Microeconomic Flexibility in Latin America

Ricardo J. Caballero
Massachusetts Institute of Technology

Eduardo M.R.A. Engel
Yale University

Alejandro Micco
Inter-American Development Bank

Latin American economies have begun to leave behind some of the most primitive sources of macroeconomic fluctuations. Policy concern is gradually shifting toward increasing microeconomic flexibility. This is a welcome trend since microeconomic flexibility, which facilitates the ongoing process of creative-destruction, 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. One way of doing this is to look directly at regulation, which is perhaps the main institutional factor hindering or facilitating microeconomic flexibility. Extensive studies examine labor market regulation, in particular. Heckman and Pagés (2000), for example, document that “even after a decade of substantial deregulation, 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 technological options, the nature of the production process, the political environment, the efficiency and

We thank Fernando Coloma for comments. We are particularly grateful to Ernesto López-Córdoba, Marcela Meléndez, and Mauricio Moreira for their help running the econometric exercises for the cases of Mexico, Colombia, and Brazil.

Labor Markets and Institutions, edited by Jorge E. Restrepo and Andrea Tokman R., Santiago, Chile.

biases of labor courts, cultural variables, and accepted practices. While studies of rules and regulation are useful for eventual policy formulation, they 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 directly at outcomes: how much factor reallocation occurs in different countries and episodes? This is also a useful exercise, but 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. It is always difficult to determine whether the observed reallocation is abnormally high or low, because the counterfactual is not part of the statistic.

A third approach, which remedies some of the main weaknesses of the previous two, 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. An economy is said to be 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 as a result of prompt adjustment. This is the approach we follow here, while extending a methodology developed in Caballero and others (2004). The main advantage of this methodology over conventional partial adjustment estimates is its ability to use limited information efficiently by correcting standard biases that are 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, this provides us with about 140,000 observations.

In the first part of the paper, we document the main features of adjustment for these economies. Our findings include the following:

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

— While more inflexible than the United States, the economies of Brazil, Chile, and Colombia exhibit a relatively high degree of microeconomic flexibility, with over 70 percent of labor adjustment taking place within a year, on average (over time). 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.
— In all the economies in our sample except Venezuela, small establishments (with fewer than the median number of employees) are substantially less flexible than large establishments (above the seventy-fifth 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 Chile, the figures are 69 percent and 78 percent, respectively; in Colombia, 68 percent and 79 percent; in Mexico, 56 percent and 61 percent; and in Venezuela, 53 percent for both.
— It follows from the previous finding that the behavior of large establishments is primarily 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 are their counterparts in Chile, Colombia, and Brazil.
— All these economies present 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.
— Increasing hazard is particularly pronounced in large establishments in the relatively more flexible economies. In fact, most of the additional flexibility experienced by large establishments in these economies is due to their rapid adjustment when gaps become large. For example, when gaps are below 25 percent in Chile, small establishments have an adjustment coefficient of 0.50, while large establishments have a coefficient of 0.51. For deviations above 25 percent, the coefficient is 0.79 for small establishments and 0.93 for large establishments. The patterns are similar in Brazil and Colombia, yet less pronounced in Mexico and Venezuela.

The second part of the paper focuses on Chile, which has the only long panel in our sample. We are thus able to explore the evolution of the economy's microeconomic flexibility over time. Our main findings are threefold:

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

— When the adjustment hazard is assumed to be constant, the decline in flexibility appears to subside toward the end of the sample. This finding is lost, however, when the hazard is allowed to be increasing, with no evidence of recovery. 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 strongly under increasing hazards.

— While it is too early to tell whether the decline we uncover is purely cyclical or whether it reflects a structural change, we can make a few interesting observations. First, much of the decline in flexibility is due to a decline in the flexibility of large establishments. Second, the speed of response to negative gaps remained fairly constant, while the speed at which establishments adjust to labor shortages slowed dramatically. This "reluctance to hire" may reflect pessimism regarding future conditions not captured in the contemporaneous gap. This is unlikely to be the only factor, however, because we do not observe a rise in the speed of firing (for a given hazard).[2] Finally, the sharpest decline in flexibility came from establishments in sectors that normally experience less restructuring, either because they experience smaller shocks than other sectors or because they are characterized by higher technological and institutional inflexibility. If either form of inflexibility is responsible for reduced restructuring, then the cost of the decline in flexibility can potentially be very large, since inflexible establishments spend a significant amount of time away from their frictionless optimum.

The last part of the paper explores 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

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

aggregate output and slows economic growth. We develop a simple framework for quantifying this effect. Our findings suggest that the aggregate consequences of microeconomic 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 total factor productivity (TFP) growth in Chile since 1997, which fell from an annual average of 3.1 percent for the preceding decade to about 0.3 percent after that date. Moreover, if the decline were to persist, it could permanently shave off almost half a percent from Chile's structural growth rate.

Section 1 presents the methodology, while section 2 describes the data. Section 3 characterizes average microeconomic flexibility in the Latin American economies in our data. Section 4 explores the case of Chile in more detail and describes the evolution of its index of flexibility. Section 5 presents a simple model with which to map microeconomic inflexibility into growth outcomes. Section 6 concludes and is followed by two appendixes.

1. Methodology and Data

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

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

where e_{ijt} and e^*_{ijt} denote the logarithm of employment and desired employment, respectively. The random variable ψ_{ijt} is assumed to be independent and identically distributed (i.i.d.) both across establishments and over time; it 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 $\alpha = 1$ represents 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) 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 below. 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 equation (1), based on the large cross-sectional size of our sample and the well-documented fact that there are significant idiosyncratic components in the realizations of the gap and ψ_{ijt}.

An important aspect of our methodology is to find an efficient method of removing 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]

$$Gap_{ijt} = (1-\lambda)\Delta e^*_{ijt} + (1-\lambda)Gap_{ij,t-1} + \varepsilon_{ijt} \ . \tag{2}$$

We report results for this specification after presenting our main results. As we will show, the results are consistent with the estimates we obtain based on equation (1) and therefore provide a useful robustness check, although they are considerably less precise. Our methodology may thus be viewed as an alternative, for the particular problem at hand, that uses data more efficiently than standard dynamic panel estimation techniques. Finally, our methodology can be adapted to the case in which flexibility evolves over time (see section 4), which is not the case with standard panel techniques.[5]

1.1 Details

Output and demand faced by an establishment are given by

$$y = \mathrm{a} + \alpha e + \beta \ \mathrm{h} \tag{3}$$

3. As documented, for example, in Arellano and Bond (1991).
4. The gap below could be either before or after adjustments take place.
5. Panel techniques along the lines of Arellano and Bond (1991) cannot be extended to the case in which the economy wide average λ varies over time, since the instruments used in these procedures are no longer valid.

and

$$p = d - \frac{1}{\eta} y \tag{4}$$

where y, p, e, a, h, and d denote firm output, price, employment, productivity, hours worked, and demand shocks, and η is the price elasticity of demand. We let $\gamma \equiv (\eta - 1)/\eta$.[6] All variables are in logs.

