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Employment Dynamics and the Structure of Labor Adjustment Costs *

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Abstract

In this paper we discuss the structure of labor adjustment costs in relationship to the dynamics of job and worker flows. Using high frequency data, we document a previously unsuspected degree of lumpiness in employment adjustment, which is characteristic of non-convex adjustment costs. By means of the statistical analysis of duration data, we relate that lumpiness to the structure of adjustment costs and not to the structure of shocks.

Keywords: Adjustment Costs, Job Flows, Worker Flows, Duration Models

JEL codes: J23, J63.

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1 Introduction

Labor adjustment costs shape the pattern of employment adjustment. Depending on their structure, the employment path observed at the micro-level will be different. If adjustment costs are convex, adjustment is frequent but employment variations in each period small. If, on the contrary, adjustment costs are non-convex (e.g., linear or fixed), adjustment is rare and employment changes large. In the latter case, but not in the former, inaction (i.e., the absence of adjustment) may be optimal even if employment is not at its long-run equilibrium.

Distinguishing between these two polar structures of adjustment costs is a major issue in the dynamic theory of labor demand (see Nickell, 1986; Hamermesh, 1993a; Hamermesh and Pfann, 1996). In fact, understanding the magnitude and structure of adjustment costs is essential to the study of the cyclical behavior of employment and the evaluation of job-security policies. Firing costs, in particular, have long been at the center of the debates over Europe's unemployment problem.

In the literature the issue is typically addressed via the estimation of structural models of labor demand that embed alternative structures of adjustment costs. What we know from the few studies that follow in this approach (Hamermesh 1989, 1993b; Anderson 1993; Rota, 1994) is that non-linear labor demand models that correspond to non-convex adjustment cost structures fit micro-data on employment well and outperform alternative specifications.

The fast growing literature on job and worker flows (surveyed by Davis et al., 1996) offers the opportunity to distinguish indirectly between competing hypotheses about the structure of adjustment costs. Establishing the empirical properties of job and worker flows permits us to fully characterize the micro-level patterns of employment adjustment and, subsequently, to compare these with those implied by different structures of adjustment costs.

There are multiple ways in which micro-level data on employment adjustment can shed light on the underlying structure of adjustment costs.

First, non-convexities in the adjustment technology translate into the distribution of employment adjustment that is expected to exhibit a high proportion of extreme events. Fixed or linear adjustment costs imply that micro-units experience episodes of sharp adjustment followed by relatively long periods of no adjustment. Such a pattern of adjustment implies that a non-trivial share of aggregate employment adjustment is accounted for by these extreme events, which result in long fat tails of the employment change distribution.

Second, non-convex adjustment costs also imply that inaction is optimal even outside equilibrium, which is contrary to what occurs if adjustment costs are convex. The fact that establishments or firms are inactive for a large number of time-periods is an indication of the importance of one type of adjustment costs over the other.

Evidence will be more conclusive if we can follow micro-units over time and see how each period's adjustment record compares with the unit's entire record over the sample period. If adjustment episodes are actually sharp and rare, and action is most likely followed by inaction, then the observed pattern of employment adjustment is consistent with non-convexities in the adjustment technology. If adjustment is smooth and persistent, convex adjustment costs must be dominant.

However, in order to successfully attribute the observed pattern of employment adjustment to one type of adjustment cost or to the other, it is essential to establish that other factors (such as the structure of shocks themselves) are not behind what we observe. If we do not observe the determinants of labor demand, this is feasible only if data on the timing of employment adjustment are available.

It has been shown that the two alternative adjustment cost structures coupled with different patterns of shocks imply differently shaped hazard functions (Power, 1994, Cooper et al., 1999). In particular, non-convex adjustment costs with autocorrelated shocks imply that the conditional probability of adjustment is increasing in the time elapsed since the previous adjustment episode, hence, upward sloping hazards. Non-convex adjustment costs coupled with the absence of a trend in the process governing the desired level of employment imply that the probability of adjustment is independent of time, as do convex adjustment costs, as well. Hence, whereas flat hazards are not necessarily produced by convex adjustment technologies, upward sloping hazards unequivocally identify non-convex adjustment costs.

The purpose of this article is to investigate the distinction between the two alternative adjustment cost structures empirically, focusing on the observed properties of job and worker flows. For this purpose, we rely on Portuguese data for two reasons. First, the Portuguese labor market is an extreme case of high firing costs (OECD, 1999, Blanchard and Portugal, 2001). Second, suitably high-frequency (quarterly) longitudinal data are available.

The paper is organised as follows. In Section 2, the dataset used is described. The concentration of job and worker flows and the frequency of inaction are analyzed in Section 3. In Section 4 individual employment series are characterized and the presence of spikes in those series is investigated. In Section 5, employment regimes are defined and transitions across regimes are described. In Section 6, the statistical analysis of duration data is used to distinguish between the roles of the structure of adjustment costs and the structure of shocks in shaping the observed micro-level patterns of employment adjustment. Section 7 concludes.

2 The Data

The data used in this article come from the *Inquérito ao Emprego Estruturado* (IEE) which is a survey run by the Portuguese Ministry of Employment. The IEE collects quarterly information on job and worker turnover at the establishment level. Establishments of all sizes and in all industries are included.[1] The corresponding sample is drawn from the universe of the respondents to *Quadros de Pessoal* (QP), which is an annual survey mandatory for all establishments with wage earners.

The probability of units with fewer than 100 employees being selected to the IEE sample is inversely related to the size of the establishment. Above that threshold,

[1] Only Agriculture, Fisheries, Public Administration and Private Services are excluded.

establishments are selected with certainty.[2] The sample is statistically representative for three-digit industries (as defined by the SIC code), region and size class. For this purpose, seven regions - five in mainland Portugal and the islands of Madeira and the Azores - were considered and six size classes were defined.

