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Re-employment Rates of Older Unemployed Workers: Decomposing the Effect of Birth Cohorts and Policy Changes

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Abstract In many European countries, re-employment probabilities of older unemployed workers are relatively low. While there is evidence that financial incentives and search obligations are effective to increase the job prospects of older workers, recent research also stresses the importance of birth cohort effects. These cohort effects may in turn stem from higher educational attainment levels and better health conditions of future generations of older workers. This paper empirically assesses the relative importance of both explanations, using a registered data set of unemployment insurance spells between 1999 and 2008 for the Netherlands. Using a Linear Probability Model, we decompose the effects of birth cohorts, age, calendar time and two policy measures that were targeted at older unemployed workers—i.e. increased job search obligations in 2004 and shorter potential benefit durations (PBD) in 2006. We find that policy effects predominantly explain the increased job return rates of unemployed of 55 years and older from 1999 to 2008. The introduction of search requirements has increased the one-year re-employment probability of eligible older men with about 5 % point, while the reduction in PBD has caused the one-year re-employment probability of eligible men to increase with 3 % point.

Keywords Older workers · Unemployment · Job search · Longitudinal data

JEL Classification J64 · J65 · C23

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1 Introduction

In many European countries, older unemployed workers have low re-employment probabilities. The Netherlands is no exception to this. At first sight, the evolution of employment rates—with an increase in employment rates of about 25 % point in 1998–2012—and the relatively low unemployment rates among workers who are older than 55 cause no major reason for concern (see Fig. 1). However, for those older workers that lose their job, unemployment is persistent. In particular, we see that in 2010 more than half of the population of unemployed older workers experienced unemployment spells that lasted longer than one year (OECD 2011). Using data on all unemployment spells between 1999 and 2008, CBS (2012) also finds that 60 % of the unemployed workers aged 45 return to work within a year's time, while this is only 20 % for 60-year old workers.

Not surprisingly, in the last decade policymakers have responded by introducing targeted activating measures to enhance the re-employment rate of older workers. In this respect, the two most important measures were the introduction of search obligations for workers who are older than 57.5 of age in 2004, and the reduction in potential benefit durations (PBD) of Unemployment Insurance (UI) in 2006. But while these policies may have been effective and well-targeted instruments to improve the labor market prospects of older workers, the question arises whether further cuts in the benefit conditions are preferred in all circumstances. Opponents may argue that the labor market position of older unemployed workers will also improve without any policy changes. The general idea is that birth cohort effects will lead to future cohorts of older workers with higher educational attainment and better health conditions. Accordingly, there will be less need for activating policies for older workers in the future. For an assessment of the use and usefulness of activating policies for older workers, one thus needs to analyze both policy and birth cohort effects.

This paper is the first to provide an integral empirical assessment of the effects of policy changes that have taken place in the Netherlands, as well as the importance of birth cohort effects to explain the re-employment probabilities of older workers. We use registered longitudinal data that are informative on all unemployment spells in the

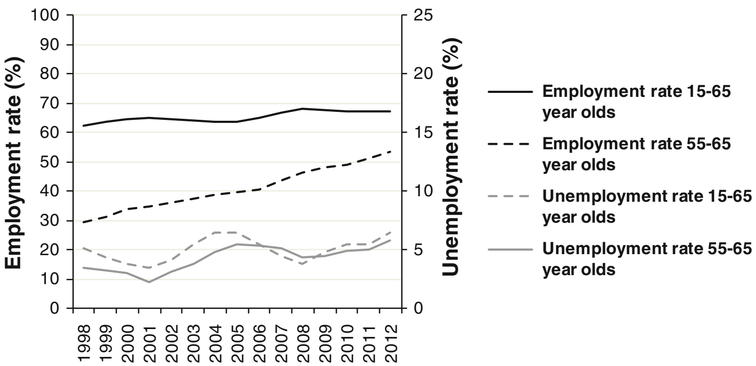

Fig. 1 Employment and unemployment rates in the Netherlands as percentage of labor supply, 1999–2012: all workers and older workers (55–65)

