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The Duration of U.S. Joblessness and the Great Recession

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The Duration of U.S. Joblessness and the Great Recession

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ABSTRACT

We develop models of duration of unemployment for the United States and apply them to the U.S. 1996-2020 biennial Displaced Worker Survey (DWS). Our approach has several notable features relevant to an analysis of unemployment during a deep, persistent recession. We estimate single-spell proportional hazard exponential and Weibull models of transitional probability to full employment, and we test the outcome for robustness to functional forms by distributional free estimation. We test for the impact of unobserved heterogeneity by modeling recession as a Gamma distributed unobservable frailty function commonly employed in biostatistics. We also estimate a non/semi-parametric proportional hazard Cox model of competing risks with three destinations out of unemployment: transition to full-time employment, part-time employment, and inactive job-search status. Our single-destination estimates demonstrate significant differences in the covariate estimates as a result of different base hazards defined for those during the Great Recession and those outside that period. The estimates show gender is the single most important determinant of the duration of unemployment with men's hazard rate of (re)employment to full-time jobs 7% to 15% lower. When interacted with Unemployment Insurance, being female also accounts for a 10% higher hazard of transition to full-time employment. Our competing risks regression results indicate that Black workers have 28% and 27% lower rates of exit to a full-time and part-time employment respectively following job displacement.

Keywords: duration of unemployment, Great Recession, hazard function, multiple transitions

JEL Codes: C1, C41, J2, J6

1. Introduction

Understanding how much time individuals spend in unemployment, how this duration changes over the business cycle and how the length of unemployment vary across individuals are important aspects of job loss because the welfare of the unemployed is closely related to the time they spend without a job. Moreover, the length of unemployment spells plays a critical role in economic theories of job search. Three aspects on the duration of unemployment stand out in most unemployment survival studies. First, with regard to provision of social security programs such as unemployment benefits, for example Katz (1986), second, with regard to gender as the main determinant of exit from the labor market, for example Mussida (2007), and the third on the impact of recession on the duration of unemployment, for instance Pissarides (2013). In this study we develop models of duration of unemployment for the United States covering the Great Recession period that addresses the above aspects of unemployment with several notable features relevant to an analysis of unemployment during a deep, persistent recession. We estimate probability of exit to full employment by first examining two models that allow the exit time to be independent of and dependent on the unemployment spell; and then test the outcome for impact of unobserved heterogeneity from effects such as periods of expansion and contraction, and robustness to HP parametric functional forms. We also estimate a multiple-exit model of unemployment duration with distinct exit destinations out of unemployment for full-time, and part-time employment and discouraged-worker status by a non/semiparametric competing risks model. We focus on gender and race of workers as the principal determinant of difference in return to employment and how that is impacted by unemployment insurance and apply the above approach to the 1996-2020 biennial Displaced Worker Survey (DWS).

2. Literature Review

Studies on unemployment duration are concerned with factors that affect an individual's probability of leaving unemployment, and these can be categorized by a few areas frequently highlighted in the literature. An understanding of these factors has widespread implications for designing unemployment public policies.

One area studied extensively is the impact of gender on the duration of unemployment. Until the 1990, in OECD countries men were expected to have lengthier unemployment spells then women, but the difference disappeared for subsequent periods (Albanesi and Sahin, 2018). Studies focusing on the 2000s found that in some countries, women are likely to have a lengthier unemployment spell than men. (Azmat, Guell and Manning, 2006, Tansel & Taşçi (2010), Mussida (2007). The common explanations to this pattern are a high reservation rate for women (possibly due to household activities) and less job opportunities. Tansel & Taşçi (2010), studying the Turkish economy, found that the probability of being unemployed after 12 months was 90% for women and only 70% for men. In particular, gender appears to be an important variable determining the exit from the labor force (Mussida, 2007). The question is which factors lengthen the searching period, and which do not? In this respect, family structures are often considered a determinant factor. The literature has found meaningful and positive relationships for marital status and the children's age (Ollikainen, 2003, Mussida, 2007). Personal characteristics, such as age, citizenship and education, were also found to be relevant. Foreign citizenship and age increase the unemployment spell (Ollikainen 2003) and (Mussada 2007); while level of education is likely to decrease it (Ollikainen, 2003). The impact of family composition was more recently outlined in Kuhlenkasper & Steinhardt (2011). The results show that the likelihood of returning to employment for German women is reduced by the presence of young children and older relatives in the household. According to Imbens & Lynch (2006) the number of children living at home has a negative effect on the chances of women finding a job. Tansel & Taşçi (2010) found that married men have higher opportunity cost of unemployment and thus search more intensively.

Another topic often considered is the impact of education on the unemployment spell (Landmesser (2011); Tansel & Taşçi (2010); Borsic & Kavkler (2009) and Babucea & Danacica (2007)). A positive impact of education on unemployment spell is not self-evident, because while highly educated workers may find job in an increasingly technology-based economy easily, low-skill jobs are usually associated with a high rate of turnover (and thus a smaller unemployment period). According to Borsic & Kavkler (2009) persons with professional college degrees or bachelor’s degrees are better off than unemployed persons with a master’s degree. Babucea & Danacica (2007) suggest a negative relationship between unemployment and education. Education can matter not only in level, but in quality. In other words, the *kind* of education, whether technical, theoretical or vocational, also affects the likelihood of finding job after unemployment. According to Landmesser (2011) vocational training is important in exiting unemployment. Tansel & Taşçi (2010) found that vocational high school graduates have higher unemployment exit probabilities compared to high school graduates.

Furthermore, one of the mostly widely studied determinant of unemployment spell is the effect of social security programs such as unemployment benefits on the job searching process (Katz, 1986); (Heath & Swann, 1999). According to Nickell (1979a) the impact of income replacement ratio (ratio of income to benefits) was lower for those in long-term unemployment, while high replacement ratio was associated with longer duration of unemployment (Lancaster, 1979). Grubb (2011) and Barro (2010) find a strong and considerable influence of UI benefits in increasing the unemployment rate. Contrary to results from Grubb and Barro, Rothstein finds that UI extensions had significant but small negative effects on the probability that the eligible unemployed would exit unemployment, the effects concentrated on the long-term unemployed. The estimates imply that UI benefit extensions raised the unemployment rate in early 2011 by only about 0.1-0.5 percentage points, much less than is implied by earlier analyses. The same small but significance influence of extended benefits was found by Farber and Valletta (2013) while studying the effects of the extended unemployment benefits on unemployment duration during the Great Recession. The small influence found is attributable to the recipients exiting the labor force, and not to reduction of job finding. However, they argue that this may have long-term consequences for the natural rate of unemployment.

