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## **JOBLESSNESS**

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# Joblessness

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## Abstract

The U.S. labor market has been experiencing unprecedented high average unemployment duration. The shift in the unemployment duration distribution can be traced back to the early nineties. In this paper, censored quantile regression methods are employed to analyze the changes in the US unemployment duration distribution. We explore the decomposition method proposed by Machado and Mata (2005) to disentangle the contribution of compositional *vis-à-vis* structural changes. The data used in this inquiry are taken from the nationally representative Displaced Worker Surveys of 1988 and 2008.

Apart from the effect of economic improvement we find that the sensitivity of joblessness duration to education and the aging of the population were the two main forces behind the increase of the unemployment duration, in the last twenty years. We tentatively argue that firms use education as a signaling device during recessions, but the signaling power of education during the recent low-unemployment environment faded significantly.

KEYWORDS: Quantile Regression, Duration Analysis, Unemployment Duration, Counterfactual Decomposition

JEL CODES: C14, C21, C41, J64

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# 1 Introduction

The U.S. labor market has changed significantly since 1985 up until the financial crisis.<sup>1</sup> Unemployment rates fell below 4 percent and it has been argued that the "natural rate of unemployment" shifted downward to 5 percent or below. This trend toward lower unemployment rates was largely driven by lower unemployment inflows. Concurrently, however, mean elapsed unemployment duration surprisingly trended up. Indeed, average unemployment duration reached 18 weeks in 2008. Figure 1 shows that the Current Population Survey (CPS) series of unemployment rates and median elapsed weeks of unemployment used to be very well-aligned until the end of the eighties. The two series began diverging significantly in the early nineties and the gap has widened ever since (see Figure 1). In a sense, the American job market resembles more the European.

The striking evolution of unemployment in the United States has not gone un-noticed. A number of studies have examined the question of why the unemployment duration became so much longer (Baumol and Wolff (1998); Valletta (1998); Abraham and Shimer (2001); Juhn et al. (2002); and, Mukoyama and Sahin (2009) ).

Explanations for the recent rising trend of average unemployment duration rely either on the compositional changes of the labor force or, more fundamentally, on the emergence of some economic mechanisms.<sup>2</sup> Examples of the former explanation include Abraham and Shimer (2001), who argue that the aging of the baby-boom generation and the increased labor force attachment of women contributed to the observed enlarged share of long-term unemployed; Juhn et al. (2002) who claim that joblessness among less-skilled men has taken the form of time spent out of the labor force rather than unemployment *per se*;<sup>3</sup> and Valletta (1998), who reports that the increase in average unemployment duration was produced by the joblessness experience of displaced workers.

Three main economic explanations have been offered for the observed lengthening of the average duration of unemployment. In the first uptake, Baumol and Wolff (1998) link average duration of unemployment to technical change, arguing that the acceleration of technical change has raised the share of the labor

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<sup>1</sup>In this paper we do not attempt to analyse the most recent developments in the US labor market, because the unemployment consequences of the financial crisis can not be taken as permanent (stable).

<sup>2</sup>The influence of methodological changes in the CPS surveys has also been studied (see, e.g., Abraham and Shimer (2001)).

<sup>3</sup>The relaxation of the Social Security Disability Insurance and Supplemental Security Income eligibility rules may also help to explain the increase in non-participation rates.

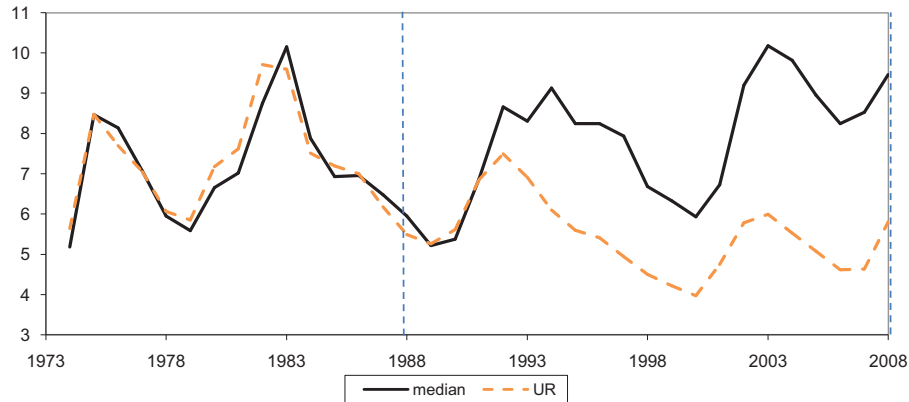


Figure 1: UNEMPLOYMENT RATE AND UNEMPLOYMENT DURATION.

force that is unemployed in any period because plants close more often. Second, Mukoyama and Sahin (2009) note that increased within-group wage inequality, which translates into higher uncertainty about wage offer distribution, is likely to lead to longer periods of job search. Finally, Juhn et al. (2002) maintain that long-term changes in joblessness have been produced by adverse shifts in labor demand.

In this paper, we rely on censored quantile regression methods to analyze the changes in the U.S. unemployment duration distribution. Quantiles seem appropriate to analyze unemployment duration for two main reasons. First, the methodology estimates the whole quantile process of duration time conditional on the attributes of interest, which constitutes a complete characterization of the distribution of duration time. Quantiles provide a natural way of characterizing important concepts such as short- or long-term unemployment, by focusing on the relevant tails of the duration distribution. Second, from a methodological standpoint, it is worth observing that quantile regression provides a unified and flexible framework for such an analysis.