The dataset used here spans over twenty quarters, from the first quarter of 1991 until the last quarter of 1995. Units in the sample were selected from the 1990 QP file. There are a total of 139,203 observations (establishments*quarters) in the sample.

The original twenty quarterly files were converted into two datasets that are, hereinafter, referred to as the pooled dataset and the longitudinal dataset.

The pooled dataset simply pools all the twenty quarterly files. No major modifications to the original files were made except that all records (127) with zero employment at both the beginning and end of period were deleted. The pooled dataset contains 139,076 records corresponding to 10,673 establishments.

The longitudinal dataset results from merging the twenty quarterly files. All records in every quarterly file have an identification code that is unique and does not change during the whole period the establishment remains in the sample. This code number served as the key for merging the two original files. As a result, an unbalanced panel of 10,673 establishments was obtained.[3] This was used to generate one balanced panel of 2,181 establishments for which information is available in each and every one of the twenty quarters surveyed.

Quarterly measures of job flows were computed using the end-of-period head-count reported in two adjacent spells of the survey. The hiring and separation rates were computed using the information on the total number of hirings and separations reported by the respondent units in each spell of the survey. The dataset also contains information on the end-of-period head-count by type of contract. Combining this information with the establishment head-count permited us to compute, for each establishment, the proportion of workers of the establishment

[2]This threshold is set at 50 employees for the Azores and Madeira regions.

[3]This is the number of establishments that were present in the sample at least once over the entire twenty-quarter period.

with fixed-term contracts and employed on a part-time basis, both of which are used in Section 6.

3 The Distribution of Job and Worker Turnover

The conventional way of checking for the relative importance of smooth versus lumpy adjustment patterns is to examine the distribution of net or gross job flows (see Boeri and Cramer, 1992).

Consider first the distribution of job creation and job destruction. In Figure 1 we depict the proportion of jobs created (destroyed) by establishments expanding (contracting) at different growth rate intervals (as measured in the horizontal axis).[4] The bars to the right of the origin correspond to job creation and those to the left to job destruction. By design, smaller episodes of job creation and destruction (those that imply an employment variation of as much as 10 percent) are concentrated in the first columns to the right and to the left of the origin.

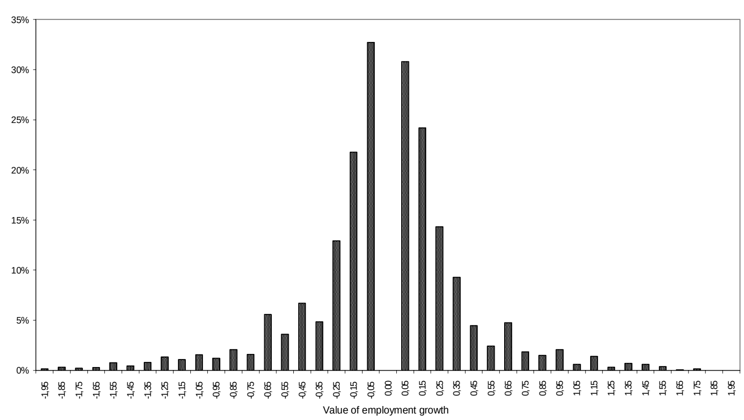

Figure 1: Distribution of Job Creation and Destruction (1991-95).

The height of these two columns indicates that establishments experiencing mild employment changes account for about 30 to 34 percent of all job creation

[4] All job and worker flows were computed according to the standard Davis and Haltiwanger definitions (see Davis et al., 1996). Over the sample period, quarterly job turnover (start-ups and shutdowns excluded) is equal to 5.4 percent (2.3 for job creation and 3.1 for job destruction) and quarterly worker turnover is equal to 8.1 percent (3.7 percent for hirings and 4.5 percent for separations). For details, see Varejão (2003).

and destruction, respectively. The complement to this information is that, on both margins, job flows are concentrated in establishments that are going through sharp employment changes: the share of job creation accounted for by establishments expanding more than 10 percent (conventionally measured) is 70 percent, whereas the corresponding figure for job destruction is 66 percent. Concentration is slightly greater for job creation than for job destruction, but both spread over the entire range of employment growth rates.

The distribution of gross job flows - hiring and separations - is represented in Figure 2. The height of each column in this figure measures the proportion of all hirings (separations) that are accounted for by establishments hiring (firing) workers at different rates (measured on the horizontal axis). The height of the two first columns in each panel indicates that only 34 percent of all hirings and 36 percent of all separations occur at establishments hiring or separating in a single period the equivalent to less than 10 percent of its average workforce in the same period. But the tails of both distributions also indicate that a non-trivial number of establishments experience extreme episodes of hiring and firing.

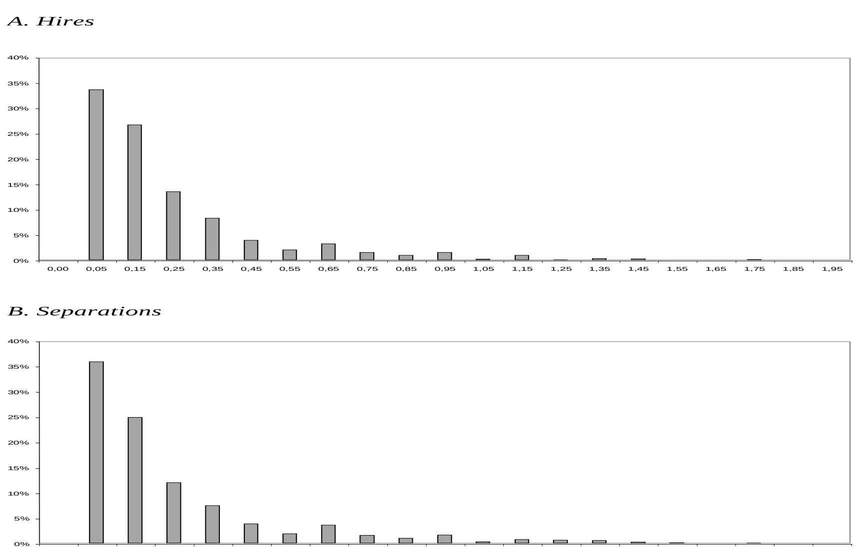

Figure 2: Distribution of Hirings and Separations

By design, establishments with stable employment are excluded from Figure 1, as are establishments reporting no hirings or separations excluded from Figure 2. The share of each type of establishment is, respectively, 74.7 percent, 83.7 percent and 80.6 percent, meaning that,*on average, at least three quarters of all observed units do not change employment, hire any worker or separate from any of its workers over an entire quarter.* This is clear evidence of the pervasiveness of inaction at the micro-level.