Finally, the literature has also considered the effects of recession periods in shaping the unemployment spell. In particular, researches have paid a great attention to the financial crisis of 2008-9 and its aftermath, also known as the Great Recession (GR). The reason is that when compared to other economic recessions, the Great Recession stand out as having the deepest and most widespread impact on the labor market (Pissarides, 2013). Recession periods, in fact, can change not only unemployment rates, but also patterns of unemployment. In other words, it raises the question of how different factors conditioning the length of joblessness spell vary during a recession period. Focusing on heterogenous effects of gender, Sahin et al (2010) investigate the specific effect of recession on the gender gap in the unemployment rate. They find that women were hit less hard – thus narrowing the gender gap – mainly because men occupy most jobs in highly procyclical sectors, and because men continued to search for jobs instead of quitting the labor force. Several studies have shown that the gender gap decreased during the recession. De la Rica and Rebollo-Sanz (2017), studying the Spanish labor market, find that the recovery-brought jobs were more equitably distributed. Kroft and al. (2014) study the effects of the Great Recession on long term unemployment. Importantly, the authors find that the exit from unemployment is less likely after longer periods of unemployment. Barnichon and Figura (2013) used a counter-factual analysis to address the effects of recession: they fitted the model for a long period before the recession (1967-2012) and used the calibrated parameters to fit the recession data. The discrepancies that arose were attributed to the recession effect.

In this study we develop models of duration of unemployment for the United States covering the Great Recession period in order to examine the factors highlighted in the literature that have several notable features relevant to an analysis of unemployment during a deep, persistent recession. We estimate probability of exit to full employment by first examining two models that allow the exit time to be independent of and dependent on the unemployment spell; and then test the outcome for impact of unobserved heterogeneity from effects such as periods of expansion and contraction. We employ a frailty Gamma function borrowed from biostatistical treatment of the unobservable. We also re-estimate our regression results by a Cox non-semiparametric hazard specification, and interpret these as two alternative methods of accounting for the impact of the recession on slope estimates, see below. Since evidence suggest inactive job seekers, as opposed to those out of the labor market, account for a large share of the unemployment, we also estimate a multiple-exit model of unemployment duration with three distinct exit destinations out of unemployment for full-time, and part-time employment and discouraged-worker status by a non/semiparametric competing risks model. Following the empirical evidence from the literature, we focus on gender of workers as the principal determinant of difference in return to employment and apply the above approach to the 1996-2020 biennial Displaced Worker Survey (DWS) covering the Great Recession period, and published as a supplement to the January Current Population Survey (CPS). However, the multiple exit results are not yet fully examined for recession frailty with alternative modeling methods and must be treated with caution.

More specifically, we contribute to the duration of unemployment literature by providing an intersectional analysis of unemployment duration, in which we examine the differences in duration of unemployment and exit probabilities across groups of White men, White women, non-white men and non-white women. Our intersectional analysis reveals a more nuanced picture of the duration of unemployment. Intersectional analysis also allows us to capture a clearer picture of the racialized impact of unemployment during a period of severe contraction of the U.S economy. In this sense our contributions are four-fold: 1. To investigate the dependency of the prospect of getting a job in the U.S. economy conditional on time spent unemployed. 2. To determine how the probabilities of exiting unemployment was affected by the Great Recession. 3.To determine the distribution of multiple transitional routes out of employment duration. 4.To determine the extent to which socio-demographics covaries influence the length of stay in unemployment.

3. Data

The data for this paper are obtained from the 1996-2020 biennial Displaced Worker Survey (DWS), published as a supplement to the January Current Population Survey (CPS). All participants older than 20 years of age are asked whether they lost a job within the three years prior to the survey date. Those responding positively to the question are then asked a series of questions about the lost job and the period of subsequent unemployment.

This data will allow us to analyze a variety of labor market outcomes following displacement, including earning losses, the duration of unemployment and unemployment benefit claiming behavior. Displacement is defined as involuntary separation based on operating decisions of the employer. Such events as a plant closing, an employer going out of business, or a layoff from which the worker was not recalled.

Calculation of the *disregard* element of wage, that is earning after adjustment for taxation purposes for lost earning due to periods of unemployment, is an important component of wage in the data sets. However, there is no uniform set of rules for disregard earning as the conditions are determined at the state-levels, and the time-consuming task of obtaining the separate amounts for each state proved beyond the limits of this study.

4. Econometric Models & their Specifications

Hazard Models

There are two main approaches in estimating the impact of exogenous variables on the duration of unemployment spells and the probability of reemployment. One is the structural form framework and the other the reduced form framework (Narendranathan & Nickell 1986). In this paper a reduced form framework is used, due to its simplicity and the fact that it permits us to estimate the impact of exogenous variables without imposing too many distributional assumptions. The reduced form framework is concerned with the specification and estimation of the conditional probability of leaving unemployment. Survival analysis is used to model unemployment duration and to assess its dependence on explanatory variables. Search theory is then adopted to choose the covariates that must be included in the model and to interpret the estimated coefficients.

