3-1-2004

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CENTER DISCUSSION PAPER NO. 884

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March 2004

Notes: Center Discussion Papers are preliminary materials circulated to stimulate discussions and critical comments.

We thank Fernando Coloma and participants at the 7th Annual Conference of the Central Bank of Chile for their comments.

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Abstract

We characterize the degree of microeconomic inflexibility in several Latin American economies and find that Brazil, Chile and Colombia are more flexible than Mexico and Venezuela. The difference in flexibility among these economies is mainly explained by the behavior of large establishments, which adjust more promptly in the more flexible economies, especially when accumulated shocks are substantial. We also study the path of flexibility in Chile and show that it declined in the aftermath of the Asian crisis. This decline can account for a substantial fraction of the large decline in TFP-growth in Chile since 1997 (from 3.1 percent per year for the preceding decade, to about 0.3 percent after that). Moreover, if it were to persist, it could permanently shave off almost half of a percent from Chile's structural rate of growth.

Keywords: Microeconomic rigidities, creative-destruction, job flows, restructuring and reallocation, productivity growth

JEL Codes: E2, J2, J6

1 Introduction

Although with varying degrees of success, Latin American economies have begun to leave behind some of the most primitive sources of macroeconomic fluctuations. Gradually, policy concern is shifting toward increasing microeconomic flexibility. This is a welcome trend since, by facilitating the ongoing process of creative-destruction, microeconomic flexibility is at the core of economic growth in modern market economies.

But how poorly are these economies doing along this flexibility dimension? Answering this question requires measuring the important but elusive concept of microeconomic flexibility. How do we do this?

One way is to look directly at regulation, perhaps the main institutional factor hindering or facilitating microeconomic flexibility. In particular, there are extensive studies of labor market regulation. Heckman and Pages (2000), for example, document that "even after a decade of substantial deregulation [in most cases], Latin American countries remain at the top of the Job Security list, with levels of regulation similar to or higher than those existing in the highly regulated South of Europe." This is important work. However, in practice microeconomic flexibility depends not only on labor market regulation, but also on a wide variety of factors, including the technological options and nature of the production process, the political environment, the efficiency and biases of labor courts, as well as cultural variables and accepted practices. Thus, while useful for eventual policy formulation, studies of rules and regulation are unlikely to provide us with the "big picture" of a country's flexibility any time soon — understanding the complex interactions of different regulations and environments is a valuable but very slow process.

At the other extreme, one can look at outcomes directly: How much factor reallocation do we see in different countries and episodes? This is also a useful exercise. However, it is equally incomplete since there is no reason to expect the same degree of aggregate flows in countries facing different idiosyncratic and aggregate shocks. Hence it is always difficult to know whether the observed reallocation is abnormally high or low, since the counterfactual is not part of the statistic.

A third approach, which remedies some of the main weaknesses of the previous ones, is to measure microeconomic flexibility by the speed at which establishments reduce the gap between their labor productivity and the marginal cost of such labor. Thus, we say an economy is inflexible at the microeconomic level if these gaps persist over time. Conversely, a very flexible economy, firm, or establishment, is one in which gaps disappear quickly due to prompt adjustment. This is the approach we follow in this paper, extending a methodology developed in Caballero, Cowan, Engel and Micco (2004) — the main advantage of this methodology over conventional partial adjustment estimates is its ability to use limited information efficiently, correcting standard biases often present when estimating such models. Our methodology also allows for nonlinearities and state-dependent responses of employment to productivity gaps, as in Caballero and Engel (1993).[1]

We use establishment level observations for all the Latin American economies for which we had access to fairly reliable data: Chile, Mexico and, to a lesser extent, Brazil, Colombia and Venezuela. All in all, about 140,000 observations.

In the first part of the paper we document the main features of adjustment for these economies. We find that:

- While more inflexible than the US, on average (over time) Brazil, Colombia and Chile exhibit a relatively high degree of microeconomic flexibility with over 70 percent of labor adjustment taking place within a year. Mexico ranks lower with about 60 percent of adjustment within a year, and Venezuela is the most inflexible of these economies, with slightly over 50 percent of adjustment within a year.

- With the only exception of Venezuela, in all our economies small establishments (below the median number of employees) are substantially less flexible than large establishments (above the 75th percentile of employees). In Brazil, the former establishments close about 67 percent of their gap within a year, while the latter close about 81 percent. In Colombia, 68 and 79, respectively; in Chile 69 and 78; Mexico 56 and 61; and Venezuela 53 percent for both.

- It follows from the previous finding that it is primarily the behavior of large establishments that is behind the substantial differences in flexibility across some of the economies we study. It may well be the case that large companies in Venezuela and Mexico are more insulated from competitive pressures than their counterparts in Colombia, Chile and Brazil.

- In all these economies there is evidence of an "increasing hazard". That is, establishments are substantially more flexible with respect to large gaps than to small ones. This points to the presence of significant fixed costs of adjustment, which may have a technological or institutional origin.

- The increasing hazard feature is particularly pronounced in large establishments in the relatively more flexible economies. In fact, most of the additional flexibility experienced by

[1] Note that our definition of microeconomic flexibility refers to the speed at which establishments react to changing conditions; *not* to whether the labor market is flexible or not in responding to aggregate shocks. Thus, a labor market regulation that makes the real wage rigid will result in a larger unemployment response to aggregate shocks —that is, it will exhibit *macroeconomic* inflexibility— yet this will not be part of our measure of *microeconomic* inflexibility.

large establishments in these economies is due to their rapid adjustment when gaps get to be large. For example, when gaps are below 25 percent in Chile, small establishments have an adjustment coefficient of 0.50 while large ones have one of 0.51. For deviations above 25%, on the other hand, small establishments have a coefficient of 0.79, while large establishments have one of 0.93. The patterns are similar in Brazil and Colombia, yet less pronounced in Mexico and Venezuela.

In the second part of the paper we specialize on Chile, which has the only long panel in our sample, and explore the evolution of its microeconomic flexibility over time. Our main findings are the following:

- Microeconomic flexibility in Chile experienced a significant decline toward the end of our sample (1997-99). From an average adjustment coefficient of 0.77 for the three years prior to the Asian/Russian crisis episode, the coefficient fell to 0.69 in the aftermath of the crisis.

- When the adjustment hazard is assumed to be constant, the decline in flexibility appears to be subsiding toward the end of the sample. However, this finding is lost and there is no evidence of recovery once the hazard is allowed to be increasing. The reason for the misleading conclusion with a constant hazard is that toward the end of the sample there is a sharp rise in the share of establishments with large (negative) gaps, to which establishments naturally react more under increasing hazards.

- While it is too early to tell whether the decline we uncover is purely cyclical, or whether there is something more structural going on, there are a few interesting observations to make:

  a) Much of the decline in flexibility is due to a decline in the flexibility of large establishments.

  b) While the speed of response to negative gaps remained fairly constant, it is the speed at which establishments adjust to shortages of labor that slowed down more dramatically. This "reluctance to hire" may reflect pessimism regarding future conditions not captured in the contemporaneous gap. But this is unlikely to be the only factor since otherwise we also should observe a rise in the speed of firing (for a given hazard).[2]

  c) Finally, the sharpest decline in flexibility came from establishments in sectors that normally experience less restructuring, either because of smaller shocks or more technological and institutional inflexibility. If either form of inflexibility is responsible for

[2]While we did see an increase in the speed of firing, as we argued above, this is accounted for by the interaction of a prolonged contraction with an increasing hazard.

reduced restructuring, then the cost of the decline in flexibility can be potentially very large, as already inflexible establishments spend significant time away from their frictionless optimum.

In the last part of the paper we explore a different metric for the degree of inflexibility and its economic impact. By impairing worker movements from less to more productive units, microeconomic inflexibility reduces aggregate output and slows down economic growth. We develop a simple framework to quantify this effect. Our findings suggest that the aggregate consequences of micro-inflexibilities in Latin America are significant. In particular, the impact of the decline in microeconomic flexibility in Chile following the Asian crisis is in itself large enough to account for a substantial fraction of the decline in TFP-growth in Chile since 1997 (from an annual average of 3.1 percent for the preceding decade to about 0.3 percent after that). Moreover, if it were to persist, it could permanently shave off almost half of a percent from Chile's structural rate of growth.