Changes over time in the distribution of unemployment duration may be

framed as resulting from changes in the distribution of the conditioning variables such as the age distribution or from changes in the conditional distribution of duration itself. We use Machado and Mata (2005) method to disentangle those effects. The basic building block is the estimation of the conditional distribution by quantile regressions; then, by resorting to resampling procedures, one estimates marginal distributions consistent with the estimated conditional model as well as with hypothesized distributions for the covariates. Comparing the marginal distributions implied by alternative distributions for the covariates one is then able to perform counterfactual exercises that isolate the different effects contributing to the overall change.

The data used in this inquiry are taken from the nationally representative Displaced Worker Surveys of 1988 and 2008. The DWS is a retrospective survey that has been conducted biennially since 1984. It contains information on the nature of the job lost and the subsequent joblessness duration of workers displaced by reason of plant closure, slack work, or abolition of shift or position. The DWS is particularly well suited to study the distributional shape of unemployment duration because, unlike the CPS, it is a representative sample of the flow of displaced workers and because it provides information on completed spells of unemployment.<sup>4</sup>

The paper is organized as follows. Section 2 describes the data set, providing a careful comparison of the two Displaced Worker Surveys used. Section 3 outlines the econometric methodology. The basic regression results are presented in Section 4. Section 5 deploys the Machado and Mata decomposition to sort out the forces behind the changes in unemployment duration. Section 6 concludes.<sup>5</sup>

## 2 Data

### 2.1 General Description

The data used in this inquiry are taken from the nationally representative, Displaced Worker Supplement to the February 1988 and 2008 Current Population Survey. The dataset - and changes in the survey, including the wording of the core displacement question and the recall period over which information on job loss is recorded - are well described elsewhere (see, for example, Kletzer, 1998; Farber (1999), so that only brief introductory remarks are required here. The

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<sup>4</sup>Is is demonstrably harder to characterize the distribution of an unemployed population based on the stock rather than the flow of the unemployed persons (Lancaster, 1990).

<sup>5</sup>The econometric details are presented in the Appendix.

DWS has been conducted biennially since 1984. It contains information on the nature of the lost job and subsequent joblessness for workers displaced by reason of plant closure, slack work, or abolition of shift or position. Such data can be supplemented by extensive information on the personal characteristics of the worker contained in the parent CPS. The choice of the 1988 and 2008 surveys was guided by the need to use a comparable framework to the greatest extent. The 1988 DWS survey was the first to provide information for a single spell of joblessness (until 1986 the recorded jobless duration included multiple spells of joblessness). The 2008 survey is the most recent available survey with adequate data on joblessness duration. Still, there remain some issues of comparability that will be discussed below.

The DWS has a number of advantages over administrative data. First, unlike the unemployment registry, the DWS survey covers both recipients and non-recipients of unemployment benefits. Second, because it is retrospective, the information on unemployment duration is not censored at the time of the exhaustion of benefits. And, third, the DWS allows the identification of transitions of displaced workers to another job without any intervening spell of unemployment.

There are inevitably some shortcomings of the DWS data. Retrospective data are subject to recall bias - individuals experiencing displacement in past years may be more likely to understate their jobless duration than are more recent job losers - and respondents are prone to round (to months and quarters) their reported spells of unemployment. Beginning with the 1994 survey, however, the period over which job loss is measured has been reduced from five to three years, which should reduce the recall bias problem.

As mentioned above, since the 1988 survey the measure of unemployment has referred to the length of the single spell of joblessness that followed the displacement event and resulted in reemployment. To be sure, the definition still does not require the unemployed individual to be engaged in active search, so that this single spell may include intervals of suspended job search/withdrawal, but it no longer includes multiple spells of joblessness. A more recent innovation which affects the 2008 survey is that the DWS unemployment data are no longer top coded (at 99 weeks of joblessness). An additional source of right censoring in the data stems from our inclusion (via the CPS) of those individuals who failed to find work after displacement but who were nevertheless economically active as of the survey date.

Although we included those who wanted but never found employment after

losing their jobs - as well as those individuals who transitioned directly into reemployment without any intervening spell of joblessness - we excluded individuals who were not economically active at the time of the survey. Further, because the nature of displacement is not well defined for certain individuals and sectors, those employed part time and in agriculture at the point of displacement were also excluded, as were those aged less than 20 years and above 61 years. These restrictions yielded a sample of 2,837 individuals for 1988 and 2,199 for 2008.

## 2.2 Comparability of the DWS Surveys

There are a number of comparability issues that need to be tackled. First, and most importantly, whereas the 1988 survey is a five-year retrospective data set of displaced workers based on the question *"In the past five years, that is since January 1983, has ...lost or left a job because of a plant closing, an employer going out of business, a layoff from which...was not recalled, or other similar reason?"*, the 2008 survey is a three-year retrospective data set based on the question *"During the last three calendar years, that is, from January of 2005 through December of 2007, did (name/you) lose a job, or leave one because a plant or company closed or moved, (your/his/her) position or shift was abolished, insufficient work, or another similar reason?"*. If the response to the job loss core question was positive, the respondent was asked whether the reason for displacement was 1) plant closing, 2) slack work, 3) position shifted or abolished, 4) seasonal job ended, 5) self-employment failed, and 6) other reasons. In line with the CPS definition of job displacement, only the first three situations will be considered in this paper.