Denote the duration of an unemployment spell, the time at which unemployment comes to an end, by $T \geq 0$ and its cumulative distribution function among the population by $\boldsymbol{F(t)}$ where t denotes a particular value of T. We define the survivor function $\boldsymbol{S(.)}$as the complement function to the *cdf*, i.e. periods of time taken to come out of the initial state, that is

$$S(t)=P[T>t]=1 - F(t)),\ t \geq 0 \qquad (1)$$

(1) is the probability of surviving unemployment status past time t. Then with a change in time $\boldsymbol{t+h}$ for a "small" $\boldsymbol{h}>0$, the probability of exiting the initial state in the time interval $\boldsymbol{[t, t+h)}$: is given by

$$P\ (t \leq T < t+h | T \ \geq t) \qquad (2)$$

For each $\boldsymbol{t}$, we define the **hazard function** $\boldsymbol{\lambda}(t)$ as the probability of instantaneous rate of exit per unit of time; given a small change in $\boldsymbol{t}$ by $\boldsymbol{h}$. The hazard function provides an approximation for the conditional probability (2) by

$$P\ (t \leq T < t+h \mid T \ \geq \ t) \approx \lambda(t)h \qquad (3)$$

Comparing Single Risk models: PH Exponential & Weibull:

The central issue in survival analysis is the shape of the hazard function., and the role of (3) in survival or duration analysis is to defined by a density function for the distribution for t. In this study we examine three such functions. The first two often provide benchmarks for comparison with other hazard distributions. The simplest is the exponential hazard function because it has a constant duration parameter, that is

$$\boldsymbol{\lambda(t) = \lambda}\ \text{, for all}\quad \boldsymbol{t \geq o} \qquad (4)$$

(4) makes the hazard function memoryless; the probability of exit in the duration period does not depend on how much time has been spent in the initial state. Hence, (1) becomes

$$S(t) = exp\left(-\int_0^t \lambda(t)dt\right) = exp(-\lambda(t)) \quad (5)$$

The exponential function is a one-parameter distribution, and that makes it too restrictive in many applications. We also employ the Weibull distribution as a second, more flexible and duration dependent alternative, based on the generalization of the exponential distribution. If $\boldsymbol{T}$ has a Weibull distribution, its *cdf* is presented by $\boldsymbol{F(t)=1- exp(-\gamma t^{\alpha})}$; its density is given by:

$$\boldsymbol{f(t) = \gamma\alpha t^{\alpha-1} exp(-\gamma t^{\alpha})}$$

, and its hazard function, by

$$\boldsymbol{\lambda(t) = \frac{f(t)}{S(t)} = \gamma\alpha t^{\alpha-1}} \quad (6)$$

The values of the nonnegative parameters $\boldsymbol{\gamma}$ and $\boldsymbol{\alpha}$ determine the exact shape of the Weibull distribution. The Weibull distribution reduces to the exponential distribution as its special case when $\boldsymbol{\alpha}=1$, demonstrating the Weibull as a more flexible generalization of the exponential. If we know the hazard to be increasing ($\boldsymbol{\alpha}>1$) or decreasing ($\boldsymbol{\alpha}<1$), then the Weibull distribution is suitable for capturing duration dependence.

The models specify the conditional hazard rate as the product of two separate functions:

$$\lambda(t|\mathbf{x}) = \lambda_0(t,\alpha)\ \varphi(\mathbf{x}\beta) \quad (7)$$

where the hazard rate, $\boldsymbol{\lambda_0(t,\alpha)}$, is a function of $\boldsymbol{t}$ alone, while $\varphi(\mathbf{x}\beta)$ is a function of $\mathbf{x}$ alone, and usually is specified by $\varphi(x\beta)=\exp(\mathbf{x}'\beta)$. Hence, both the exponential and Weibull distributions can be defined as the proportional hazard (PH) models since their hazards are, respectively, $\exp(\mathbf{x}'\beta)$ and $\exp(\mathbf{x}'\beta)\alpha t^{\alpha-1}$. We note that since the baseline is scaled by the vector of covariates, if all covariates are zero, then $\boldsymbol{\lambda_0(t,\alpha)\beta_0}$ where $\boldsymbol{\beta_0}$; the intercept constant, becomes a part of the baseline.

Recession and Functional Form Effects

Estimation based on exponential and Weibull densities are vulnerable to misspecification errors for two different reasons. First, we need to account for the impact of the various sources of unobservable heterogeneity on the estimation results, most importantly the impact of the recession and expansion periods on employment. Second, any functional form misspecification can also lead to inconsistent estimates. We next address each one of these two issues in turn.

To address heterogeneity, we employ a mixture survival model that has unobserved heterogeneity explicitly entered into the hazard function. The mixture function is based on the following assumptions: *a.* the heterogeneity is independent of the observed covariates, starting and censoring times, *b.* the distribution of heterogeneity variable is known; *c.* heterogeneity enters the hazard function multiplicatively. We employ a Weibull hazard function conditional on observed covariates $\mathbf{x_i}$ and an unobserved heterogeneity $\boldsymbol{v_i}$ by

$$\lambda(t|x_i, v_i) = v_i exp(x_i\beta)\alpha t^{\alpha-1} \tag{8}$$

where $\boldsymbol{x_{i1}}\equiv 1$ for the intercept, and the multiplicative hazard is $\boldsymbol{v_i} > 0$; the identification of $\boldsymbol{\beta}$ & $\boldsymbol{\alpha}$ requires conditioning on normalization for the distribution of frailty $\boldsymbol{v_i}$, the applications usually adopt a *gamma*-distributed heterogeneity function, $\boldsymbol{v_i}\sim$ *gamma* (δ, δ), and $E(\boldsymbol{v_i})$=1 and Var=1/δ. This method is one way to capture the unobservable effects of a recessionary period on estimations obtained, and is commonly employed in biostatistics, for example Hongaard (1986). In practice, it may be hard to distinguish between duration dependence parameter α and heterogeneity $\boldsymbol{v_i}$ without repeated spell observations. However, if the test of Ho**:** α=1 is confirmed, then that outcome can remove the effects of Weibull flexibility parameter and hence isolate that of heterogeneity, particularly that for a recession, on the duration function. But the assumption of α=1 may not be confirmed, and in any case unobserved heterogeneity is still vulnerable to functional form misspecification error. Hence, we also employ an alternative distribution-free method to measure the recession effect.

The PH models are known to provide consistent estimate of regression parameter vector by non-semiparametric methods without specification of the functional form for λ_0. We employ a PH Cox semi-parametric survival function that offers robustness against (4) and (6) to order to estimate β, that unlike HB (7) and (8) models without estimating the base hazard at the same time. The Cox model is semi-parametric in that the hazard rate λ_0 drops out of the estimation of β as a consequence of the HP assumption, though λ_0 estimate can be recovered once β is estimated. Hence, the Cox model with the exponential covariant specification leaves λ_0 in (7) unspecified, now denoted as $h_o(t)$

$$\lambda(t|\mathbf{x}) = h_0(t)\ \exp(\mathbf{x}\beta) \tag{9}$$

Specifically, the estimates by this model are presented in Table 2 below. The Cox proportional hazards regression model states that the survival function for the *i*th person in the data is estimated by:

β: a vector of unknown parameters to be estimated from the data,

$h_0(t)$: the baseline hazard function at time t, that is hazard function when all predictors are equal to zero,

x: an independent vector of predictor variables.