Firms are competitive in the labor market but pay wages that are increasing in the average number of hours worked, according to[7]

$$w = w^0 + \mu(h - \bar{h}) \tag{5}$$

where $\bar{h}$ is constant over time and interpreted below.[8]

A key assumption is that firms only face adjustment costs when they change employment levels, not when they change the number of hours worked.[9] 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 for hours.

In a frictionless labor market, the firm's employment level also satisfies a first-order condition for employment. Our functional forms then imply that the optimal choice of hours does not depend on the employment level.[10] We denote the corresponding employment level by $\hat{e}$ and refer to it as the static employment target.[11] This leads to the following relation between the employment gap and the hours gap:

$$\hat{e} - e = \frac{\mu - \beta\gamma}{1 - \alpha\gamma}(h - \bar{h}) \, . \tag{6}$$

6. To ensure interior solutions, we assume $\eta > 1$ and $\alpha\gamma < 1$.

7. The expression below should be interpreted as a convenient approximation for $w = k^0 + \log(H^{\mu} + \Omega)$, with w^o and μ determined by k^o and Ω.

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

9. See Sargent (1978) and Shapiro (1986).

10. A patient calculation shows that

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

11. We have

$$\hat{e} = C + \frac{1}{1 - \alpha\gamma}\left[d + a - w^0 \right]$$

where $\boldsymbol{C}$ is a constant that depends on μ, α, β, and γ.

This is the expression used by Caballero and Engel (1993). It is not useful in our case, since we do not have information on hours worked. However, the argument used to derive equation (6) can also be used to express the employment gap in terms of the marginal labor productivity gap:

$$\hat{e} - e = \frac{\phi}{1 - \alpha\gamma}(v - w^0)$$

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 equation (5). This result is intuitive: the employment response to a given deviation of wages from marginal product will be larger if the marginal cost of the alternative adjustment strategy—changing hours—is higher. Also note that $\hat{e} - e$ is the difference between the static target, $\hat{e}$, and realized employment, not the dynamic employment gap, $e^*_{ijt} - e_{ijt}$, related to the term on the right-hand side of equation (1). We assume, however, that demand, productivity, and wage shocks follow a random walk.[12] Consequently, e^*_{ijt} is equal to $\hat{e}_{ijt}$ plus a constant, δ_τ.[13] It follows that

$$e^*_{ijt} - e_{ij,t-1} = \frac{\phi}{1 - \alpha\gamma_j}(v_{ijt} - w^0_{ijt}) + \Delta e_{ijt} + \delta_t \tag{7}$$

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^0_{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 worker compensation, where the wage may not represent the actual marginal cost of labor.[14] We performed our estimations

12. Given the preceding footnote, it suffices that $d + \gamma\alpha - w^o$ follows a random walk.

13. To allow for variations in future expected growth rates of a and d, the constant δ is allowed to vary exogenously over time.

14. While we have assumed a simple competitive market for the base wage within each firm (i.e., the wage for regular hours), our procedure can easily accommodate rent sharing as part of the wage-setting mechanisms (with a suitable reinterpretation of some parameters, but not λ).

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 large magnitude of the idiosyncratic shocks faced by establishments. We report the more robust alternative and approximate w^o by the average marginal productivity, which leads to

$$e^*_{ijt} - e_{ij,t-1} = \frac{\phi}{1-\alpha\gamma_j}(v_{ijt} - v_{.jt}) + \Delta e_{ijt} + \delta_t \equiv Gap_{ijt} + \delta_t \ . \tag{8}$$

The above expression 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 $(v_{ijt} - v_{.jt})$ in equation (8) a moving average of lagged relative productivity by establishment, $\hat{\theta}_{ijt}$.[15] The resulting expression for the estimated employment gap is[16]

$$e^*_{ijt} - e_{ij,t-1} = \frac{\phi}{1-\alpha\gamma_j}(v_{ijt} - \hat{\theta}_{ijt} - v_{.jt}) + \Delta e_{ijt} + \delta_t \equiv Gap_{ijt} + \delta_t \ . \tag{9}$$

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

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

where κ is a year dummy, $\Delta e_{ijt}*$ is the change in the desired level of employment, and

$$z_{ijt} = (\Delta v_{ijt} - \Delta v_{.jt}) / \left(1 - \alpha\gamma_j\right) .$$

15. Where

$$\hat{\theta}_{ijt} \equiv \frac{1}{2}\left[\left(v_{ijt-1} - v_{.jt-1}\right) + \left(v_{ijt-2} - v_{.jt-2}\right)\right].$$

The alternative specification, with relative wages instead of relative marginal productivities, leads to almost identical results.

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

By assumption, Δe^*_{ijt} is i.i.d. and independent of lagged variables. To avoid endogeneity and measurement error bias, we estimate equation (10) using $(\Delta w_{ij,t-1} - \Delta w_{.jt-1})$ as an instrument for $(\Delta v_{ijt} - \Delta v_{.jt})$.[17] Table 1 reports the estimation results of equation (10) across the countries in our sample.[18] We report estimates both with and without the one percent of extreme values for the independent variable. Based on the estimates reported in table 1, we chose a common value of ϕ equal to 0.40, to facilitate comparison across countries.

Table 1. Estimating ϕ[a]

	Country			
Parameter	*Chile*	*Colombia*	*Mexico*	*Venezuela*
ϕ (with extreme values)	0.460	0.414	0.372	0.336
	(0.028)	(0.035)	(0.033)	(0.108)
ϕ (without extreme values)	0.495	0.394	0.365	0.317
	(0.037)	(0.035)	(0.037)	(0.118)
No. observations	21,149/20,938	20,268/20,065	27,752/27,475	2,906/2,877

a. The parameter ϕ is estimated using equation 10 in the text. The regression is run both with and without the one percent of extreme values for the independent variable. Brazil is excluded from the analysis because wage data are not available. Robust standard errors in parentheses.