Even though the slight change of wording is unlikely to raise any major comparison problems, the reduction of the retrospective period is potentially more serious. Since there is information on the year of displacement of the worker, one can minimize this problem excising from the 1988 sample the individuals displaced in 1983 and 1984.<sup>6</sup> But this procedure does not completely solve the issue. If an individual experienced multiple spells of joblessness (which affects a fraction of displaced workers) the interviewer has instructions to record the episode where the worker lost the job with the longest duration. It may well occur that after losing a long-tenure job during 1983 or 1984 an individual was displaced again during the 1985-1987 period. In this case, this displacement

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<sup>6</sup>Displacements that occurred during January of 1988 were also excluded. The 2008 survey does not include, by construction, workers displaced in 2008.



from a short-duration job is not registered. There is a clear implication for distortion of the distribution of job duration, with short job durations being likely to be under represented in the 1988 survey in comparison with the 1988 survey. But there is no unambiguous implication for the distribution of unemployment duration.<sup>7</sup>

Second, even though unemployment rates were falling and labor market conditions were improving over the survey periods, the cyclical conditions were not identical. In fact, the average state unemployment rate at the time of displacement is 2.4 percentage points lower in the 2008 survey than the 1988 survey. We expect that by conditioning the unemployment duration distribution on the unemployment rate, we will be able to isolate the impact of the business cycle.

Third, in both surveys the displaced workers are asked whether they received advance notice of impending their lay-off, but in the 2008 survey this question is restricted to written notice, where in the 1988 survey the individuals distinguish between informal and written notice. In order to make this variable as comparable as possible we will consider a notified only those workers who received written notice at least two months before the date of displacement.

Apart from these three comparability issues, which can be partially overcome, we are convinced that the two DWS surveys provide an adequate framework for characterizing the the evolution of the unemployment experience of displaced workers throughout the period 1985 up to 2007.

### 3 Composition and Structure

The basic pieces of information to our counterfactual analysis are the changes in the attributes (covariates) of the jobless population and the changes in the distribution of duration for any given level of those attributes ("structure" or coefficient changes). The latter are estimated by censored quantile log-linear regressions (Koenker and Bassett (1978) and Powell (1984, 1986)).

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<sup>7</sup>Some checks can, however, be implemented. First, one can compare the job duration distribution for the 1983-1984 period with the 1985-1987 period. Second, one can exclude from both samples workers with fewer than two years of tenure in the pre-displacement job. And third, one can use our decomposition methodology to simulate the 2008 unemployment distribution with the 2008 job duration distribution. In all cases we arrive to the conclusion that the issue of multiple job spells does not significantly affect the comparison of the two unemployment duration distributions.

### 3.1 Covariates

Descriptive information on the two samples is provided in Table 1 and Figure 2. The composition of the 2008 sample differs from that of 1988 in some important ways.

	1985-1987	2005-2007
Age	36	41
Male	0,65	0,60
White	0,86	0,82
Married	0,60	0,52
Married female	0,17	0,18
Schooling (years)	12,5	13,2
Tenure (years)	4,54	4,81
Close	0,46	0,35
Written notice	0,05	0,10
Unemployment rate	7,06	4,67
Unemployment insurance	0,61	0,46
Proportion censored	0,17	0,19
Proportion duration is zero	0,12	0,12
Unemployment Duration (median in weeks)	8	8
Number observations	2496	1944
Total number observations	2837	2199

Table 1: SAMPLE DESCRIPTIVE STATISTICS

- The median unemployment duration is stable between the 1985-87 period and the 2005-07 period. This indication is best understood in the empirical survival functions (Kaplan-Meier estimates) exhibited in Figure 2. Although this leftward shift is noticeable at both tails of the joblessness distribution, upper quantiles increased relative to the mean unemployment rate, as pointed out by Abraham and Shimer (2001). This indication is stronger if one considers the conventional measure of unemployment duration, where direct transitions without an intervening unemployment spell are excluded.
- The proportion of direct job-to-job transitions (joblessness spells with duration equal to 0) did not change. In both periods these individuals involved in job-to-job transitions were not significantly different from the rest of the displaced group (see Tables 1).

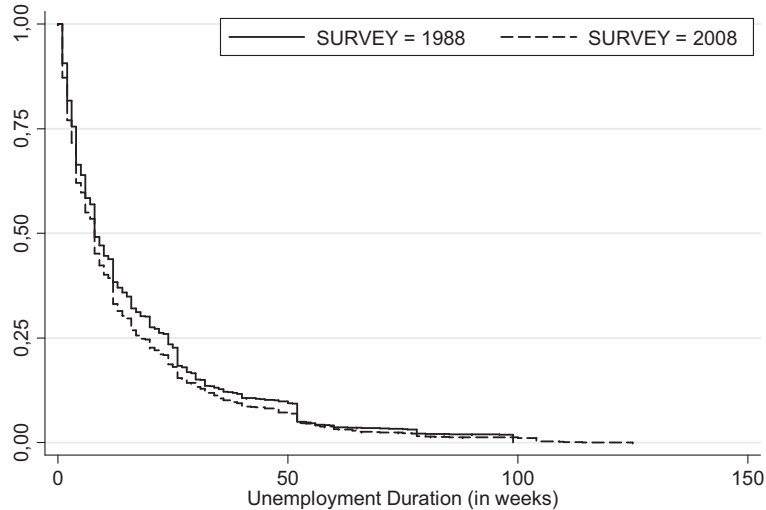


Figure 2: KAPLAN-MEIER SURVIVAL FUNCTIONS.