period 1999–2008, amounting to more than 1.9 million individual entries. With the data, we estimate Linear Probability Models (LPM) for the one-year re-employment probability as a function of birth cohort, age category, year of UI inflow and some observed individual characteristics (i.e., the ‘Age Period Cohort’ model). We also include two policy parameters that are particularly relevant for older workers, namely the (re-)introduction of search requirements for individuals aged 57.5 and older in 2004, and the reduction in the PBD from 60 to 38 months in 2006. In the time period under consideration, these measures were the most prominent ones to increase the job prospects of older workers in the Netherlands.

With our analysis, we essentially contribute to two strands of literature. First, there is a small yet growing literature that analyzes the responsiveness of older unemployed workers to activating policies and incentives. So far, evidence of the effects of the PBD reduction on older workers is limited to the case of Austria (Lalive 2008). Lalive (2008) finds that an extension of the UI benefit length in Austria with 179 weeks prolongs the UI duration of eligible workers by 15 weeks. To study the effects of the search requirements on older workers, however, we use similar data and a roughly similar design as Bloemen et al. (2013) and Hullegie and van Ours (2013). Bloemen et al. (2013) study how the introduction of search requirements for unemployed workers aged 57.5 and older in the Netherlands in 2004 affected their re-employment rates. The authors use a difference-in-difference approach and find a 6 % point increase in the outflow to jobs, but also an increase in the inflow in Disability Insurance by 2.5–4 % point. Hullegie and van Ours (2013) find that the introduction of the search requirement not only increases the re-employment rate of unemployed workers that are older than 57.5 years of age, but also those that are younger. Thus, it seems that the search exemption also resulted in anticipation effects among workers that had not reached the age of 57.5 yet. In a similar vein, Bennmarker et al. (2012) also find anticipation effects for older workers trying to avoid the inflow in mandatory activation programs.

The second strand of literature we contribute to concerns the role of birth cohort effects. While our focus is on the re-employment rates of older workers, this literature typically addresses participation rates. With data from the Dutch Labor Force Survey, Deelen and van Vuuren (2009) stress the importance of birth cohort effects in explaining the future rise in participation rates of older workers. Likewise, Euwals et al. (2011) find birth cohort effects to be the main driver of the increased participation rates of women in the Netherlands. Gárcia-Gómez et al. (2010) focus on improvements in health among subsequent birth cohorts as an explanation for the increase in participation rates among older workers in the Netherlands. They conclude that mortality rates of cohorts born after 1950 are much lower at any given age than for cohorts born before 1950, which hints at improvements in health among the working age population.

Combining these two strands of literature, the picture that emerges is that both activating policies and birth cohort effects may contribute to the job prospects of older unemployed workers. The evidence so far suggests that older unemployed workers are responsive to search obligations and financial incentives, just like workers of younger ages. And although the birth cohort studies explain participation rates and not the re-employment rates of older workers, it is likely that these are related. This suggests that the labor market position of older workers may improve in the near future, with educational attainment and health as its most probable drivers.

According to our analysis, the effects of these policy changes largely explain the increased job return rates of unemployed of 55 years and older that is observed in the time period under investigation. In particular, the introduction of search requirements for unemployed workers aged 57.5 and older in 2004 has increased the one-year re-employment probability of eligible men with 5 % point, while the reduction in the maximum duration of UI has caused the one-year re-employment probability of eligible men to increase with 3 % point. Birth cohort effects also seem to matter, with effect estimates amounting to 4 % point.

The remainder of this paper is organized as follows. Section 2 presents our data along with some descriptive statistics. Section 3 lays out our empirical approach, the estimation results and presents a range of sensitivity analyses to check the robustness of our findings. Finally, Section 4 concludes.