Inaction pervasiveness and fat tails in the distribution of job flows indicate how important non-convex adjustment costs are. However, the concentration of these very same distributions also documents the importance of convex adjustment costs. Put differently, for both net and for gross employment flows, strong evidence of a pattern of adjustment consistent with a non-convex adjustment cost function coexists with signs of smooth adjustment.

4 Spikes in Individual Employment Series

The distinctive characteristic of the pattern of employment adjustment implied by non-convex adjustment cost structures is the presence of spikes - i.e. infrequent moments of sharp employment adjustment - in the establishment-level employment record.

Therefore, discriminating empirically between convex and non-convex adjustment cost structures also implies an investigation of the importance of adjustment spikes.

An appropriate way of doing this is to put each establishments adjustment record in one quarter against the background of its entire record over the whole sample period (twenty periods). If adjustment was lumpy, the employment record over this period should have a few spikes, which, in the context of the whole period of observation, would show as moments of sharp adjustment.

To check for the existence of these spikes we compute, establishment-by-establishment, quarterly rates of turnover (net and gross). Three series for each unit in the panel - one for the net employment change, another for the hiring rate and the third for the separation rate - are thus obtained. Each individual series is then ordered from its highest value (rank 1) to the lowest (rank 20).[5] If the employment series exhibits a few spikes, the first ranks of each individual series should be of a magnitude greater than that of the remaining ranks. For example, if only one spike occurred during the whole sample period, then rank 1 corresponds to the sole period with a spike and will be much greater than rank 2 and all the subsequent ranks, which, in this case, will not correspond to spikes. If no spikes occurred (as we expect if adjustment costs are convex), then all the 20 ranks of each series will be of a similar magnitude.

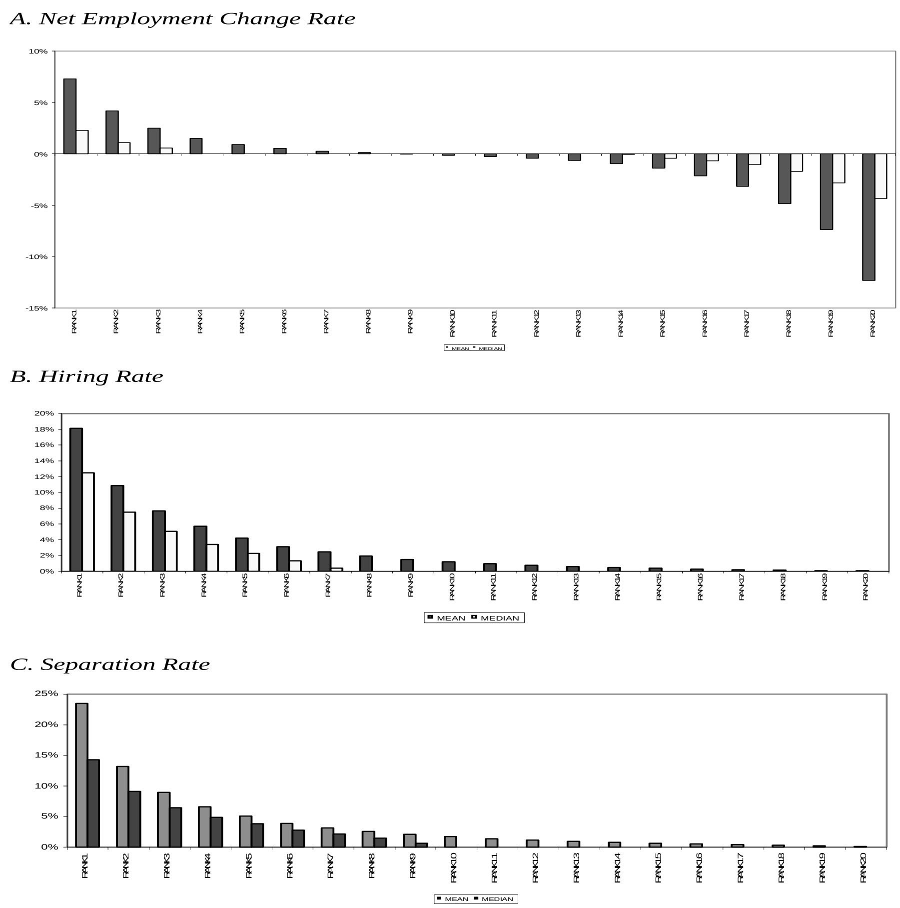

Figure 3: Net and Gross Employment Flow Rates, by Rank

[5]Here we follow the approach adopted by Doms and Dunne (1998) in their study of plant-level patterns of capital accumulation.

The next step is to compute, across establishments, the average of each rank. It is the figures thus obtained that are represented in the three panels of Figure 3. What they tell us is, for example, for hiring rates, that the sharpest hiring episode, which for different plants may have occurred in different calendar quarters, corresponds to an average hiring rate of 18 percent (rank 1). The figures corresponding to the remaining ranks should be interpreted similarly, remmembering that rank 2 corresponds to the second sharpest espisode, and so on.

