such ceteris-paribus comparison can raise, as well as to the impact of the linear aggregate technology assumption on the growth versus levels claim, but the point of the

[20] There also is a (static) jump in the level of aggregate productivity when λ increases, given by:

$$\frac{A_2 - A_1}{A_1} \cong \left[\frac{1}{\lambda_1} - \frac{1}{\lambda_2}\right]\xi.$$

See Appendix C for the proof.

[21] This is the average across the five countries considered in Caballero et al. (2004).

table is simply to provide an alternative metric of the potential significance of observed levels of effective labor protection.

5 Concluding Remarks

Many papers have shown that, in theory, job security regulation depresses firm level hiring and firing decisions. Job security provisions increase the cost of reducing employment and therefore lead to fewer dismissals when firms are faced with negative shocks. Conversely, when faced with a positive shock, the optimal employment response takes into account the fact that workers may have to be fired in the future, and the employment response is smaller. The overall effect is a reduction of the speed of adjustment to shocks.

However, conclusive empirical evidence on the effects of job security regulation has been elusive. One important reason for this deficit has been the lack of information on employment regulation for a sufficiently large number of economies that can be integrated to cross sectional data on employment outcomes. In this paper we have developed a simple empirical methodology that has allowed us to fill some of the empirical gap by exploiting: (a) the recent publication of two cross-country surveys on employment regulations (Heckman and Pages (2000) and Botero et al. (2004)) and, (b) the homogeneous data on employment and production available in the UNIDO dataset. Another important reason for the lack of empirical success is differences in the degree of regulation enforcement across countries. We address this problem by interacting the measures of employment regulation with different proxies for law-enforcement.

Using a dynamic labor demand specification we estimate the effects of job security across a sample of 60 countries for the period from 1980 to 1998. We consistently find a relatively lower speed of adjustment of employment in countries with high legal protection against dismissal, especially when such protection is likely to be enforced.

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APPENDICES

A Firm's Optimization Problem

In this appendix we derive the representative firm's optimal policy described in (8) and (9).

The Bellman equation for the problem described in Section 2 is:

$$V(x,k) = \min_{w} \left\{ (x-w)^2 + kw^2 + \rho \int\int V(x-w+u,k')dF(u)dH(k') \right\}, \tag{25}$$

where k denotes the current draw of the adjustment cost (it can take the values 0, K and ∞), $x = e_t^* - e_{t-1}$, $w = \Delta e_t$ and $u = \Delta e_{t+1}^*$. As mentioned in the main text, e^* follows a random walk, with innovations distributed according to $F(u)$. For simplicity we set the corresponding drift (and thus the mean of F) equal to zero, yet the proof that follows can be extended to the case with non-zero drift, resulting in the same expression we derive for ν below. The variance of F is denoted by σ^2. Finally, H denotes the distribution of adjustment costs, which was characterized in (6) in the main text.

Existence and uniqueness of a solution to the Bellman equation is proved along similar lines to the standard case of quadratic adjustment costs. We therefore concentrate on characterizing the optimal policy. We posit a value function of the form :

$$V(x,k) = A + C(k)x^2, \tag{26}$$

where both A and C are positive and C depends on the current draw of the adjustment cost, k.

Substituting (26) into the r.h.s. of (25) and minimizing over w yields:

$$w(k) = \psi(k)x,$$

with

$$\psi(k) = \frac{1+\rho\mu(C)}{1+\rho\mu(C)+k}, \tag{27}$$

$$\mu(C) = \int C(k)dH(k). \tag{28}$$

Substituting the expression for $w(k)$ back into (25) and equating both sides yields:

$$A = \rho[A+\mu(C)\sigma^2] \tag{29}$$

$$C(k) = \frac{[1+\rho\mu(C)]k}{1+\rho\mu(C)+k}. \tag{30}$$

If follows that

$$A = \frac{\rho}{1-\rho}\mu(C)\sigma^2.$$

To complete the proof we next find a solution to the fixed-point problem:

$$\mu(C) = \int \frac{[1+\rho\mu(C)]k}{1+\rho\mu(C)+k} dH(k).$$

Equivalently, letting $G(k) \equiv 1+\rho C(k)$ it suffices to find a value of $\mu(G)$ that satisfies:

$$\mu(G) = \int \left[1 + \frac{\rho k \mu(G)}{k+\mu(G)}\right] dH(k). \tag{31}$$