1.2 Summary

Our methodology has four advantages over previous specifications used to estimate cross-country differences in speed of adjustment. First, it only requires data on nominal output and the 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.[19] Second,

17. 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.

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

19. Abraham and Houseman (1994), Hamermesh (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 concerns associated with real output, it requires constructing an adequate sectoral demand shock for every country. Burgess and Knetter (1998) and Burgess, Knetter, and Michelacci (2000), for example, 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.

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, as well as to study the determinants of the speed of adjustment using interaction terms. Third, in contrast with standard panel techniques, the methodology we develop can be used when the average microeconomic flexibility is evolving over time (see section 4). Finally, our approach can be extended to incorporate nonlinearities in the adjustment function—that is, the possibility that ψ in equation (1) depends on the gap before adjustments take place. This feature also turns out to be useful.

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

$$\Delta e_{ijt} = \left(Gap_{ijt} + \delta_t\right) + \varepsilon_{ijt} \tag{11}$$

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 it can adapt to slow time variations in heterogeneity.

2. Data and Basic Facts

This section describes the data and sources used in the empirical analysis. These data are from manufacturing censuses and surveys conducted by national government statistical 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 (from the International Standard Industrial Classification, or 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. In all countries, we include only plants that existed during the full sample period (continuous plants).

For Brazil, the data are from the annual manufacturing survey *(Pesquisa Industrial Anual)* conducted by the *Instituto Brasileiro de Geografia e Estatística*. This survey started in 1967 but underwent a severe methodological change in 1996, so we only use observations from 1996 to 2000. In the case of Chile, the data are from the Chilean manufacturing census *(Encuesta Nacional Industrial Anual)* conducted by the country's *Instituto Nacional de Estadísticas*. In principle, the survey covers all manufacturing plants in Chile with more than ten employees during the period 1979–97. However, the years before 1985 are characterized by large macroeconomic shocks and structural adjustments that introduce too much noise for our methodology to handle properly. We thus use only continuous plants from the period 1985–97. For Colombia, we use the Colombian manufacturing census *(Encuesta Anual Manufacturera y Registro Industrial)* conducted by the *Departamento Administrativo Nacional de Estadística*. 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 limit the sample to continuous plants in the period 1992–99 because of a methodological change in the survey in 1992.

For Mexico, we use the annual manufacturing survey *(Encuesta Industrial Anual)* conducted by the *Instituto Nacional de Estadística, Geografía e Informática*. The survey covers a random sample of establishments in the manufacturing sector during the period 1993–2000. Finally, the data for Venezuela are from the manufacturing survey *(Encuesta Industria Manufacturera)* conducted by the country's *Instituto Nacional de Estadística*. The survey covers all plants with more than fifty employees, and it has a yearly random sample for plants with less than fifty employees. As a result of changes in the methodology, we are only 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 relevant sample period. The coverage of plants by size differs across countries. Chile and Colombia have the largest coverage of small plants (fewer than fifty employees), whereas Venezuela's survey mainly covers large establishments.

In table 3, we compute the average job creation and job destruction for each country. We also report the simple average over time of

Table 2. Descriptive Statistics: Breakdown of Establishments by Size[a]

	Country				
Statistic	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
No. observations	42,525	24,450	27,440	37,384	4,950
No. establishments	8,505	1,630	3,430	4,673	990
Observations by plant size (percent)					
Under 50 employees	15.9	56.7	45.1	21.0	9.9
50–99 employees	28.5	17.9	22.8	21.4	31.5
100–249 employees	28.9	15.4	19.5	29.4	33.7
250 employees and above	26.6	9.9	12.7	28.2	24.9
Period	1996–2000	1985–99	1992–99	1993–2000	1995–99

a. Only continuous plants are considered.

Table 3. Descriptive Statistics: Job Creation and Destruction[a]

	Country				
Statistic	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Total jobs in sample	2,555,035	169,813	461,441	1,214,776	233,746
Net change	–0.024	0.021	–0.013	0.018	–0.023
Job creation	0.074	0.080	0.072	0.071	0.069
Job destruction	0.098	0.059	0.086	0.053	0.091
Reallocation	0.173	0.139	0.158	0.123	0.160
Excess reallocation	0.135	0.099	0.124	0.086	0.125
Period	1997–2000	1986–99	1993–99	1994–2000	1996–99

a. Quantities reported are yearly averages over the sample period. Definition of all variables follows Davis, Haltiwanger, and Schuh (1996).

the net change in employment and the excess turnover (that is, the sum of job flows net of the change in employment stemming from cyclical factors). All statistics are defined following Davis, Haltiwanger, and Schuh (1996). These numbers suggest that microeconomic flexibility in these countries is limited: they are of the same order of magnitude as those of developed economies—which presumably need less restructuring than emerging economies that are still catching up—and substantially below economies such as Taiwan.[20]

20. See, for example, Caballero and Hammour (2000) and references therein.

3. 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 the following equation (we now drop the sector j subscript):

$$\Delta e_{it} = d_t + \lambda Gap_{it} + \varepsilon_{it} \quad . \tag{12}$$

The first apparent result is that microeconomic flexibility is more limited in our economies than in the very flexible United States. In the latter, estimates of λ using annual data are much closer to one.[21]

Table 4. Average Flexibility Estimates[a]

	Country				
Explanatory variable	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Gap	0.701	0.724	0.722	0.581	0.539
	(0.004)	(0.005)	(0.005)	(0.004)	(0.014)
Summary statistic					
R^2	0.50	0.50	0.53	0.47	0.37
No. observations	25,260	20,979	20,375	27,757	2,941
Period	1998–2000	1988–99	1995–99	1995–2000	1997–99

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

Although comparisons must be interpreted with caution since the samples differ in number of observations, time periods, establishments' demographics, and so forth, we can identify a discernible pattern. Within the region, Brazil, Chile, and Colombia 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 least

21. For example, Caballero, Engel, and Haltiwanger (1997) find a quarterly λ exceeding 0.4 for U.S. manufacturing, which implies an annual λ of approximately 0.90.

flexible of these economies, with slightly more than 50 percent of adjustment within a year.