The purpose of constructing these series is to compare the magnitude of the highest rank of each series (the highest and the lowest in the case of the net employment change series) to the remaining ranks. As discussed, the difference between the highest and the second highest ranks in each series measures the relative importance of the two sharpest adjustment episodes, large differences indicating the presence of spikes and lumpy adjustment processes.

For the net employment change series the mean of rank 1 is 7.3 percent and of rank 20 is -12.3 percent (Figure 3). This means that the greatest (least) net employment change corresponds to 7.3 (-12.3) percent of the average employment in the quarter in which the change occurred. Figure 3 also indicates that rank 1 is more than 70 percent higher than rank 2 and rank 19 is 66 percent higher than rank 20. This indicates clearly that large episodes of net employment adjustment are, indeed, extreme events in the history of employment adjustment of the individual establishments. The pattern of employment adjustment these data convey is consistent with non-convexities in the adjustment cost function.[6]

Evidence in favor of lumpy adjustment is even clearer with gross flows. At its maximum, the hiring rate represents about 18 percent of the establishments workforce and this percentage drops off significantly after rank 1. An even more pronounced picture is obtained on the separations side. The average of the establishment-level maximum separation rate rounds off to 23 percent and it also drops off significantly after rank 1.[7]

[6]When interpreting the mean values of the net employment change associated with each rank, it should be remembered that some units may be expanding or shrinking throughout all of the 20 quarters surveyed, in which case they artificially reduce the mean value of each rank.

[7]The pattern of net employment adjustment depicted in Figure 3 could also imply that adjustment is lumpier

Both hiring and separation rates are below 5 percent after rank 5 and remain above zero until rank 20, indicating how infrequent even mildly large adjustment episodes are.

5 Quarterly Transition Rates across Employment Regimes

The importance of large and infrequent episodes of employment adjustment was documented in previous sections. However, if this pattern of adjustment is to be attributed to non-convexities in the adjustment cost technology, it is essential to also analyze the sequence of events. Convex adjustment costs imply that one period of adjustment is followed by yet another period of adjustment, the intensity of episodes decreasing over time. On their side, non-convex adjustment costs imply that one period of adjustment is followed by periods of inaction.

To distinguish between these two adjustment patterns, all establishments in each period were classified into one employment regime and their situation one period ahead was recorded. Six alternative employment regimes as defined in Table 1, were considered. This information was then used to compute the probabilities of transition across regimes. With convex adjustment costs, we should observe high probabilities of false transitions (transitions into the same regime). With non-convex adjustment costs we should observe high probabilities of transition to the inaction regime, which should be resilient.

Hence, the focus of our analysis is on the principal diagonal of the matrix of probabilities of transition across employment regimes. High values on this diagonal must be taken as signals of smooth adjustment, except for the cell corresponding to the inaction regime. Signals of lumpy adjustment must show as high values in the third column of the matrix, where the probabilities of transition from action to inaction and of inaction persistence are documented.

Looking first at the main diagonal of the matrix in Table 1, it becomes clear that with the exception of the expansion regimes (regimes 4 and 5), establishments

in cases of workforce reduction, although this would have to be controlled for by using data on the possible asymmetry of shocks.

	Regime					
Regime	**1**	**2**	**3**	**4**	**5**	**6**
1 (Δ $E<0$, $H=0$)	29.7	15.0	27.3	5.3	12.3	10.0
2 (Δ $E<0$, $H>0$)	21.0	34.0	7.3	7.3	7.2	23.2
3 (Δ $E=0$, $H=0$)	14.5	1.9	70.0	3.2	8.5	2.0
4 (Δ $E=0$, $H>0$)	20.7	15.5	24.9	11.2	13.0	14.8
5 (Δ $E>0$, $S=0$)	24.8	8.9	36.0	7.6	15.4	7.3
6 (Δ $E>0$, $S>0$)	15.1	30.0	7.3	7.4	8.9	31.4

Table 1: Transition Across Employment Adjustment Regimes.

in each regime are likely to be in that same state one period ahead. Expanding establishments (regimes 4 and 5) will most likely move into the inaction regime (regime 3). Seventy percent of all establishments that visit the inaction regime in one quarter will still be in that same regime in the subsequent quarter.

Column three tells us that those establishments that make a true transition move primarily to the inaction regime (the exceptions being transitions originating in regimes 2 - employment declining, but hiring-, and in regime 6 - employment declining but with separations).

The resilience of the inaction regime and, particularly, the importance of this regime as a destination of all establishments that make a transition from one quarter to the next are consistent with fixed adjustment costs. This result should be emphasized as the regime definition that has been used biases the results in favor of smooth adjustment. Here, for an establishment to be classified as inactive, not a single worker is allowed to move in or out of the establishment. Complete inaction is even less likely because natural attrition, which is observed as a visit to one action regime (regime 1 or 2), implies positive separations. [8]

[8] Arguments in favor of using a relative criterion, as opposed to the absolute zero criterion used here to define the inaction regime, can be found. Relative measures could be preferred if, as it is here, we cannot discriminate between employer and employee-initiated separations. However, in a strict formulation, fixed adjustment costs imply that the costs borne by the establishment by hiring/firing (expanding/contracting by) one worker are exactly the same of larger similar actions. This is, of course, why we expect lumpy rather than smooth adjustment when this strucuture of adjusmtent costs dominates.

6 A Duration Model of Employment Adjustment

6.1 Motivating the approach

The importance of establishment-level discrete patterns of adjustment was largely documented in the previous sections. However, no link between the observed micro-level adjustment pattern and the underlying structure of adjustment costs has yet been established. So far, we have established only that the pattern of employment adjustment observed is consistent with non-convex adjustment costs. But we could not conclude that it is the result of non-convex adjustment costs because other factors may be shaping what we observe. Shocks impacting establishments are the most obvious alternative explanation. Concern is that lumpy adjustment processes reflect a skewed distribution of idiosyncratic shocks, rather than the structure of adjustment costs.[9] Hence, it is essential for our purpose to distinguish between these two causes.