- Displaced workers in the latter survey are older and better educated than during the eighties, reflecting the aging of the baby-boom generation (see Figure 3).
- The proportion of female workers among displaced also increased sizably, probably because labor market participation rates of women increased and so did the risk of being displaced over the relevant period.
- The likelihood of receiving formal notice of job lay-off more than doubled in the 2008 survey, probably due to the introduction of the Worker Adjustment and Retraining Notification Act, which was enacted in 1988, which made pre-notification of displacements mandatory for mass-layoffs or shut-downs generated by large firms (Addison and Blackburn (1994)).
- Interestingly, despite the change in the reference period of job displacements (from five to three years), there are no significant changes in the distribution of job duration in the pre-displacement job (see Figure 3). It may still happen, however, that workers that are now displaced have longer tenure than before.

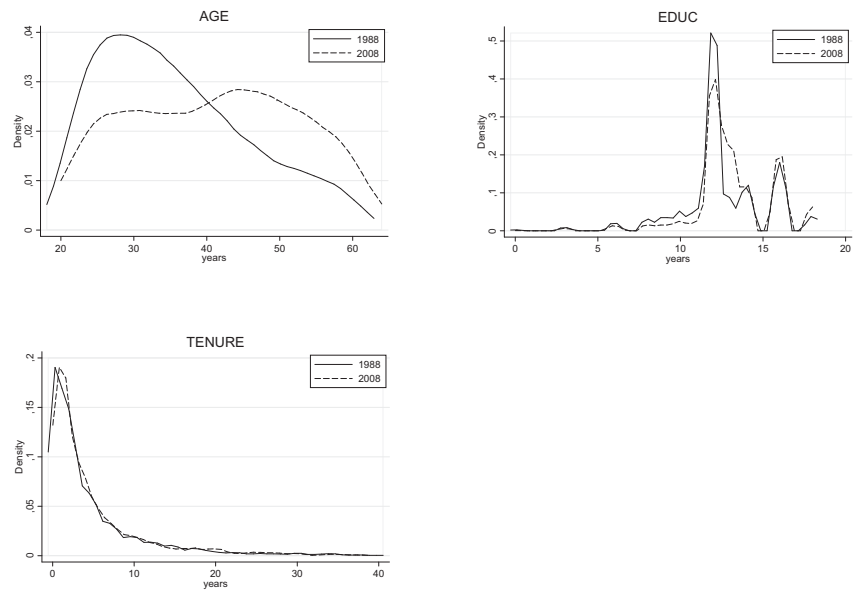


Figure 3: KERNEL DENSITIES FOR AGE, TENURE, SCHOOLING, AND THE UNEMPLOYMENT RATE.

In a nutshell, displaced workers are older, more educated and experienced and more likely to be female than before.

### 3.2 Coefficients

We characterize the conditional distributions of jobless duration by quantile regression (QR) models.

Empirical results for selected quantiles from fitting the QR model are given in Tables (2) and (3) for both surveys. Focusing on the 1985-1987 survey, the regression coefficient estimates are fairly conventional:

- Age reduces escape rates proxying the reduced arrival rate of job offers with age.
- The impact of Tenure is statistically significant only at high quantiles. Its effect may capture the elevated reservation wages of long-serving workers.
- The result for race is familiar and captures the poorer opportunities facing non-whites as a result of both objective and discriminatory factors.
- The familiar (opposing) effects of marital status on reemployment probabilities - positive for males and negative for females - are also obtained. The result for married males presumably picks up a household head effect, and thus likely reflects the higher opportunity cost of unemployment for married males and their greater search intensity.
- Schooling enhances the chances of getting a job, but much more so for low durations. It can be argued that larger human capital endowments are associated with greater job opportunities and higher opportunity costs of unemployment that necessarily erode with the progression of the unemployment spell. A number of explanations can be suggested here: human capital depreciation, unobserved individual heterogeneity correlated with the measures of human capital, or employers' stigmatization of long-term unemployed, would lead to a fading human capital effect on the transition rate out of unemployment.
- Like schooling, written pre-notification (defined as written notice of at least three months) and job loss by reason of plant closure have significantly higher effects during the early phase of the unemployment spell. This pattern reflects the influence of on-the-job search. Faced with the

	Quantile Regression					
	Q20	Q50		Q80		
Age	0,012	(0,006)	0,022	(0,005)*	0,014	(0,004)*
Male	0,031	(0,188)	0,335	(0,142)*	0,307	(0,116)*
White	-0,282	(0,176)	-0,269	(0,135)*	-0,377	(0,112)*
Married	-0,332	(0,155)*	-0,281	(0,119)*	-0,160	(0,098)
Married female	0,634	(0,249)*	0,818	(0,190)*	0,470	(0,154)*
Schooling	-0,122	(0,025)*	-0,050	(0,018)*	-0,047	(0,013)*
Tenure	-0,001	(0,011)	0,011	(0,008)	0,018	(0,006)*
Close	-0,730	(0,118)*	-0,433	(0,090)*	-0,182	(0,074)*
Written notice	-0,643	(0,262)*	0,234	(0,199)	-0,004	(0,161)
Constant	2,206	(0,418)*	2,065	(0,313)*	3,519	(0,253)*
Observations	2818		2674		2522	

Table 2: UNEMPLOYMENT DURATION (LOGS) REGRESSION RESULTS FOR 1985-1987. NOTE: 2.837 OBSERVATIONS.

prospect of an imminent discharge, the worker will engage in on-the-job search. If successful, he or she will experience a short spell of unemployment (Addison and Portugal (1992)).

- Identically, workers displaced by reason of plant closing — in comparison with workers dismissed due to slack work or position shifted or abolished — benefit from an essentially short-term advantage conveyed by job search assistance and early (and unmistakable) warning of displacement.