The list can be found in Table 1 below.

Note that the effect of recession in both methods is defined with reference to the base hazard function, multiplicatively for unobserved heterogeneity; unspecified and removed semi-parametrically from the regression equation based only on the individual covariates. That may still leave unaccounted the direct recession effects on the covariates. If there are doubts regarding the presence of such direct effects on the covariates, we can also move a covariate of interest inside the hazard function, then re-estimate the model separately in terms of that covariate. That is, if deemed necessary we can define the hazard function to have the same shape for all individuals, but separately estimated in terms of a critical variable, e.g. by gender or race. This method, however, becomes unyielding with more than one variable if several separate questions have to be estimated for several different hazard functions.

Estimation of the duration models is complicated by the presence of right-censoring of the length of a possible incomplete spells. The survival model data are typically right-censored from above. With right-censored data, we observe spells from time 0 until a censoring time c; some spells are completed by the time c, while all we know for the other spells is that they will end at some time in the interval (c, ∞), since in such cases we only know that the duration exceeded t. Hence, in application, we need to specify a hazard functions that accounts for right-censoring. For the hazard function (7), the uncensored contribution to the log *ML* function and the Weibull survival function lead to the Weibull distributed log *ML* as

$$\ln \mathrm{L} = \sum_{i=1}^{N} [\delta_i \{\mathrm{x}_\mathrm{i}' \beta + \ln \alpha + (\alpha - 1) \ln t_i \exp(x_i' \beta) t_i^{\alpha}\} - (1 - \delta_i) \exp\,(\mathrm{x}' \beta) t_i^{\alpha}] \quad (10)$$

The differentiation of (10) with respect to β and α leads to the first-order conditions that solve for their estimates.

Competing Risk Model

Competing risk is a class of survival models designed for multiple transitional exit destinations. McCall (1996) developed a model in which workers could be reemployed either in full-time or part-time jobs, and for each outcome the considered factors were allowed to have varying effects. Exit from the labor force can also be thought of a potential unemployment outcome. After a period searching unsuccessfully for a job, the person may quit the job market altogether. The distinction between transition to employment and that of out of the labor market by the discouraged workers is particularly important during a long recession. Analyzing the Great Recession, Elsbery et al (2010) found the non-participation exit was particularly important after the recession, while the flow to employment was essentially flat. Moreover, unemployment benefits may affect the outcome in making possible to the unemployed to search for job for a longer period than otherwise would be possible (Rothstein 2011). It is clear that this case is an alternative outcome to full or part-time job, because though still in the job market, the job search ceases. Farber and Valletta (2013) found a negative influence of unemployment benefit when the outcome is labor force exit, suggesting that those recipients stick to the job market, perhaps with a lower search level. Krueger, Cramer, and Cho (2004) suggest that the likelihood of exiting the work force is highly influenced by the duration of the unemployment spell, with longer periods being associated with higher labor force withdrawals. It is important to investigate exit from the work force because it may be associated with specific and particularly strong economic effects, such as hysteresis (DeLong and Summers, 2012), skill mismatches and housing performance (Estevão and Tsounta, 2011).

We expand the above models to account for the probability survival when there are more than one exit destinations, i.e., moving from unemployment to either full-time or part-time employment. A survival function with this type of multivariant transitional probabilities will involve estimating a joint distribution of durations. We employ the competing risk model (CRM)to estimate a multiple hazards version of the single-spell model (10) where each exit destination provides one complete duration m, and m -1 censored durations, thus competing risks determine the destination state. We denote destination-specific covariates by $\mathrm{x}_j (j = 1,2, \ldots, m)$, and note that only one duration, the shortest, is observed at the end of the spell $\tau = min_j(t_j)$, t_j>0. If the risks are independent, then the multiple-spell survival function is:

$S_\tau(t)$=Pr[$\tau > t$]= Pr[$t_1 > t$]x . . . x Pr[$t_m > t$].

Let $g_j(t)dt$ be the probability of risk *j* materializing over a small interval change by *dt*, then the total hazard rate of all durations is:

$\lambda_\tau(t) \equiv -d/dt \ln S_\tau(t) = \sum_{j=1}^{m} g_j(t)$.

That is, the probability of exit remains the same regardless of *j* being one of the risks or the only risk.

Given hazard functions independence, the HP Cox duration model provides probability estimates for the integrated hazard over different destinations by a PH model of the form

$$\lambda_j(t;\mathbf{x}) = \lambda_{0j}(t) \exp[(\mathbf{x}'(t)\beta_j], \ j=1, \ldots, m \qquad (11)$$

where both the base hazard and parameter are specific to *j* -type hazard $t_{j1} < \ldots < t_{jk_j}$; $k_j$denotes the ordered destination of type *j*. For instance, with *m*=2, k_1=full-time work and k_2=part-time work.

Specifications of the Econometric Equations:

Below we specify equations (4), (7), (8)-(14) for the single-spell duration, and equations (15)-(16) for the triple-spell competing risks Cox model.

Exponential & Weibull Models:

$$\lambda(t|\mathbf{x}) = \lambda_0(t,\alpha)\ \varphi(\mathbf{x}\beta) \qquad (12)$$

where β is the vector of unknown parameters to be estimated from the data; $\lambda_0(t,\alpha)$ = baseline hazard function at time t, that is hazard function when all predictors are equal to zero, and x = independent predictor variables (see Table 1 for the list of the variables). For exponential and Weibull models their hazards are, respectively, exp($\mathbf{x}'\beta$) and exp($\mathbf{x}'\beta$)$\alpha t^{\alpha-1}$. Specification for our model and estimates are presented in Table 2.