Section 2 presents the methodology while Section 3 describes the data. Section 4 characterizes average microeconomic flexibility in the Latin American economies in our data. Section 5 explores the case of Chile in more detail, and describes the evolution of its index of flexibility. Section 6 presents a simple model to map microeconomic inflexibility into growth outcomes. Section 7 concludes and is followed by several appendices.

2 Methodology and Data

2.1 Overview

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

$$\Delta e_{ijt} = \psi_{ijt}(e^*_{ijt} - e_{ijt-1}), \tag{1}$$

where e_{ijt} and e^*_{ijt} denote the logarithm of employment and desired employment, respectively. The random variable ψ_{ijt}, which is assumed i.i.d. both across establishments and over time, takes values in the interval $[0,1]$ and has mean λ and variance $\alpha\lambda(1-\lambda)$, with $0 \leq \alpha \leq 1$. The case $\alpha = 0$ corresponds to the standard quadratic adjustment model, the case $\alpha = 1$ to the Calvo (1983) model. The parameter λ captures microeconomic flexibility. As λ goes to one, all gaps are closed quickly and microeconomic flexibility is maximum. As λ decreases, microeconomic flexibility declines.

Equation (1) also hints at two important components of our methodology: We need to find a measure of the employment gap, $(e^{*}_{ijt} - e_{ijt-1})$, and an estimation strategy for the mean of the random variable ψ_{ijt}, λ. We describe both ingredients in detail in what follows. In a nutshell, we construct estimates of e^{*}_{ijt}, the only unobserved element of the gap, by solving the optimization problem of the firm, as a function of observables such as labor productivity and a suitable proxy for the average market wage. We estimate λ from (1), based upon the large cross-sectional size of our sample and the well documented fact that there are significant idiosyncratic components in the realizations of the gaps and the ψ_{ijt}'s.

An important aspect of our methodology is to find an efficient method to remove fixed effects while, at the same time, avoiding the standard biases present in dynamic panel estimation.[3] The model we develop also leads to a standard dynamic panel formulation, namely:[4]

$$\mathrm{Gap}_{ijt} \; = \; (1-\lambda)\Delta e^{*}_{ijt} + (1-\lambda)\mathrm{Gap}_{ijt-1} + \varepsilon_{ijt}. \tag{2}$$

We report results for this specification as well, using dynamic panel techniques, in Table 12. They are consistent with the estimates we obtain based on (1) and therefore provide a useful robustness check. Yet they are considerably less precise. Thus our methodology may be viewed as an alternative, for the particular problem at hand, that uses data more efficiently than standard dynamic panel estimation techniques.

2.2 Details

Output and demand for establishment are given by:

$$\begin{aligned} y &= a + \alpha e + \beta h, & (3)\\ p &= d - \frac{1}{\eta} y, & (4) \end{aligned}$$

where y, p, e, a, h, d denote firm output, price, employment, productivity, hours worked and demand shocks, and η is the price elasticity of demand. We let $\gamma \equiv (\eta - 1)/\eta$.[5] All variables are in logs.

Firms are competitive in the labor market but pay wages that are increasing in the average

[3]As documented, for example, in Arellano and Bond (1991).

[4]The "Gap" below could be the gap *before* or *after* adjustments take place.

[5]In order to have interior solutions, we assume $\eta > 1$ and $\alpha\gamma < 1$.

number of hours worked, according to:[6]

$$w = w^o + \mu(h - \overline{h}), \tag{5}$$

where $\overline{h}$ is constant over time and interpreted below.[7]

A key assumption is that firms only face adjustment costs when they change employment levels, not when they change the number of hours worked.[8] It follows that the firm'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 (FOC) for hours.

In a frictionless labor market the firm's employment level also satisfies a FOC for employment. Our functional forms then imply that the optimal choice of hours does not depend on the employment level.[9] We denote the corresponding employment level by $\widehat{e}$ and refer to it as the *static employment target*.[10] The following relation between the employment gap and the hours gap then follows:

$$\widehat{e} - e \;=\; \frac{\mu - \beta\gamma}{1 - \alpha\gamma}(h - \overline{h}). \tag{6}$$

This is the expression used by Caballero et Engel (1993). It is not useful in our case, since we do not have information on worked hours. Yet the argument used to derive (6) 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 (5). This result is intuitive: the employment response to a given deviation of wages from marginal product will

[6]The expression below should be interpreted as a convenient approximation for:

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

with w^o and μ determined by k^o and Ω.

[7]To ensure interior solutions, we assume $\alpha\mu > \beta$ and $\mu > \beta\gamma$.

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

[9]A patient calculation shows that

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

.

[10]We have:

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

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

be larger if the marginal cost of the alternative adjustment strategy —changing hours— is higher. Also note that $\widehat{e}-e$ is the difference between the static target $\widehat{e}$ and realized employment, not the dynamic employment gap $e^*_{ijt}-e_{ijt}$ related to the term on the right hand side of (1). However, we assume that demand, productivity and wage shocks follow a random walk.[11] We then have that e^*_{ijt} is equal to $\widehat{e}_{ijt}$ plus a constant δ_t.[12] It follows that

$$e^*_{ijt}-e_{ijt-1} \;=\; \frac{\phi}{1-\alpha\gamma_j}\left(v_{ijt}-w^o_{ijt}\right)+\Delta e_{ijt}+\delta_t, \tag{7}$$

where we have allowed for sector-specific differences in γ.

We estimate the marginal productivity of labor (v_{ijt}) using output per worker multiplied by an industry-level labor share, assumed constant over time.

Two natural candidates to proxy for w^o_{ijt} are the average (across each industry, at a given point in time) of either observed wages or observed marginal productivities. The former is consistent with our assumption of a competitive labor market, the latter may be expected to be more robust in settings with long-term contracts and multiple forms of rewards, where the salary may not represent the actual marginal cost of labor.[13] Our estimations were performed using both alternatives and we found no discernible differences. This suggests that statistical power comes mainly from the cross-section dimension, that is, from the well documented and large magnitude of idiosyncratic shocks faced by establishments. In what follows we report the more robust alternative and approximate w^o by the average marginal productivity, which leads to:

$$e^*_{ijt}-e_{ijt-1} \;=\; \frac{\phi}{1-\alpha\gamma_j}(v_{ijt}-v_{\cdot jt})+\Delta e_{ijt}+\delta_t \;\equiv\; \mathrm{Gap}_{ijt}+\delta_t. \tag{8}$$

The expression above ignores systematic variations in labor productivity that may occur across establishments, which would tend to bias estimates of the speed of adjustment downward. In Appendix A we provide evidence in favor of incorporating this possibility by subtracting from $(v_{ijt}-v_{\cdot jt})$ in (8) a moving average of relative productivity by establishment, $\widehat{\theta}_{ijt}$.[14] The resulting

[11] Fron the preceding footnote if follows that it suffices that $d+\gamma a-w^o$ follows a random walk.

[12] In order to allow for variations in future expected growth rates of a and d, the constant δ is allowed to vary over time.

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

[14] Where $\widehat{\theta}_{ijt}\equiv\frac{1}{2}[(v_{ijt-1}-v_{\cdot jt-1})\;+\;(v_{ijt-2}-v_{\cdot jt-2})]$. The alternative specification, with relative wages instead of relative marginal productivities, leads to almost identical results.

expression for the estimated employment-gap is:[15]

$$e^{*}_{ijt} - e_{ijt-1} \;=\; \frac{\phi}{1-\alpha\gamma_j}(v_{ijt} - \widehat{\theta}_{ijt} - v_{\cdot jt}) + \Delta e_{ijt} + \delta_t \;\equiv\; \text{Gap}_{ijt} + \delta_t, \tag{9}$$

Finally, we estimate ϕ (related to the substitutability between hours worked and employment) using

$$\Delta e_{ijt} \;=\; -\frac{\phi}{1-\alpha\gamma_j}(\Delta v_{ijt} - \Delta v_{\cdot jt}) + \kappa_t + \upsilon_{it} + \Delta e^{*}_{ijt} \;\equiv\; -\phi z_{ijt} + \kappa_t + \varepsilon_{ijt}, \tag{10}$$

where κ is a year dummy, Δe^{*}_{ijt} is the change in the desired level of employment and $z_{it} \equiv (\Delta v_{ijt} - \Delta v_{\cdot jt})/1-\alpha\gamma_j)$. By assumption Δe^{*}_{ijt} is i.i.d. and independent of lagged variables. In order to avoid endogeneity and measurement error bias we estimate (10) using $(\Delta w_{ijt-1} - \Delta w_{\cdot jt-1})$ as an instrument for $(\Delta v_{ijt} - \Delta v_{\cdot jt})$.[16] Table 1 reports the estimation results of (10) across the countries in our sample.[17] We report estimates both with and without the one percent of extreme values for the independent variable. For ease of comparison across countries, based on the estimates reported in Table 1 we choose a common value of ϕ equal to 0.40.

