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ABSTRACT

Does Part-Time Work Help Unemployed Workers to Find Full-Time Work? Evidence from Spain*

This paper examines whether part-time work acts as a bridge towards full-time work for unemployed workers in Spain. We follow the timing-of-event approach and estimate the causal effect of part-time work on the exit rate to full-time work using a multivariate duration model. Our findings show that the exit rate to full-time work declines when working part time (lock-in effect) but increases afterwards (stepping-stone effect), implying a trade-off between the two opposite effects. The resulting net effect of part-time work on the expected time until full-time work is positive in most cases, leading to longer spells without full-time work. This undesirable effect has increased over time, so that the value of temporary part-time work as a pathway to full-time work for the unemployed has reduced.

JEL Classification: J64, J65

Keywords: part-time employment, work trajectories, unemployment duration, mixed proportional hazard model

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1. Introduction

Improving competitiveness of firms, increasing labour market flexibility and fighting persistent and high unemployment have been at the fore of the policy debate in OECD countries in the last four decades or so. Although a number of different strategies have been pursued to achieve those intentions, one of the most ubiquitous has been to favour the use of atypical forms of employment. Therefore, temporary and part-time contracts have gradually been gaining importance in many labour markets. In principle, both firms and workers can benefit from part-time work, since it is intended not only to provide flexibility to employers to better adapt to changes in product demand but also to help workers achieve a balance between professional and private life, enhance labour market entry, and increase employment when full-time jobs are not available. However, policies promoting part-time work has also been criticized on the grounds of the disadvantageous situation of part-timers who often earn lower wages, receive fewer fringe benefits, participate less in training, have more limited career prospects, alternate more frequently jobs and have a higher risk of non-employment than full-time workers (Ermisch and Wright, 1993; Dekker et al., 2000; Connolly and Gregory, 2008; McDonald et al., 2009; Lyonette et al., 2010).

Despite the increasing prevalence of part-time work, there has been relatively little empirical research on its role as a means to combat unemployment, especially in the Southern European countries, where its share has been traditionally low. The objective of this paper is to investigate the role that temporary part-time jobs may play in improving the prospects of finding full-time employment afterwards for unemployed workers in Spain. Our paper is the first to examine this issue focusing on the Spanish case.

In Spain, unlike in most other European countries, part-time employment is predominantly involuntary among both women and men. The share of part-time employment remained rather low by European standards until the 2008 recession, but since then it has been increasing steadily. While part-time contracts amounted to about one fourth of all employment contracts signed in 2004–2007, their share was more than one third in 2012–2014. Also the share of part-timers who would prefer to work full time has increased substantially since 2008. This development has been boosted by an intentional labour market policy. As part of a series of measures to fight the impact of the 2008 financial crisis, the Government approved a reform to encourage part-time employment amidst the recession by lowering employers' social security contributions for part-timers well below those for full-time workers.

We address the question whether part-time employment (PTE) acts as a bridge towards full-time employment (FTE) for unemployed job seekers. The answer to this question helps to assess whether the recent policy of promoting part-time work is beneficial from the viewpoint of millions of Spanish unemployed workers. We use a rich longitudinal dataset obtained from the Social Security records that covers the labour market histories of private-sector workers from 2005 to 2013. The empirical analysis is conducted on a flow sample of workers who started receiving unemployment benefits between 2005 and 2013 after a full-time employment spell of at least six months. Given their history of full-time employment, we assume these workers are looking for full-time work even though some of them took up part-time jobs. As such, we consider possible periods of part-time work as part of the unemployment spell, which we define as consisting of all consecutive spells of unemployment benefits and part-time work.

In the econometric analysis, our outcome variable is the expected time until the unemployed individual finds a full-time job. In order to model this duration outcome, we specify a hazard model for exits to full-time employment. Episodes of PTE within the unemployment spell are then viewed as "treatments", which possibly affect the expected unemployment duration through an impact on the exit rate to FTE. We ask what would have happened to part-timers if they had not worked part time but instead continued their search of FTE on unemployment benefits. We distinguish between the effect of PTE on the exit rate to FTE during the period when the individual is working part time and the effect following the completion of the part-time job. Since the former effect is negative and the latter positive, we refer to these effects as the lock-in and stepping-stone effect, respectively.

To deal with the endogeneity of the timing and duration of part-time jobs, we specify hazard models also for transitions from unemployment benefits into PTE and for transitions from PTE into unemployment benefits within the unemployment spell. This leads to a multivariate duration model where the three hazard rates are interrelated through observed and unobserved characteristics. Provided there is some randomness in the timing and duration of part-time jobs, the causal effects of such jobs on the exit rate to FTE can be distinguished from the selection effects without exclusion restrictions or parametric assumptions about the shape of the hazard functions (i.e. duration dependence) or the distribution of unobservables. This approach for identifying the causal treatment effects in the context of duration outcomes is known as the timing-of-events analysis (Abbring and Van den Berg, 2003).

We find evidence of a notable lock-in effect and stepping-stone effect, so that there is a trade-off between the two opposite effects. We show that, except for relatively short part-time jobs, the resulting net effect on the expected time until full-time work is generally positive for both women and men. In other words, taking up a part-time job prolongs the expected time without a full-time job in most cases. We also find that the net effect of part-time work on the expected unemployment duration has increased over time, and hence the value of temporary part-time work for the unemployed who would like to work full time has reduced. This development has been, in part, driven by the policy reforms to foster the use of part-time employment contracts implemented in the aftermath of the 2008 financial crisis. On the other hand, the very high unemployment level and stagnant economy since 2008 are likely to have played an important role as well.

The paper proceeds as follows. Section 2 provides an overview of the literature. Section 3 describes some features of the Spanish institutional context. Section 4 presents the data, the sample selection process and reports some descriptive statistics. Section 5 outlines the econometric methodology employed in the empirical analysis. Section 6 discusses the estimation results, and Section 7 concludes.

2. Literature review

Although not similar in scope to the debate concerning the extension of temporary employment, the increasing use of part-time contracts has brought about two opposing views in this development among economists and policy makers. On the one hand, the existence of part-time jobs has been seen as a means to improve labour market flexibility and to reduce labour costs in industries subject to large seasonal or cyclical variation, thereby increasing the overall labour demand. Moreover, these jobs may help currently unemployed workers since they provide them with opportunities to gain work experience and maintain and/or acquire human capital when no full-time jobs are available. They may also provide a bridge into full-time work especially for women who have been out of the labour force for family reasons and for labour market entrants.[1] For these reasons part-time work may serve as a *stepping stone* into full-time employment.

On the other hand, critics are sceptical about the potential of PTE to enhance job creation and stress the danger of regular full-time jobs being substituted by part-time jobs. Furthermore, part-time jobs are often characterised by having less work-related training and worse labour conditions, such as lower wages and greater labour insecurity, than full-time jobs. Many workers seem to be trapped in a sequence of part-time jobs and non-employment spells. Therefore, some authors consider part-time jobs as *dead ends* which do not lead to regular full-time employment and stable work trajectories.

The "stepping-stone" vs. "dead-end" hypothesis of temporary work has been studied empirically by several authors, but these studies do not offer a clear conclusion. Hagen (2003) and Boockmann and Hagen (2008) for Germany, De Graaf-Zijl et al. (2011) for the Netherlands, and Engellandt and Riphahn (2005) for Switzerland find that temporary jobs increase the probability of getting a permanent job. However, Amuedo-Dorantes (2000) obtains the opposite result for Spain, while Booth et al. (2002) for the UK, Gagliarducci (2005) for Italy, Güell and Petrongolo (2007) for Spain and D'Addio and Rosholm (2005) for the European Union Member States find evidence supporting both hypotheses. Some studies distinguish different groups of workers with non-standard contracts and investigate, for instance, whether temp agency workers have a higher or lower probability of being hired on a permanent basis than direct-hire workers (Amuedo-Dorantes et al., 2008).

In the case of part-time work, much of the empirical literature has focused on its effect on the labour market trajectories of women (Hakim, 1998; Grimshaw and Rubery, 2001; Connolly and Gregory, 2008; Manning and Petrongolo, 2008; Blázquez and Moral, 2014). Part-time jobs have been seen as a voluntary choice for many women to combine labour market involvement with household responsibilities, particularly during childcare years. From a life-cycle perspective, part-time work would then be a temporary alternative to full-time work or non-participation, constituting a "maintenance" role for women who would otherwise be working full time. However, it is widely documented that many part-time jobs are poorly paid and offer little opportunity for career progression. As such, for many workers part-time work can be

[1] Part-time jobs may also increase employment by postponing retirement among older workers with reduced working capacity.

part of an “exclusionary” cycle, where insecure part-time jobs alternate with spells of non-employment.

O’Reilly and Bothfeld (2002) find, using the British Household Panel Survey for 1990–1995, that only a small number of women used part-time work as a bridge back into a full-time job after a spell of non-employment. On the contrary, a large share of all spell sequences implied women transiting through part-time work from non-employment back to non-employment, giving support to the view of part-time work in an “exclusionary” pattern. For the USA, Blank (1998) identifies two leading patterns in transitions through part-time work. For the majority, a part-time job serves as a temporary alternative to full-time work, to which they subsequently return: this is the “maintenance” role of part-time work, supporting continued labour market participation within a stable working career. The other major group enters part-time work from non-employment and then leaves the labour market again, forming part of an “exclusionary” cycle of weak labour market attachment. Connolly and Gregory (2010), following a cohort of women until they reached age 42 using the British National Child Development Survey, reach similar conclusions, with part-time work serving two different functions. Women whose past history predominantly involves full-time work, possibly in conjunction with spells of part-time work or non-employment, tend to revert to full-time work. But women whose labour market history combines spells in part-time work with non-employment are subsequently unlikely to take up full-time work. In sum, women are persistent workers or persistent marginal/non-workers. Both groups engage in part-time work but in different ways. Part-time work is both a support and a trap for women’s future careers.