Our ranking is essentially uncorrelated with the ranking obtained by Heckman and Pagés (2000) and Djankov and others (2003) based on measuring labor market regulations (see table 5). For example, the Djankov and others (2003) index of job security places Venezuela at a level of flexibility similar to that of Brazil and Chile, while Colombia is significantly more flexible than all of the above.[22] These contrasting results lend support to our earlier motivation for adopting our approach in constructing a broad measure of microeconomic inflexibility.

Table 5. Comparing Flexibility Measures[a]

	Country				
Measure	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Job security index (Heckman and Pagés, 2000)	3.04	3.38	3.79	3.16	4.54
Job security index (Djankov and others, 2003)	0.69	0.62	0.31	0.71	0.64
Excess reallocation (from table 3)	0.135	0.099	0.124	0.086	0.125
Microeconomic flexibility index (this paper)	0.701	0.724	0.722	0.581	0.539

a. Flexibility is decreasing in the index for the first two measures and increasing for the remaining two measures. Yearly values for 1990–99 are available for the Heckman-Pagés index only; 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 reports the results from repeating the estimation of regression 1, but conditioning on whether establishments are small or large. Small establishments are defined as those with a number of employees below the median in the preceding year, while large establishments are those above the seventy-fifth 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

22. According to the Heckman and Pagés (2000) index, the most flexible countries in our sample are Brazil and Mexico—not Chile and Colombia, as suggested by our index.

Table 6. Average Flexibility Estimates by Plant Size[a]

Explanatory variable and plant size	*Country*				
	Brazil	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Gap					
Small plants	0.670	0.685	0.675	0.561	0.529
	(0.006)	(0.007)	(0.007)	(0.006)	(0.020)
Large plants	0.808	0.783	0.790	0.607	0.529
	(0.009)	(0.010)	(0.010)	(0.007)	(0.026)
Summary statistic					
R^2					
Small plants	0.47	0.49	0.52	0.44	0.35
Large plants	0.57	0.54	0.56	0.53	0.39
No. observations					
Small plants	12,560	10,404	10,087	13,784	1,469
Large plants	6,340	5,265	5,131	7,008	741
Period	1998–2000	1988–99	1995–99	1995–2000	1997–99

a. Small plants are below the fiftieth percentile of the lagged employment distribution; large plants are above the seventy-fifth percentile of the lagged employment distribution. All regressions include year dummies. All estimates in this table are significant at the 1 percent level. Observations corresponding to extreme values (0.5 percent in right tail and 0.5 percent in left tail) of regressor excluded. Robust standard errors are in parentheses.

about 81 percent. In Chile, the figures are 69 and 78, respectively; in Colombia, 68 and 79; in Mexico, 56 and 61; and in Venezuela, 53 percent for both.

The behavior of large establishments primarily explains the substantial differences in flexibility across some of these economies. Again, this need not come from differences in labor market regulation, in which case it would not be captured by indices based on this variable. It could also reflect, for example, barriers to entry or social objectives assigned to large firms by the government.

Table 7 further splits the observations by the magnitude of the employment gap. Small gaps are defined as gaps of less than 25 percent in absolute value, while large gaps are above 25 percent. That is, we reestimate equation (1) for each country-size/size-of-gap combination (*jsg*):

$$\Delta e_{ijsgt} = d_{jsgt} + \lambda_{jsg} Gap_{ijsgt} + \varepsilon_{ijsgt} \quad . \tag{13}$$

Table 7. Average Flexibility Estimates by Plant Size and Gap Size[a]

Explanatory variable, plant size, and gap size	*Country*				
	Brazil	*Chile*	*Colombia*	*Mexico*	*Venezuela*
Gap					
Small plants, small gap	0.473	0.499	0.440	0.330	0.275
	(0.010)	(0.009)	(0.010)	(0.009)	(0.033)
Small plants, large gap	0.722	0.790	0.752	0.626	0.570
	(0.013)	(0.016)	(0.012)	(0.010)	(0.031)
Large plants, small gap	0.541	0.513	0.551	0.418	0.222
	(0.011)	(0.013)	(0.014)	(0.010)	(0.044)
Large plants, large gap	0.870	0.927	0.890	0.682	0.540
	(0.018)	(0.023)	(0.020)	(0.015)	(0.040)
Summary statistic					
R^2					
Small plants, small gap	0.21	0.27	0.22	0.14	0.08
Small plants, large gap	0.56	0.65	0.65	0.57	0.41
Large plants, small gap	0.28	0.29	0.29	0.26	0.06
Large plants, large gap	0.64	0.68	0.65	0.68	0.40
No. observations					
Small plants, small gap	9,204	8,844	7,493	9,812	886
Small plants, large gap	3,356	1,560	2,594	3,972	583
Large plants, small gap	4,903	4,342	4,052	5,729	441
Large plants, large gap	1,437	923	1,079	1,279	300
Period	1998–2000	1988–99	1995–99	1995–2000	1997–99

a. Small plants are below the fiftieth percentile of the lagged employment distribution; large plants are above the seventy-fifth percentile of the lagged employment distribution. A small gap has an absolute value less than or equal to 25 percent; a large gap has an absolute value larger than 25 percent. All regressions include year dummies. All estimates in this table are significant at the 1 percent level. Observations corresponding to extreme values of regressors (0.5 percent in right tail and 0.5 percent in left tail) are excluded. Robust standard errors are in parentheses.

Several significant conclusions follow from this table. First, all the economies we study show evidence of an increasing hazard.[23] In other words, establishments are substantially more flexible with respect to large gaps than to small ones. This hints 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 an institutional origin, as when dismissals require approval by a government agency or are likely to be litigated in court.

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

Second, the increasing hazard 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 is the case, since they adjust more frequently than their counterparts in inflexible economies. Rather, it means that the benefits of adjustment overcome fixed costs sooner in large establishments in flexible economies vis-à-vis inflexible economies and that the adjustment decisions of large establishments in inflexible economies are less predictable (that is, they are not correlated with the size of the gap).

Finally, 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 percent 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, the Latin American economies present evidence of microeconomic inflexibility, and in some cases, such as Mexico and Venezuela, the problem is quite severe. Studies based only on quantifying job flows are unable to detect either of these facts: gross job flows are comparable in magnitude to those in the United States and across all the economies we study, or they yield the wrong ranking (for example, 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 based solely on studying labor market regulation.[24]

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 in which 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.

24. Our remarks refer only to the measurement of a broad concept of microeconomic flexibility, not to the general merit of such studies.