2 Data

In our analysis, we use registered longitudinal unemployment data from CBS (2012), with the re-employment probability one year after the start of UI benefits as our primary variable of interest. The data set covers all unemployment spells between 1999 and 2008, which amounts to 1,903,955 yearly individual observations in total. In order to obtain sufficiently accurate control variables, we choose to exclude unemployment spells of individuals who: (i) also had a (small) job at the time of UI inflow; (ii) received sickness or disability insurance payments prior to UI; and (iii) were younger than 17. These limitations result in a loss of 527,430 observations. Our selected sample contains 793,196 unemployment spells of men and 538,329 unemployment spells of women.

For each unemployment spell in our sample, variables are observed over the period 1999–2008, with the exception of gross personal yearly income (only as of 2003).[1] Table 1 presents summary statistics of our sample, stratified by gender and age category. To start with, we see that re-employment rates among older workers are much lower than among prime age and younger workers. After one year, only 18 percent (16 percent) of the older men (women) have found re-employment, and the shares hardly increase when extending the time window to 18 months.

Figures 2 and 3 illustrate the (strong) negative relationship between age and the re-employment rate for men and women, respectively. Interestingly, the one-year re-employment probability of men decreases monotonously with age, whereas the one-year re-employment probability of women declines somewhat faster after their mid-twenties. As to the latter, the decline in the one-year re-employment may be the result of mothers returning to the labor market that have lost part of their network and their skills. For higher ages, however, the slopes are rather similar, that is, for both men and

[1] See Table 5 in the “Appendix” for a complete list of variables.

Table 1 Summary statistics for selected sample of unemployment spells (1999–2008)

Age category	Men			Women		
	17–24	25–54	55–64	17–24	25–54	55–64
Fraction						
Immigrant	0.27	0.29	0.18	0.25	0.27	0.18
Married	0.04	0.44	0.76	0.11	0.45	0.62
Single	0.15	0.26	0.14	0.16	0.19	0.25
Children in household	0.72	0.49	0.31	0.58	0.55	0.21
Re-employment						
Within 6 months	0.60	0.45	0.16	0.54	0.39	0.13
Within 12 months	0.70	0.54	0.18	0.69	0.52	0.16
Within 18 months	0.75	0.61	0.21	0.72	0.57	0.19
Mean						
UI spells per worker	1.16	1.37	1.32	1.14	1.31	1.30
UI: hours per week	35	36	36	30	28	24
Maximum UI duration (in months)	8	19	37	7	17	33
Duration of last job (in months)	12	34	119	12	32	81
Gross personal yearly income (in euro's)	17,938	38,589	47,823	15,236	25,677	26,564
Observations	107,123	605,937	80,136	95,998	453,387	33,944

Immigrants are defined to include individuals who were born in the Netherlands but have at least one parent who was born abroad. Gross personal yearly income is documented as of 2003 so the averages are based on a smaller sample size

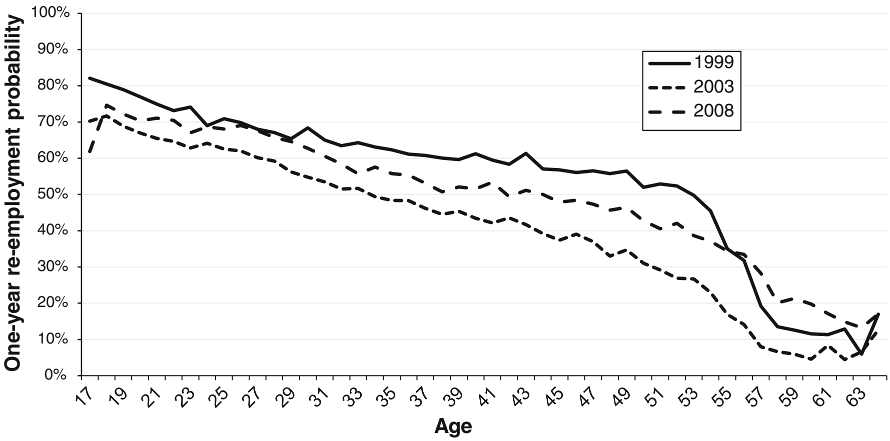

Fig. 2 One-year re-employment probability of men by age for 1999, 2003 and 2008

women it holds that the one-year re-employment probability decreases rapidly after age 50. Finally, at age 64, when men and women are close to the statutory retirement age of 65, the one-year re-employment probability experiences a short, albeit small revival. This phenomenon can be due to selection effects: at age 64, only motivated