Another strand of empirical research has focused on the labour market effects of a particular type of part-time jobs, the so-called “mini-jobs” or “marginal employment” (employment with low working hours and earnings not or only partially subject to social security contributions). The labour supply effects of the German “mini-job” reform that took place in 2003 have been analysed by ex-ante simulation studies (Steiner and Wrohlich, 2005) and ex-post evaluations (Caliendo and Wrohlich, 2010; Caliendo et al., 2016). In general, these studies find that the reform had only small labour supply effects, especially among the target group of the long-term unemployed. Freier and Steiner (2007) find that marginal employment does not affect time spent in regular employment within a three-year observation period, reduces future unemployment and slightly increases cumulated future earnings on average among German men (although a small negative cumulative earnings effect is obtained for older workers in western Germany). Caliendo et al. (2016) find heterogeneity in the effect of taking up a mini-job on the exit rate from unemployment benefits to self-supporting employment. The mini-jobs appear to be helpful for the long-term unemployed and for those who live in regions with a high unemployment rate, whereas job seekers who take up a mini-job during the first six months of unemployment tend to collect unemployment benefits for a longer time. Other studies have focused on the influence of similar types of job on the trajectories of workers in other countries. In a study for Austria, Böheim and Weber (2011) find that the unemployed who take up marginal employment end up with less regular employment, more unemployment and lower wages after three years compared to the control group of unemployed who did not enter marginal employment.

Finally, other related studies have examined the implications of subsidies for atypical employment (partial unemployment benefits or wage subsidies to the unemployed who accept a part-time or low-pay full-time job in the regular labour market) for unemployed workers' subsequent labour market outcomes. Evidence for Switzerland is mixed. Gerfin and Lechner (2002) and Gerfin et al. (2005) conclude that the temporary wage subsidy is a successful programme in terms of increasing job seekers' chances of getting an unsubsidized job and reducing the time spent on benefits. Lalive et al. (2008) find the same result when they use a matching estimator but no significant effects when they apply a timing-of-events duration approach which allows for selection on unobserved characteristics.

Kyyrä (2010) finds that short full-time working on partial unemployment benefits facilitates the transitions to regular employment but subsidized part-time jobs are much less effective in the Finnish labour market. Kyyrä et al. (2013) study the effects of partial unemployment benefits in Denmark, where such benefits can be received when working hours over a week are below a given threshold level. Unlike the Finnish study, this study finds evidence of a significant lock-in effect: working part time on partial benefits reduces the unemployment exit rate. However, after returning to full-time unemployment from subsidized part-time work, the exit rate is larger compared to the counterfactual case of having been full-time unemployed for the whole time. As such, there is a trade-off between a negative lock-in effect and a positive stepping-stone effect afterwards, so that the net effect of subsidized part-time work on unemployment duration depends on the relative magnitudes of these two effects. Kyyrä et al. (2013) show that collecting partial benefits and working part time reduces, on average, the remaining unemployment duration in the Danish labour market. Yet they emphasize impact heterogeneity: while subsidized part-time work tends to reduce the expected unemployment duration of young individuals and immigrants, it can prolong unemployment spells of married women.

In this vein, Fremigacci and Terracol (2013) find that working part time on partial benefits is associated with a lock-in effect also in France, but Godøy and Røed (2016) find no such effect in Norway. Both studies find however that the exit rate to full-time employment increases after a period of part-time work. As a result, subsidized part-time work unambiguously reduces the expected time until a full-time job in Norway, and does so in most cases also in France.

3. Institutional background

3.1. Part-time work

Part-time work has traditionally been considered secondary or marginal employment in Spain. During the 1970s and 1980s, less than 5% of the employees were part-timers. The passing of the Workers' Statute (*Estatuto de los Trabajadores*) in 1980 removed the social security costs' penalty (social security costs were higher for part-timers) and the 1984 labour reform eliminated the hiring restrictions (part-time contracts were only legal for certain types of workers considered at risk of social exclusion, such as workers with disabilities, new job seekers, older workers with family responsibilities and long-term unemployed). The 1994 reform aimed at increasing part-time work as a flexible work arrangement by reducing the social security contributions and the access to unemployment benefits of those who worked less than 12 hours per week or less than 48 hours per month. As a consequence of the combined effect of this legal change and the crisis of the early-1990s, the share of part-time work increased slightly. In 1998, the previous legal change was undone as the Spanish part-time regulation converged to that of the European Union. This did not alter the proportion of part-time work, which remained remarkably stable in subsequent years.

At the beginning of 2009, the Spanish government approved a Royal Decree-Law to foster part-time work which reduced employers' social security contributions in the case of hiring part-time workers.[2] In 2011 the Government passed an “Emergency (short-term) Plan”, which included a programme for improving transitions toward stable employment while promoting part-time work by further reducing employers’ social security contributions for part-time workers.[3] The exact size of the cost reductions varied in a complex way, depending on personal and firm characteristics, working hours etc., but in some cases, the employers were exempted from social security contributions altogether. All the social security cost reductions were time-limited, being available up to three years at a maximum.

As it happened during the crisis of the 1990s, the financial crisis and the reforms in 2009 and 2011 brought about a new rise in the share of part-time employment: after having remained around 11% in 2005–2008 without any trend in previous years, it increased steadily in 2009–2013 up to 16%. Differences between men and women are substantial though, part-time work being less frequent among male workers than their female counterparts: the share is currently about 8% for the former and 26% for the latter (after having increased from 4% and 22%, respectively). These figures are still low compared with the UK, the Netherlands and the Nordic countries, while they are more similar to the ones in France and the Southern European countries (see the top panel of Table 1).

Part-time employment is often involuntary and concentrated in low-pay occupations. In Spain, the share of part-timers who declared that they would prefer working full time has increased during the recession: from about one third in 2006 to nearly two thirds in 2014 (see the bottom panel of Table 1). A similar change has happened in nearly all European countries with the exception of Germany and, to some extent, Denmark. The rise has been relatively large in Italy, France, Portugal and Spain, although only Italy exhibits similar figures to those of Spain. In addition, the proportion of involuntary part-time work has increased more for men than for women.

Finally, part-time workers show a higher probability than full-time workers of entering non-employment; this difference is particularly large in Spain compared to other

[2] “Royal Decree-Law 2/2009 of urgent measures for the maintenance and promotion of employment and the protection of the unemployed persons”. Moreover, not only unemployed workers hired with a part-time contract but also those working part-time who moved into a part-time job in another firm were eligible for the cut in social security contributions.

[3] “Royal Decree-Law 1/2011 of urgent measures to promote the transitions into stable employment and the professional qualifications of the unemployed persons”.

European countries (see OECD, 2010; Horemans and Marx, 2013; Blázquez and Moral, 2014).

[Insert Table 1]

3.2. Unemployment compensation schemes

One important aspect concerning part-time employment that has to be taken into account is its relationship with the unemployment compensation system. The Spanish system (like in many other OECD countries) comprises two schemes: unemployment insurance (UI) and unemployment assistance (UA). UI benefits are paid to workers who lost their job or whose temporary contract came to an end, who can and want to work, and who have paid UI contributions for at least 12 months during the past 72 months (excluding civil servants and workers hired by households). The length of UI entitlement varies between 4 and 24 months, depending on the number of the months contributions were made. The gross replacement rate is 70% for the first six months of UI receipt and 50% thereafter (60% before July 2012), though the benefit level is subject to a certain upper limit. Moreover, workers who are not eligible for UI or who have exhausted their benefits may qualify for flat-rate UA benefits. The UA benefit is means tested, and its level and duration depend on the number of family dependents and the age of the recipient.

The unemployed worker who takes up a part-time job does not necessarily lose all of his or her unemployment benefits but may receive some benefits on the top of the wage income. In that case, the benefit amount is reduced in the proportion to the full-time working time (i.e. by 50% for a half-day job). The only requirement for partial benefits is that the worker makes a formal request. If the unemployed worker takes up a full-time job but returns to unemployment within one year, he or she is entitled to unused benefits from the previous unemployment spell.

4. Data and descriptive analysis

4.1. Description of the dataset and sample

The dataset used in this paper is the “Continuous Sample of Working Life” (*Muestra Continua de Vidas Laborales*, MCVL hereinafter), which is based on the administrative records of the Spanish Social Security. The population of reference in the MCVL includes employed workers who are registered with the Social Security, pensioners and unemployment benefit recipients.[4] Of this population, 4% are selected by means of a simple random sampling system every year. The individual sampled in a given year will be included in the data also in the following years provided he or she remains registered with the Social Security. Thanks to the existence of a unique identification number for each individual, the individuals can be tracked through the files of different editions of the data. Due to its longitudinal design, the MCVL remains representative of the target population over time. The resulting database thus provides longitudinal information on over one million people who were registered with the Social Security between 2004 and 2013.

The dataset contains information on individual characteristics (gender, age, nationality and province of residence) as well as on job and employer attributes (job category, type of contract, starting and ending dates of job matches, reason for termination of a job, working time, and employer's size, industry and region). For unemployment individuals we observe the starting and ending days of benefit receipt, but not the benefit level nor the length of the entitlement period. The longitudinal nature of the dataset makes it possible to follow individuals across different labour market states over time.

For purposes of this study, we select a subsample of individuals aged 18–60 who started to receive UI or UA benefits between the years 2005 and 2013 after a spell of full-time employment of at least six months. We assume these workers are looking for a full-time job. This seems a plausible assumption as they became unemployed after working full time for a relatively long time. Because of this assumption, we treat periods of PTE that follow receipt of unemployment benefits as part of the unemployment spell. More precisely, the unemployment spell is defined as a sequence of days during which the worker receives either UI or UA benefits, or works part time (with or without partial unemployment benefits) provided that the gap between these periods is no longer than 28 days. To eliminate a few outliers we censor the unemployment spells at 120 weeks (2% of the spells). The unemployment spell is completed if it was followed by a full-time job that started within four weeks after unemployment exit and lasted for at least one week. Otherwise the spell is treated as right-censored, which happens if the benefits

[4] Job seekers not receiving benefits and the inactive population (as distinct from pensioners) are not included. The same applies to workers with a social welfare system other than the Social Security (civil servants who decide so) and those with none (such as those working in the informal or submerged economy or some marginal activities). All in all, the MCVL is broadly representative of the private sector of the economy (see Arranz and García-Serrano, 2011).

expired or were suspended due to a sanction, the worker left the labour force or found a full-time job shorter than one week, the spell was longer than 120 weeks, or the spell was still ongoing on December 31, 2013.