Despite broad qualitative agreement between the regression coefficient estimates from the two surveys, there are, nevertheless, some differences. For their magnitude and potential impact on the unemployment duration distribution (see section 4.3), two are most striking.

- First, the sharp decrease in the sensitivity of duration to education throughout the distribution. One may speculate that as displaced workers became more educated and experienced, the signaling power of education faded significantly.
- Second, the intercept also dropped sharply which reflects an overall shift to the left of the distribution of durations. The intercept will capture (among other things) all the time-varying common factors and, so, will certainly reflect the improved business cycle conditions in 2008.

Also worth noticing, but of limited quantitative impact, are the following

	Quantile Regression					
	Q20		Q50		Q80	
Age	0,018	(0,005)*	0,023	(0,004)*	0,018	(0,004)*
Male	-0,043	(0,161)	-0,051	(0,116)	0,036	(0,134)
White	-0,538	(0,146)*	-0,243	(0,106)*	-0,433	(0,122)*
Married	-0,135	(0,147)	-0,231	(0,106)*	-0,269	(0,122)*
Married female	0,351	(0,225)	0,208	(0,163)	0,292	(0,188)
Schooling	-0,048	(0,024)*	-0,015	(0,017)	-0,028	(0,018)
Tenure	0,004	(0,010)	0,030	(0,007)*	0,033	(0,008)*
Close	-0,425	(0,118)*	-0,283	(0,084)*	-0,176	(0,096)
Written notice	-1,203	(0,186)*	-0,204	(0,127)	-0,049	(0,142)
Constant	0,844	(0,410)*	1,437	(0,288)*	3,030	(0,328)*
Observations	2159		2102		1943	

Table 3: UNEMPLOYMENT DURATION (LOGS) REGRESSION RESULTS FOR 2005-07. NOTE: 2.199 OBSERVATIONS.

findings. The jobless distribution became independent of gender: the market treats female and male displaced workers similarly. (some reference would be nice here). Although the being displaced by plant closing still significantly reduces the spell duration, the effect is much more attenuated in 2008.

## 4 Changes in the unemployment duration distribution

### 4.1 An overall view

The law of total probability implies that changes over time in the distribution of unemployment duration may result from changes in the distribution of the conditioning variables (e.g., labor force characteristics such as the age distribution) or from changes in the conditional distribution of duration itself (which may be thought of as changes in the way those labor force characteristics impact duration, the “coefficients”). The first is a composition effect and the second may be thought of as a “structural effect” (as in Autor et al. (2008)). Machado and Mata (2005) proposed a method (hereafter, M&M decomposition) for disentangling those effects. The method is based on the estimation of marginal distribution of the variable of interest consistent with a conditional distribution estimated by quantile regression, as well as with any hypothesized distribution for the covariates. Comparing the marginal distributions implied by different distributions for the covariates one will then able to perform counterfactual ex-

ercises and identify the sources of the changes in the distribution of duration over the ten-year period (see Appendix for further details).

Between the 1988 and 2008 survey, the distribution of unemployment duration shifted to the left, most notably at higher percentiles. Whereas unemployment duration decreased by 1.2 weeks at the median it decreased by 4.2 weeks at the 8th decile (see the third column of table 4). It is clear from columns 4th and 5th that (aggregate) changes in the coefficients were more influential driving the overall displacement of the unemployment duration distribution than (aggregate) changes in the covariates. “Coefficient changes” are everywhere larger, in absolute magnitude, than “covariates changes”. Interestingly, whereas “coefficient changes” led to shorter durations above the median duration, “covariates changes” generated longer durations at the highest percentiles. At the 8th decile, unemployment duration increased by 3.1 weeks due to changes in covariates but decreased by 6.8 weeks due to changes in the coefficients.

	Marginals			Aggregate contributions		
	1988	2008	Change	Covariates	Coefficients	Residual
10 th quant.	0,293	0,251	-0,042	0,001	-0,145	-0,102
	0,266;0,320*	0,214;0,289*	-0,087;0,003	-0,038;0,041	-0,182;-0,107*	
20 th quant.	1,460	1,186	-0,274	0,126	-0,515	-0,116
	1,405;1,514*	1,141;1,231*	-0,342;-0,205*	0,068;0,183*	-0,542;-0,488*	
30 th quant.	2,945	2,494	-0,451	0,353	-0,949	-0,144
	2,881;3,010*	2,428;2,560*	-0,542;-0,360*	0,266;0,441*	-0,978;-0,919*	
40 th quant.	4,969	4,215	-0,753	0,646	-1,633	-0,234
	4,854;5,083*	4,122;4,309*	-0,897;-0,609*	0,521;0,771*	-1,690;-1,576*	
50 th quant.	7,727	6,523	-1,204	1,031	-2,488	-0,253
	7,574;7,880*	6,400;6,645*	-1,388;-1,020*	0,855;1,206*	-2,564;-2,412*	
60 th quant.	11,613	9,796	-1,817	1,571	-3,584	-0,196
	11,392;11,834*	9,628;9,964*	-2,077;-1,558*	1,342;1,800*	-3,713;-3,456*	
70 th quant.	17,519	14,433	-3,086	2,320	-5,442	-0,036
	17,188;17,849*	14,153;14,714*	-3,487;-2,684*	1,967;2,674*	-5,602;-5,282*	
80 th quant.	26,392	22,197	-4,195	3,137	-6,786	0,546
	26,039;26,745*	21,796;22,597*	-4,720;-3,670*	2,628;3,646*	-6,978;-6,594*	
90 th quant.	41,963	37,981	-3,983	3,815	-8,125	-0,327
	41,333;42,594*	37,291;38,670*	-4,980;-2,985*	2,813;4,818*	-8,498;-7,752*	

Table 4: CONTRIBUTIONS TO CHANGES IN THE QUANTILES OF THE UNEMPLOYMENT DISTRIBUTION (WEEKS). MEDIAN AND 95% INTERVAL ESTIMATES (IN WEEKS) OF THE CHANGES IN THE QUANTILES (2008“MINUS” 1988) OF THE MARGINAL AND OF THE COUNTERFACTUAL DISTRIBUTIONS (BASED ON 500 REPLICATIONS).