The standard approach to this problem is to estimate structural labor demand models with non-convex adjustment costs and see how they fit the data (and, possibly, compare with the results obtained for models with alternative convex adjustment cost structures).

An alternative approach focuses on the timing of employment adjustment, especially the time between two consecutive episodes of adjustment. Intuition is simple and may be easily rationalized by a simplified (S,s) model which implies a micro-level path of adjustment of the same type as the one resulting from non-convex adjustment costs.

Assume, as in Caballero (1992), that at time t each individual establishment i has a desired level of employment l^*_{it}, which, because of non-convex adjustment costs, is not necessarily equal to the actual level of employment, l_{it}. Denote by s_{it} the difference between desired and actual employment levels ($s_{it} = l^*_{it}$-l_{it}). With non-convex adjustment costs, firms adjust the employment level when s_{it} hits

[9] Notice, however, that the observed pattern of adjustment cannot be reconciled with convex adjustment costs. Convex adjustment technologies imply continuous adjustment activity and cannot match either the evidence of pervasive inaction or the observed transience of action regimes.

the upper or lower bounds of its inaction region (denoted, respectively, S_i and s_i). Assume further, that l_i^* is governed by a Brownian motion with drift μ and instantaneous variance σ^2.

Under these assumptions, for each individual firm, there is a probability density function for the firms location within its inaction regime. The exact shape of this density function depends on the value of $\frac{\mu}{\sigma^2}$ (for details, see Foote, 1998). Firms adjust the level of employment whenever they hit one barrier control (S_i or s_i), which may occur either because they received a shock to l^* great enough or because they received a series of smaller shocks of the same sign. Such shocks change the distance (s_{it}) between a firm's current location and the position of the inaction bands.

Hence, in the non-convex case, the probability of employment adjustment depends on how firms approach the inaction band, if they actually do. If $\mu \neq 0$ the probability of adjustment may be thought of as a positive function of time elapsed since the previous adjustment, as each shock received brings the firm a little closer to one control barrier. [10] If $\mu = 0$, the hazard function is constant and the duration of the inaction regime is exponentially distributed.

Statistical duration analysis offers a methodology well-suited for distinguishing the roles that non-convexities in the adjustment technology and the structure of shocks play in shaping micro-level patterns of employment adjustment. Only if adjustment costs are non-convex (and shocks are autocorrelated) can we expect to see a positive relationship between the conditional (on time) probability of adjustment and the duration of the no-adjustment period, that is, a positively sloped hazard function. In all remaining cases no such relationship (positive or negative) exists, and hazard functions are expected to be flat.[11]

[10] Higher values of σ reduce the probability that either control barrier is hit in each period.

[11] If adjustment costs are convex, inaction is not expected outside the steady-state. Starting from the equilibrium, the probability of adjustment depends only on receiving a shock, independently of its size. Hazard-functions based approaches were used before in the study of micro-level patterns of investment (Power, 1994, Cooper et al., 1999) and, with some modifications, to describe the patterns of employment adjustment (Caballero et al., 1997). The approach adopted here makes this study closer to the studies on investment than to the one on employment.

6.2 Estimation Procedure

The estimation of the hazard function, as applied to the context of employment adjustment, starts with the definition of the duration variable (t) that measures the establishments time of stay in the inaction regime. For that purpose, a flow-sampling scheme was adopted. According to this scheme, each establishment is selected upon entry to the inaction regime (at which point its individual clock is set to zero) and followed until exit time. All units are observed over a fixed time interval (from the first quarter of 1991 to the fourth quarter of 1995). Hence, left censoring is eliminated by construction, but right censoring may exist and must be accommodated.

A useful concept in statistical analysis of a duration phenomenon is the *hazard function*. In the study of inaction duration, the hazard function gives the instantaneous probability of adjusting employment at t, given that the establishment stayed inactive until t

$$h(t) = \lim_{\Delta t \to 0} \frac{(P(t \leq T < t + \Delta t \mid T \geq t)}{\Delta t} = \frac{f(t)}{1 - F(t)} = \frac{f(t)}{S(t)}. \qquad (1)$$

where $f(t)$ is the probability density function, $F(t)$ is the distribution function, $S(t)$ is the survival function. A useful function is the integrated hazard function

$$\Lambda(t) = \int_0^t h(u)du \qquad (2)$$

which relates to the survivor function simply by

$$S(t) = exp\left(- \int_0^t h(u)du\right) = exp(-\Lambda(t)) \qquad (3)$$

In this paper we employ a conventional Weibull hazard model

$$h(t) = \rho\lambda^\rho t^{\rho-1} \qquad (4)$$

which implies the following survival function:

$$S(t) = exp[-(\lambda t)^\rho] \qquad (5)$$

and the corresponding cumulative hazard function

$$\Lambda(t) = (\lambda t)^{\rho} \tag{6}$$

The Weibull distribution function is a natural choice since it allows a direct test of duration dependence based solely on its shape parameter ρ. A ρ parameter lower than 1 indicates negative duration dependence. Symmetrically, $\rho > 1$ implies monotonic increasing hazard rates through time. An exponential duration distribution (and a constant hazard function) is implied by $\rho = 1$.

In this paper we shall also distinguish between two exit modes out of the inaction regime: employment increase or decrease. Thus, we define cause-specific hazard functions to destination j

$$h(t)_r = \lim_{\Delta t \to 0} \frac{(P(t \leq T < t + \Delta t, R = r \mid T \geq t)}{\Delta t} \tag{7}$$

which yield the aggregate hazard function

$$h(t) = \sum_{j=1}^{2} h_j(t) \tag{8}$$

and the survivor function

$$S(t) = \prod_{j=1}^{2} S_j(t) \tag{9}$$

where $S_j(t) = e^{-\Lambda_j(t)}$ and $\Lambda_j(t) = \int_0^t h_j(u)du$.