Many individuals who found a job or whose spell was censored returned to unemployment – possibly many times – during our observation period. We include these subsequent spells in the analysis provided they started after a full-time job. For these spells we apply a less strict criterion for the length of the preceding FTE spell. Namely, a new unemployment spell (the second, third and so on up the tenth) starts when the individual returns to UI or UA benefits after a spell of a full-time job that lasted for at least four weeks.[5]

All analyses will be conducted by gender, since part-time work is likely to play a different role for women and men. Our final sample includes 79,312 women who experienced 140,612 unemployment spells, and 130,030 men who experienced 258,459 unemployment spells. 59% of the individuals experienced only a single spell of unemployment, 20% experienced two spells, 9% three spells and 4% more than five spells. On average, women experienced 1.8 spells and men 2.0 spells.

4.2. Main variables and descriptive statistics

Figure 1 depicts the unemployment inflow for the period 2005–2013, distinguishing between men and women. Several findings are worth nothing. First, the unemployment inflow varies in a countercyclical manner, showing a large increase in 2008 as a consequence of the global financial crisis. Second, there are important differences by gender. The inflow is larger for men than for women, especially in the later years. Although the inflow increased substantially in 2008 for men, the rise was comparatively small for women, suggesting that men were more strongly hit by the 2008 recession. Since then the inflow, at least for men, has been declining, although the average level remained higher than before the recession. Finally, the inflow owns a clear seasonal component, related to the activity of certain sectors, such as agriculture and those linked to tourism.

[Insert Figure 1]

[5] We only include the first ten spells for each individual because a large number of spells may cause numerical difficulties in the maximum likelihood estimation of duration models involving unobserved heterogeneity. This restriction reduces the number of spells by 0.8%.

Table 2 provides sample statistics by gender and PTE status (i.e. whether the worker experienced at least one episode of PTE during his or her unemployment spell or not). The observations refer to spells, not to individuals. The share of immigrants is somewhat higher among unemployed men than women (15% vs. 11%). For both sexes the largest age group is 30–39 years old, the average age being 37 for women and 38 for men. There are some differences between sexes in industry and type of the past job. While many women were employed in trade, hotels and restaurants, manufacturing and education and health, men often worked in financial intermediation, construction, manufacturing and trade. Of women's past jobs, one third were white-collar low-skilled jobs, whereas 39% of men held a blue-collar high-skilled job before becoming unemployed.

[Insert Table 2]

The average unemployment duration is 32 and 29 weeks for women and men, respectively. About one half of the spells (45% for women and 53% for men) are uncensored, i.e. followed by a full-time job. As expected, part-time work is more common among women than men: while 10% of women's unemployment spells contain at least one episode of PTE, only 4% of men's spells do. This can be due to a variety of reasons: women may have stronger preference for working part time; employers of women may have higher demands for part-time workers and put women into those jobs; or women are more likely to experience PTE because they remain on average unemployed for longer time. Accordingly, PTE is more common in white-collar low-skilled occupations (administrative and service jobs) and service sectors (trade, hotels and restaurants, education and health and real estate and renting), where women have been traditionally over-represented.

Unemployed workers experiencing PTE are about 2 years younger than their counterparts without such experiences. Among the former group workers under age 30 are over-represented while workers above age 49 are under-represented, especially in the case of female workers. Finally, unemployment history is fairly similar for both groups as measured by the fraction of time spent on unemployment benefits one, two or three years before the current spell of unemployment.

Only one third of the spells involving PTE (31% for women and 34% for men) ended with a transition to full-time work compared to about one half of the spells without episodes of PTE (47% for women and 53% for men). At glance, this seems to be in contrast to the hypothesis that PTE acts as a stepping stone to FTE. But note that the unemployment spells involving PTE are also much longer on average than other spells.

Table 3 offers some summary statistics for the episodes of PTE within the unemployment spells. PTE spells tend to be concentrated on certain occupational groups and industries. White-collar low-skilled jobs and blue-collar low- and medium-skilled jobs make up the majority of PTE episodes for women (44% plus 34%), while blue-collar jobs account for the majority for men (nearly 70%). This is closely related to sectoral differences: trade, hotels and restaurants, education and health and real estate and renting account for 77% of women's PTE spells, whereas trade, hotels and restaurants, real estate and renting and transport make up 61% of men's PTE spells.

PTE episodes are rather long on average (24 weeks), which partly explains why the unemployment spells involving PTE are relatively long. The mean duration of PTE episodes is slightly longer for women (26 weeks) than men (22 weeks). As seen in Figure 2, PTE spells shorter than three months are more common for men, whereas women are more likely to experience very long spells that last around 40 weeks or exactly one year, i.e. 52 weeks. However, the importance of different exit routes is rather similar for both sexes, with more than one half of part-timers returning to unemployment benefits and about one sixth exiting to FTE.

[Insert Table 3]

[Insert Figure 2]

5. Econometric model

As pointed out earlier, we can think of the episodes of PTE as "treatments" occurring within the unemployment spells and then estimate the effects of these treatments on the exit rate to FTE. To identify the causal effects we follow the timing-of-events approach developed by Abbring and Van den Berg (2003). Formally, let T_u be a continuous random variable for the time from unemployment entry until finding a full-time job. Its distribution can be characterized by a hazard function, for which we assume the following mixed proportional hazard (MPH) form:

$$h_u(t|v_u) = \lambda_u(t)\exp\left\{x\beta_u + z(\tau_u + t)\delta_u + \theta_u ur(\tau_u + t) + \alpha_1 d_1(t) + \alpha_2 d_2(t) + \mu w(t) + v_u\right\},$$

where t is the elapsed duration of unemployment, τ_u is the calendar time at unemployment entry, x is a vector of observed characteristics, $z(\tau_u+t)$ is a vector of time-varying quarter-by-year dummies, $ur(\tau_u+t)$ is the regional unemployment rate varying on a quarterly basis, $d_1(t)$ is the time-varying indicator for being part-time employed at unemployment duration t, $d_2(t)$ is the corresponding indicator for having completed at least one episode of PTE during the ongoing spell by unemployment duration t, $w(t)$ is the number of PTE weeks from the completed PTE episodes at unemployment duration t (note that $w(t) > 0$ if and only if $d_2(t) = 1$), and v_u is an unobserved heterogeneity term that captures the effect of unobserved skills, preferences and motivation.[6] The hazard function is the product of a baseline hazard function, $\lambda_u(t)$, describing the duration dependence, and a scaling function that captures the effects of both observed and unobserved characteristics, calendar time effects and the effect of PTE episodes.

The parameters of primary interest are α_1, the lock-in effect of PTE, and α_2 and μ, which jointly capture the stepping-stone effect of PTE. The stepping-stone effect depends on the "amount of treatment received" as measured by the number of weeks the worker has been part-time employed during the current spell of unemployment. In some specifications, we allow α_1, α_2 and μ to vary with the elapsed duration of unemployment t and/or calendar time $\tau_u+ t$.

In order to interpret α_1, α_2 and μ as causal effects, we need to take into account the potential endogeneity of the timing and duration of PTE episodes. Therefore we define T_p as the time until the beginning of the (next) PTE episode and T_d as the duration of PTE until return to unemployment benefits, both of which are continuous random variables. Workers may have several PTE episodes within the single unemployment spell, in which case T_p is measured from the end of the previous PTE episode. To deal with the endogeneity problem, we must base our statistical inference on the joint distribution of T_u, T_p and T_d. In order to do so, we specify MPH models also for T_p and T_d.

The hazard rate for transitions from unemployment benefits into PTE is specified as

[6] The hazard function $h_u(t|v_u)$ is conditional on the observed characteristics, calendar time at unemployment entry and PTE experiences in addition to the unobserved heterogeneity term. For ease of exposition, we emphasize conditioning on the unobserved heterogeneity term as it must be integrated out of the likelihood function while we ignore the other conditioning variables in the equations as we are always conditioning on them.

$$h_p(t|v_p) = \lambda_p(t)\phi_p(\tau_p - \tau_u + t)\exp\left\{x\beta_p + z(\tau_p + t)\delta_p + \theta_p ur(\tau_p + t) + \psi_p n(\tau_p - \tau_u) + v_p\right\},$$

where t is either the time since unemployment entry (for the first PTE episode) or the time since the end of the previous PTE episode (for the subsequent PTE episodes within the same unemployment spell), τ_p is either the calendar time at unemployment entry (for the first PTE episode, in which case $\tau_p = \tau_u$) or the calendar time at the end of the previous PTE episode (for the subsequent PTE episodes), and $n(\tau_p - \tau_u)$ is the number of past PTE episodes within the current unemployment spell. Note that $\lambda_p(t)$ describes the duration dependence in the time until the next PTE episode, whereas $\phi_p(\tau_p - \tau_u + t)$ captures the effect of elapsed unemployment duration.

Similarly, the hazard rate for transitions from PTE back into unemployment benefits is given by

$$h_d(t|v_d) = \lambda_d(t)\phi_d(\tau_d - \tau_u + t)\exp\left\{x\beta_d + z(\tau_d + t)\delta_d + ur(\tau_d + t)\theta_d + \psi_d n(\tau_d - \tau_u) + v_d\right\},$$

where t is the elapsed duration of PTE, τ_d is the calendar time at the beginning of the PTE episode, $n(\tau_d - \tau_u)$ is the number of PTE episodes preceding the ongoing PTE episode, and $\phi_d(\tau_d - \tau_u + t)$ is the effect of the elapsed duration of unemployment.

The unobserved heterogeneity terms, v_u, v_p and v_d, are allowed to be arbitrarily correlated to control for the potential endogeneity of PTE experiences. Abbring and Van den Berg (2003) show that random variation in the timing of treatments identifies the causal treatment effects without any exclusion restrictions under the assumptions that (1) the hazard rates are of the MPH form and that (2) individuals do not know their *exact* treatment times (starting and ending dates of PTE episodes in our case) in advance. Under these assumptions, the model is non-parametrically identified in the sense that no functional form assumptions on the baseline hazards or the distribution of unobservables are needed.

The first assumption is required to distinguish the effects of unobserved heterogeneity from other effects. In our application this assumption can be relaxed to some extent as we observe multiple unemployment spells for many individuals and our model includes time-varying covariates (time-varying quarter-by-year dummies and regional unemployment rate), both of which aid the identification, and thereby the identification does not hinge so much on the MPH assumption (Abbring and Van den Berg, 2003; Brinch, 2007; Gaure et al., 2007).