Table 8. Flexibility Estimates Based on Equation (2)[a]

	Country			
Explanatory variable	*Brazil*	*Chile*	*Mexico*	*Venezuela*
Gap	0.855	0.675	0.592	0.401
	(0.048)	(0.034)	(0.037)	(0.184)
No. observations	8,322	17,631	18,368	968
Period	1998–2000	1988–99	1995-2000	1997–99

a. The dependent variable is the change in the gap after adjustments. Second and third lags are used as instruments. All estimates in this table are significant at the 1 percent level, with the exception of Venezuela, which is significant at the 5 percent level. All estimates are 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 of regressors (0.5 percent in right tail and 0.5 percent in left tail) are excluded. Robust standard errors are in parentheses.

Finally, as mentioned in section 1, our model can also be estimated with standard dynamic panel methods. Table 8 shows that our basic conclusions remain unchanged when we use this procedure, but the precision of the estimates falls significantly, as expected.

4. The Evolution of Flexibility

Has microeconomic flexibility improved over time? Unfortunately, we only have a long time dimension for the case of Chile. In this section, we focus our analysis on this case, and we conclude that the answer to this question is negative. Flexibility has declined significantly since the Asian crisis.

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

$$\begin{aligned}\Delta e_{ijt} = {} & \left[\lambda_{0jt} + \lambda_{1j}\left\{\left|Gap_{ijt}\right| > 0.25\right\} + \lambda_{2j}\left\{Gap_{ijt} < -0.05\right\}\right] Gap_{ijt} + \\ & d_{1j}\left\{\left|Gap_{ijt}\right| > 0.25\right\} + d_{2j}\left\{Gap_{ijt} < -0.05\right\} + \varepsilon_{ijsgt}\end{aligned} \tag{14}$$

where we include, but do not report, constants, time dummies, and group dummies (for example, $\left|Gap_{ijt}\right| > 0.25$). The results of these variants are reported in table 9.

Table 9. Evolution of Flexibility by Plant Size: Chile, 1987–99[a]

Explanatory variable	Constant hazard			Increasing hazard			Increasing and asymmetric hazard		
	All plants (1)	Small plants (2)	Large plants (3)	All plants (4)	Small plants (5)	Large plants (6)	All plants (7)	Small plants (8)	Large plants (9)
Gap 1987	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 1988	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 1989	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 1990	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 1991	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 1992	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 1993	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 1994	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 1995	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 1996	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 1997	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 1998	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 1999	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)

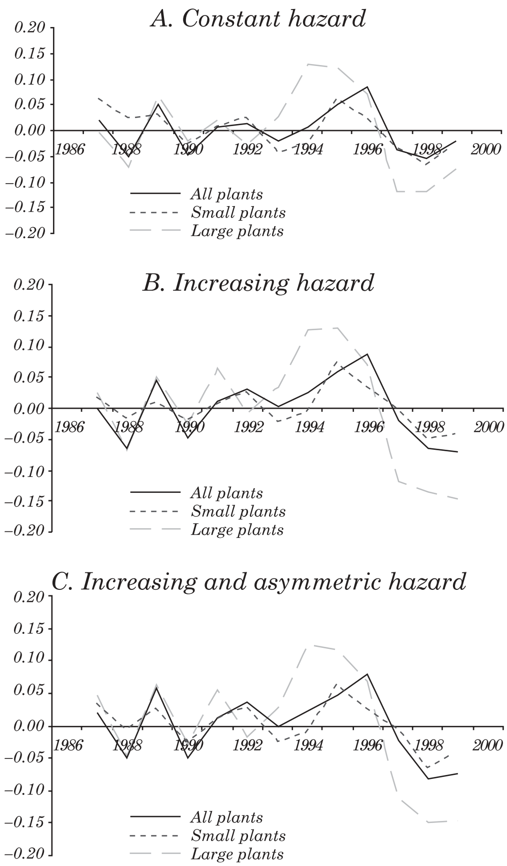
Table 9. (continued)

Explanatory variable	Constant hazard: All plants (1)	Constant hazard: Small plants (2)	Constant hazard: Large plants (3)	Increasing hazard: All plants (4)	Increasing hazard: Small plants (5)	Increasing hazard: Large plants (6)	Increasing and asymmetric hazard: All plants (7)	Increasing and asymmetric hazard: Small plants (8)	Increasing and asymmetric hazard: Large plants (9)
Gap (|Gap| > 0.25)				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 < –0.05)							–0.095	–0.172	–0.012
							(0.031)	(0.420)	(0.062)
|Gap| > 0.25				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 < –0.05							–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

a. Small plants are below the fiftieth percentile of the lagged employment distribution; large plants are above the seventy-fifth percentile of the lagged employment distribution. All regressions include year and group dummies. Observations corresponding to extreme values of regressors (0.5 percent in right tail and 0.5 percent in left tail) are excluded. Robust standard errors are in parentheses.

Figure 1 plots the path of the estimated values of l_{0jt}, with their mean subtracted. 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 context of increasing hazard, 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

Figure 1. Time-varying Adjustment Hazards



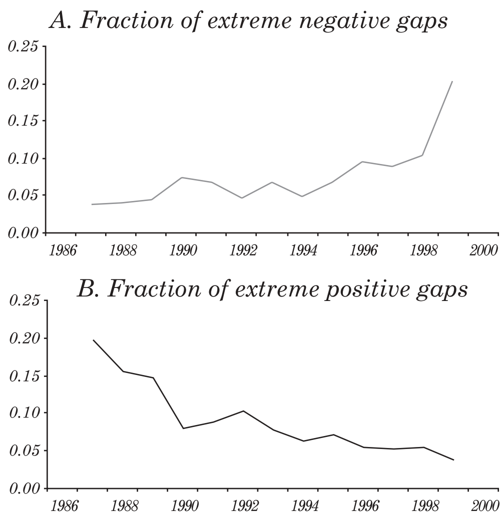
Source: Authors' calculations.

is not one for one, since the interactions with the cross-sectional distribution of gaps complicates the mapping.

The first column in table 9 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 exhibits fluctuations in the second half of the 1980s and early 1990s, settles at a fairly high value in the mid-1990s, and then declines sharply during the 1997–99 period. This coefficient fell from an average adjustment coefficient of 0.77 for the three years prior to the Asian/Russian crisis episode to 0.69 in the aftermath of the crisis.