To solve this fixed point problem, we note that:[22]

$$\int \frac{k}{k+x} dH(k) = \pi_\infty + \frac{k}{k+x}\pi_k.$$

Substituting this expression in (31) yields the following second degree equation for $x \equiv \mu(G)$:

$$(1-\rho\pi_\infty)x^2 + [k-1-\rho(\pi_\infty+\pi_k)]x - k = 0.$$

Solving this equation,[23] leads to an expression for $\mu(G)$, which combined with (27) yields:

$$\frac{1}{1-\nu(k)} = \frac{\mu(G)+k}{k} = \frac{1+k+\rho k(\pi_k-\pi_\infty) + \sqrt{[1+k+\rho k(\pi_k+\pi_\infty)]^2 - 4k\rho[(k+1)\pi_\infty + k\pi_k]}}{2k(1-\rho\pi_\infty)}.$$

This extends the well known result for quadratic adjustment costs to incorporate lumpy adjustment as well.

B Endogeneity of the Gap Measure

The model is the one described in Section 2.1. Ignoring country differences and fixing t,[24] we have that

$$\Delta e_j = \psi_j x_j, \tag{32}$$

where $x_j \equiv e^*_{jt} - e_{j,t-1}$ and the ψ are i.i.d., independent of the x_i, with mean λ and variance $\zeta\lambda(1-\lambda)$, $0 \le \zeta \le 1$. It follows from (32) that:

$$\Delta e_j = \lambda x_j + u_j, \tag{33}$$

with $u_j \equiv (\psi_j - \lambda)x_j$ satisfying all properties of an error term in a standard regression setting, except for homoscedasticity. Thus, if we observe the Δe_j and x_j and estimate (33), we obtain an unbiased and consistent estimator for λ.[25]

Removing Δe from the gap measure is equivalent to replacing x_j by $z_j \equiv x_j - \Delta e_j$ in (33). It is obvious that $z_j = 0$ when $\psi = 1$ and $\Delta e_j = 0$ when $\psi_j = 0$. Since the two values ψ_j takes in the Calvo case ($\zeta = 1$) are 0 and 1, we have that when estimating the regression coefficient via OLS the covariance in the numerator

[22] So far we have not used our assumptions on the distribution of adjustment costs. The approach followed here extends easily, for example, to the case where the parameter K can take more than one value.

[23] Only the positive root implies an economically meaningful positive value for $\mu(G)$.

[24] The latter is justified by the fact that most of our identification comes from cross-sectional variation.

[25] This, in a nutshell, is the essence of the estimation procedure described in detail in Section 2 of the main text, with x_j corresponding to the gap measure defined in (13).

will be equal to minus the product of the average of Δe_j and the average of the z_j. Since both averages have the same sign, it follows that the regression coefficient will be negative (or zero if both averages are equal to zero).

Since we are considering sectoral data, the case $\zeta = 1$ may seem somewhat extreme. The following proposition shows that even for small departures from the partial adjustment case, the bias is likely to be significant.