## 4.2 Composition Effects

As hinted above, the composition of the displaced workers (and the underlying economic environment) changed significantly between surveys: displaced workers became older and more educated; the proportion of females increased;



written pre-notification of impending lay-off became more common; and the macroeconomic conditions improved. Overall, these changes produced longer jobless durations for all percentiles. A finer analysis, one that would enable us pinpoint the most influential regressors, requires the estimation of the impact of each of those changes on the conditional distribution of durations.

Using the techniques described in Appendix it is possible to isolate the contribution of the changes in the distribution of each covariate to the changes in the distribution of durations of joblessness spells. As it turns out, solely one explanatory variable exhibit a statistically significant composition effect: age (see Table 5). The results displayed in table 5 are obtained from the difference between predicted duration under 2008 covariates and coefficients and predicted duration for 2008 covariates and 2008 coefficients, except the covariate under examination which will take its 1988 values.

The ageing of the population translated into longer durations most notably, for the long-term unemployed (that is, those in the right tail of the unemployment duration distribution). Here, we estimate that at the 90th quantile duration is 2.5 weeks (11%) longer in 2008 than it would have been if the age of the population had been distributed as in 1988.

	Age	Male	White	Married	Married female	Schooling	Tenure	Close	Written notice
20 th quant.	0,10	0,00	0,03	0,05	0,01	-0,03	0,01	0,07	-0,08
	0,087;0,121*	-0,006;0,007	0,017;0,046*	0,036;0,061*	-0,001;0,025	-0,040;-0,015*	0,003;0,022*	0,053;0,089*	-0,098;-0,063*
50 th quant.	0,68	-0,02	0,05	0,14	-0,06	-0,04	-0,02	0,22	-0,09
	0,612;0,742*	-0,039;-0,002*	0,012;0,081*	0,098;0,192*	-0,094;-0,021*	-0,060;-0,020*	-0,066;0,033	0,170;0,269*	-0,115;-0,066*
80 th quant.	2,49	0,03	0,35	0,43	0,08	-0,27	0,31	0,41	-0,19
	2,283;2,695*	-0,048;0,111	0,192;0,515*	0,254;0,604*	-0,059;0,215	-0,382;-0,162*	0,110;0,503*	0,264;0,564*	-0,265;-0,109*

Table 5: CONTRIBUTION OF SELECTED COVARIATES TO THE CHANGE IN THE QUANTILES OF THE UNEMPLOYMENT DISTRIBUTION. Median and 95% interval estimates (in weeks) of the changes in the quantiles (2008 “minus” 1988) of the marginal and of the counterfactual distributions (based on 500 replications).

### 4.3 Changes in the conditional duration

The major changes in the conditional distribution were a fall in the sensitivity of duration to the level of education of the displaced workers, the attenuation of the gender effect, and a sharp downturn in the intercept (see table 6). The values exhibited in table 6 are computed as the difference between estimated duration for the 2008 population and all 2008 coefficients, except the coefficient under scrutiny, which will take its 1988 value.

A one point percent increase in the male population generates a much larger unemployment duration decrease in the 2008 survey than in the 1988 survey. Indeed, if the male population regression coefficient of 1988 prevailed, the median unemployment duration would be 1.4 shorter (28%).

The increase in the tenure and the age coefficients implied an increase in median duration of 1 week and an increase at the 8th decile of 3.6 weeks (16%). It appears that, in the most recent displacement survey, being older translates into a even slower transition into employment than it was the case in 1988.

The fall in the education coefficient implied an increase in median duration of 2.7 weeks (54%). It appears that, in the most recent displacement survey, being more educated no longer translate into a faster transition into employment as it was the case in 1988. With some trepidation, we offer the tentative explanation that schooling is relatively more helpful in high unemployment than in low unemployment environments. It can be argued that under low unemployment regimes there is less heterogeneity among unemployed individuals (a higher proportion of truly unemployable workers), which will mean longer durations for a given (lower)unemployment rate (As predicted in our simple statistical model, below).

Using an argument similar to Blanchard and Diamond (1994) (footnote 6, page 423) being more educated is a weaker correlate of good quality when the unemployment is low.