In this case, the decline in flexibility appears to be subsiding toward the end of the sample. However, columns 4 and 7 in table 9 (corresponding to all plants under increasing hazard) and the continuous lines in the middle and lower panels of figure 1 show that this finding is lost, with no evidence of recovery, once the hazard is allowed to be nonlinear. The reason for the misleading conclusion with a constant hazard is that the share of establishments with large negative gaps rises sharply toward the end of the sample (see figure 2), and establishments naturally react more strongly to this situation

Figure 2. Fraction of Extreme Gaps



Source: Authors' calculations.

under increasing hazards.[25] That is, the average speed of adjustment rises—even if the hazard does not change—when a large number of establishments accumulate substantial negative gaps.

While it is too early to tell whether this decline in microeconomic flexibility is purely cyclical or whether it reflects some structural change, we can make a few interesting observations. First, the remaining columns in table 9 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).

Second, table 10 shows that while the speed of response to negative gaps remained fairly constant, the speed at which establishments adjust to shortages of labor slowed dramatically.[26] This reluctance to hire may reflect pessimism with respect to future conditions not captured in the current gap. This is unlikely to be the only factor, however, since we do not observe the rise in the speed of firing that should accompany it. The increasing nature of the adjustment hazard partly explains the asymmetry seen in the decline of the speed of

Table 10. Evolution of Flexibility and Asymmetric Hazards[a]

	Gap		*Gap < –0.05*		
Year	*Coefficient*	*Standard error*	*Coefficient*	*Standard error*	*No. observations*
1987	0.689	0.030	0.227	0.062	1,300
1988	0.720	0.030	–0.079	0.058	1,216
1989	0.729	0.033	0.155	0.061	1,248
1990	0.702	0.036	0.016	0.060	1,155
1991	0.815	0.036	–0.097	0.061	1,153
1992	0.752	0.035	0.061	0.067	1,151
1993	0.721	0.037	0.034	0.064	1,124
1994	0.831	0.039	–0.135	0.066	1,073
1995	0.891	0.036	–0.152	0.060	1,134
1996	0.859	0.039	–0.040	0.063	1,139
1997	0.710	0.039	0.028	0.062	1,146
1998	0.734	0.046	–0.078	0.069	1,144
1999	0.698	0.052	0.031	0.070	1,252
Simple average	0.758		–0.002		

a. Estimates for parameters in equation (14) in the main text. All regressions include year and gap size dummies. Observations corresponding to extreme values of regressors during the whole sample (0.5 percent in right tail and 0.5 percent in left tail) are excluded.

25. Large negative gaps are defined as being smaller than –0.25, and large positive gaps are gaps larger than 0.25.

26. 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.

adjustment with respect to positive and negative gaps. As we mentioned above, a substantial number of establishments developed large negative gaps (excess labor) during the slowdown; the increasing hazard implies that their adjustment did not slow as much as the decline in the average speed of adjustment.

Table 11 illustrates that the sharpest decline in flexibility came from establishments in sectors that normally experience a low degree of restructuring, either because they experience smaller shocks or because they are characterized by more technological and institutional inflexibility. Normal restructuring for sectors with a high or low degree of restructuring is measured by the excess reallocation above or below the median in Chile prior to 1997.[27] If inflexibility rather than shocks explains the ranking, then the cost of the increase in flexibility can potentially be very large, as inflexible establishments spend a significant amount of time away from their frictionless optimum.

Table 11. Evolution of Flexibility and Ex Ante Restructuring[a]

	Type of sector					
	High restructuring			*Low restructuring*		
Year	*Coefficient*	*Standard error*	*No. observations*	*Coefficient*	*Standard error*	*No. observations*
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		

a. High- and low-restructuring sectors are defined using the median sector excess reallocation in Chile prior to 1997. All regressions include year and gap size dummies. Observations corresponding to extreme values of regressors during the whole sample (0.5 percent in right tail and 0.5 percent in left tail) are excluded.

27. Similar results are obtained when sectors are classified according to the excess reallocation in the corresponding U.S. sectors (which serves as an instrumental variable for technological factors).

In conclusion, we clearly identified a decline in microeconomic flexibility toward the end of the 1990s, although we cannot pinpoint a specific reason for the decline. We also found that the increasing nature of the hazard is important for showing that the recovery in average flexibility around 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 in the case of large establishments. We also found that the decline in flexibility is more pronounced in sectors that normally have a low degree of restructuring. If the latter is a consequence of large technological or institutional adjustment costs, then their relative slowdown is worrisome since the cost of further reducing their restructuring is particularly large.

5. Gauging the Costs of Microeconomic Inflexibility

By impairing worker movements from less productive to more productive units, microeconomic inflexibility reduces aggregate output and slows 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. 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, which fell from an annual average of 3.1 percent for the preceding decade to about 0.3 percent after the crisis. If the decline were to persist, it could permanently shave about 0.4 percent off Chile's structural growth rate.

5.1 The 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 values of B_{it} follow geometric random walks, which 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's are i.i.d. $N(\mu_A, \sigma_A^2)$ and the v_{it}'s are i.i.d. (across productive units, over time, and with respect to shocks) $N(0, \sigma_I^2)$. We set $\mu_A = 0$, since we are interested in the interaction between rigidities and idiosyncratic shocks, rather than in Jensen's inequality effects associated with aggregate shocks.

The price-elasticity of demand is $\eta > 0$. Aggregate labor is assumed constant and set equal to 1. We define aggregate productivity, A_t, as

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

so that aggregate output,

$$Y_i \equiv \int Y_{it} di ,$$

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.[28] 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: [29]

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

where s denotes the savings rate (which we assume to be exogenous) and δ the capital depreciation rate.

When microeconomic flexibility decreases from λ_0 to λ_1, aggregate productivity decreases, reflecting slower reallocation of workers

28. 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 values of ξ_{it} are independent across units and over time.