Proposition 1 *Consider the setting described above. Denote by $\hat{\beta}$ the OLS estimate of the coefficient β in*

$$\Delta e_j = const + \beta z_j + error.$$

Denote by μ and σ^2 the (theoretical) mean and variance of the x_j's. Then:

$$\text{plim}_{N\to\infty}\hat{\beta} = \frac{(1-\zeta)\sigma^2 - \zeta\mu^2}{[1-(1-\zeta)\lambda]\sigma^2 + \zeta\lambda\mu^2}\lambda. \tag{34}$$

Proof From $\Delta e_j = \psi_j x_j$ and $z_j = (1-\psi_j)x_j$ it follows that:

$$\text{Cov}(\Delta e, z) = \frac{1}{N}\sum_i \psi_i(1-\psi_i)x_i^2 - \left(\frac{1}{N}\sum_i \psi_i x_i\right)\left(\frac{1}{N}\sum_i (1-\psi_i)x_i\right),$$

and

$$\text{Var}(z) = \frac{1}{N}\sum_i (1-\psi_i)^2 x_i^2 - \left[\frac{1}{N}\sum_i (1-\psi_i)x_i\right]^2.$$

Taking expectations over the ψ_i, conditional on the x_i, and letting N tend to infinity leads to:

$$\text{E}[\hat{\beta}] = \frac{\text{E}[\psi(1-\psi)](\sigma^2+\mu^2) - \lambda(1-\lambda)\mu^2}{\text{E}[(1-\psi)^2](\sigma^2+\mu^2) - (1-\lambda)^2\mu^2}.$$

The result now follows from the expression above and the fact that:

$$\begin{aligned}
\text{E}[\psi(1-\psi)] &= (1-\zeta)\lambda(1-\lambda), \\
\text{E}[(1-\psi)^2] &= [1-(1-\zeta)\lambda](1-\lambda). \quad \blacksquare
\end{aligned}$$

It follows that $\text{plim}_{N\to\infty}\hat{\beta}$ is decreasing in ζ, varying from $\lambda/(1-\lambda)$ when $\zeta = 0$ to $-\lambda\mu^2/(\sigma^2+\lambda\mu^2)$ when $\zeta = 1$. It also follows that $\text{plim}_{N\to\infty}\hat{\beta}$ is decreasing in $|\mu|$, so that:

$$\text{plim}_{N\to\infty}\hat{\beta} \leq \frac{(1-\zeta)\lambda}{1-(1-\zeta)\lambda}.$$

C Gauging the Costs

In this appendix we derive (24). From (23) and (24) it follows that it suffices to show that under the assumptions in Section 4 we have:

$$\frac{A_2 - A_1}{A_1} \cong \left[\frac{1}{\lambda_1} - \frac{1}{\lambda_2}\right] \xi, \tag{35}$$

where we have dropped the subindex c from the λ and

$$\xi = \frac{\alpha\gamma(2-\alpha\gamma)}{2(1-\alpha\gamma)^2}(\sigma_I^2 + \sigma_A^2), \tag{36}$$

with $\gamma = (\eta - 1)/\eta$.

The intuition is easier if we consider the following, equivalent, problem. The economy consists of a very large and fixed number of firms (no entry or exit). Production by firm i during period t is $Y_{i,t} = A_{i,t}L_{i,t}^{\alpha}$,[26] while (inverse) demand for good i in period t is $P_{i,t} = Y_{i,t}^{-1/\eta}$, where $A_{i,t}$ denotes productivity shocks, assumed to follow a geometric random walk, so that

$$\Delta \log A_{i,t} \equiv \Delta a_{i,t} = v_t^A + v_{i,t}^I,$$

with v_t^A i.i.d. $\mathcal{N}(0, \sigma_A^2)$ and $v_{i,t}^I$ i.i.d. $\mathcal{N}(0, \sigma_I^2)$. Hence $\Delta a_{i,t}$ follows a $\mathcal{N}(0, \sigma_T^2)$, with $\sigma_T^2 = \sigma_A^2 + \sigma_I^2$. We assume the wage remains constant throughout.

In what follows lower case letters denote the logarithm of upper case variables. Similarly, *-variables denote the frictionless counterpart of the non-starred variable.