	Constant	Age	Male	White	Married	Married female	Schooling	Tenure	Close	Written notice
20 th quant.	-4,03	-0,12	-0,01	-0,27	0,09	-0,07	0,59	0,06	0,15	-0,09
	-4,075;-3,994*	-0,144;-0,093*	-0,020;-0,003*	-0,294;-0,240*	0,071;0,100*	-0,085;-0,057*	0,564;0,612*	0,051;0,070*	0,135;0,165*	-0,101;-0,073*
50 th quant.	-8,24	0,50	-1,41	-0,10	0,08	-0,56	2,73	0,50	0,23	-0,19
	-8,347;-8,142*	0,466;0,528*	-1,466;-1,349*	-0,124;-0,079*	0,054;0,110*	-0,608;-0,515*	2,660;2,792*	0,462;0,534*	0,203;0,262*	-0,219;-0,158*
80 th quant.	-13,14	1,97	-3,52	0,17	-0,83	-1,29	6,44	1,67	0,06	-0,32
	-13,384;-12,891*	1,834;2,112*	-3,697;-3,343*	0,074;0,259*	-0,935;-0,718*	-1,430;-1,144*	6,299;6,588*	1,541;1,805*	-0,002;0,124	-0,401;-0,238*

Table 6: IMPACT ON DURATION (IN WEEKS) OF CHANGES IN QR COEFFICIENTS.

## 5 Illustrating the results in a simple mixture model

Suppose that job-offers arise as a Poisson process with rate  $\lambda$ , and that there are two types of workers,  $A$  and  $B$ , with

$$\lambda_A > \lambda_B$$

The proportion of types  $A$  in the unemployment stock at  $t$  is denoted by  $p(t)$ . If all job offers are taken, the unemployment duration survivor function at  $t$  is

$$S(t) = p(t) \exp\{-\lambda_A t\} + (1 - p(t)) \exp\{-\lambda_B t\}$$

It may be instructive to learn how in such a simple model one can generate the global patterns highlighted by the empirical analysis. Our empirical model identified two chief culprits:

- A composition effect: the ageing of the jobless population (and consequently increasing experience and schooling);
- A structural effect: the reduced sensitivity of unemployment duration to education.

The ageing of the baby boomers may be captured by a decrease in  $p$ , the proportion of individuals with higher exit rates.

$$\frac{d(1 - S(t))}{dp(t)} = \exp\{-\lambda_B t\} - \exp\{-\lambda_A t\} > 0.$$

Therefore,

$$\frac{dQ(\tau)}{dp} < 0,$$

That is, a decrease in  $p$  would increase the quantiles and, in particular the median duration. How does this impact vary over the distribution?

$$\frac{d^2(1 - S(t))}{dt dp(t)} = f_B(t) - f_A(t)$$

where  $f(t)$  denotes the (exponential) p.d.f. of the two subpopulations. Therefore, there is a value of  $t$  ( $t^* = \ln(\lambda_B/\lambda_A)/(\lambda_B - \lambda_A)$ ), such that,

$$\frac{d^2(1 - S(t))}{dt dp(t)} < 0, \text{ for } t < t^* \text{ and } \frac{d^2(1 - S(t))}{dt dp(t)} > 0, \text{ for } t > t^*.$$

The impact of changes in  $p$  on duration quantiles is thus predicted to be U-shaped. For  $1/\lambda_B = 16$  weeks and  $1/\lambda_A = 2$  weeks,  $t^* \approx 2$  weeks. So in the range that QR can estimate it is natural to find an increasing effect.

In this exceedingly simple framework, the structural shock identified by the empirical analysis must be modeled by a reduction in the arriving rates of job offers, namely of  $\lambda_A$  (identifying A as the group with more schooling). Blanchard and Diamond (1994) argue that the exit rate from unemployment would be a decreasing function of unemployment duration. According to their "ranking assumption" (firms prefer to hire individuals that are unemployed for the least time), it is natural to infer that unemployment duration erodes the role of education as a signal of (unobserved) worker quality. In this story, groups A and B will become more similar. The impact of such a change is

$$dS(t) = -t[\theta(t)d\lambda_A + (1 - \theta(t))d\lambda_B]S(t)$$

where

$$\theta(t) = p(t)S_A(t)/S(t)$$

Thus, if  $d\lambda_A < 0$  and/or  $d\lambda_B < 0$ ,  $dS(t) > 0$ , and, consequently, the duration quantile function will shift to the right.

## 6 Conclusions

The starting point of this paper was the evidence that measured unemployment duration in the U.S. increased substantially relative to unemployment rates. Here, the decomposition method proposed by Machado and Mata (2005) was employed in order to disentangle the contribution of the changes generated by covariates distribution and the conditional distribution. The estimation indicates that structural changes in the labor market played a pivotal role.

Composition effects related to age (but not gender) played a significant role. But, apart from this rather mechanical impact, important structural changes, captured in the changes of the regression coefficients, were at play. We have identified a major force reshaping the unemployment duration distribution: the change in sensitivity to education, increased the median unemployment duration 2.7 weeks.

We tentatively argue that the signaling power of education during the recent low-unemployment environment faded significantly. When the unemployment rate is low, the information that is passed to the employer through the education signal does not promote more job offers to the more educated unemployed.

Finally, a note of caution is in order. These results rely solely on the joblessness experience of displaced workers and may not apply to other unemployment experiences, for example, the unemployment experience of job market incomers and re-entrants or job quitters.

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## Appendix: Econometric methodology

### Censored quantile regressions

Let  $T_i$  represents the duration of the “most representative” unemployment spell of individual  $i$  and  $x_i$  ( $x_{1i} \equiv 1$ ) be the vector of covariates for the  $i$ th observation. We consider statistical models specifying , the  $p$ th ( $p \in (0, 1)$ ) quantile of  $T$  as

$$Q_{y(T)}(p|x) = x'\beta(p) \quad (1)$$

where  $y(\cdot) \equiv \log$  and  $\beta(p)$  is a vector of QR parameters, varying from quantile to quantile.