29. Here we use $g_A = 0$, since we assumed $\mu_A = 0$.

from less productive to more productive units. Equation (16) implies that the reduction in aggregate productivity is given by

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

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

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

where

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

and

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

Using equation (17) to eliminate A_0 yields our main result:

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

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

We choose parameters to apply equation (18) as follows. The markup is set at 20 percent. Parameters $g_{Y,0}$, σ_P, and σ_A are set at their average values for Chile over the 1987–96 period, namely 7.9 percent, 19 percent, and 4 percent, respectively. We also set δ equal to 6 percent. The microeconomic flexibility parameters are set at their average values for large establishments in 1994–96 and 1997–99, since these firms account for most production.[30] From this exercise we conclude that the reduction in flexibility has lowered structural output growth by 0.4 percent. This permanent cost is due to the

30. The values are 0.688 in 1994–96 and 0.892 in 1997–99; see table 9.

effect of reduced productivity on capital accumulation. One must add to this the initial direct effect of a decline in productivity on output growth, which amounts to 2.7 percent.[31] 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 in the decade preceding the Asian crisis to 0.3 in the 1997-99 period.

Table 12 reports the potential gain in structural growth that each country in our sample could obtain from raising microeconomic flexibility to U.S. levels. Our estimates indicate that, on the low end, Chile and Colombia would have an initial gain in the range of 2 to 4 percent and a permanent increase in their structural growth rate of approximately 0.3 percent. On the high end, Venezuela would see an initial gain of 22.2 percent, although the impact on its growth rate is less pronounced because it had the lowest growth rate in our sample. Mexico could expect an initial gain of 7.4 percent and an impressive permanent increase in growth of 0.7 percent, 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 comparisons carry, but the table provides an alternative metric of the potential significance of observed levels of inflexibility in the region.

Table 12. Gains from Acquiring U.S. Level of Flexibility[a]
Percent

	Country				
Indicator	*Brazil*	*Chile*	*Colombia*	*Mexico*	*Venezuela*
σ_1	27.6	19.3	25.8	24.1	38.1
g_{Y0}	2.7	6.6	2.7	3.5	2.0
Additional growth on impact	5.0	2.1	3.8	7.4	22.2
Increase in growth rate	0.43	0.27	0.33	0.70	0.18

a. The volatility of idiosyncratic shocks by country is computed using equation (10) in the text and $\phi = 0.4$. Observations corresponding to extreme values of gaps (0.5 percent in right tail and 0.5 percent in left tail) are excluded.

31. This is equal to $\frac{\Delta A}{A_0} \cong \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1}\right]\theta$.

6. Concluding Remarks

Policymakers and observers seem to have a nagging feeling that the microeconomic structure of the Latin American economies is rather inflexible and that this is a significant obstacle to growth. Not surprisingly, structural reforms aimed at heightening flexibility are extensive in most of the countries in the region. Despite this widespread belief, formal and systematic evidence on the extent of inflexibility and its costs is scarce. The data and methodological obstacles to producing this evidence are significant.

For this paper, we collected extensive data sets for several Latin American countries. We then developed a methodology suitable for extracting an answer to the inflexibility questions from these data sets.

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

Appendix A
Estimating the Speed of Adjustment

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

$$\Delta e_{it} = \psi_{it}(e^{*}_{it} - e_{i,t-1}) \ , \tag{A1}$$

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

$$\Delta e^{*}_{it} = \Delta e^{*}_{At} + \varepsilon_{it} \ \ ,$$

where Δe^{*}_{At}'s are i.i.d. with mean μ_A and variance $\sigma_A{}^2$ and where ε_{it} are i.i.d. independent from the Δe^{*}_{it}'s, with zero mean and variance $\sigma_I{}^2$.

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

$$z^{M}_{it} = \frac{1}{M}\sum_{K=1}^{M} z_{it-k} \ . \tag{A2}$$

The central idea is that with plant-specific fixed effects (for example, systematic differences in labor force composition), we do not observe the z's implicit on the right-hand side of equation (A1), but only observe the difference, $z_{it} - z_{it}{}^{M}$ (since the fixed effects cancel out once we subtract z^M). We therefore fix t and estimate equation (A1) with $z - z^M$ on the right-hand side 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 in the sample, as is the case when fixed effects are estimated using the time average over the whole sample.

Denote $\sigma_t{}^2 \equiv \text{var}\ (z_{it})$, 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

$$E\left[\hat{\lambda}_t\right] = \lambda_t \left\{1 + \frac{\sigma^2_{t-1} - \sigma^2_{t-2}}{4Var\left(z_{it} - z^{\mathrm{M}}_{it} + \Delta l_{it}\right)}\right\} \ , \tag{A3}$$

with

$$\sigma^2_t = \frac{1-\lambda_t}{\lambda_t\left[\alpha + (1-\alpha)\lambda_t\right]}\left\{\left[1-(1-\alpha)\lambda_t\right]Var(\Delta e_{it}) + \frac{\alpha(2\lambda_t - 1)}{\lambda_1}(\Delta e_{At})^2\right\}, \tag{A4}$$

where Δe_{At} denotes the average (over i) of Δe_{it}.

It follows from equation (A3) that the time average of the estimates for λ_t will be unbiased, since $\sigma_{t-1}{}^2$ is equal to $\sigma_{t-2}{}^2$, on average. The estimator may be biased for any particular t, but the expression in equation (A4) can be used to correct the bias in equation (A3), 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 equation (A3) can be obtained for values of M larger than 2. Surprisingly, the result of an unbiased average described above holds only for $M = 2$.[32] 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 stemming from time-varying fixed effects. In this sense, $M = 2$ provides a good compromise.

32. As M tends to infinity, the estimator is (asymptotically) unbiased, and there is thus no need to average over time.

APPENDIX B
Gauging the Costs

Here we show that for the model in section 5,

$$\frac{\Delta A}{A_0} \cong \left[\frac{1}{\lambda_0} - \frac{1}{\lambda_1}\right]\theta \ , \tag{B1}$$

with

$$\theta = \frac{\alpha\gamma(2-\alpha\gamma)}{2(1-\alpha\gamma)^2}(\sigma_1^2 + \sigma_A^2) \tag{B2}$$

and

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

The intuition becomes evident on considering the following equivalent problem. The economy consists of a very large, fixed number of firms (no entry or exit). Production by firm i in period t is $Y_{it} = A_{it}L_{it}^{\alpha}$, where A_{it} denotes productivity shocks, which are assumed to follow a geometric random walk.[33] Consequently,

$$\Delta \log A_{it} \equiv \Delta a_{it} = v_t^A + \mathrm{v}_{it}^I$$

where v_t^A is i.i.d. $N(0, \sigma_A^2)$ and v_{it}^I is i.i.d. $N(0, \sigma_I^2)$. Hence, $\Delta\alpha_{it}$ follows $N(0, \sigma_T^2)$, with $\sigma_T^2 = \sigma_A^2 + \sigma_I^2$. The (inverse) demand for good i in period t is

$$P_{it} = Y_{it}^{-1/\eta} \ .$$

Finally, we assume the wage remains constant throughout.