2.3 Summary

Our methodology has three advantages when compared with previous specifications used to estimate cross-country differences in speed of adjustment. First, it only requires data on nominal output and employment level, two standard and well-measured variables in most industrial surveys. Most previous studies on adjustment costs require measures of real output or an exogenous measure of sector demand.[18] Second, it summarizes in a single variable all shocks faced by a firm. This feature allows us to increase precision, and therefore the power of hypothesis testing, and to study the determinants of the speed of adjustment using interaction terms. Finally, our approach can be extended to incorporate non-linearities in the adjustment function. That is, the possibility that the ψ in (1) depend on the gap before adjustments take place. This feature also turns out to be

[15] Where $\alpha\gamma_j$ is constructed using the sample median of the labor share for sector j across year and countries (Brazil, Chile, Colombia, Mexico and Venezuela).

[16] We lag the dependent variable because it is correlated with the error term, and we use lagged wages to instrument lagged labor productivity to avoid measurement errors.

[17] We do not have wage data for Brazil, so we cannot estimate the parameter for this country.

[18] 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 on employment are usually marginally significant, and often of the opposite sign than expected.

useful.

Summing up, in our basic setup we estimate the microeconomic flexibility parameter λ from

$$\Delta e_{ijt} = \lambda(\text{Gap}_{ijt} + \delta_t) + \varepsilon_{ijt}, \tag{11}$$

where Gap_{ijt} is proportional to the gap between marginal labor productivity and the market wage. To correct for labor heterogeneity across establishments, a fixed effect is also included in the gap-measure. This fixed effect is estimated by the average labor productivity in the two preceding periods. As shown in Appendix A, the resulting estimator is unbiased (on average). It forces us to discard only two time periods, and can adapt to slow time variations in heterogeneity.

3 Data and basic facts

This section describes the source and data used in the empirical analysis. These data are from manufacturing censuses and surveys conducted by national statistical government agencies in five Latin American countries: Brazil, Chile, Colombia, Mexico and Venezuela. The variables used in our analysis are nominal output, employment, total compensation and industry classification within the manufacturing sector (ISIC at three digits). For the case of Chile, we also use capital stock and a measure of cash flow defined as sales minus total input costs.

For Brazil, the data are from the Manufacturing Annual Survey (Pesquisa Industrial Anual) conducted by the Instituto Brasileiro de Geografia e Estatística. This survey started in 1967 but experienced a severe methodological change in 1996, thus we only use observations from 1996 to 2000. In this, as well as in all other countries, we only include plants that existed during the full period (continuous plants). In the case of Chile the data are from the Chilean Manufacturing Census (Encuesta Nacional Industrial Anual) conducted by the Instituto Nacional de Estadísticas. In principle, the surveys covers all manufacturing plants in Chile with more than ten employees during the period 1979-97. In the empirical section we only use continuous plants during the period 1985-97. We do not use the years before 1985 because they are characterized by large macroeconomic shocks and structural adjustments that introduce too much noise and complications to our methodology. For Colombia we use the Colombian Manufacturing Census (Encuesta Anual Manufacturera y Registro Industrial) conducted by the Departamento Administrativo Nacional de Estadísticas. The survey covers all manufacturing plants with more than twenty employees during the period 1982-99. For plants with less than twenty employees only a random sample is covered. Again, we only use continuous plants during the period 1992-99 due to a methodological change in the survey in 1992.

For Mexico we use the Mexican Manufacturing Annual Survey (Encuesta Industrial Anual) conducted by the Instituto Nacional de Estadística, Geografía e Informática. The survey covers a random sample of firms in the manufacturing sector during the period 1993-2000. Finally, for Venezuela the data are from the Manufacturing Survey (Encuesta Industria Manufacturera) conducted by the Instituto Nacional de Estadistica. The survey covers all plants with more than 50 employees and it has a yearly random sample for plants with less than 50 employees. Due to changes in the methodology, we only are able to follow firms during the 1995-1999 period.

Table 2 presents the number of observations per size bracket (measured by the number of employees) for each of the five countries, for the sample period at hand. The coverage of plants by size differs across countries. Chile and Colombia have the largest coverage of small plants (less than 50 employees), whereas Venezuela's survey mainly covers large establishments.

In table 3 we compute the average job creation and job destruction for each country. In addition we report the simple average over time of net change in employment and the excess turnover (i.e., the sum of job flows net of the change in employment due to cyclical factors). All statistics are defined following Davis et al. (1996). It is already apparent in these numbers that microeconomic flexibility in these countries is limited: they are of the same order of magnitude of those of developed economies —which presumably need less restructuring than catching-up emerging economies— and substantially below economies such as Taiwan.[19]

4 Microeconomic Flexibility

In this section we report our average (over time) flexibility findings. The basic results are reported in Table 4. All of our regressions include year-dummies, d_t. That is, for each country, we estimate (we drop the sector j subscript):

$$\Delta e_{it} = d_t + \lambda \text{Gap}_{it} + \varepsilon_{it}. \tag{12}$$

The first apparent result is that microeconomic flexibility is more limited in our economies than in the very flexible US. In the latter, estimates of λ using annual data are much closer to one.[20]

Although comparisons must be interpreted with caution since the samples differ in number of observations, time-periods, establishments' demographics, etc., there is a discernible pattern.

[19] See e.g., Caballero and Hammour (2000) and references therein.

[20] For example, Caballero, Engel and Haltiwanger (1997) find a *quarterly* λ for US manufacturing exceeding 0.4, which implies an annual λ of approximately 0.90.

Within the region, Brazil, Colombia and Chile exhibit a relatively high degree of microeconomic flexibility with over 70 percent of labor adjustment taking place within a year. Mexico ranks lower with about 60 percent of adjustment within a year, and Venezuela is the most inflexible of these economies, with slightly more than 50 percent of adjustment within a year.

Lending support to our earlier motivation for adopting our approach in constructing a broad measure of microeconomic inflexibility, our ranking is essentially uncorrelated with the ranking obtained by Heckman and Pages (2000) and Botero et al. (2003) based on measuring labor market regulations (see Table 5). For example, and in contrasts to our results, the Botero et al. (2003) index of job security places Venezuela at a level of flexibility similar to that of Brazil and Chile, and Colombia as significantly more flexible than all of the above.[21]

Table 6 reports the results from repeating estimation of regression (12), but conditioning on whether establishments are small or large. The former are defined as those with a number of employees below the median in the preceding year, large ones are those above the 75th percentile in number of employees (also in the preceding year).

In all our economies but Venezuela, small firms are substantially less flexible than large establishments. In Brazil, the former close about 67 percent of their gap within a year, while the latter close about 81 percent. In Colombia, 68 and 79, respectively; in Chile 69 and 78; Mexico 56 and 61; and Venezuela 53 percent for both.

It also follows from this table that it is primarily the behavior of "large" establishments that explains the substantial differences in flexibility across some of these economies. Again, this need not come from differences in labor market regulation — and hence it would not be captured by such indices — but it could also reflect, for example, barriers to entry or social objectives assigned to large firms.

In addition to splitting by size, Table 7 splits observations by the magnitude of the employment-gap. Small gaps are defined as gaps of less than 25 percent, in absolute value, while large ones are for gaps above 25 percent. That is, we re-estimate (12) for each country-size/size-of-gap combination (jsg):

$$\Delta e_{ijsgt} = d_{jsgt} + \lambda_{jsg}\mathrm{Gap}_{ijsgt} + \varepsilon_{ijsgt}. \tag{13}$$

There are several significant conclusions that follow from this table:

1. In all the economies we study there is evidence of an *increasing hazard.*[22] That is, establishments are substantially more flexible with respect to large gaps than to small ones. This hints

[21] Also, according to the Heckman and Pages (2000) index, the most flexible countries in our sample are Brazil and Mexico; not Chile and Colombia as suggested by our index.