The model has a conventional competing risks interpretation. In this framework, a latent duration (T_j) attaches to each exit mode. We only observe the minimum of each latent variable. If risks are assumed to be independent, with continuous duration, this model simplifies to two separate single-cause hazard models.

A common way to accommodate the presence of observed individual heterogeneity is to specify a proportional hazards model

$$h(t \mid x) = h_{0j}(t) exp(x'\beta_j) \tag{10}$$

where $h_{0j}(t)$ denotes the baseline specific hazard function, that is, the hazard function corresponding to zero values for the covariates x. In this case, the covariates affect the hazard function proportionally (i.e. $\frac{dh(x)}{dx_k} = \beta_k h(x)$). An implication of this assumption is that the impact of the covariates does not change (in relative terms) with the progression of the spell of inaction.

Our information on the elapsed duration of inaction is grouped into quarterly intervals (while transitions can only be identified over a fixed interval of one quarter). Let $M = m$ denote the occurrence of an exit in a given month $[c_{t-1}, c_t]$, where m is the realization of a discrete random inaction duration variable $M \in (1, ..., K)$. The probability that an event occurs in the m^{th} interval (i.e. that an exit occurs over the course of the 3-month window), and that such an exit is to destination r, will be given (neglecting, for the sake of parsimony, the t and x variables) by

$$f_j(m) = \frac{S_j(m-1) - S_j(m-1)}{S_j(m-1)} S(m-1) = h_j(m)S(m-1). \tag{11}$$

The functions $f_j(m)$ and $[1 - S_j(m)]$ provide a convenient characterization of the probability density and the cumulative functions associated with the marginal distribution for each latent duration, T_j , in terms of the specific hazard function $h_j(m)$. A censored observation (i.e. a spell of inaction that is still in progress after the 3-month window) occurs with probability $S(m) = \prod_{j=1}^{2} S_j(m)$, which is simply the product of the two specific survivor functions.

With our sampling plan, where we collect the information of inaction duration for the flow of entrants into the inaction regime, the contribution of observation i for the likelihood function is simply

$$L(\theta|t, j, x) = \prod_{m=1}^{K-1} \prod_{j=1}^{2} [S_j(m-1) - S_j(m)]^{\delta_{mj}} [\prod_{m=2}^{K} S(m)]^{1-\delta_m} \tag{12}$$

where θ is a vector of parameters that include regression coefficients and baseline hazard parameters, and δ_{mj} is an indicator that assumes the value 1 if the individual exits to destination j during the m^{th} interval, and 0 otherwise. The indicator $\delta_m = \sum_{j=1}^{2} \delta_{mj}$ identifies complete durations, so that $1-\delta_m$ equals 1 for a censored observation. The contribution to the likelihood function from a censored observation is simply the product of the two specific survival terms ($\prod_{j=1}^{2} S_j(m)$), that is, the probability of not exiting to either employment growth or employment decline.

We also attempt to accommodate the presence of unobserved individual heterogeneity by assuming, as conventional, a multiplicative error term associated with each specific hazard function

$$h(t \mid x) = h_{0j}(t)exp(x'\beta_j)v_j \tag{13}$$

We further assume that the errors v_j are gamma distributed, with mean 1 and variance σ_j^2, and are uncorrelated across destinations (independent competing risks).

We proceed by redefining the specific "survivor" function using the well-known result for gamma mixtures $S_j(m) = (1 + \sigma_j^2 \Lambda_j(m))^{-\frac{1}{\sigma_j^2}}$ (see Lancaster, 1990, p.66). After this transformation, the likelihood function is derived as above

$$L(\theta, \sigma_j^2 | t, j, x) = \prod_{m=1}^{K-1} \prod_{j=1}^{2} [(1 + \sigma_j^2 \Lambda_j(m-1))^{-\frac{1}{\sigma_j^2}} - (1 + \sigma_j^2 \Lambda_j(m))^{-\frac{1}{\sigma_j^2}}]^{\delta_{mj}} \tag{14}$$

$$[\prod_{m=2}^{K} 1 + \sigma_j^2 \Lambda_j(m))^{-\frac{1}{\sigma_j^2}}]^{1-\delta_m}$$

Empirical implementation of the model implies a definition of the no-adjustment regime. There is no such obvious definition. For that reason three alternative criteria were used.[12] These criteria are defined as follows:

- Absolute zero threshold on net employment adjustment: the establishment is classified as inactive if, during the period, there was no change in the level of employment;

- Relative 10 percent threshold on net employment adjustment: the establishment is classified as inactive if, during the period, the change in the level of

[12] An additional definition setting a 2.5 percent threshold was also used but the corresponding results are not reported as they do not differ substantially from those corresponding to the first criterion listed here.

employment is less than 10 percent of the employment count at the beginning of the period;

- Collective dismissal criterion: the establishment is classified as inactive whenever its employment level varies by less than two or less than five depending on whether the establishment employs fewer or more than fifty workers.[13]

The first criterion corresponds to the strictest definition of inaction as applied to net employment adjustment. This would be the most appropriate for investigating the importance of, say, fixed adjustment costs, as these imply that the same cost is borne independently of the number of individuals joining or leaving the establishment and of the size of the establishment itself. However, because data on separations do not permit us to distinguish between firings and other separations (temporary separations, voluntary quits, retirements, or deaths) a relative threshold - the second criterion - may also be adequate. Finally, the third criterion listed follows the legal rule applying to the definition of collective dismissals, which are submitted to a number of obligations and imply considerably higher firing costs than those applying to individual dismissals.[14]

Being able to use alternative definitions for the no-adjustment regime has the obvious advantage of permitting us to contrast the corresponding predictions and evaluate the sensitivity of the results to necessarily arbitrary thresholds.