The second assumption implies that the unemployed do not know the exact starting dates of their future part-time jobs, nor the exact ending dates of ongoing part-time jobs in advance. In immediate proximity of the starting and ending days of part-time jobs this assumption is obviously violated, but as long as the individuals do not know these starting and ending dates too much in advance, this should not be a major problem. It should be stressed that the assumption does not require complete randomness, nor it rules out forward-looking behaviour. The only requirement is that there is *some* uncertainty in the timing of these events. In other words, the individuals can know at which probability they will find part-time work in the future and at which probability their ongoing part-time job will end at a given day in the future, and they can react on this information.

The contribution of a single individual with N unemployment spells to the likelihood function is given by

$$L = \iiint \left[\prod_{i=1}^{N} L_i \left(v_u, v_p, v_d \right) \right] dG\left(v_u, v_p, v_d \right), \tag{3}$$

where $L_i \left(v_u, v_p, v_d \right)$ is the likelihood of the i-th spell conditional on the unobserved heterogeneity terms and G is the joint distribution function of the heterogeneity terms. The unobserved heterogeneity terms are assumed to remain constant over different unemployment spells for the same individual. The structure of the conditional likelihood for a given spell depends on experiences of PTE within that spell. If the worker did not work part time during his or her i-th spell that ended at time t_{ui}, the conditional likelihood is

$$L_i \left(v_u, v_p, v_d \right) = h_u (t_{ui} | v_u)^{C_{ui}} \exp \left\{ - \int_0^{t_{ui}} h_u (u | v_u) du - \int_0^{t_{ui}} h_p (u | v_p) du \right\},$$

where C_{ui} equals 1 if the worker found a full-time job at time t_{ui}, and 0 otherwise (the spell was censored at that time). Instead, if the worker experienced two episodes of PTE during the i-th unemployment spell: the one that started at time $t_{p1,i}$ (measured from unemployment entry) and lasted for $t_{d1,i}$ days, and another that started at time $t_{p2,i}$ (measured from the end of the first PTE period) and lasted for $t_{d2,i}$ days, the conditional likelihood becomes

$$L_i\left(v_u, v_p, v_d\right) = h_p(t_{p1,i}|v_p)h_d(t_{d1,i}|v_d)\exp\left\{-\int_0^{t_{p1,i}} h_p(u|v_p)du - \int_0^{t_{d1,i}} h_d(u|v_d)du\right\}$$

$$\times h_p(t_{p2,i}|v_p)h_d(t_{d2,i}|v_d)^{C_{d2,i}}\exp\left\{-\int_0^{t_{p2,i,}} h_p(u|v_p)du - \int_0^{t_{d2,i}} h_d(u|v_d)du\right\}$$

$$\times h_u(t_{ui}|v_u)^{C_{ui}}\exp\left\{-\int_0^{t_{ui}} h_u(u|v_u)du\right\},$$

where $C_{d2,j}$ equals 1 if the worker moved back into unemployment benefits after the second part-time job, and 0 if the part-time job was still in progress at time t_{ui} when the unemployment spell terminated either because the worker found a full-time job (C_{ui} = 1) or because the spell was censored (C_{ui} = 0). Other possible cases are constructed in the similar manner.

For the baseline hazards we specify piecewise constant functions using 16 duration intervals for $\lambda_u(t)$ and $\lambda_p(t)$, and 10 for $\lambda_d(t)$. Similarly, we model the effects of elapsed unemployment duration in the hazard functions for PTE episodes, $\phi_p(\tau_p - \tau_u + t)$ and $\phi_d(\tau_d - \tau_u + t)$, using piecewise constant functions with 7 duration intervals.

Since the joint distribution of the unobserved heterogeneity terms G is not known, we approximate it in a non-parametric fashion using a trivariate discrete distribution. This is a very flexible approach because the discrete distribution can approximate any distribution arbitrarily well as the number of the points of support increases. In practice, we re-estimate the model many times, starting with 2×2×2 points of support and then adding support points until the likelihood function stabilizes. On the basis of the Akaike information criterion, we end up to the heterogeneity distributions with three or four points of support for each heterogeneity term.

6. Results

6.1. Empirical hazard rates

Before turning to the estimation results of the hazard models, it is useful to compare empirical exit rates to full-time employment between unemployed workers with different PTE experiences. Figure 3 shows smoothed weekly exit rates for those who have not been part-time employed by the week in question, for those who are currently part-time employed, as well as for those who have been part-time employed during the current spell but are not anymore. Differences in these hazard rates give us a hint about the sign and magnitude of the effect of PTE on the exit rate to FTE, and how this effect is likely to change over the course of the unemployment spell, and whether it differs between sexes.

[Insert Figure 3]

The three hazard rates show a similar pattern, without strong differences by gender. Within the first couple of weeks the hazard rates reach their highest values, implying that roughly 2% to 3% of the individuals who are not currently holding a part-time job exit to full-time employment each week. After the first ten weeks of unemployment or so, the hazard rates start to decline. They decline quite smoothly, and after one year of unemployment the hazard rates are only about one third of their peak values. It is also evident that those workers who have been part-time employed in the past exit to FTE at the highest rate among the three groups. Their exit rate over the first two years of unemployment is on average 64% (for men) or 99% (for women) higher than the exit rate of those who have not been part-time. Furthermore, their exit rate drops less steeply with unemployment duration, so that the relative difference in the exit rates between these two groups increases over the course of the unemployment spell (from 23% during the first two to six months to 96% during the second year of unemployment for men, and from 49% to 134% for women between the same periods). These findings suggest that having completed a spell of PTE improves chances of finding a full-time job, and that the long-term unemployed are likely to benefit from past PTE episodes the most. Stated differently, there seems to be a stepping-stone effect that is larger for the long-term unemployed.

On the other hand, part-timers move to full-time employment at a lower rate than unemployed who have not been part-time employed. This is true for both sexes, albeit the difference is rather small after six months of unemployment for women. During the first two years of unemployment, male part-timers exit to FTE at a 41% lower rate on average than men who have not been part-time employed. The exit rate of female part-timers is only 29% lower during the same interval. In other words, working part time seems to be associated with the lock-in effect, which is likely to be larger for men than women.

These differences in the empirical hazard rates cannot be interpreted as causal effects, since they can be driven by differences in observed and unobserved characteristics between the groups which stem from dynamic selection over the course of the unemployment spell. Next we turn to the results from the estimation of the multivariate mixed proportional hazard models presented in the previous section. Using these models we can control for differences in the background characteristics of the workers and take into account non-random selection in and out of part-time employment within the unemployment spells.

6.2. Baseline results

Columns 1 and 4 of Table 4 show estimates for the hazard rates from unemployment to FTE (outcome hazard), columns 2 and 5 from unemployment benefits to PTE (time to treatment hazard), and columns 3 and 6 from PTE to unemployment benefits (hazard for treatment duration). We have excluded from the table some parameter values due to the large number of them but they are available upon request.

[Insert Table 4]

Individuals below age 30 have a relatively high transition rate to both PTE and FTE, while individuals aged 50 and above have the lowest. In fact, the hazard rate to PTE declines more or less monotonically with age for both sexes, although the negative effect seems to be a bit larger for 30 to 39 years old women (as compared to men). Something similar happens with the hazard rate to FTE; in this case, it is even clearer, since the negative impact of age is non-existent for men until the age of 40. These findings may suggest that having young children in the family increases the value of women's non-market time or that women sometimes substitute unemployment benefits for family leave benefits. Also, it may be difficult to arrange day care, which can explain part of the larger negative effect on the hazard rates to FTE and PTE for women aged 30–39.

Immigrants are characterized by lower exit rates to both FTE and PTE than native Spanish workers. There are also clear gender differences in the impact of occupation. Compared to individuals in white-collar high-skilled occupations, male workers in (either blue- or white-collar) low-skilled occupations typically enter PTE at higher rates; for women, these categories have slightly lower exit rates to PTE than the reference group. For both genders, white-collar medium-skilled workers (basically, in administrative jobs) are the least prone to make a transition to PTE. As regards transitions to FTE, male workers who entered unemployment from blue-collar jobs have the highest hazard rates, while female workers in all other categories are less prone to exit than the ones in white-collar high-skilled occupations.

The impact of industry on exits to FTE and PTE is quite similar for both sexes with a few exceptions. Unemployed workers who were previously employed in education and health, other services, hotels and restaurants, real estate and renting, public administration and trade show particularly high hazard rates towards PTE, as compared to individuals working in manufacturing and energy. In the case of the transitions to FTE, the highest rates are found for education and health, real estate and renting, and hotels and restaurants, with public administration (for women) and construction (for men) showing also relatively high rates. Men with a financial intermediation background leave unemployment for FTE at a relatively high rate but are less likely to take up a part-time job, whereas their female counterparts exhibit the opposite pattern, being more likely exit to PTE but less likely to FTE.

The number of weeks spent in unemployment one, two and three years before the current unemployment spell have positive effects on the exit rates to both PTE and FTE, the effect being clearly higher for the most recent experience of unemployment. This somewhat strange finding is perhaps reflecting the large degree of worker turnover that characterizes the Spanish labour market (see García-Serrano and Malo, 2013). That is, there can be certain groups of workers who are moving between unemployment and short-time jobs on a regular basis. On the other hand, the unemployed who have spent more time in unemployment during the past year are likely to be entitled to UI benefits for a shorter time, so their higher exit rate to both FTE and PTE may also reflect the incentives associated with the potential duration of benefits.

From the last row of Table 4 we see that the number of past PTE episodes (within the ongoing unemployment spell) increases the transition rate from unemployment into PTE, and this effect is larger for men than for women. Therefore, workers previously involved in part-time work are more likely to repeat the experience in the future, perhaps because they are somehow trapped in a chain of part-time jobs and unemployment. Male part-timers with several PTE episodes in the past are also likely to return to unemployment benefits sooner than those with fewer PTE episodes. No such effect is found for women.

Not surprisingly, a higher local unemployment rate (conditional on the calendar time fixed effect) is associated with the lower exit rate to FTE. The higher unemployment rate induces unemployed men to take up part-time jobs but has no effect on women's exit rate to PTE. This may suggest that men consider part-time jobs as an alternative only when no full-time jobs are available. Both women and men experience shorter PTE spells when unemployment is at a high level.

In Figure 4 we show the effect of calendar time on job finding rates by plotting estimated coefficients on the time-varying quarter-by-year dummies, which are omitted from Table 4. These effects are expressed as proportional changes from the level of the hazard rates in the first quarter of 2005 (the omitted calendar time category in the model). Since the hazard models also include the time-varying regional unemployment rate as a control for the effect of local economic conditions, the calendar time effects aim to capture the effect of the general time trend in relative supply and demand of part-time and full-time jobs, which are partly driven by the policy reforms in 2009 and 2011.