[22] See Caballero and Engel (1993) for a description of increasing hazard models and their aggregate implications.

at the presence of significant fixed costs (increasing returns) in the adjustment technology. These fixed costs may have a technological origin, as when there are strong complementarities in production or fixed proportion with sunk capital, or institutional, as when dismissals require approval by a government agency or are likely to be litigated in court.

2. The increasing hazard feature is particularly pronounced in large establishments in the relatively more flexible economies. This does not mean that these firms face larger fixed costs than the same establishments in less flexible economies. Quite the opposite, since they still adjust more frequently than their counterparts in inflexible economies. It means that the benefits of adjustments overcome fixed costs sooner in large establishments in flexible economies and that there are more elements of randomness (i.e., not correlated with the size of the gap) in the adjustment decisions of large establishments in inflexible economies.

3. In fact, most of the additional flexibility experienced by large establishments in the more flexible Latin American economies is due to their rapid adjustment when gaps get to be very large (over 25 percent). For example, both small and large establishments have an adjustment coefficient of approximately 0.50 for gaps below 25% in Chile. For large deviations, on the other hand, small establishments have a coefficient of 0.79, while large establishments have one of 0.93. The patterns are similar in Brazil and Colombia, and less pronounced in Mexico and Venezuela.

In conclusion, there is evidence of microeconomic inflexibility in the Latin American economies, and in some cases, such as Mexico and Venezuela, the problem is quite severe. Studies based only on quantifying job flows would be unable to detect either of these facts: Gross job flows are comparable in magnitude to those in the US, and across all the economies we study, or yield the wrong ranking (e.g., Chile would be the second most inflexible of these economies, according to the excess reallocation numbers presented in Table 3); the same remark applies to studies solely based on studying labor markets regulation.[23]

We also find that allowing for an increasing hazard is important: There is clear evidence of increasing hazards, especially for large establishments in the more flexible economies. To a substantial extent, more inflexible economies seem to be those where large imbalances go uncorrected for sustained periods of time. Conversely, large establishments in the more flexible economies seldom tolerate (or can afford to tolerate) large microeconomic imbalances.

[23] Of course there is plenty of merit and usefulness in such studies. Our remarks only refer to our attempt of measuring a broad concept of microeconomic flexibility.

5 The Evolution of Flexibility

Has microeconomic flexibility improved over time? Unfortunately, we only count with a long time dimension for the case of Chile. In what follows we specialize our analysis to this case, and conclude that the answer to this question is negative. Quite the opposite, flexibility has declined significantly since the Asian crisis.

All our results in this section are obtained from running variants of the regression:

$$\begin{aligned}\Delta e_{ijt} \;=\; & [\lambda_{0jt}+\lambda_{1j}\{|\mathrm{Gap}_{ijt}|>0.25\}+\lambda_{2j}\{\mathrm{Gap}_{ijt}<-0.05\}]\mathrm{Gap}_{ijt}\;+ \\ & + d_{1j}\{|\mathrm{Gap}_{ijt}|>0.25\}\,+\,d_{2j}\{\mathrm{Gap}_{ijt}<-0.05\}\,+\,\varepsilon_{ijsgt}, \qquad (14)\end{aligned}$$

where we include, but do not report, constants, time and group (e.g., $|\mathrm{Gap}_{ijt}| > 0.25$) dummies. The results of these variants are reported in Table 8.

Figure 1 plots the path of the λ_{0jt}'s, with their mean subtracted. The solid lines represent the results for all firms, the dashed lines those for large firms, and the dotted lines those for small firms. A high value represents an upward shift in the adjustment hazard. We focus on the shift in the hazard itself as an index of flexibility rather than on the average speed of adjustment, because in the realistic increasing hazard context the latter depends on the endogenous path of the cross section. When the hazard is constant, its shift also represents an equal shift in the speed of adjustment. When the hazard is increasing, on the other hand, the mapping from a vertical shift in the hazard to a change in the average speed of adjustment is not one-for-one, since the interactions with the cross sectional distribution of gaps complicates the mapping.

Column 1 in Table 8 and the continuous line in the upper panel of Figure 1 show the results for the constant hazard case. Under this assumption, the index of flexibility exhibited fluctuations in the second half of the 1980s and early 1990s, eventually settled at a fairly high value in the mid 90s, but then declined sharply during the 1997-99 period. From an average adjustment coefficient of 0.77 for the three years prior to the Asian/Russian crisis episode, this coefficient fell to 0.69 in the aftermath of the crisis.

Note also that in this case the decline in flexibility appears to be subsiding toward the end of the sample. However columns 4 and 7 in Table 8, and the continuous lines in the middle and lower panels of Figure 1, show that this finding is lost and there is no evidence of recovery once the hazard is allowed to be nonlinear. The reason for the misleading conclusion with a constant hazard is that toward the end of the sample there is a sharp rise in the share of establishments with large negative gaps (see Figure 2), to which establishments naturally react more under increasing hazards.[24] That is, the average speed of adjustment rises even if the hazard does not change, due to substantial negative gaps accumulated by a large number of establishments.

While it is too early to tell whether this decline in microeconomic flexibility we uncover is purely cyclical, or whether there is something more structural going on, there are a few interesting observations we can make at this time. We begin by noting that the remaining columns in Table 8 and series in Figure 1 show that much of the decline in flexibility is due to a decline in the flexibility of large establishments (as measured by their lagged employment).

Continuing with the characterization of the decline in microeconomic flexibility, Table 9 shows that while the speed of response to negative gaps remained fairly constant, it is the speed at which establishments adjust to shortages of labor that slowed down more dramatically.[25] This "reluctance to hire" may reflect pessimism respect to future conditions not captured in the current gap. But this is unlikely to be the only factor since otherwise we also should observe a rise in the speed of firing, which we do not. In fact, the increasing hazard nature of the adjustment hazard partly explains the asymmetry seen in the decline of the speed of adjustment with respect to positive and negative gaps. As we mentioned above, since there was a substantial number of establishments that developed large negative gaps (excess labor) during the slowdown, the increasing hazard implied that their adjustment did not slow down as much as the decline in the average speed of adjustment.

However, Table 10 illustrate that the sharpest decline in flexibility came from establishments in sectors that normally experience less restructuring, either because of smaller shocks or more technological and institutional inflexibility. Normal restructuring for high and low restructuring sectors is measured by the excess reallocation above/below median in Chile prior to 1997.[26] If it is not shocks but inflexibility that explains the ranking, then the cost of the increase in flexibility can be potentially very large, as already inflexible establishments spend significant time away from their frictionless optimum.

In conclusion, while we cannot pinpoint to a specific reason for why microeconomic flexibility declined toward the end of the 1990s, we clearly identified such a decline. Moreover, we found that the increasing nature of the hazard is important to show that the recovery in average flexibility toward 1999 does not seem to correspond to a real increase in flexibility. Instead, it simply reflects the interaction between an increasing hazard and a depressed phase of the business cycle. Flexibility declined in 1997 and remained down until the end of our sample, particularly so for large establishments. We also found that the decline in flexibility is more pronounced in sectors that

[24] Where *large negative gaps* are gaps smaller than -0.25 and *large positive gaps* are gaps larger than 0.25.

[25] Between 1994-96 and 1997-99, the latter fell from 0.86 to 0.71, while the former fell from 0.75 to 0.71.

[26] Similar results are obtained when sectors are classified according to the excess reallocation in the corresponding US sectors (a sort of instrumental variables for technological factors).

normally restructure less. If the latter is a consequence of larger adjustment costs (technological or institutional), then their relative slowdown is worrisome since the cost of reducing their restructuring further is particularly large. In the next section we turn to gauging some of the potential costs of microeconomic inflexibility.