Our sample provides information on complete unemployment durations, but there are some incomplete spells (right-censoring). Moreover, to avoid problems with taking logs of very short spells (0 or close to 0 weeks) we, arbitrarily, censored durations inferior to 0.25 at 0.25 weeks. The sample information we consider may thus be represented by  $(y_i^*, x_i)$ ,  $i = 1, \dots, n$  where  $y_i^* = \min[\max(y_i, l), u_i]$ ,  $u_i$  denotes the upper threshold for  $y_i$  and  $l$  the left-censoring point ( $l = \log(0.25)$ ). When observation  $i$  is not censored  $u_i$  was taken to be the potential censoring duration (for instance, for a spell of six weeks starting in March 2007,  $u_i$  was 44 weeks). The QR estimator minimizes the sample objective function

$$\sum_{i=1}^n \rho_p(y_i - \min[u_i, \max(x_i'b, l)])$$

with,

$$\rho_p(z) = \begin{cases} pz & \text{for } z \geq 0 \\ (p-1)z & \text{for } z < 0, \end{cases}$$

(Powell (1984, 1986)). Estimation was performed iteratively using Buchinsky (1994) ILPA procedure with the modification suggested by Fitzenberger (1997). The censored quantile algorithm is programmed in STATA as an adofile. If you are interested in using it yourself, the genetic algorithm is available upon request. For the estimation of standard errors for the individual coefficients we resort to the bootstrap. Since the “errors” from the QR equation are not necessarily homogeneously distributed, to achieve robustness we resample  $(y, x, l, u)$  following the method of Billias et al. (2000).

Due to censoring, it may not be possible to identify the whole quantile process. Let  $(p_l, p_u)$  represent the range of quantiles that can be consistently estimated. Technically, any  $p$  in that range must be such that

$$M_n(p) = E\left\{\frac{1}{n} \sum_{i=1}^n I(l + \xi < x'_i \beta(p) < u_i - \xi) x_i x'_i\right\}$$

is uniformly positive definite in  $n$  for some  $\xi > 0$  (Fitzenberger (1997), Theorem 2.1).

## Machado and Mata decomposition

The conditional quantile process – i.e.,  $Q_y(p | x)$  as a function of  $p \in (0, 1)$  – provides a full characterization of the conditional unemployment duration in much the same way as ordinary sample quantiles characterize a marginal distribution. The resampling procedures proposed in Machado and Mata (2005) (henceforth, M&M) provide an easy way of simulating a random sample,  $\{T_i^*, i = 1, \dots, m\}$ , from a conditional distribution of duration times that is consistent with the restrictions imposed on the conditional quantiles by the QR model. For completeness we outline here the procedure:

1. Generate  $m$  random draws from a Uniform distribution on  $(p_l, p_u)$ ,  $\pi_i$ ,  $i = 1, \dots, m$ ;
2. For each  $\pi_i$  estimate the QR model (1), thereby obtaining  $m$  vectors  $\hat{\beta}(\pi_i)$ ;
3. For a given value of the covariates,  $x_0$ ,

$$T_i^* \equiv \hat{Q}_T(\pi_i | x_0) = g(x'_0 \hat{\beta}(\pi_i)) \quad i = 1, \dots, m,$$

is a random sample from the estimated conditional c.d.f.  $F_T(t | X = x_0)$  censored at  $p_l$  and  $p_u$ .

The sample generated by the procedure above is drawn from the conditional distribution. In many instances it is important to integrate out the conditioning covariates. This integration or marginalization can be performed with respect to different joint distributions,  $g(x)$ , of the covariates. The approach in M&M may be described as follows:

1. As described before, generate  $\pi_i$ ,  $i = 1, \dots, m$  and estimate the corresponding  $\hat{\beta}(\pi_i)$ ;
2. Generate a random sample of size  $m$  from a given  $g(x)$ ; let it be denoted by  $\{x_i^*\}$ ,  $i = 1, \dots, m$ .

### 3. Obtain

$$T_i^* \equiv \hat{Q}_T(\pi_i | x_i^*) = g(x_i^{*\prime} \hat{\beta}(\pi_i)),$$

which is a random sample from the marginal distributions of durations times implied by the model postulated for the quantile process and by the assumed joint distribution of the covariates.

When  $g(x)$  is an estimate of the actual distribution of the covariates in the population, the resulting sample of durations is drawn from the actual marginal distribution. In this case,  $\{x_i^*\}$  may be obtained by drawing with replacement from the rows of  $\mathcal{X}$ , the regressors' data matrix. But, in reality,  $g(x)$  may be any distribution of interest. If it is an estimate of the distribution of the covariates in 1988 ( $g(x(1988))$ ), the resulting durations will constitute a simulated sample from the marginal distribution of durations that would have prevailed in 2008 if all covariates had been distributed as in 1988, (assuming, of course, that the  $\beta$  vector was estimated with 2008 data).

Comparing this counterfactual sample with samples of durations from the actual marginals for 2008 and 1988, it is possible to derive Oaxaca type decompositions for the entire distribution, rather than for just its mean. Specifically, it is possible to decompose the observed changes in those due to changes in the conditional distribution of durations (the  $\beta$ 's) and those stemming from changes in the joint distribution of the covariates. Other decompositions of interest often involve isolating the contribution of a single covariate. For further details on how to implement this decomposition, see M&M.)

In the implementation of the method in this paper we made  $p_l = 0.10$  and  $p_u = 0.95$  and estimated the quantile regression coefficients at equally spaced intervals of length 0.005. We then draw 1000 ( $= m$ ) of such estimates with replacement. A code in STATA with the whole procedure is available on request.

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