A number of findings in Figure 4 are worth noting. First, within calendar years the exit rates are typically lowest in the first or fourth quarter and highest in the second quarter. This kind of seasonal variation is very useful from the methodological viewpoint as it aids identification, suggesting that our results do not hinge so much on the assumption of the MPH structure for the hazard functions (e.g. Brinch, 2007; Gaure et al., 2007). Second, the exit rate to FTE declined sharply in the second half of 2008, coinciding with the onset of the global financial crisis. Men's exit rate to FTE dropped by some 50% at that time, showing no signs of recovery by the end of 2013. Women's exit rate to FTE in 2009–2013 is very close to the reference level of the early 2005. Yet one should note that women's exit rate, unlike men's, exhibits an increasing trend over the period 2005–2007, so when compared to the average exit rate in 2007, which is about 22% above the 2005 level, women's chances to find full-time work were also hit by the 2008 crisis. Finally, among both women and men the exit rate to PTE compared to the exit rate to FTE has increased notably over time. This change has been more dramatic for men as their exit rate to PTE did not rise until 2009, whereas women's exit rate exhibits an increasing trend from 2005 onwards (yet there was a temporary drop in 2008 and 2009).

The key lesson from Figure 4 is that the gap between the hazard rates to PTE and FTE has widened over the years, but especially since 2010 and mainly due to the increase of the hazard rate to PTE. Also the hazard rate from PTE to unemployment benefits includes the time-varying year dummies.[7] Their coefficients are not reported in Table 4, nor plotted in Figure 4, but they imply that the exit rate from PTE back to unemployment benefits (holding the regional unemployment rate constant) increased by some 30% during the observation period for both sexes. Altogether these estimates imply that while finding a full-time job has become more difficult, part-time employment has become more common but less stable since the 2008 crisis.

[Insert Figure 4]

The parameters of primary interests – the lock-in effect (captured by a dummy variable indicating whether the individual is currently working part time) and the stepping-stone effect (captured by a dummy variable for having completed at least one episode of PTE and the cumulative duration of those episodes in weeks) of PTE on the hazard rate to FTE – are reported in the last rows in columns 1 and 4. It can be observed that part-time working causes a large reduction in the transition rate to FTE (compared to staying on unemployment benefits) which is of similar magnitude for both sexes (a reduction of 62% for women and 58% for men). This suggests that taking a part-time job crowds out the search for full-time work. However, it is also evident that having participated in PTE earlier in the unemployment spell enhances changes of finding a full-time job, suggesting that part-time working increases human capital, reduces the stigma of being unemployed and/or that part-time jobs are used by employers to screen potential applicants for full-time jobs. For men the stepping-stone effect does not depend on the time spent in PTE, and hence having been part-time employed in the past implies a constant increase of 18% in the exit rate to FTE. Somewhat surprisingly, the effect of the number of completed PTE weeks is negative for women. As such, one month in PTE increases women's exit rate to FTE by 13% whereas six months has a much smaller effect of 6%.

Recall that the empirical hazard rates in Figure 3 implied that the effect of PTE may differ between job seekers who have been unemployed for different lengths of time. The model specification in Table 4 ignores such a possibly. Moreover, the pattern of calendar time effects in Figure 4 and the policy reforms to foster the use of part-time

[7] Unlike for the other two hazard functions, we use the year dummies (not the quarter-by-year dummies) because of a much smaller number of observations on part-timers.

employment contracts suggest the possibility that the effect of PTE may not have been constant over time. The issue of impact heterogeneity is investigated in depth next.

6.3. Heterogeneous effects

We have estimated augmented models with heterogeneous treatment effects. These models are otherwise similar to those discussed above but now the time-varying PTE variables, $d_1(t)$, $d_2(t)$ and $w(t)$, are interacted with the linear time trend and/or the elapsed duration of the unemployment spell.[8] In this set of results, the effect of PTE can change over calendar time and/or differ between the long-term unemployed and those who entered unemployment quite recently. The heterogeneous effects of PTE on the exit rate to FTE from three different specifications are reported in Table 5. We do not show other parameter estimates as they are similar to those reported in Table 4 (yet they are available upon request). The models in columns 1, 2, 4 and 5 include the interactions either with the time trend or with the dummies for the elapsed duration of unemployment (6 to 12 months and over 12 months, so that those who have been unemployed for less than 6 months are the reference group), whereas the models in columns 3 and 6 include both sets of the interactions.

[Insert Table 5]

As seen in panel A, the lock-in effect of PTE has become slightly stronger over time. This change has been roughly the same for both women and men. According to the estimates in column 1 and 3, working part time reduced the exit rate to FTE by some 50% in the first quarter of 2005, and by 67% (for women) or by 62% (for men) in the last quarter of 2013. At the same time the stepping-stone effect of PTE has become weaker (panel B). For women the interaction term with the time trend does not differ significantly from 0 in column 1 but it becomes significant at the 5% level once the effects of PTE are allowed to vary also with the elapsed unemployment duration in column 3. The estimates in column 4, for instance, imply that a completed spell of PTE led to an increase of 45% in men's exit rate to FTE in the beginning of 2005, but that this stepping-stone effect dropped to below 10% by the last quarter of 2013. The larger lock-in effect together with the smaller stepping-stone effect in the later years implies that the potential of part-time jobs for acting as a bridge towards FTE for the unemployed has reduced over time.

[8] The time trend variable is 0 for the first quarter of 2005 and increases by 0.25 at the beginning of each subsequent quarter, reaching the value of 8.75 in the last quarter of 2013.

The effects of PTE also vary with the elapsed duration of unemployment among both women (columns 2 and 3) and men (columns 5 and 6). The lock-in effect is somewhat larger for those who have been unemployed over one year compared to those who have been unemployed for a shorter time. In the case of women, this is compensated by a larger stepping-stone effect for the long-term unemployed. According to the estimates in column 3, for a woman who has been unemployed over one year the stepping-stone effect is about twice the effect for an otherwise identical woman who has been unemployed for less than six months. There is no clear evidence that the stepping-stone effect for men would vary with the elapsed unemployment duration, albeit the coefficients on the interaction term with the dummy for unemployment spells longer than 12 months are relatively large in columns 5 and 6. Unlike in the baseline model for men in Table 3, the stepping-stone effect declines with the number of completed PTE weeks in columns in 4 and 5 of Table 4, but only over the first six months of unemployment.

6.4. The net effect of part-time employment

The hazard estimates above imply that the net effect of PTE depends on the relative magnitudes of the two opposite effects, both of which vary with the elapsed unemployment duration and over time. As such, the net effect of taking up a part-time job on the expected time until full-time employment is ambiguous, depending on the timing and duration of part-time work. To determine the size of the net effect in certain cases, we compute the difference in the expected remaining unemployment durations in counterfactual situations with and without a period of PTE as

$$E(T_u - t_p | \tau_u, t_p, t_d, T_u > t_p) - E(T_u - t_p | \tau_u, T_p = \infty, T_u > t_p),$$

where T_u is the duration of the unemployment spell until FTE, t_p denotes the beginning of a part-time job measured in weeks from unemployment entry at calendar time τ_u, and t_d is the (intended) duration of that job in weeks.[9] The first term in the equation is the expected remaining unemployment duration when the individual enters PTE at unemployment duration t_p and holds that part-time job for a maximum duration of t_d weeks. The second term is the expected remaining unemployment duration in the

[9] The individual may exit to FTE while working part time, in which case the realized duration of the part-time job will be shorter than t_d weeks.

counterfactual case of no PTE during the unemployment spell. Both these expected durations are conditional on the unemployment spell lasting over t_p weeks.

Using the model with the heterogeneous effects (columns 3 and 6 of Table 4), we compute the net effect of PTE in 18 different cases by varying calendar time of unemployment entry (January 1, 2005, or January 1, 2010), the timing of the part-time job within the unemployment spell (after 13, 26 or 52 weeks of unemployment) and the intended duration of the part-time job (7, 15 or 30 weeks). We compute these effects for each individual in the subsample of those who actually experienced PTE during their unemployment spell,[10] and then report the average of these individual-specific effects. That is, we focus on a relevant sample of those who truly received the "treatment" but consider the average effects of hypothetical treatments. In this subsample, the mean time until the first PTE episode is 26 weeks, and the mean duration of "completed" PTE spells followed by receipt of UI or UA benefits is 15 weeks,[11] so that the chosen values for t_p and t_d represent variation around the typical timing and duration of PTE episodes observed in the data.

[Insert Table 6]

The results are shown in Table 6 where the net effect of PTE on the expected remaining duration is reported in columns 3 and 6 for women and men, respectively. Assuming that all the unemployment spells started on January 1, 2005, we find that the part-time job with the intended duration of 7 weeks reduces the expected remaining unemployment duration by 2 to 5 weeks depending on the elapsed duration of unemployment at the beginning of the part-time job (panel A). In absolute terms, these effects are quite similar for women and men, but in relative terms, PTE has a larger impact for men as their counterfactual unemployment durations without PTE are shorter (column 4 vs. 1). In 2005, the relatively short spells of PTE thus seem to enhance finding a full-time job for both sexes; for example, if a part-time job begins after 26 weeks of unemployment, the reduction in the expected remaining time until FTE is 9% for women and 13% for men.

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[10] When computing the expected unemployment durations, we ignore the possibility of exiting to inactivity because we did not model such exits in econometric analysis but treat the spells followed by inactivity as censored observations. Moreover, we restrict the maximum unemployment duration to two years, so that we actually consider the expected value of min(T_u - t_p, 104 - t_p) given $T_u > t_p$. This is because very long unemployment spells were right-censored in the econometric analysis.

[11] The mean duration of all PTE spells, i.e. including also censored spells and those that were followed by FTE or inactivity, is about 24 weeks, as seen in Table 3.