6 Gauging the Costs of Microeconomic Inflexibility

By impairing worker movements from less to more productive units, microeconomic inflexibility 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 microeconomic inflexibilities in Latin America are significant. In particular, the impact of the decline in microeconomic flexibility in Chile following the Asian crisis accounts for a substantial fraction of the large decline in TFP-growth in Chile since 1997 (from an annual average of 3.1 percent for the preceding decade to about 0.3 percent after that). Moreover, if it were to persist, it could permanently shave off about 0.4 percent from Chile's structural rate of growth.

6.1 Model

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:

$$Y_{it} \;=\; B_{it} K_t L_{it}^{\alpha}, \tag{15}$$

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}'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 $\eta > 0$. Aggregate labor is assumed constant and set equal to one. We define *aggregate productivity*, A_t, as:

$$A_t = \int B_{it} L_{it}^{\alpha} di, \qquad (16)$$

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

$$Y_t = A_t K_t.$$

Units adjust with probability λ in every period, independent of their history and of what other units do that period.[27] The parameter that captures microeconomic flexibility is λ. Higher values of λ 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: [28]

$$g_Y = sA - \delta, \qquad (17)$$

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

Consider now what happens when microeconomic flexibility decreases from λ_0 to λ_1. Aggregate productivity decreases, reflecting slower reallocation of workers from less to more productive units. Indeed, from (16) we have that :

$$\Delta A = \int B_{it} \Delta L_{it}^{\alpha} di,$$

where ΔL_{it}^{α} denotes the difference between the value of L_{it}^{α} for the new value of λ and the value it would have had under the old λ. A tedious, but straightforward calculation relegated to Appendix B shows that:

$$\Delta A \simeq \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1}\right] \theta A_0,$$

with

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

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

Using (17) to get rid of A_0 yields our main result:

$$\Delta g_Y \simeq (g_{Y,0} + \delta)\left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1}\right]\theta, \qquad (18)$$

[27] More precisely, whether unit i adjusts at time t is determined by a Bernoulli random variable ξ_{it} with probability of success λ, where the ξ_{it}'s are independent across units and over time.

[28] Here we use that $g_A = 0$, since we assumed $\mu_A = 0$.

where $g_{Y,0}$ denotes the growth rate of output before the change in λ.

We choose parameters to apply (18) as follows: The mark-up is set at 20%. Parameters $g_{Y,0}$, σ_I and σ_A are set at their average values for Chile over the 1987–96 period, namely 7.9%, 19% and 4%. We also set $\delta = 6\%$. The microeconomic flexibility parameters are set at their average values during 1994-96 and 1997-99 for large establishments,[29] since they concentrate most production. From this exercise we conclude that the reduction in flexibility has reduced structural output growth by 0.4%. This *permanent* cost is due to the effect of reduced productivity on capital accumulation. One must add to this the initial direct effect of a decline in productivity on output growth,[30] which amounts to 2.7 percent. The sum of these two *structural* costs is very relevant. As mentioned earlier, it can account for a significant share of the decline in Chilean TFP growth from an annual average of 3.1 percent during the decade preceding the Asian crisis to 0.3 during the 1997-99 period.

Going back to the average results presented in Section 3, Table 11 reports the potential gain in structural growth that each country could obtain from raising microeconomic flexibility to US levels. Our estimates indicate that, on the low end, Chile and Colombia would have an initial gain in the range between 2 and 4% and a permanent increase in the structural rate of growth of approximately 0.3%. On the high end, Venezuela would see an initial gain of 22.2%, even the impact on its growth rate is less pronounced, due to it having had the lowest growth rate in our sample. By contrast, Mexico could expect an initial gain of 7.4% and an impressive permanent rise of growth of 0.7%, while the corresponding percentages for Brazil are 5.0 and 0.43. These numbers are large. We are fully aware of the many caveats that such ceteris-paribus comparison can raise, but the point of the table is to provide an alternative metric of the potential significance of observed levels of inflexibility in our region.

7 Concluding Remarks

There is the nagging feeling among policymakers and observers that the microeconomic structure of the Latin American economies is rather inflexible, and that this is a significant obstacle to growth. Not surprisingly, pro-flexibility structural reforms are high in most of the countries in the region.

[29] Equal to 0.688 and 0.892, respectively, see Table 8.

[30] This is equal to:

$$\frac{\Delta A}{A_0} \simeq \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1}\right]\theta.$$

Despite this widespread belief, there is very little in terms of formal and systematic evidence, both on the extent of inflexibility and on its costs. The data and methodological obstacles to produce this evidence are significant.

In this paper we collect extensive data sets for several Latin American countries. We then develop a methodology suitable to extract an answer to the inflexibility questions from these data sets.

Our estimates confirm the above fears. Microeconomic inflexibility is significant and very costly in our region. Moreover, in Chile, where we could measure the time path of flexibility with some precision, the trend does not seem to be pointing in the right direction. Our initial estimates suggest that the decline in flexibility observed at the end of the 1990s, if it were to persist, could shave off near half of a percent from Chile's potential growth rate.

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APPENDIX

A Estimating λ

Our starting point is (1) in the main text, where for simplicity we ignore sectors and time-variation in the target's drift:

$$\Delta e_{i,t} = \psi_{i,t}(e^*_{it} - e_{i,t-1}), \tag{19}$$

with $\psi_{i,t}$: i.i.d., with mean λ and variance $\alpha\lambda(1-\lambda); \alpha \in [0,1]$. We denote by z_{it} the gap *after* period t adjustments; that is, $z_{i,t} \equiv e^*_{it} - e_{it}$. We assume

$$\Delta e^*_{i,t} = \Delta e^*_{A,t} + \varepsilon_{i,t},$$

with $\Delta e^*_{A,t}$ i.i.d. with mean μ_A and variance σ^2_A and $\varepsilon_{i,t}$ i.i.d. independent from the $\Delta e^*_{i,t}$"s, with zero mean and variance σ^2_I.

Given an integer $M = 2, 3, \dots$ we define:

$$z^M_{i,t} = \frac{1}{M}\sum_{k=1}^{M} z_{i,t-k}. \tag{20}$$

The central idea is that with plant-specific fixed effects (e.g., systematic differences in labor force composition) we do not observe the z's implicit on the r.h.s. of (19), but only observe the difference $z_{i,t} - z^M_{i,t}$ (since the fixed effects cancel out once we subtract z^M). We therefore fix t and estimate (19) with $z - z^M$ on the r.h.s. instead of z. One advantage of this approach is that the estimated values of λ_t do not vary with the length of the time period considered, as is the case when estimating fixed effect using the time-average over the whole sample.

Denote $\sigma^2_t \equiv \mathrm{Var}[z_{i,t}]$, where the variance is calculated over i, keeping t fixed. Also denote by $\hat{\lambda}_t$ the OLS estimator of λ_t, again keeping t fixed and regressing over i. A calculation from first principles then shows that for $M = 2$ we have:

$$\mathrm{E}[\hat{\lambda}_t] = \lambda_t \left\{ 1 + \frac{\sigma^2_{t-1} - \sigma^2_{t-2}}{4\mathrm{Var}(z_{i,t} - z^M_{i,t} + \Delta l_{i,t})} \right\}, \tag{21}$$

with

$$\sigma^2_t = \frac{1-\lambda_t}{\lambda_t[\alpha + (1-\alpha)\lambda_t]} \left\{ [1-(1-\alpha)\lambda_t]\mathrm{Var}(\Delta e_{i,t}) + \frac{\alpha(2\lambda_t - 1)}{\lambda_t}(\Delta e_{A,t})^2 \right\}, \tag{22}$$

where $\Delta e_{A,t}$ denotes the average (over i) of $\Delta e_{i,t}$.

It follows from (21) that, the time average of the estimates for λ_t will be unbiased, since on average σ^2_{t-1} is equal to σ^2_{t-2}. Of course, for any particular t, the estimator may be biased. Yet the expression in (22) can be used to correct the bias in (21), since it expresses the bias in terms of observables. We calculated the actual bias for the Chilean data and it is rather small, for all periods.

Expressions analogous to (21) can be obtained for values of M larger than 2 and, surprisingly, the "average unbiasedness result" described above holds only for $M=2$.[31] An additional advantage of the $M=2$ case is that, if the fixed effect changes slowly over time, then the added precision associated with larger values of M comes at the expense of a larger bias due to time-varying fixed-effects. In this sense, $M=2$ provides a good compromise.