Unobserved Heterogeneity Model:

Effective estimation of the recession unobservable frailty effect requires a recession dummy inside the frailty Gamma function in eq. (7).

$$\boldsymbol{V_{id}} \sim gamma\ (\boldsymbol{\delta},\ d) \qquad (13)$$

where $\boldsymbol{d}$=1 for the Great Recession, and zero otherwise. The frailty implementation with this procedure, however, proves difficult with a large sample size of this study, and we have to adopt a modification discussed below.

Cox semiparametric model:

$$\lambda(t|\mathbf{x}) = h_0(t)\ \exp(\mathbf{x}\beta) \qquad (14)$$

The Cox proportional hazards regression model states that the hazard rate for the i^{th} person in the data is: β is the vector of unknown parameters to be estimated from the data $h_0(t)$ = baseline hazard function at time t, that is hazard function when all predictors are equal to zero, and x = independent predictor variables. The list of variables can be found in Table 1, and specification for our model and estimates are presented in Table 2.

Competing Risk Model:

The estimates in Table 3 (competing risk model) assume that the full-time and part-time re-employment hazards are independent and that the regressors coefficients are time-constant. Moreover, it is assumed that the probability of being re-employed into a type *w* job in period k is conditional on remaining jobless for more than k - 1 periods, indicators for *UI* receipt (*UI*), gender and their interactions and a *J*-dimensional vector of other regressors **x**, satisfies:

$$P(T_w = k|UI, SEX, RACE, x, T > k-1) = 1 - \exp\big(-exp(\alpha_{wk} + g_w(UI, race, gender) + \beta_w x)\big) \qquad (15)$$

Where *w* = full-time, or part-time. Here, α_{wk} are the baseline hazard parameters for risk *w* and β_w is a J-dimensional dimensional vector of parameters measuring the effect of the other regressors on the conditional probability of re-employment into a type *w* job.

Moreover, g_{wk} is assumed to satisfy

$$g_w(UI, race, gender) = \gamma_{1w}UI + \gamma_{2w}race + \gamma_{3w}gender +$$

$$\gamma_{4w} genderUI + \gamma_{5w}\text{raceUI}$$

where *genderUI* and *raceUI* denote the interaction of the gender with *UI* receipt and the interaction of race with *UI* receipt, respectively, and γ_{iw} represents the effect of the ith regressor on risk *w*,

i =1,…,5 ; *w*= full/ part time

Furthermore, we model unobserved heterogeneity as follow:

$$\text{P}(T_w = k|UI, SEX, RACE, x, \theta_w, T > k-1) = 1 - \exp(-exp(\alpha_{wk} + g_w(UI, race, gender) + \beta_w x)\theta_w) \tag{16}$$

The univariate random variable θ_w which affects the conditional probability of re-employment into a type w job is assumed to be unobserved.

We present our competing risks results as provisional, and although they still to be completed in combination with the unobserved recession frailty analysis, we feel the tentative results are suggestive and sufficiently interesting despite their limitation.

5. Results

Single Risk Model

The Current Population Survey allows us to distinguish between three different exits after a period of unemployment: partial job, full-time job and exit from the labor market. While we keep track of alterative post-displacement outcomes through four binary variables (CENSOR1, CENSOR2, CENSOR3, and CENSOR4), our initial focus is on the transition to full-time job. We leave the part-time and exit from the labor force transitions to the next section. Thus, a spell is complete if person is re-employed at a full-time job. We analyze the factors conditioning it through the variables specified in Table 1. Table below provides information about the specification of variables used in the analysis. Although not exhaustive, the variables selected address the most important characteristics with which we are presently concerned.

Additionally, we consider several specifications to our regression models. We start with the PH versions of exponential and Weibull parametric functional forms. The results are reported in Table 2. Hence, this suggests the duration of unemployment spell is not a critical influence on transition to a full-time job. It can be seen that, for almost all variables, the results by exponential and Weibull hazard densities are very close to each other. However, the third rows for Cox regression estimates are critical to support the robustness of these estimates to the unobserved recession impact, and on the face of it, those too support the parametric estimates; apparently, they are not vulnerable to either functional form misspecification, or to the impact of the Great Recession period.

Table 1: Summary a of Variables Used in the Analysis

Variable Label	**Variable Description**	**Mean or Share**
Weeks Unemployed	Number of weeks unemployed	13.1 (mean)
Censor 1	1 if re-employed full time	0.62
Censor 2	1 if re-employed part time	0.24
Censor 3	1 if Unemployed	0.07
Censor 4	1 if not in Labor Force	0.04
UI	1 if received Unemployment benefits	0.45
ColDegree	1 if individual has a college degree	0.28
Black	1 if individual identifies as Black	0.09
Female	1 if individual identifies as male	0.43
Tenure	Length of time worked at lost job in years	4.7 (mean)
Log Wages	Logarithm of weekly earnings at lost job	6.3 (mean)
Age	Age of individual	39.5 (mean)
Young Child	1 if individual has Children under the age of 8	0.23
Married	1 if individual is married	0.56

In all three specifications, gender is the single most important factor affecting the length of unemployment spell. Men are found to take 7% to 15% more time before finding a new full-time job. The result is in line with findings for other countries Ollikainen (2003), Mussid (2007). However, the fact that we are considering long-term data (since the 1996) does not allow us to say much about the shrinkage of this factor over time (as reported, for instance, in Albanesi and Sahin (2018)). On the opposite side, we found that UI diminishes the hazard of exiting the unemployment to 64%. The literature has for a long time considered this result as a shift in the worker's incentives. Our result so far then is apparently in line with not only classical studies on this question (Katz, 1986), but also with recent research Grubb (2011), and Barro (2010). However, since this variable is most sensitive to parametric form the Cox model produces a more significant result than the exponential one. Taken together with gender and race, however, the *UI* variable exhibits a *weaker* influence. Interacting UI with being Black or female (Variable RUI and GENUI respectively), increases the hazard function to 73% and 74% respectively.

Table2: Unemployment duration: Estimated hazard ratios from three parametric models.