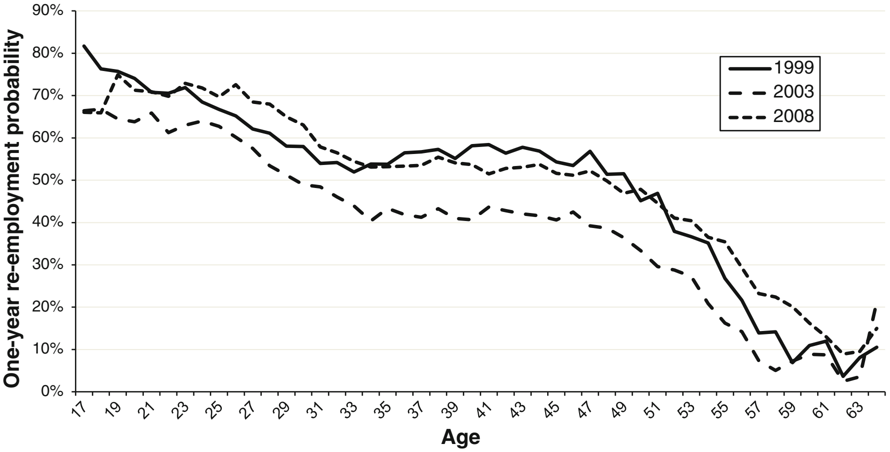

Fig. 3 One-year re-employment probability of women by age for 1999, 2003 and 2008

individuals are still active on the labor market, while others have already withdrawn from it.[2]

When comparing men and women across different years of UI inflow, we find that the one-year re-employment probability in 1999 was higher than in 2003 at every age. This is most likely the result of business cycle effects. In 2008, the one-year re-employment probability of unemployed men and women had risen again compared to 2003, but now especially among older workers. In total, the difference in re-employment probabilities of male workers of 45–55 and 55–65 years of age has decreased by 15 % points between 1999 and 2008. One explanation may be that the policy changes that were implemented have been targeted exclusively at older workers. But as we argued earlier, another explanation is the existence of birth cohort effects, i.e. the future generation of older workers may have been better able to find re-employment, just because their characteristics—such as their educational attainment or their health levels—make them more attractive to the labor market.

As a first investigation on the influence of birth cohort on the one-year re-employment probability, we divide our sample in eleven five-year birth cohorts. Table 2 gives an overview of the resulting five-year birth cohorts, the observed age interval, and the number of men and women that belong to each birth cohort. The table shows that the oldest workers in the population were born before 1940, while the youngest workers were born after 1984. Most cohorts are well-represented in the population, though the size of the pre-1940 cohort is relatively small because of early retirement and—to a lesser extent—mortality. The cohorts whose age interval is not restricted because of age can be observed over a 15 year age interval, which enables

[2] Our selected sample of unemployment spells only counts 1,929 unemployment spells of individuals aged 64.