B Gauging the Costs

Here we show that, for the model in Section 6:

$$\frac{\Delta A}{A_0} \simeq \left[\frac{1}{\lambda_0}-\frac{1}{\lambda_1}\right]\theta, \tag{23}$$

with

$$\theta = \frac{\alpha\gamma(2-\alpha\gamma)}{2(1-\alpha\gamma)^2}(\sigma_I^2+\sigma_A^2), \tag{24}$$

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

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 $Y_{i,t}=A_{i,t}L_{i,t}^{\alpha}$,[32] 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:

$$\Delta l_{i,t}^* = \frac{\gamma}{1-\alpha\gamma}\Delta a_{i,t}, \tag{25}$$

and hence

$$\Delta y_{i,t}^* = \frac{1}{1-\alpha\gamma}\Delta a_{i,t}. \tag{26}$$

Denote by Y_t^* aggregate production in period t if there were no frictions. It then follows from (26)

[31] Of course, as M tends to infinity the estimator is (asymptotically) unbiased, without the need of averaging over time.

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

that:

$$Y_{i,t}^* = e^{\tau \Delta a_{i,t}} Y_{i,t-1}^*, \tag{27}$$

with $\tau \equiv 1/(1-\alpha\gamma)$, Taking expectations (over i for a particular realization of v_t^A) on both sides of (27) and noting that both terms being multiplied on the r.h.s. are, by assumption, independent (random walk), yields

$$Y_t^* = e^{\tau v_t^A + \frac{1}{2}\tau^2\sigma_I^2} Y_{t-1}^*, \tag{28}$$

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

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

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

$$Y_t^* = e^{\frac{1}{2}k\tau^2\sigma_T^2} Y_{t-k}^*. \tag{29}$$

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 to 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 (29) 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 (29) gives:

$$Y_{t,t-k} = e^{-k\theta} Y_t^*, \tag{30}$$

with θ defined in (24).

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

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

Substituting (30) in the expression above yields:

$$Y_t = \frac{\lambda}{1-(1-\lambda)e^{-\theta}} Y_t^*. \tag{31}$$

It follows that the production gap, defined as:

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

is equal to:

$$\text{Prod. Gap} = \frac{(1-\lambda)(1-e^{-\theta})}{1-(1-\lambda)e^{-\theta}}. \tag{32}$$

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

$$\text{Prod. Gap} \simeq \frac{(1-\lambda)}{\lambda}\theta. \tag{33}$$

Subtracting this gap evaluated at λ_0 from its value evaluated at λ_1, and noting that this gap difference corresponds to $\Delta A/A_0$ in the main text, yields (23) and therefore concludes the proof.

Table 1: ESTIMATING ϕ

COUNTRY:	Colombia	Chile	Mexico	Venezuela
$\hat{\phi}$ with extreme values:	0.414	0.460	0.372	0.336
	(0.035)	(0.028)	(0.033)	(0.108)
$\hat{\phi}$ without extreme values:	0.394	0.495	0.365	0.317
	(0.035)	(0.037)	(0.037)	(0.118)
Observations:	20,268/20,065	21,149/20,938	27,752/27,475	2,906/2,877

Robust standard errors in parenthesis.

Table 2: DESCRIPTIVE STATISTICS I

COUNTRY:	Brazil	Colombia	Chile	Mexico	Venezuela
Observations:	42,525	27,440	24,450	37,384	4,950
Establishments:	8,505	3,430	1,630	4,673	990
Employment (% obs.):					
(0 , 50):	15.9	45.1	56.7	21.0	9.9
[50 , 100):	28.5	22.8	17.9	21.4	31.5
[100 , 250):	28.9	19.5	15.4	29.4	33.7
$\geq$ 250:	26.6	12.7	9.9	28.2	24.9
Period:	1996-2000	1992-1999	1985-1999	1993-2000	1995-1999

'Employment' reports the percentage of observations with employment below 50, between 50 and 100, between 100 and 250, and larger than 250. Only continuous plants are considered.

Table 3: DESCRIPTIVE STATISTICS II

COUNTRY:	Brazil	Colombia	Chile	Mexico	Venezuela
Employment:	2,555,035	461,441	169,813	1,214,776	233,746
Net Change:	−0.024	−0.013	0.021	0.018	−0.023
Job Creation:	0.074	0.072	0.080	0.071	0.069
Job Destruction:	0.098	0.086	0.059	0.053	0.091
Reallocation:	0.173	0.158	0.139	0.123	0.160
Excess Reallocation:	0.135	0.124	0.099	0.086	0.125
Period:	1997-2000	1993-1999	1986-1999	1994-2000	1996-1999

Quantities reported are yearly averages over the sample period. Defition of all variables follows Davis et al. (1996).

Table 4: AVERAGE FLEXIBILITY ESTIMATES

COUNTRY:	Brazil	Colombia	Chile	Mexico	Venezuela
Gap:	0.701	0.722	0.724	0.581	0.539
	(0.004)	(0.005)	(0.005)	(0.004)	(0.014)
R-squared:	0.50	0.53	0.50	0.47	0.37
Observations:	25,260	20,375	20,979	27,757	2,941
Period:	1998-2000	1995-1999	1988-1999	1995-2000	1997-1999

Robust standard errors in parenthesis. All estimates in this table are significant at the 1% level. All regressions have year dummies. All estimates based on one regression per country, using all available observations. Observations corresponding to extreme values (0.5% in right tail and 0.5% in left tail) of regressors excluded.

Table 5: COMPARING FLEXIBILITY MEASURES

COUNTRY:	Brazil	Colombia	Chile	Mexico	Venezuela
Job Security Index (Heckman and Pages, 2000):	3.04	3.79	3.38	3.16	4.54
Job Security Index (Botero et al., 2003):	0.69	0.31	0.62	0.71	0.64
Excess Reallocation:	0.135	0.124	0.099	0.086	0.125
Microeconomic flexibility index (this paper):	0.701	0.722	0.724	0.581	0.539

Flexibility is decreasing in the index for the first two measures, and increasing for the remaining two measures. Since yearly values for 1990–1999 are available for the Heckman-Pages index (this is not the case for the remaining indices), the numbers reported for this index are the average over the sample period (years before 1990 are proxied by the 1990 value, and years after 1999 by the 1999 value).

Table 6: AVERAGE FLEXIBILITY ESTIMATES BY PLANT SIZE

		COUNTRY				
	Plant Size	Brazil	Colombia	Chile	Mexico	Venezuela
Gap:	Small	0.670	0.675	0.685	0.561	0.529
		(0.006)	(0.007)	(0.007)	(0.006)	(0.020)
	Large	0.808	0.790	0.783	0.607	0.529
		(0.009)	(0.010)	(0.010)	(0.007)	(0.026)
R^2:	Small	0.47	0.52	0.49	0.44	0.35
	Large	0.57	0.56	0.54	0.53	0.39
Obs.:	Small	12,560	10,087	10,404	13,784	1,469
	Large	6,340	5,131	5,265	7,008	741
Period:		1998-2000	1995-99	1988-99	1995-2000	1997-99

Small: below 50th percentile of the lagged employment distribution. Large: above the 75th percentile of the lagged employment distribution. Robust standard errors in parenthesis. All estimates in this table are significant at the 1% level. All regressions have year dummies. Observations corresponding to extreme values (0.5% in right tail and 0.5% in left tail) of regressor excluded.