	Exponential		Weibull		Cox PH	
	B	Z	B	Z	B	Z
COLDEGREE	1.35	11	1.25	11.13	1.20	10.01
MALE	0.86	-3.5	0.92	-2.77	0.94	-2.18
BLACK	0.71	-6.33	0.75	-6.44	0.75	-6.66
UI	0.65	-14.27	0.68	-16.9	0.65	-21.61
TENURE	0.98	-10.71	0.99	-10.34	0.99	-10.66
LOGWAGE	1.60	26.65	1.46	28.82	1.41	28.87
OVER55	0.41	-26.44	0.46	-27.27	0.48	-27.71
GENUI	1.16	3.24	1.13	3.5	1.13	3.77
RUI	1.14	1.7	1.11	1.7	1.12	2.04
YOUNG CHILD	1.06	1.66	1.06	2.22	1.06	2.67
MARRIED	1.41	10.66	1.35	12.18	1.33	12.75
*GENDER*CHILD*	0.66	-7.5	0.70	-8.46	0.70	-8.96
*MARRIED*GENDER*	0.60	-10.97	0.64	-12.05	0.65	-12.58

Notes: Variable *RUI* is defined as an interaction term between unemployment benefits and being Black. Variable *GENUI* is defined as an interaction term between unemployment benefits and being female. GENDER*CHILDERN and MARRIED*GENDER are defined as interaction terms between female and a dummy variable for having children and being female and a dummy variable for being married.

The DWS does not collect information about UI benefit eligibility or the amount of benefit that displaced workers receive. Ul benefit eligibility depends on both the earnings history of an individual and the reason for job loss. No individual in our sample would be disqualified for UI benefits based on why they lost their job. It is difficult to accurately determine whether a displaced worker satisfied the earnings requirements for UI benefit eligibility using the earnings information in the DWS. Moreover, not all individuals eligible for UI benefits apply to receive them. Elements that might impact workers' take-up decision, need to be accounted for in the estimations since the unobservables determining UI receipt may be correlated with other covariates leading to selectivity bias.

Numerous factors impact UI take-up decisions. Individual's contribution to total household income has the most significant effect on the take-up decision. Other income variables appear to be less impactful. Total monthly household income matters, with those in the middle range (monthly income between $2, 000 and $10, 000) significantly more likely to take up UI benefits relative to those with the lowest total monthly income (less than $2, 000). The highest income category (more than $10, 000) is not significantly different from the lowest group. Additionally, the lowest income group may apply for other programs, such as Food Stamps or other welfare programs. In this case, the household may not expect to be eligible for UI benefits even when they are (Auray, Fuller, Lepage-Saucier, 2018). Given these challenges regarding take-up decisions our estimates might underestimate the true impact of UI on the duration of unemployment.

Age also appears as a strong variable determining the unemployment spell. People over 55 years old has an 52% lower probability of leaving the unemployment status to a fulltime job after experiencing displacement. Being married is found to increase unemployment hazard ratios. A possible explanation is that having a partner act as an additional source of income, allowing individuals to wait longer before accepting a full-time job. We find the opposite effect when it comes to young child. The urgency and the level of financial commitments associated with children may explain a more intense job search behavior.

Recession

Recession periods have a strong impact on the job market, affecting both the ability of the economy in demanding jobs and the behavior of workers in the supply side. Consequently, the covariates being studied in this work are likely to have considerably different values outside the recession period. We aim at capturing such differences by fitting the unobserved heterogeneity with a *Gamma* frailty function. Doing so, will allow our base hazard function to vary multiplicatively with the unobservable effects, principally for the Great Recession period, thus capturing the recession effect. Additionally, we performed the same task, but this time restricting the frailty function to estimate two separate survival equations, one for the recession years and another for the non-recession years. The attempt to estimate the model with frailty function based on a single dummy does not easily converge, hence we adopt an alternative of treating recession/expansion as an ancillary variable. The results are reported in Table 3 below, where we consider workers who have held one (full-time) job since displacement.

Our results indicate that recession periods significantly alter most coefficients. Even the more general estimates with a non-specific unobservable display notable differences for based on the PH Weibull distribution between the middle columns of table 2 and the first columns of table 3. In general, the effects are much smaller when we control for the recession. Such differences become sharp when we compare the results restricted to the recession years (the last two columns) against those restricted for non-recession periods (middle two columns). Increases in the chances of finding a job take place for high-education (0.42, from 0.28), male (0.42, from 0.30), and married (0.26 from 0.16) workers. The education effect is as expected, inasmuch as the positions that get closed during a recession are usually low-skilled ones, affecting disproportionately the low-skill workforce. The result for gender contrasts with the findings of Azmat, Guell and Manning (2006) that pointed the recession as an important factor for closing the gender gap in unemployment rates. A higher transition hazard for men, as in our results, imply the opposite: an increase in the gender gap in unemployment rates.

Table3: Unemployment duration: Estimated hazard Coefficients considering heterogeneity and recession effects

	Gamma frailty		Gamma frailty for non-recession years		Gamma frailty for recession years	
	B	z	B	z	B	z
COLDEGREE	0.28	4.37	0.28	3.07	0.42	2.53
MALE	0.29	4.90	0.30	3.45	0.42	2.75
BLACK	-0.14	-2.51	-0.18	-2.29	-0.07*	-0.58
UI	-0.15	-3.91	-0.16	-2.94	-0.17*	-1.93
TENURE	0.01	3.36	0.01	2.48	0.01	1.97
LOGWAGE	0.11	4.02	0.11	2.86	0.16	2.22
OVER55	-0.36	-5.06	-0.36	-3.62	-0.48	-2.90
YOUNG CHILD	-0.04*	-1.49	-0.04*	-1.18	-0.04*	-0.56
MARRIED	0.17	5.00	0.16	3.75	0.26	2.61
RUI	-0.08*	-1.18	-0.03*	-0.31	-0.23*	-1.35
GENUI	0.17	1.70	0.15*	1.17	0.28*	1.15

Notes: Variable *RUI* is defined as an interaction term between unemployment benefits and race. Variable GENUI is defined as an interaction term between unemployment benefits and gender. (*) Indicates non-significant coefficients.