6.3 Estimation Results

Results of fitting the duration model with no covariates added for the three alternative definitions of the no-adjustment regime are depicted in Table 2.

The first result to notice is that, independently of how inaction is defined, the estimated values of the λ parameters of the Weibull distribution are low, indicat-

[13]The legal definition applies to separations only. In the absence of an obvious corresponding criterion to apply on the hiring margin, the option was to set the same absolute thresholds on this margin as well, and work with an inaction region symmetric about zero.

[14]The Dismissals Act (Law 64-A/89) makes the distinction between individual and collective dismissals on the grounds of the number of individuals being dismissed within a three month period. For a dismissal to be termed collective and subject to the corresponding legislation, the law requires a minimum of two or five workers to be dismissed simultaneously (i.e., within a three month period), depending on the firm having fewer or more than 50 workers.

	Employment Expansion		Employment Decline	
Inaction (Absolute Zero Threshold)				
	coefficient estimate	t statistic	coefficient estimate	t statistic
λ	0.036	27.506	0.052	39.718
ρ	1.069	54.636	1.073	66.976
n		14710		14710
Log likelihood		-8864.6		-12004.1
Inaction (10 percent Threshold)				
λ	0.011	13.802	0.015	17.777
ρ	1.089	36.706	1.118	42.151
n		17121		17121
Log likelihood		-5473.1		-6787.0
Inaction (Collective Dismissal Criterion)				
λ	0.015	17.968	0.021	24.055
ρ	1.048	44.701	1.020	53.756
n		18123		18123
Log likelihood		-7387.1		-10024.8

Table 2: Weibull Distribution Duration Model.

ing a low conditional probability of abandoning the no-adjustment regime. This is consistent with the evidence discussed previously that documented a substantial degree of inactivity and few quarterly transitions from inaction to action. Furthermore, low hazard rates are observed for the two exit modes: employment expansion and decline. The probability of exiting the inaction regime is just slightly higher in the case of the employment decline destination in comparison with employment expansion. This is the first indication that the timing of employment adjustment is fairly symmetric.

In all cases, the Weibull hazard function exhibits a positive duration dependence. The estimated ρ parameters are always greater than one and, with one exception, the hypotheses of constant hazard rates ($\rho = 1$) are soundly rejected. As explained above, this is clear evidence of the presence of non-convexities in the adjustment cost function.

However, in this very stylized model the estimated shape parameters (of the Weibull distribution) are barely above one. It is, in fact, well known from the duration analysis literature that failure to properly account for individual characteristics biases the results towards negative duration dependence.

Control for individual specific effects was, thus, implemented via the use of regression analysis. A set of covariates contemporaneous to the timing of events was added to the model. These covariates control for the size of the establishment (as measured by the log of total employment) and for the proportion of workers with fixed-term contracts, and part-time contracts at the establishment level. Results are reported in Table 3.

The estimated ρ parameters show that controlling for observed individual heterogeneity increases the slope of the hazard function, for all inaction regime definitions. Remarkably, the estimated ρ parameters suggest, again, that the shape of the hazard functions is very similar across destination states. This is a second indication of symmetry between employment expansion and decline.

Results in Table 3 further tell us that for all the criteria but the second, larger firms are those facing the highest probability of exiting inaction (a one percent increase in total employment increases the hazard rate by 0.51 to 0.80 percent).

However, it is interesting to look more carefully at the estimated coefficient of the log of employment in the second specification. Remember that this estimate corresponds to the more demanding definition of action for variations of employment as large as 10 percent of the beginning-of-period count, establishments are still considered inactive according to this criterion. Put differently, it takes a net employment variation as large as 10 percent of the beginning-of-period count for an establishment to be considered as exiting the no-adjustment regime. What the estimate of the employment coefficient tells us is that larger establishments are the least likely to exit the no-adjustment regime so defined, although they are more likely to do so for alternative inaction definitions. This necessarily implies that most of the action we observed among larger establishments corresponds to relatively small adjustment episodes.

The proportion of the workforce with fixed-term contracts was included as a regressor because costs of adjusting labor are lower if establishments employ temporary workers. To the extent that part-time work can also be considered, as it usually is, as a contingent or flexible form of work, it is also appropriate to include a variable measuring the proportion of part-time workers.

Results in Table 3 also indicate that the proportion of fixed-term contracts has, indeed, a strong positive effect on the establishments conditional probability of exiting inaction. Raising the proportion of temporary workers by one percentage point increases the conditional probability of exiting the inaction regime between 0.9 and 1.8 percent, depending on the regime definition and exit mode. The fact that the largest estimates for the coefficient of the fixed-term contract variable were obtained when the most demanding definitions of inaction (the 10 percent employment threshold) indicates that fixed-term contracts are particularly instrumental in facilitating severe employment adjustment, which indicates that, to some extent at least, they play the role of buffer-stocks.

For the part-time variable, results indicate that part-time work either has no statistically significant effect on the probability of adjusting employment or it has a negative effect. This is contrary to what we would expect if establishments employed workers on a part-time basis, because they are less costly to hire and fire, but is consistent with the fact that all of the existing legislation applying to full-time work also applies to part-time work on the basis of the proportion of normal full-time hours that part-timers work.

In an attempt to control for the conditions that initiated the current spell of inaction, a dummy variable that indicates whether employment decline or employment expansion preceeded the current inaction episode was also included. No clear pattern emerges from this exercise, suggesting that state dependence (generated by persistent product demand shocks, for example) does not appear to play an important role.