Table 7: AVERAGE FLEXIBILITY ESTIMATES BY PLANT SIZE AND GAP SIZE

			COUNTRY				
			Brazil	Colombia	Chile	Mexico	Venezuela
	Plant Size	Gap Size					
Gap:	Small	Small	0.473	0.440	0.499	0.330	0.275
			(0.010)	(0.010)	(0.009)	(0.009)	(0.033)
		Large	0.722	0.752	0.790	0.626	0.570
			(0.013)	(0.012)	(0.016)	(0.010)	(0.031)
	Large	Small	0.541	0.551	0.513	0.418	0.222
			(0.011)	(0.014)	(0.013)	(0.010)	(0.044)
		Large	0.870	0.890	0.927	0.682	0.540
			(0.018)	(0.020)	(0.023)	(0.015)	(0.040)
R^2:	Small	Small	0.21	0.22	0.27	0.14	0.08
		Large	0.56	0.65	0.65	0.57	0.41
	Large	Small	0.28	0.29	0.29	0.26	0.06
		Large	0.64	0.65	0.68	0.68	0.40
Obs.:	Small	Small	9,204	7,493	8,844	9,812	886
		Large	3,356	2,594	1,560	3,972	583
	Large	Small	4,903	4,052	4,342	5,729	441
		Large	1,437	1,079	923	1,279	300
Period			1998-2000	1995-99	1988-99	1995-2000	1997-99

Plant size can be small (below 50th percentile of the lagged employment distribution) or large (above the 75th percentile of the lagged employment distribution). Gap size can be small (absolute value less than 0.25) or large (absolute value larger than 0.26). Robust standard errors in parenthesis. All estimates in this table are significant at the 1% level. All regressions have year dummies. Observations corresponding to extreme values (0.5% in right tail and 0.5% in left tail) of regressors excluded.

Table 8: EVOLUTION OF FLEXIBILITY: CHILE 1987–99

	1	2	3	4	5	6	7	8	9
	Constant hazard				Increasing (and asymmetric) hazard				
Plant size:	all	small	large	all	small	large	all	small	large
Gap 87:	0.745	0.742	0.782	0.490	0.514	0.537	0.343	0.384	0.365
	(0.030)	(0.036)	(0.068)	(0.030)	(0.038)	(0.064)	(0.030)	(0.039)	(0.063)
Gap 88:	0.674	0.707	0.716	0.424	0.481	0.445	0.272	0.344	0.270
	(0.031)	(0.041)	(0.059)	(0.031)	(0.040)	(0.058)	(0.031)	(0.040)	(0.060)
Gap 89:	0.776	0.714	0.854	0.533	0.504	0.564	0.381	0.377	0.381
	(0.038)	(0.042)	(0.054)	(0.034)	(0.043)	(0.054)	(0.035)	(0.043)	(0.055)
Gap 90:	0.677	0.656	0.765	0.441	0.478	0.488	0.274	0.326	0.289
	(0.031)	(0.039)	(0.072)	(0.030)	(0.039)	(0.068)	(0.032)	(0.041)	(0.072)
Gap 91:	0.731	0.688	0.806	0.501	0.503	0.578	0.335	0.362	0.374
	(0.033)	(0.053)	(0.058)	(0.032)	(0.050)	(0.055)	(0.034)	(0.051)	(0.058)
Gap 92:	0.740	0.705	0.758	0.520	0.522	0.503	0.359	0.380	0.302
	(0.039)	(0.063)	(0.065)	(0.036)	(0.058)	(0.063)	(0.038)	(0.062)	(0.064)
Gap 93:	0.706	0.640	0.812	0.492	0.474	0.547	0.322	0.327	0.347
	(0.034)	(0.047)	(0.066)	(0.032)	(0.046)	(0.060)	(0.033)	(0.047)	(0.065)
Gap 94:	0.730	0.656	0.913	0.515	0.487	0.639	0.345	0.339	0.443
	(0.036)	(0.050)	(0.071)	(0.035)	(0.049)	(0.066)	(0.036)	(0.050)	(0.070)
Gap 95:	0.775	0.743	0.907	0.547	0.569	0.641	0.370	0.415	0.434
	(0.034)	(0.048)	(0.072)	(0.032)	(0.044)	(0.065)	(0.033)	(0.046)	(0.069)
Gap 96:	0.808	0.706	0.856	0.577	0.531	0.582	0.402	0.378	0.386
	(0.035)	(0.055)	(0.059)	(0.034)	(0.054)	(0.056)	(0.035)	(0.055)	(0.059)
Gap 97:	0.686	0.648	0.667	0.469	0.495	0.395	0.301	0.346	0.206
	(0.033)	(0.043)	(0.073)	(0.032)	(0.042)	(0.072)	(0.034)	(0.046)	(0.074)
Gap 98:	0.669	0.614	0.667	0.425	0.446	0.377	0.242	0.285	0.168
	(0.040)	(0.051)	(0.095)	(0.038)	(0.051)	(0.091)	(0.040)	(0.052)	(0.092)
Gap 99:	0.705	0.655	0.712	0.418	0.455	0.367	0.250	0.309	0.172
	(0.034)	(0.045)	(0.076)	(0.035)	(0.048)	(0.075)	(0.038)	(0.050)	(0.080)
Gap($\|Gap\| > .25$):				0.371	0.295	0.407	0.479	0.410	0.508
				(0.016)	(0.023)	(0.031)	(0.016)	(0.023)	(0.032)
Gap(Gap $< -.05$):							−0.095	−0.172	−0.012
							(0.031)	(0.420)	(0.062)
$\|Gap\| > .25$:				0.002	0.027	−0.023	0.004	0.019	−0.012
				(0.004)	(0.006)	(0.009)	(0.005)	(0.007)	(0.010)
Gap $< -.05$:							−0.093	−0.097	−0.087
							(0.003)	(0.004)	(0.007)
R^2:	0.50	0.49	0.54	0.53	0.51	0.57	0.55	0.54	0.59

Plant size can be small (below 50th percentile of the lagged employment distribution) or large (above the 75th percentile of the lagged employment distribution). Robust standard errors in parenthesis. All regressions have year dummies. Observations corresponding to extreme values (0.5% in right tail and 0.5% in left tail) of regressors excluded.

Table 9: Evolution of Flexibility and Asymmetric Hazards

	Gap		$(\text{Gap} < -.05)$		
Year	Coeff.	St. Error	Coeff.	St. Error	No. Obs.
1987	0.689	0.030	0.227	0.062	1300
1988	0.720	0.030	−0.079	0.058	1216
1989	0.729	0.033	0.155	0.061	1248
1990	0.702	0.036	0.016	0.060	1155
1991	0.815	0.036	−0.097	0.061	1153
1992	0.752	0.035	0.061	0.067	1151
1993	0.721	0.037	0.034	0.064	1124
1994	0.831	0.039	−0.135	0.066	1073
1995	0.891	0.036	−0.152	0.060	1134
1996	0.859	0.039	−0.040	0.063	1139
1997	0.710	0.039	0.028	0.062	1146
1998	0.734	0.046	−0.078	0.069	1144
1999	0.698	0.052	0.031	0.070	1252
Simple Average:	0.758		−0.002		

Table 10: Evolution of Flexibility and Ex-ante Restructuring

	High Restructuring			Low Restructuring		
Year	Coeff.	St. Error	No. Obs.	Coeff.	St. Error	No. Obs.
1987:	0.745	0.024	902	0.749	0.030	709
1988:	0.750	0.023	898	0.552	0.029	712
1989:	0.824	0.023	904	0.698	0.031	705
1990:	0.704	0.025	911	0.640	0.026	706
1991:	0.722	0.023	902	0.748	0.030	710
1992:	0.722	0.025	908	0.768	0.031	709
1993:	0.786	0.024	909	0.575	0.027	713
1994:	0.767	0.025	913	0.689	0.029	711
1995:	0.765	0.023	904	0.788	0.030	717
1996:	0.824	0.024	906	0.788	0.029	705
1997:	0.722	0.026	912	0.634	0.027	702
1998:	0.723	0.026	911	0.580	0.029	705
1999:	0.733	0.027	895	0.664	0.029	700
Simple Average:	0.753			0.682		

Table 11: GAINS FROM ACQUIRING US-TYPE FLEXIBILITY

COUNTRY:	Brazil	Colombia	Chile	Mexico	Venezuela
σ_I (%):	27. 6	25.8	19.3	24.1	38.1
$g_{Y,0}$ (%):	2.7	2.7	6.6	3.5	2.0
Additional Growth Upon Impact (%):	5.0	3.8	2.1	7.4	22.2
Increase in Growth Rate (%):	0.43	0.33	0.27	0.70	0.18

Table 12: FLEXIBILITY ESTIMATES BASED ON (2)

COUNTRY:	Brazil	Chile	Mexico	Venezuela
Gap:	0.855	0.675	0.592	0.401
	(0.048)	(0.034)	(0.037)	(0.184)
Observations:	8,322	17,631	18,368	968
Period:	1998-2000	1988-1999	1995-2000	1997-1999

Robust standard errors in parenthesis. The dependent variable is the change in the gap (after adjustments). Second and third lag are used as instruments. All estimates in this table are significant at the 1% level, with the exception of Venezuela, which is significant at the 5% level. All estimates based on one regression per country, using all available observations. Colombia was not included because we did not have access to the data. All regressions that consider more than one year (Chile and Mexico) use year dummies. Observations corresponding to extreme values (0.5% in right tail and 0.5% in left tail) of regressors excluded.