By contrast, the longer PTE episodes of 15 or 30 weeks generally have an opposite effect, leading to longer unemployment spells on average. The only exception is a small reduction of 0.5 weeks in the expected unemployment duration for women whose part-time job with the intended duration of 15 weeks begins after 26 weeks of unemployment in column 3. The larger net effects on the expected unemployment duration for longer PTE episodes arise because longer part-time jobs come with the cost of larger cumulative lock-in effects without improving job finding prospects afterwards in terms of the larger stepping-stone effect. If anything, the stepping-stone effect declines with the time spent in PTE. It is noteworthy that the relative net effects of PTE episodes with the intended durations of 15 and 30 weeks are larger for men. For example, a part-time job with the potential duration of 30 weeks starting after 26 weeks of unemployment increases the expected remaining time until FTE by 9% for women and 18% for men. Thus, while women benefit less than men (in terms of reduction in expected unemployment duration) from short part-time jobs of 7 weeks, they suffer less (in terms of an extension in expected unemployment duration) from longer part-time jobs of 15 and 30 weeks.

Panel B presents the results when the unemployment spells are assumed to begin in 2010, i.e. in the aftermath of the financial crisis and after the 2009 policy reform to foster the use of part-time employment contracts. In this case, the net effects of PTE on the expected unemployment duration are uniformly larger compared to the 2005 cases in panel A and almost always positive owing to larger lock-in effects and smaller stepping-stone effects as both of these effects were found to decline with calendar time. Taking up a part-time job enhances the chances of finding a full-time job only for those who have been unemployed for at least six months and only if the part-time job in question is very short, lasting for 7 weeks at maximum; and even in these cases, the net effect on the expected unemployment duration is very close to zero. In all other cases, working part time increases the expected time until FTE; and this effect is roughly of the same size for men and women (in absolute and relative terms), and relatively large when the intended duration of the part-time job is 15 or 30 weeks. As such, PTE has not helped the unemployed to obtain a full-time job faster in the later years of our observation period.

It is noteworthy that the increase in the expected time until FTE due to part-time working is always much smaller than the intended length of the part-time job, being one third or less of the potential part-time job duration. Also, the increase is typically relatively small compared to the counterfactual unemployment duration. It follows that taking up a part-time job is likely to increase the number of working hours for the unemployed (assuming that part-time hours are no less than one third of full-time hours) even in the cases of longer part-time jobs in the later period. As such, it might be beneficial for unemployed workers to accept a part-time job when no full-time jobs are readily available even though that would increase the expected time until FTE somewhat. And this argument for part-time working is stronger when partial unemployment benefits can be received on top of wage income from the part-time job.

7. Conclusions

This study has provided evidence on the impact of taking up a part-time job on the expected time until full-time employment among unemployed job seekers in the Spanish labour market. Our findings show that current participation in PTE causes a reduction in the exit rate to FTE (lock-in effect) but that having participated in PTE earlier in the unemployment spell brings about an increase in the exit rate (stepping-stone effect), implying a trade-off between the two opposite effects. Both of these effects are roughly of similar magnitude for women and men. The size of the stepping-stone effect does not increase (or may even decrease) with the time spent in PTE. As a consequence, the net effect of PTE on the expected time until FTE increases with the duration of PTE episodes due to the increasing cumulative lock-in effect. Indeed, our results show that only relatively short PTE spells reduce the expected time until FTE, whereas longer PTE spells prolong the expected time without a full-time job.

We found that both the lock-in and stepping-stone effect vary with the elapsed duration of unemployment and that they have also changed over the years. In particular, the lock-in effect gets stronger after one year of unemployment but, on the other hand, the long-term unemployed (at least in the case of women) also benefit more from PTE spells afterwards, which in large part mitigates the effect of the larger lock-in effect for the long-term unemployed. It follows that the net effect of PTE on the expected time until FTE does not vary much between individuals who have been unemployed for different lengths of time.

Furthermore, our results reveal that the lock-in effect has got stronger but the stepping-stone weaker over the years. By implication, the net effect of PTE on the expected unemployment duration has increased over time. In recent years, the net effect has been almost uniformly positive, implying that taking up a part-time job generally increases the expected time until full-time employment. To some extent the deterioration in the potential of PTE as a bridge towards FTE is likely to be due to the policy reforms implemented in 2009 and 2011 to foster the use of part-time employment contracts. On the other hand, mass unemployment and a long period without economic growth that followed the 2008 financial crisis may have contributed to the outcome as well.

The main lesson of our analysis is that only relatively short PTE spells tend to reduce the expected time until FTE. This gives support for a policy that encourages unemployed workers to take up short part-time jobs when no full-time jobs are available and employers to offer such jobs when they are unable to offer full-time work. But it seems important that these employment arrangements are short-lived, temporary responses to poor economic conditions. As such, the current reductions in the employer's social security contributions for part-time workers up to three years might be too generous, leading to too long periods of part-time work. A tighter time-limit for the social security cost reduction might be in order.

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Table 1. Share of part-time employment and involuntary part-time work in selected European countries (Source: Eurostat)

	Total			**Women**			**Men**		
	2006	2010	2014	2006	2010	2014	2006	2010	2014
A. Part-time employment as percentage of total employment									
EU (28)	17.5	18.5	19.6	30.5	31.3	32.2	6.9	7.9	8.8
Denmark	23.0	25.6	24.6	35.0	38.1	35.0	12.3	14.0	15.2
Germany	25.2	25.5	26.5	45.4	45.0	46.3	8.5	8.7	9.2
Spain	11.6	12.9	15.8	22.4	22.6	25.5	4.2	5.2	7.7
France	17.1	17.6	18.6	30.2	30.0	30.5	5.6	6.4	7.4
Italy	13.1	14.8	18.1	26.3	28.8	32.1	4.3	5.1	7.8
Netherlands	45.8	48.3	49.6	74.5	76.2	76.7	22.1	24.2	26.1
Portugal	8.2	8.5	10.1	12.8	12.4	12.6	4.2	5.0	7.6
Sweden	24.3	25.8	24.6	39.7	40.3	37.3	10.6	12.7	12.8
UK	24.2	25.7	25.3	41.6	42.3	41.3	9.1	11.0	11.2
B. Involuntary part-time employment as percentage of part-time employment									
EU (28)	22.7	27.0	29.6	20.4	24.3	26.3	31.1	36.2	40.2
Denmark	15.2	15.6	16.9	16.5	15.7	18.3	11.9	15.4	14.0
Germany	23.1	21.9	14.5	20.1	18.7	12.8	38.5	37.8	22.3
Spain	33.8	50.1	64.0	33.9	48.7	61.8	33.6	55.1	70.0
France	30.8	34.8	42.4	29.8	33.9	40.8	35.6	38.8	48.9
Italy	37.8	50.2	65.4	34.3	46.6	60.4	52.3	64.4	80.6
Netherlands	6.2	5.7	10.9	5.8	5.1	9.4	7.8	7.5	15.2
Portugal	34.5	42.1	49.3	36.3	43.8	53.7	29.5	38.3	42.3
Sweden	24.9	28.1	29.8	24.8	27.7	29.1	25.3	29.3	31.8
UK	9.5	18.8	18.8	7.1	13.9	14.0	18.6	35.3	35.1

Note: The figures for the UK in 2010 correspond to 2011.

Table 2. Sample means for unemployment spells by gender and part-time employment status (Source: MCVL database)

	Women			Men		
	All	w/o PTE	w/ PTE	All	w/o PTE	w/ PTE
	(1)	(2)	(3)	(4)	(5)	(6)
Age group, %						
< 30	25.5	24.6	33.4	22.8	22.6	28.3
30-39	37.2	37.0	39.1	35.1	34.9	39.0
40-49	23.9	24.3	20.3	25.7	25.8	22.9
50+	13.5	14.1	7.1	16.4	16.7	9.8
Immigrant, %	10.5	10.6	9.5	14.7	14.8	13.4
Occupation group, %						
White-collar high-skill	10.4	10.5	9.8	5.2	5.2	6.0
White-collar medium-skill	5.2	5.3	4.7	5.8	5.9	5.2
White-collar low-skill	36.4	35.9	41.2	12.2	11.9	17.7
Blue-collar high-skill	11.5	11.6	10.6	39.4	39.7	31.8
Blue-collar medium-skill	14.3	14.3	13.7	14.6	14.6	15.1
Blue-collar low-skill	22.2	22.5	20.0	22.8	22.7	24.3
Industry, %						
Manufacturing and energy	13.1	13.4	9.6	17.4	17.6	13.9
Construction	1.9	1.9	1.7	18.9	19.1	12.6
Trade	21.2	20.8	25.3	12.1	12.0	15.8
Hotels and restaurants	16.4	16.5	15.2	7.5	7.2	12.9
Transport	5.3	5.4	4.6	8.5	8.4	9.1
Financial intermediation	5.4	5.4	5.6	19.3	19.7	11.7
Real estate and renting	10.5	10.3	11.8	8.0	7.8	11.2
Public administration	8.0	8.2	6.3	3.6	3.6	4.3
Education and Healt	12.9	12.8	13.6	2.2	2.1	4.5
Other services	5.3	5.2	6.3	2.5	2.5	3.8
Regional unemployment rate	16.5	16.5	16.8	17.6	17.5	19.2
Fraction of time unemployed, %						
0-1 years ago	10.4	10.4	10.1	12.8	12.8	13.3
1-2 years ago	10.1	10.2	9.4	11.0	11.0	11.2
2-3 years ago	8.7	8.8	7.7	8.6	8.6	8.4
At least one part-time job, %	9.8	0.0	100.0	4.2	0.0	100.0
Number of part-time jobs	0.1	0.0	1.2	0.0	0.0	1.2
Part-time work, weeks	3.1	0.0	31.8	1.1	0.0	26.3
Unemployment duration, weeks	32.1	28.5	65.2	28.8	27.3	61.1
Exit to full-time work, %	45.0	46.5	30.8	52.5	53.3	34.2
Number of spells	140,612	126,811	13,801	258,459	247,503	10,956

Notes: Columns 2 and 5 show means for spells that do not include episodes of part-time employment and columns 3 and 6 for spells with at least one episode of part-time employment. Data includes 209,342 individuals, of whom 79,312 are females and 130,030 are males.