Conversely, for other variables, recession has the opposite effect, namely decreasing the hazard. That happens with *Black* variable (insignificant for recession and -0.18 otherwise), *UI* (insignificant for recession and -0.16 otherwise), and *Over 55* dummy (-0.48 against -0.36 in non-recession years). We note that our estimate for UI for the Great Recession period observations is no longer significant at 5%, in contrast to our earlier result and in contrast with Grubb (2011) and Barro (2010), as we find that benefits do not significantly impact the transition-to-employment likelihood during recession periods. Our results are more in line with Rothstein (2011) findings that UI extensions had small negative effects on the probability that the eligible unemployed would exit unemployment (Rothstein 2011)

As for age, we find a disproportionate burden of recession on older workers. Negative coefficients indicate decreased hazard and increased survival times, which translates to a prolonged spell of unemployment for older workers compared to younger workers in non-recession years and especially in times of economic downturn.

Multiple Risk Model

Next, we consider alternative post-displacement outcomes. That means the individuals may transition not only to full-time jobs (Risk 1), but also to part-time jobs (Risk 2) , Risk 3, is to remain in unemployment, which is treated in our model as a right-censoring case, the last outcome (Risk 4) is to leave the labor market altogether. Plot 1 compares the cumulative hazard estimates for each transitional destination for the PH Cox competing risks model; note the part-time destination curve displays a steep rise compared to other relatively flat flow of return to work. Black workers have 28% and 27% lower rates of exit to a job (both full-time and part-time), while having a higher probability of remaining in unemployment (35%). Men also have a lower exit to (re)employment rate by 7% to a full-time job and 15% to a part-time job.

Interaction with child and marital status shed light on the gender role. For full-employment outcome, having a young child lowers the probability by 34%. The risk of exiting the labor force after displacement is also higher for women with children. Possible explanations include uneven household responsibilities and imbalanced intra-household bargaining power that imposes unpaid care work on women (Folbre 2021), difficulties finding a suitable job due to gender discrimination, and possibly higher reservation wages due to mothers' preference for spending time with children. (Youderian 2014). Being married is also associated with an increased (28%) rate of transition to a full-time job. Older workers have a lower rate of (-6%) transitioning to full-time re(employment) as well.

Table 4: Competing Risks Estimates for Multiple Transitional Probabilities

Covariates	**Risk 1**	**Risk 2**	**Risk 3**	**Risk 4**
COLDEGREE	0.184	0.112	-0.157	-0.111
	(0.018)	(0.031)	(0.021)	(0.029)
Male	-0.066	-0.154	0.193	-0.117
	(0.03)	(0.042)	(0.032)	(0.04)
Black	-0.282	-0.277	0.35	-0.146
	(0.042)	(0.056)	(0.034)	(0.049)
UI	-0.437	-0.569	0.189	-0.5
	(0.02)	(0.038)	(0.023)	(0.037)
Tenure	-0.014	-0.009	-0.023	0.013
	(0.001)	(0.002)	(0.001)	(0.001)
Log wage	0.346	-0.134	0.22	-0.118
	(0.012)	(0.017)	(0.012)	(0.014)
Black*UI	-0.729	-0.276	-0.225	0.734
	(0.026)	(0.037)	(0.023)	(0.026)
female*UI	0.121	0.186	0.138	0.092
	(0.032)	(0.05)	(0.035)	(0.046)
Young Child	0.116	0.097	-0.267	0.049
	(0.057)	(0.084)	(0.044)	(0.076)
Over 55	-0.064	0.041	-0.039	-0.247
	0.024	0.047	0.027	0.062
Married	0.285	-0.087	-0.14	-0.099
	(0.022)	(0.039)	(0.022)	(0.038)
Young Child*female	-0.353	-0.121	-0.143	0.733
	(0.039)	(0.061)	(0.042)	(0.068)
Married*female	-0.424	0.057	-0.062	0.311
	(0.034)	(0.052)	(0.034)	(0.047)

Note: Standard errors in brackets.

Finally, our results also lead to some considerations on the unemployment benefit. Irrespective of the outcome, there is strong evidence unemployment benefits decrease the probability of unemployment transition, rendering the spell longer. The strongest effect happens in the part-time transition, which can be explained in terms of the *UI* allowing workers more job search time. However, in view of the *UI* estimate given in the last columns of table 3, this general effect may be negligible in a deep recession. The regression coefficient for the interaction term of female and UI is positive, meaning an increase of the exit to a job hazard for women. While the coefficient associated with the risk of leaving the labor force is negative indicating increased labor market attachment in presence of UI benefits.

We end this section by noting the relatively sharp contrast between the estimates from the single and multiple estimation Cox estimates of tables 2 and 4. For instance, gender and unemployment insurance estimates are 0.86 (z = -3.5) and 0.64 (z = -14.2) respectively in table 2, but 0.06 (*s.e.* = 0.03) and -0.43 (*s.e.* =0.03). We discuss this outcome in the next section.

6. Discussion

The study suggests that UI benefits impact unemployment spells and exit destinations differently across gender and racial groups. We also presented evidence that UI benefits do not prolong unemployment spells in recessions. One issue not examined in sufficient depth is the effect of age on this outcome. A nearly 50% lower probability of leaving the unemployment status for over 55 years old workers from our single-spell models is very high, the chances are in the prolonged period of the Great Recession most of them might well have been permanently pushed out of the labor market and never able to return once the economy rebounded. Retirement of these older workers earlier than required for adequate retirement saving, has serious implications for old-age poverty and wellness. In future research, we plan to address the impact of age as a recessionary duration of unemployment influence.

Our treatment of the recession relies on what is commonly employed in biostatistics but there may be more effective alternatives available from a state-space econometric modeling that specify a model for recession/expansion states and estimate the system of equations subject to unobservable state specifications. We intend to explore this approach in future work. However, an important caveat is the limitation of our multiple exit probability estimates without accounting for the recession frailty impact. Despite their limitation, the multiple destination transitional estimates are suggestive. A comparison of single-spell and multiple-spell of transition to full-time status by the PH Cox model reveals notably different estimates; the multiple exit model differs significantly not just in magnitudes of covariates effects but also in the *direction* of those effects. Until further research, our findings in this regard must be considered as provisional.