Figure 1:

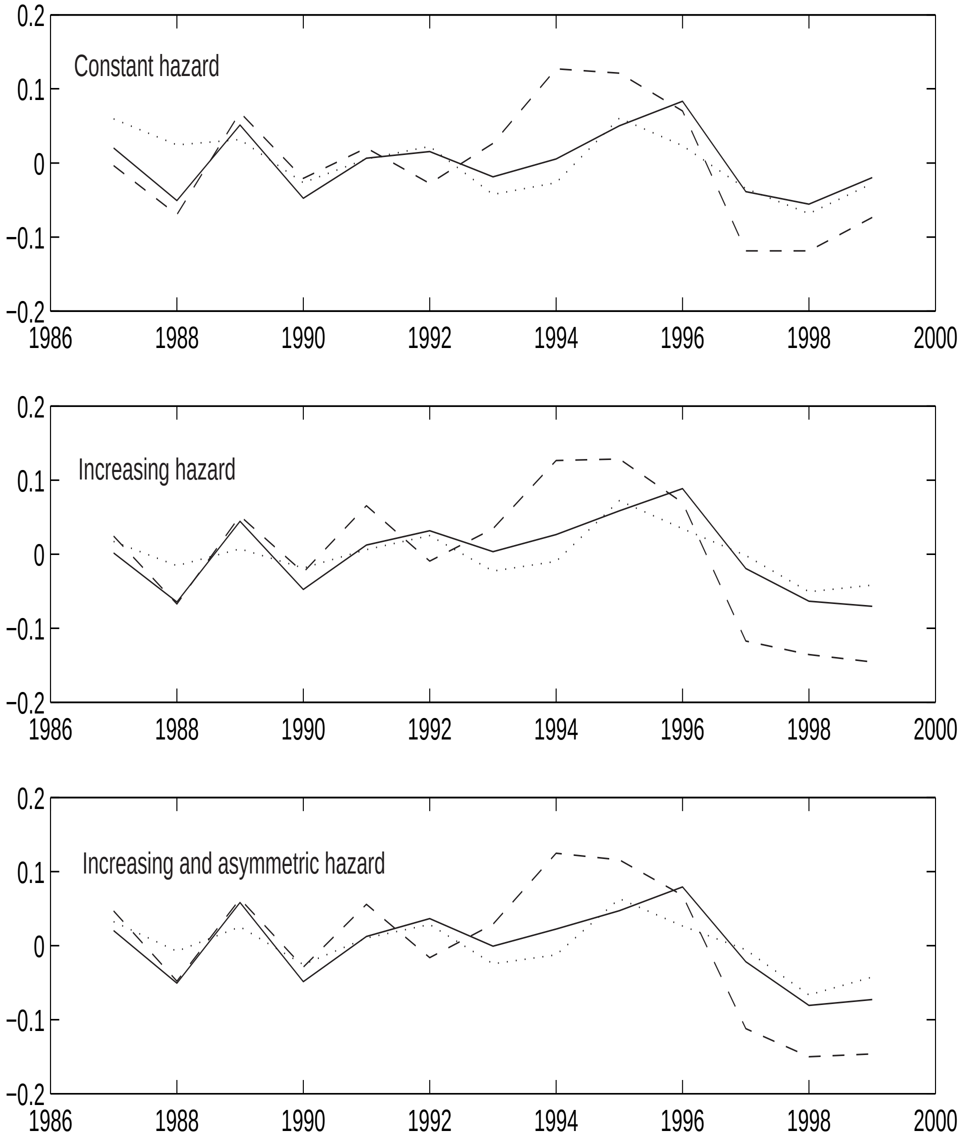
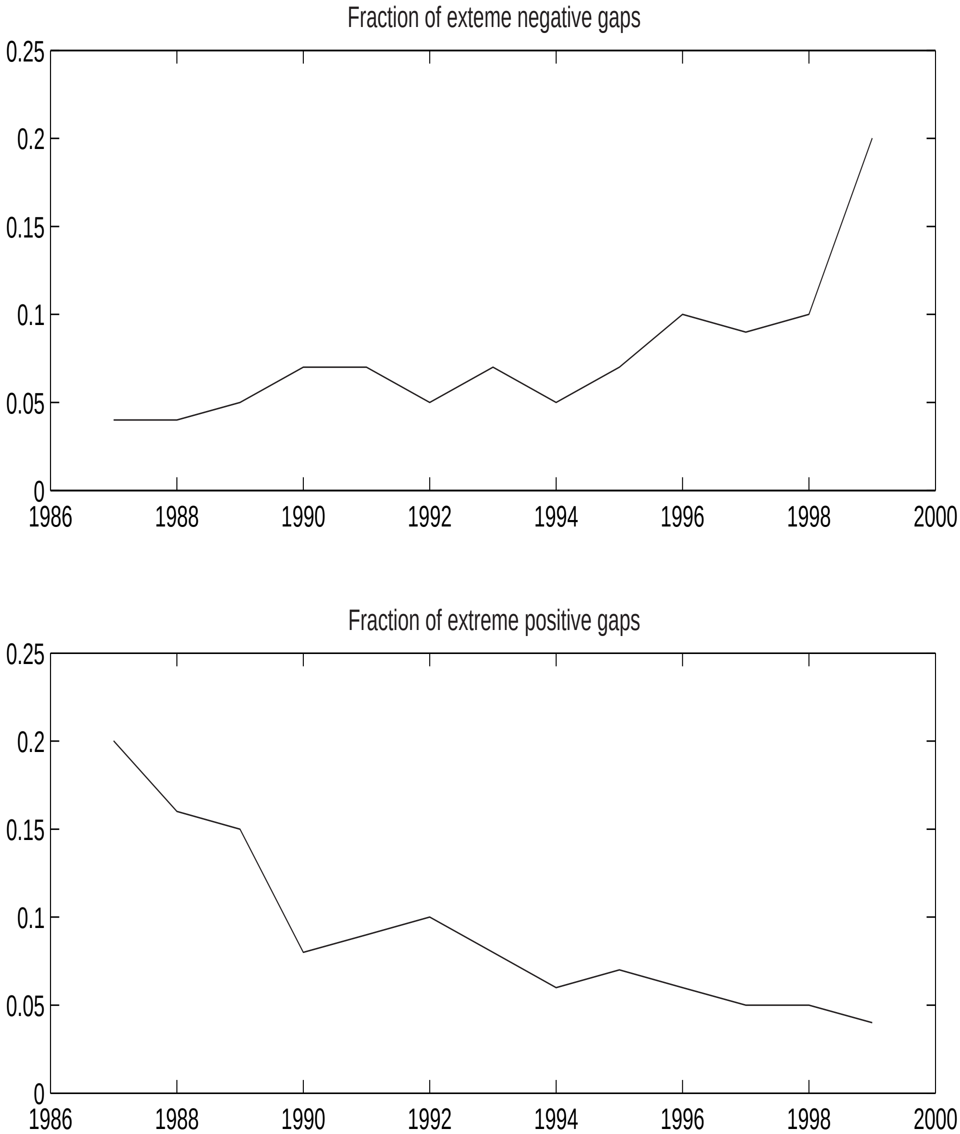

Constant hazard
Increasing hazard
Increasing and asymmetric hazard
0.2
0.1
0
−0.1
−0.2
1986
1988
1990
1992
1994
1996
1998
2000

Figure 2:


Fraction of exteme negative gaps
0.25
0.2
0.15
0.1
0.05
0
1986
1988
1990
1992
1994
1996
1998
2000
Fraction of extreme positive gaps
0.25
0.2
0.15
0.1
0.05
0
1986
1988
1990
1992
1994
1996
1998
2000