Table 3. Sample means for part-time employment spells by gender (Source: MCVL database)

	All (1)	**Women** (2)	**Men** (3)
Occupation group, %			
White-collar high-skill	7.6	9.0	5.9
White-collar medium-skill	4.1	4.1	4.0
White-collar low-skill	34.2	43.5	22.0
Blue-collar high-skill	16.7	9.9	25.7
Blue-collar medium-skill	15.1	14.2	16.2
Blue-collar low-skill	22.3	19.4	26.1
Industry, %			
Manufacturing and energy	4.8	3.5	6.4
Construction	3.7	0.9	7.2
Trade	18.8	22.2	14.3
Hotels and restaurants	18.1	17.2	19.2
Transport	6.4	3.6	10.1
Financial intermediation	4.3	3.3	5.6
Real estate and renting	19.2	20.6	17.3
Public administration	5.7	5.3	6.3
Education and Healt	12.8	16.9	7.4
Other services	6.0	6.3	5.6
Part-time work duration, weeks	24.1	25.6	22.1
Exit to UI or UA, %	49.3	48.3	50.4
Exit to full-time work, %	15.2	14.9	15.7
Number of spells	29,284	16,613	12,671

Table 4. Baseline estimates

	Women			Men		
	U -> FTE	UB -> PTE	PTE -> UB	U -> FTE	UB -> PTE	PTE -> UB
	(1)	(2)	(3)	(4)	(5)	(6)
Age (vs. below 30)						
30-39	-0.258***	-0.468***	0.019	-0.009	-0.302***	0.040
	(0.014)	(0.022)	(0.033)	(0.010)	(0.027)	(0.039)
40-49	-0.274***	-0.677***	0.086**	-0.169***	-0.681***	0.050
	(0.016)	(0.027)	(0.038)	(0.011)	(0.030)	(0.043)
50+	-0.597***	-1.285***	-0.034	-0.679***	-1.287***	0.003
	(0.019)	(0.036)	(0.056)	(0.013)	(0.038)	(0.057)
Immigrant	-0.128***	-0.071**	-0.062	-0.151***	-0.059*	-0.068
	(0.022)	(0.033)	(0.050)	(0.012)	(0.032)	(0.049)
Occupation (vs. White-collar high-skill)						
White-collar medium-skill	-0.535***	-0.326***	-0.121	-0.025	-0.207***	0.281***
	(0.029)	(0.052)	(0.080)	(0.022)	(0.062)	(0.098)
White-collar low-skill	-0.488***	-0.082**	0.116**	-0.080***	0.203***	0.270***
	(0.019)	(0.034)	(0.053)	(0.019)	(0.050)	(0.077)
Blue-collar high-skill	-0.442***	-0.078*	0.249***	0.286***	0.042	0.447***
	(0.024)	(0.043)	(0.066)	(0.018)	(0.048)	(0.075)
Blue-collar medium-skill	-0.415***	-0.060	0.264***	0.184***	0.111**	0.446***
	(0.023)	(0.042)	(0.062)	(0.019)	(0.053)	(0.080)
Blue-collar low-skill	-0.455***	-0.074*	0.218***	0.089***	0.199***	0.472***
	(0.021)	(0.038)	(0.058)	(0.018)	(0.050)	(0.077)
Industry (vs. Manufacturing and energy)						
Construction	-0.273***	-0.031	-0.054	0.194***	-0.037	0.205***
	(0.044)	(0.074)	(0.116)	(0.011)	(0.037)	(0.059)
Trade	-0.081***	0.422***	-0.025	-0.080***	0.347***	-0.025
	(0.018)	(0.033)	(0.050)	(0.013)	(0.037)	(0.057)
Hotels and restaurants	0.256***	0.567***	0.048	0.225***	1.094***	0.076
	(0.020)	(0.036)	(0.055)	(0.016)	(0.042)	(0.059)
Transport	0.084***	0.181***	-0.008	0.156***	0.440***	-0.009
	(0.026)	(0.050)	(0.076)	(0.014)	(0.042)	(0.064)
Financial intermediation	-0.207***	0.172***	-0.150**	0.183***	-0.210***	0.220***
	(0.028)	(0.047)	(0.072)	(0.011)	(0.039)	(0.060)
Real estate and renting	0.261***	0.585***	0.136**	0.257***	0.679***	0.169***
	(0.019)	(0.037)	(0.055)	(0.014)	(0.040)	(0.060)
Public administration	0.342***	0.383***	-0.046	0.003	0.526***	0.166**
	(0.022)	(0.045)	(0.070)	(0.019)	(0.054)	(0.083)
Education and Health	0.650***	0.823***	-0.073	0.380***	1.293***	0.041
	(0.020)	(0.039)	(0.059)	(0.023)	(0.058)	(0.085)
Other services	-0.053**	0.488***	-0.121*	-0.009	0.704***	0.114
	(0.027)	(0.045)	(0.068)	(0.023)	(0.057)	(0.084)
Fraction of time unemployed						
0-1 years ago	0.925***	0.834***	0.277***	0.720***	0.684***	0.263***
	(0.024)	(0.047)	(0.065)	(0.015)	(0.047)	(0.065)
1-2 years ago	0.525***	0.427***	0.115	0.357***	0.360***	0.048
	(0.026)	(0.049)	(0.071)	(0.017)	(0.052)	(0.074)
2-3 years ago	0.601***	0.406***	0.142**	0.363***	0.330***	0.215***
	(0.026)	(0.048)	(0.072)	(0.018)	(0.051)	(0.077)
Regional unemployment rate	-0.020***	-0.003	0.017***	-0.011***	0.014***	0.023***
	(0.002)	(0.003)	(0.005)	(0.001)	(0.004)	(0.006)
Part-time employed	-0.968***			-0.860***		
	(0.023)			(0.025)		
At least one part-time job	0.133***			0.163***		
	(0.035)			(0.035)		
Part-time work in weeks	-0.003**			-0.001		
	(0.001)			(0.002)		
Number of part-time jobs		0.152***	0.011		0.225***	0.056***
		(0.007)	(0.017)		(0.013)	(0.020)

Notes: Columns 1 and 4 show estimates for hazard rates from unemployment to FTE, columns 2 and 5 from unemployment benefits to PTE, and columns 3 and 6 from PTE to unemployment benefits. All hazards also include regional dummies, time-varying quarter-by-year effects, and parameters for duration dependence and unobserved heterogeneity. Standard errors in parentheses. Significance levels: *** 1%, ** 5% and * 10%.

Table 5. Heterogeneous effects of part-time employment on the hazard rate from unemployment to full-time employment

	Women			Men		
	(1)	(2)	(3)	(4)	(5)	(6)
A. Lock-in effect						
Part-time employed	-0.707***	-0.943***	-0.753***	-0.702***	-0.816***	-0.669***
	(0.049)	(0.034)	(0.053)	(0.060)	(0.038)	(0.065)
x Linear trend	-0.047***		-0.041***	-0.030***		-0.027***
	(0.009)		(0.009)	(0.010)		(0.010)
x Unemployed 6-12 months		0.062	0.077		-0.026	-0.020
		(0.051)	(0.051)		(0.055)	(0.055)
x Unemployed > 12 months		-0.119**	-0.093*		-0.169***	-0.145**
		(0.054)	(0.054)		(0.058)	(0.059)
B. Stepping-stone effect						
At least one part-time job	0.328***	0.128*	0.256**	0.375***	0.199***	0.450***
	(0.088)	(0.073)	(0.105)	(0.110)	(0.072)	(0.123)
x Linear trend	-0.023		-0.032**	-0.035**		-0.046***
	(0.015)		(0.015)	(0.017)		(0.017)
x Unemployed 6-12 months		0.199**	0.187*		0.026	0.028
		(0.098)	(0.098)		(0.094)	(0.094)
x Unemployed > 12 months		0.252***	0.256***		0.138	0.152*
		(0.094)	(0.094)		(0.091)	(0.091)
Part-time work in weeks	-0.001	-0.025**	-0.022*	-0.004	-0.029**	-0.035**
	(0.004)	(0.012)	(0.012)	(0.005)	(0.013)	(0.014)
x Linear trend	-0.000		-0.000	0.001		0.001
	(0.001)		(0.001)	(0.001)		(0.001)
x Unemployed 6-12 months		0.014	0.013		0.025*	0.027**
		(0.013)	(0.013)		(0.014)	(0.014)
x Unemployed > 12 months		0.019	0.018		0.026*	0.028**
		(0.012)	(0.012)		(0.013)	(0.013)

Notes: Standard errors in parentheses. Significance levels: *** 1%, ** 5% and * 10%.

Table 6. The effect of a part-time job on expected remaining time until full-time employment

		Women			Men		
Timing of PTE spell	Duration of PTE spell	Remaining time until FTE w/o PTE	Remaining time until FTE w/ PTE	Difference	Remaining time until FTE w/o PTE	Remaining time until FTE w/ PTE	Difference
t_p	t_d	(1)	(2)	(3)	(4)	(5)	(6)
A. Unemployment starting in 2005							
13	7	52.1	48.2	-3.8	33.9	30.1	-3.7
13	15	52.1	52.5	0.5	33.9	35.1	1.2
13	30	52.1	58.5	6.5	33.9	42.2	8.3
26	7	50.0	45.6	-4.4	36.0	31.2	-4.8
26	15	50.0	49.5	-0.5	36.0	36.2	0.2
26	30	50.0	54.3	4.3	36.0	42.4	6.4
52	7	38.0	35.6	-2.4	30.0	26.6	-3.4
52	15	38.0	38.8	0.8	30.0	31.3	1.3
52	30	38.0	43.4	5.4	30.0	37.6	7.6
B. Unemployment starting in 2010							
13	7	56.3	57.0	0.7	47.8	47.9	0.2
13	15	56.3	61.4	5.0	47.8	52.1	4.4
13	30	56.3	66.9	10.5	47.8	58.3	10.5
26	7	54.9	54.0	-0.9	48.5	47.4	-1.1
26	15	54.9	57.5	2.6	48.5	51.3	2.8
26	30	54.9	61.6	6.7	48.5	55.8	7.4
52	7	41.4	40.9	-0.5	38.6	37.9	-0.8
52	15	41.4	43.4	2.0	38.6	40.8	2.2
52	30	41.4	46.8	5.3	38.6	44.7	6.1

Notes: t_d denotes the start of a part-time job spell measured in weeks from the beginning of the unemployment spell, and t_d denotes the potential duration of the part-time job in weeks. The expected remaining time until full-time employment without part-time working (i.e. the counterfactual unemployment duration) is shown in columns 1 and 4. The expected duration when a part-time job starts at unemployment duration t_p that potentially lasts for t_d weeks is shown in columns in 2 and 5. Difference in columns 3 and 6 is the effect of such a part-time job on the expected remaining duration until full-time employment. All the unemployment durations are conditional on the unemloyment spell being no shorter than t_p and subject to an overall maximum of two years. Panel A reports the results when all spells begin on January 1, 2005, and Panel B when the spells begin on January 1, 2010. The average unemployment durations are computed over a subgroup of those workers who experienced at least one PTE episode during their unemployment spell. The calculations are based on results from model specifications with heterogeneous effects of PTE reported in columns 3 and 6 of Table 4.