Solving the firm's maximization problem in the absence of adjustment costs leads to:

$$\Delta l_{i,t}^* = \frac{\gamma}{1-\alpha\gamma}\Delta a_{i,t}, \tag{37}$$

and hence

$$\Delta y_{i,t}^* = \frac{1}{1-\alpha\gamma}\Delta a_{i,t}. \tag{38}$$

Denote by Y_t^* aggregate production in period t if there were no frictions. It then follows from (38) that:

$$Y_{i,t}^* = e^{\tau \Delta a_{i,t}} Y_{i,t-1}^*, \tag{39}$$

with $\tau \equiv 1/(1-\alpha\gamma)$, Taking expectations (over i for a particular realization of v_t^A) on both sides of (39) and noting that both terms being multiplied on the r.h.s. are, by assumption, independent (random walk), yields

$$Y_t^* = e^{\tau v_t^A + \frac{1}{2}\tau^2\sigma_I^2} Y_{t-1}^*, \tag{40}$$

Averaging over all possible realizations of v_t^A (these fluctuations are not the ones we are interested in for the calculation at hand) leads to

$$Y_t^* = e^{\frac{1}{2}\tau^2\sigma_T^2} Y_{t-1}^*,$$

[26] That is, we ignore hours in the production function.

and therefore for $k = 1, 2, 3, ...$:

$$Y_t^* = e^{\frac{1}{2}k\tau^2\sigma_T^2} Y_{t-k}^*. \tag{41}$$

Denote:

- $Y_{t,t-k}$: aggregate Y that would attain in period t if firms had the frictionless optimal levels of labor corresponding to period $t-k$. This is the average Y for units that last adjusted k periods ago.

- $Y_{i,t,t-k}$: the corresponding level of production of firm i in t.

From the expressions derived above it follows that:

$$\frac{Y_{i,t,t-1}}{Y_{i,t}^*} = \left(\frac{L_{i,t-1}^*}{L_{i,t}^*}\right)^\alpha = e^{-\alpha\gamma\tau\Delta a_{i,t}},$$

and therefore

$$Y_{i,t,t-1} = e^{\Delta a_{i,t}} Y_{i,t-1}^*.$$

Taking expectations (with respect to idiosyncratic and aggregate shocks) on both sides of the latter expression (here we use that $\Delta a_{i,t}$ is independent of $Y_{i,t-1}^*$) yields

$$Y_{t,t-1} = e^{\frac{1}{2}\sigma_T^2} Y_{t-1}^*,$$

which combined with (41) leads to:

$$Y_{t,t-1} = e^{\frac{1}{2}(1-\tau^2)\sigma_T^2} Y_t^*.$$

A derivation similar to the one above, leads to:

$$Y_{i,t,t-k} = e^{\Delta a_{i,t} + \Delta a_{i,t-1} + \ldots + \Delta a_{i,t-k+1}} Y_{t-k}^*,$$

which combined with (41) gives:

$$Y_{t,t-k} = e^{-k\xi} Y_t^*, \tag{42}$$

with ξ defined in (36).

Assuming Calvo-type adjustment with probability λ, we decompose aggregate production into the sum of the contributions of cohorts:

$$Y_t = \lambda Y_t^* + \lambda(1-\lambda)Y_{t,t-1} + \lambda(1-\lambda)^2 Y_{t,t-2} + \ldots$$

Substituting (42) in the expression above yields:

$$Y_t = \frac{\lambda}{1-(1-\lambda)e^{-\xi}} Y_t^*. \tag{43}$$

It follows that the production gap, defined as:

$$\text{Prod. Gap} \equiv \frac{Y_t^* - Y_t}{Y_t^*},$$

is equal to:

$$\text{Prod. Gap} = \frac{(1-\lambda)(1-e^{-\xi})}{1-(1-\lambda)e^{-\xi}}. \tag{44}$$

A first-order Taylor expansion then shows that, when $|\xi| << 1$:

$$\text{Prod. Gap} \cong \frac{(1-\lambda)}{\lambda}\xi. \tag{45}$$

Subtracting this gap evaluated at λ_1 from its value evaluated at λ_2, and noting that this gap difference corresponds to $(A_2 - A_1)/A_1$ in the main text, yields (35) and therefore concludes the proof.