7. Conclusion

Our single-destination estimates show that gender is the single most important determinant of the duration of unemployment, with men hazard rate of unemployment status exit 7% to %15 lower than women. While unemployment insurance lowers the probability of transition to full employment by 64%. However, gender also accounts for this small hazard transition; an interactive term of unemployment insurance and gender lower that transition by 5%-10%. Age is the other major effect on the unemployment transition with over 55 years of age experiencing 50% lower hazard transitional probability, while being married increases that transition by nearly 30%. We also attempted to control for the impact of recession in estimations and the outcome apparently indicate robust parametric estimates in that regard.

The estimates from our competing risks regression for multiple hazard by the Cox non/semiparametric also provide evidence for the critical role played by gender. Our results indicate that race increases the length of the unemployment spell substantially when the individual transitions to full-time employment (-27%). Since the spells mainly differ by gender effect, the evidence points to gender of the worker a crucial factor in the decision to leave the labor force. Interaction with child and marital status also shed light on the gender role. For full-employment outcome, having a young child increases the probability by 12%. For part-time and exit of out of the labor force, however, there is no statistically meaningful relationship. There is also strong evidence that unemployment benefits decrease the probability of unemployment transition, rendering the spell longer. The strongest effect happens in the part-time transition.

Finally, we suggested our control for unobserved heterogeneity and functional form misspecification can be interpreted as control for the impact of the recession. We presented the impact of the Great Recession on our covariate estimation with a frailty function for the unobservables, and with separate frailty-multiplied hazard function for recession and non-recession observations separately; the results point to large impacts of the recession on the estimates obtained, esp. for the sample of the Great Recession period observations. The task of extending this procedure to our multiple destination probability estimation remains an important future work.

Bibliography

Albanesi, Stefania, and Ayşegül Şahin. “The gender unemployment gap.” *Review of Economic Dynamics* 30 (2018): 47-67.

Auray, S., David L. Fuller and Nicolas Lepage-Saucier. “Why do Half of Unemployment Benefits Go Unclaimed?” (2018).

Azmat, Ghazala, Maia Güell, and Alan Manning. “Gender gaps in unemployment rates in OECD countries.” *Journal of Labor Economics* 24.1 (2006): 1-37.

Barnichon, Regis and Figura, Andrew, Declining Labor Force Attachment and Downward Trends in Unemployment and Participation (2013). FEDS Working Paper No. 2013-88.

Barro, Robert. “The folly of subsidizing unemployment.” *Wall Street Journal* (August 30, 2010).

Boršič, Darja, and Alenka Kavkler. “Modeling unemployment duration in Slovenia using Cox regression models.” *Transition Studies Review* 16.1 (2009): 145-156.

Dănăcică, Daniela, and Ana-Gabriela Babucea. “The role of education for the duration of unemployment in Gorj Country.” *Studia Universitatis Babes Bolyai–Oeconomica* 1 (2007): 67-82.

De la Rica, Sara, and Yolanda F. Rebollo-Sanz. “Gender differentials in unemployment ins and outs during the great recession in Spain.” *De Economist* 165.1 (2017): 67-99.

Elsby, Michael W., Bart Hobijn, and Aysegul Sahin. *The labor market in the Great Recession*. No. w15979. National Bureau of Economic Research, 2010.

Folbre, Nancy. *Gender inequality and bargaining in the U.S. labor market*. Economic Policy Institute, 2021

Grubb, David. “Assessing the impact of recent unemployment insurance extensions in the United States.” *draft presented at IZA, Berlin* 25 (2011).

Heath, Alexandra, and Troy Swann. “Reservation Wages and the Duration of Unemployment.” *Reserve Bank of Australia Research Discussion Papers*, February (1999).

Hougaard, Philip. “Survival models for heterogeneous populations derived from stable distributions.” *Biometrika* 73, no. 2 (1986): 387-396.

Imbens, Guido W., and Lisa M. Lynch. “Re-employment probabilities over the business cycle.” *Portuguese Economic Journal* 5.2 (2006): 111-134.

Katz, Lawrence F. “Efficiency wage theories: A partial evaluation.” *NBER macroeconomics annual* 1 (1986): 235-276.

Kroft, Kory, Fabian Lange, Matthew J. Notowidigdo, and Lawrence F. Katz. “Long-term unemployment and the Great Recession: the role of composition, duration dependence, and nonparticipation.” *Journal of Labor Economics* 34, no. S1 (2016): S7-S54.

Kuhlenkasper, Torben, and Max Friedrich Steinhardt. *Unemployment duration in Germany: A comprehensive study with dynamic hazard models and P-Splines*. No. 111. HWWI Research Paper, 2011.

Lancaster, Tony. “Econometric methods for the duration of unemployment.” *Econometrica: Journal of the Econometric Society* (1979): 939-956.

Landmesser, Joanna. “The impact of vocational training on the unemployment duration.” *International Advances in Economic Research* 17.1 (2011): 89-100.

Mussida, Chiara. “Unemployment duration and competing risks: A regional investigation.” *DISCES- Quaderni del Dipartimento di Scienze Economiche e Sociali, (0749)* (2007).

Nickell, Stephen. “Estimating the probability of leaving unemployment.” *Econometrica: Journal of the Econometric Society* (1979): 1249-1266.

Ollikainen, Virve. “The determinants of unemployment duration by gender in Finland.” (2003).

Pissarides, Christopher A. “Unemployment in the great recession.” *Economica* 80.319 (2013): 385-403

Rothstein, Jesse. “Unemployment Insurance and Job Search in the Great Recession.” *Brookings Papers on Economic Activity* (2011): 143 - 213.

Sahin, Aysegul, Joseph Song, and Bart Hobijn. “The unemployment gender gap during the 2007 recession.” *Current Issues in Economics and Finance* 16.2 (2010).

Tansel, Aysit, and H. Mehmet Taşçı. “Hazard analysis of unemployment duration by gender in a developing country: The case of Turkey.” *Labour* 24.4 (2010): 501-530.

Valletta, Robert, and Henry S. Farber. “Do Extended Unemployment Benefits Lengthen Unemployment Spells? Evidence from Recent Cycles in the US Labor Market.” Federal Reserve Bank of San Francisco, 2013.

Xiaoyan Chen Youderian, 2014. “The motherhood wage penalty and non-working women,” Economics Bulletin, AccessEcon, vol. 34(2), pages 757-765.