Figure 1. Overall unemployment inflow composed by gender (Source: MCVL database)

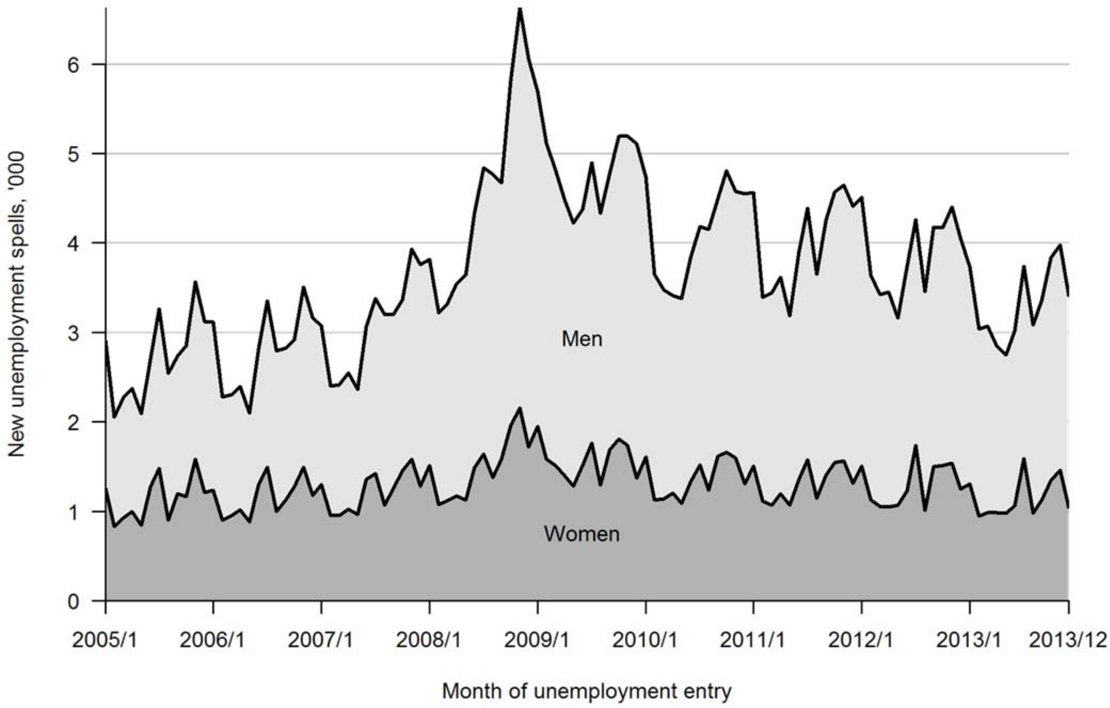


Figure 2. Distribution of part-time job duration by gender (Source: MCVL database)

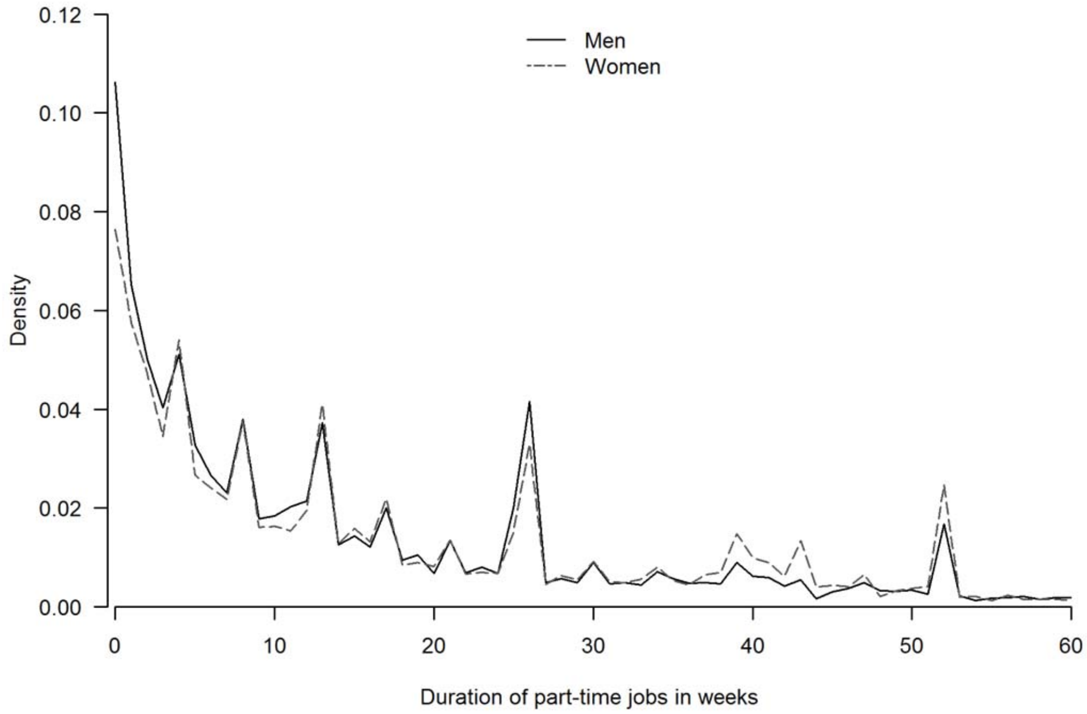
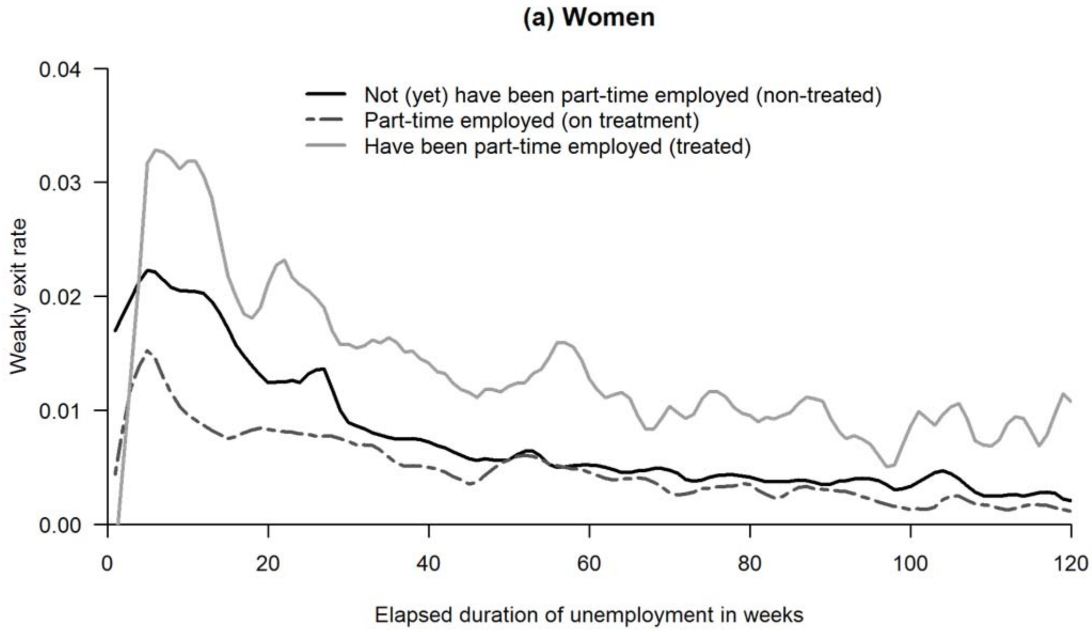


Figure 3. Smoothed job finding rate by gender and part-time employment status (Source: MCVL database)



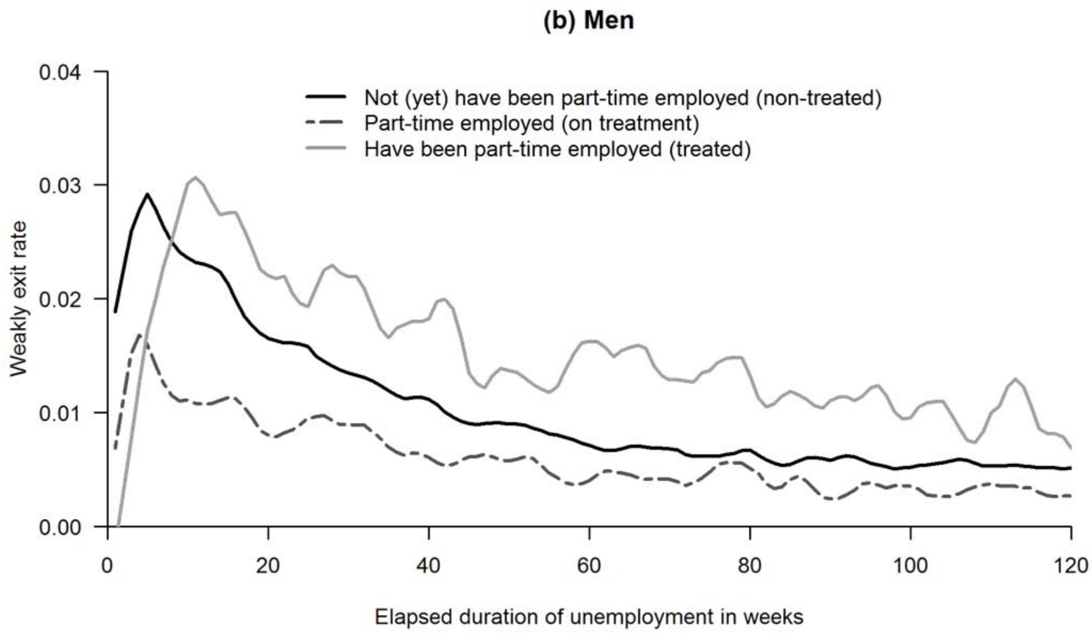


Figure 4. Proportional effect of calendar time on hazard rates from unemployment to full-time and part-time employment along 95% confidence interval by gender