Table 2 Distribution of birth cohorts over the selected sample of individuals

Birth cohort	Observed age interval	Number of men	Number of women
<1940	59–64	3,329	1,027
1940–1944	54–64	26,868	9,026
1945–1949	49–63	61,026	29,316
1950–1954	44–58	70,071	46,295
1955–1959	39–53	84,519	60,393
1960–1964	34–48	105,912	74,931
1965–1969	29–43	122,156	89,342
1970–1974	24–38	125,690	100,274
1975–1979	19–33	105,109	90,361
1980–1984	17–28	71,183	65,424
>1984	17–23	17,333	16,940
	Total	793,196	583,329

us to compare the one-year re-employment probability of subsequent birth cohorts of men and women over a common age interval.[3]

Figures 4 and 5 present the re-employment probability age profiles of unemployed workers, stratified by birth cohorts and gender. Comparing the one-year re-employment probability of subsequent birth cohorts, a recurrent pattern emerges: the subsequent cohort performs better on the higher end of the common age interval, but poorer on the lower end. If we had seen that the plots of subsequent birth cohorts make discrete 'jumps', this would have confirmed our hypothesis that new generations of men and women have a larger one-year re-employment probability at every age. The figures however do not account for period effects that may explain these patterns.

3 Empirical analysis

3.1 Empirical strategy

The key motivation of this paper is to provide an integrative analysis of the importance of age effects, birth cohort effects, period effects and the effects of policy changes on the one-year re-employment rates of (older) workers. Given the large number of observations in our sample, a linear specification is preferred (Wooldridge 2002). We therefore adopt a Linear Probability Model (LPM) to explain re-employment probabilities of pooled data, using standard estimation techniques to obtain robust standard errors.[4] As a baseline specification, this yields the following Age Period Cohort (APC)

[3] The age interval of the cohorts <1940 and 1940–1944 is cut off at age 64 because of the regulatory retirement age, whereas the age interval of the cohorts 1980–1984 and >1984 is cut off at age 17 because school is compulsory before that age.

[4] In our data, some individuals are observed more than once during the period of observation. In principle, this would allow for estimating individual fixed effects instead of birth cohort effects. When using such a

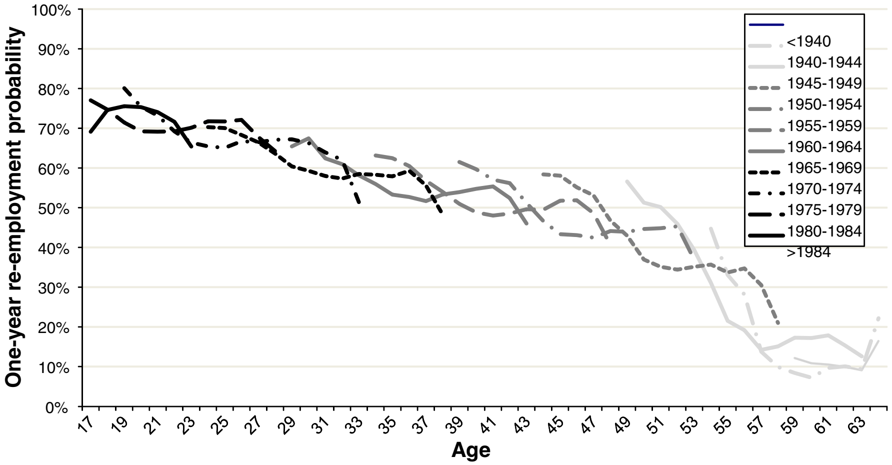

Fig. 4 One-year re-employment probability of men by age category and birth cohorts (1999–2008)

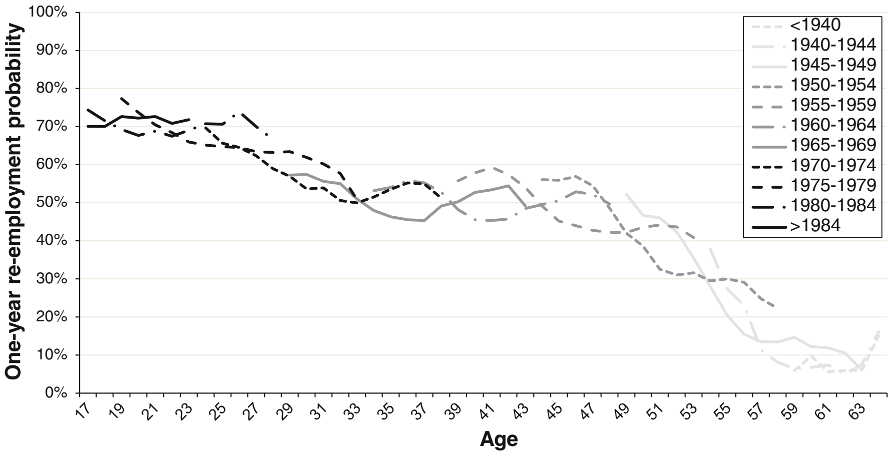

Fig. 5 One-year re-employment probabilities of women by age category and birth cohorts (1999–2008)