Finally, an attempt was made to control parametrically for the unobserved individual heterogeneity assuming a gamma distributed error term (Table 4). This estimation suggested that individual unobserved heterogeneity may be present in five regressions, further reinforcing the empirical evidence of increasing hazard rates. [15]

7 Conclusions

If adequate matched employer-employee longitudinal data are available, it is possible to assess the relative importance of convex and non-convex adjustment cost structures. Signs of both patterns of adjustment were investigated by checking the importance of extreme events (jumps in employment processes), their frequency and sequel. Unequivocal signs of discrete adjustment consistent with fixed adjustment costs were found at all these different levels. Large employment changes (larger than 10 percent of the establishments workforce) account for two-thirds of total job creation and destruction, and also of all gross employment flows. Inaction is pervasive with 75 percent of all units surveyed not changing the level of employment over one quarter, and 72 percent also not hiring or firing a single individual over the same period. Visits to any regime implying some sort of adjustment are frequently followed by a transition to the inaction regime, which emerges as highly resilient.

Distinguishing between the effects of the pattern of shocks and those of the structure of adjustment costs is essential for the purpose of this chapter. Statistical analysis of duration data successfully permitted us to attribute the observed pattern of employment adjustment to the underlying structure of adjustment costs, not to shocks. If individual heterogeneity is properly accounted for, the hazard function (defined as the instantaneous conditional probability of abandoning the inaction regime) is unmistakably upward sloping. This indicates that there is a non-convex component to the adjustment cost technology and that it is important enough to shape the whole process of adjusting employment. This result survives changes in the definition of the inaction regime.

Independently of the regime definition used, the proportion of workers with

[15] Alternatively, individual heterogeneity was also controlled for by the establishments pre-sample history of employment adjustment (Blundell et al., 1995). Results of estimating the model accounting for each establishments previous labor market history (not reported here) further reinforced the indication of increasing hazard function. The downside of this approach is that eight cross sections of observations are lost for estimation.

fixed-term contracts emerged as a major determinant of the micro-level conditional probability of a transition from the inaction regime to one implying some sort of adjustment. Establishments with a larger share of temporary workers have a higher probability of abandoning the inaction regime. This result highlights the fact that the existence of low-firing-cost contracts influences the way employers respond to shocks on both margins.

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	Employment Expansion		Employment Decline	
Inaction (Absolute Zero Threshold)				
	coefficient estimate	t statistic	coefficient estimate	t statistic
Employment (log)	0.650	34.643	0.806	51.486
Fixed-term contracts (% of total)	1.153	9.898	0.925	8.656
Part-time workers (% of total)	-0.560	-2.545	-0.205	-1.158
From employment declining regime	0.084	1.848	0.067	1.794
λ	0.081	20.173	0.103	26.230
ρ	1.480	54.687	1.571	68.440
industry and year dummies present				
number of observations		14710		14710
Log likelihood		-7944.5		-10157.3
Inaction (10 percent Threshold)				
Employment (log)	-0.205	-9.033	-0.129	-6.530
Fixed-term contracts (% of total)	1.520	10.545	1.819	14.821
Part-time workers (% of total)	-0.119	-0.451	0.414	1.953
From employment declining regime	0.323	4.854	-0.095	-1.638
λ	0.035	12.101	0.036	13.277
ρ	1.340	36.414	1.357	41.786
industry and year dummies present				
number of observations		17121		17121
Log likelihood		-5169.3		-6424.3
Inaction (Collective Dismissal Criterion)				
Employment (log)	0.512	21.701	0.688	33.612
Fixed-term contracts (% of total)	1.674	12.409	1.712	13.965
Part-time workers (% of total)	0.257	0.854	0.938	3.858
From employment declining regime	-0.037	-0.663	0.051	1.116
λ	0.042	14.953	0.045	17.978
ρ	1.348	44.544	1.400	54.654
industry and year dummies present				
number of observations		18123		18123
Log likelihood		-6831.9		-8788.6

Table 3: Regression Model Estimates (Weibull distribution) .

	Employment Expansion		Employment Decline	
Inaction (Absolute Zero Threshold)				
	coefficient estimate	t statistic	coefficient estimate	t statistic
Employment (log)	0.650	34.640	1.263	30.332
Fixed-term contracts (% of total)	1.153	9.898	1.928	9.582
Part-time workers (% of total)	-0.560	-2.545	-0.509	-1.844
From employment declining regime	0.084	1.849	0.027	0.422
Constant	-3.717	-48.533	-4.175	-38.774
σ	0.036	1.000	1.989	20.355
ρ	1.480	54.684	2.564	27.695
industry and year dummies present	YES		YES	
number of observations		14710		14710
Log likelihood		-7944.4		-9986.34
Inaction (10 percent Threshold)				
Employment (log)	-0.340	-7.327	-0.227	-5.853
Fixed-term contracts (% of total)	2.654	7.848	3.242	10.193
Part-time workers (% of total)	0.039	0.084	1.250	2.728
From employment declining regime	0.758	5.817	-0.217	-2.078
Constant	-4.714	-27.045	-4.982	-28.957
σ	4.241	9.794	3.817	10.912
ρ	2.076	13.710	2.141	14.891
industry and year dummies present	YES		YES	
number of observations		17121		17121
Log likelihood		-5138.8		-6378.3
Inaction (Collective Dismissal Criterion)				
Employment (log)	0.768	14.280	0.874	23.784
Fixed-term contracts (% of total)	2.823	8.598	2.191	11.244
Part-time workers (% of total)	0.025	0.047	-0.002	-0.009
From employment declining regime	0.023	0.234	0.024	0.363
Constant	-4.487	-28.499	-4.595	-41.373
σ	3.780	10.251	2.019	12.702
ρ	2.230	13.227	1.877	24.501
industry and year dummies present	YES		YES	
number of observations		18123		18123
Log likelihood		-6787.63		-8752.34

Table 4: Regression Model Estimates (Weibull distribution with gamma heterogeneity) .

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