In what follows, lower case letters denote the logarithm of upper case variables, while an asterisk on a variable denotes the frictionless counterpart of the same variable without an asterisk.

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

$$\Delta l_{it}^* = \frac{\gamma}{1-\alpha\gamma}\Delta a_{it} \ , \tag{B3}$$

33. We ignore hours in the production function.

and hence

$$\Delta y_{it}^{*} = \frac{\gamma}{1-\alpha\gamma}\Delta a_{it} \ . \tag{B4}$$

Denote by Y_t^* aggregate production in period t if there were no frictions. It then follows from equation (B4) that

$$Y_{it}^{*} = e^{\tau\Delta a_{it}} Y_{it-1}^{*} \ , \tag{B5}$$

with $\tau \equiv 1/(1 - \alpha\gamma)$. Taking expectations (over i for a particular realization of v_t^A) on both sides of equation (B5) and noting that both terms being multiplied on the right-hand side are, by assumption, independent (random walk), yields

$$Y_{t}^{*} = e^{\tau v_t^A + \frac{1}{2}\tau^2\sigma_I^2} Y_{t-1}^{*} \ . \tag{B6}$$

We then average over all possible realizations of v_t^A, since these fluctuations are not the ones we are interested in for the calculation at hand. This leads to

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

and therefore, for $k = 1, 2, 3,\ldots,$

$$Y_{t}^{*} = e^{\frac{1}{2}k\tau^2\sigma_T^2} Y_{t-k}^{*} \ , \tag{B7}$$

where $Y_{t,t-k}$ represents the 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) and where $Y_{it,t-k}$ is the corresponding level of production of firm i in t.

From the expressions derived above it follows that

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

and therefore

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

Taking expectations (with respect to idiosyncratic and aggregate shocks) on both sides of this last expression (here we use the fact that Δa_{it} is independent of $Y_{t,t\text{-}1}{}^*$) yields

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

which, combined with equation (B7), leads to

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

A derivation similar to the one above leads to

$$Y_{it,t-k} = e^{\Delta a_{it}+\Delta a_{it-1}+\ldots+\Delta a_{it-k+1}}\mathrm{Y}_{t-k}^* \;,$$

which, combined with equation (B7), gives

$$Y_{t,t-k} = e^{-k\theta}\mathrm{Y}_t^* \quad , \tag{B8}$$

with θ defined in equation (B2).

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

$$Y_t = \lambda\mathrm{Y}_t^* + \lambda(1-\lambda)\mathrm{Y}_{t,t-1} + \lambda(1-\lambda)^2\mathrm{Y}_{t,t-2} + \ldots$$

Substituting equation (B8) into the expression above yields

$$\mathrm{Y}_t = \frac{\lambda}{1-(1-\lambda)e^{-\theta}}\mathrm{Y}_t^* \;. \tag{B9}$$

It follows that the production gap, defined as

$$\text{Prod.Gap} = \frac{\mathrm{Y}_t^* - \mathrm{Y}_t}{\mathrm{Y}_t^*},$$

is equal to

$$\text{Prod.Gap} = \frac{(1-\lambda)(1-e^{-\theta})}{1-(1-\lambda)e^{-\theta}} \;. \tag{B10}$$

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

$$\text{Prod.Gap} \cong \frac{(1-\lambda)}{\lambda}\theta \;. \tag{B11}$$

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 equation (B1). This concludes the proof.

References

Abraham, K. and S. Houseman. 1994. "'Does Employment Protection Inhibit Labor Market Flexibility? Lessons From Germany, France and Belgium." *In Protection Versus Economic Flexibility: Is There A Tradeoff?*, edited by R. M. Blank. University of Chicago Press.

Arellano, M. and S. Bond. 1991. "Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations." *Review of Economic Studies* 58(2): 277–98.

Burgess, S. and M. Knetter. 1998. "An International Comparison of Employment Adjustment to Exchange Rate Fluctuations." *Review of International Economics* 6(1): 151–63.

Burgess, S., M. Knetter, and C. Michelacci. 2000. "Employment and Output Adjustment in the OECD: A Disaggregated Analysis of the Role of Job Security Provisions." *Economica* 67(267): 419–35.

Caballero, R., K. Cowan, E. Engel, and A. Micco. 2004. "Effective Labor Regulation and Microeconomic Flexibility." Massachusetts Institute of Technology and Yale University. Mimeographed.

Caballero, R. and E. Engel. 1993. "Microeconomic Adjustment Hazards and Aggregate Dynamics." *Quarterly Journal of Economics* 108(2): 359–83.

Caballero, R., E. Engel, and J. Haltiwanger. 1997. "Aggregate Employment Dynamics: Building from Microeconomic Evidence." *American Economic Review* 87(1), 115–37.

Caballero, R. and M. Hammour. 2000. "Creative Destruction and Development: Institutions, Crises, and Restructuring." Paper presented at the Annual Bank Conference on Development Economics (ABCDE). World Bank, Washington, 18–20 April.

Calvo, G. 1983. "Staggered Prices in a Utility Maximizing Framework." *Journal of Monetary Economics* 12(3): 383–98.

Davis, S., J. Haltiwanger, and S. Schuh. 1996. Job Creation and Destruction. MIT Press.

Djankov, S., R. La Porta, F. Lopez-de-Silanes, A. Shleifer and J. Botero. 2003. "The Regulation of Labor." Working paper 9756. Cambridge, Mass.: National Bureau of Economic Research.

Heckman, J. and C. Pagés. 2000. "The Cost of Job Security Regulation: Evidence from Latin American Labor Markets." Working paper 7773. Cambridge, Mass.: National Bureau of Economic Research.

Hamermesh, D. 1993. *Labor Demand*. Princeton University Press.

Nickell, S. and L. Nunziata. 2000. "Employment Patterns in OECD Countries." Discussion paper 448. London School of Economics, Center for Economic Performance.

Sargent, T. 1978. "Estimation of Dynamic Labor Demand Schedules under Rational Expectations." *Journal of Political Economy* 86(6): 1009–44.

Shapiro, M. 1986. "The Dynamic Demand for Capital and Labor." *Quarterly Journal of Economics* 101(3): 513–42.