specification for the dummy Y, which equals 1 if workers are re-employed within one year after the start of their UI-spell, and zero otherwise:

$$Y_{it} = g_0 + g_c(c) + g_a(a_{it}) + g_t(t) + \beta X_{it} + \varepsilon_{it} \tag{1}$$

where i denotes the individual $(i = 1, \ldots I)$, t denotes the time period of the year of UI inflow $(t = 1, \ldots T)$, c denotes the birth cohort of individual i $(c = 1, \ldots C)$ and ε is an error term that is assumed to be independent and identically distributed. The

Footnote 4 continued
fixed effects specification, however, estimates would be based on an overly selective sample of individuals with multiple unemployment spells, having a weak labor market position that typically deteriorates over time. We therefore opt for the estimation of model forms with (pooled) birth cohort effects.

(transformation) function values g_c, g_a and g_t denote the effects of birth cohorts c, age a and time effects t, respectively. The functions $\boldsymbol{g}$ are all specified as piecewise constants, with intervals of five years for age and birth cohort and one year for calendar time.[5] Matrix $\boldsymbol{X}$ includes a dummy for search requirement exemption of older workers, the PBD of UI, various personal characteristics, and some characteristics of the last job prior to UI inflow. For a complete overview of all variables that are used as controls, we refer to Table 5 in the "Appendix" to this paper.

A well-known problem with APC models is that age, period and cohort effects are linearly dependent, and thus not identified without making further assumptions. Following Deaton and Paxson (1994), we therefore impose normalization restrictions on the period effects in our LPM specification. The intuition behind this approach is that period effects in our model are transitory. This means we impose that the period dummies sum to zero and are orthogonal to a linear time trend.[6] As a result, any trends in re-employment rates should be picked up by birth cohort effects or changes in the age composition of workers.

Like in most APC models, the second key assumption we make is that time period effects are similar for all age and birth cohort groups—the so called 'common trends assumption'. As we control for age, birth cohort, and time period effects, the identification of policy coefficient variables thus follows from a 'difference-in-difference design': the search exemption variable distinguishes between a treatment group (eligible workers aged 57.5 and older) and a control group (non-eligible workers aged 57.5 and older), with the job search exemption being removed in 2004. Likewise, the effect of the other major UI reform in our period of observation—the reduction of the PBD—can be inferred from the comparison of groups that were affected differently by the reform, before and after 2006. Next, for older workers the effect of this reduction can be calculated by taking the difference between the effect of a PBD of 60 months and the effect of a PBD of 38 months on the one-year re-employment probability. In order to extrapolate this difference, the PBD (and two higher order polynomials) is added as a continuous variable rather than a categorical variable.

3.2 Estimation results

Table 3 presents the parameter estimates of Eq. (1) for men and women, respectively. In the table, we distinguish between two model variants, depending on the use of control variables. First, we estimate the model without any individual characteristics, other than age. As such, the birth cohort estimates reflect the average impact of *all* observed and unobserved time-constant personal characteristics on this group. Next, we include the variables in matrix $\boldsymbol{X}$, causing birth cohort effect estimates to stem from (remaining) unobserved characteristics only.

5 We also have estimated model variants with spline functions of cohort and/or age effects. This yielded model outcomes that were virtually equivalent to step function specifications (the model outcomes are available on request). For expository arguments, we preferred to present the latter specification. As such, compare the size of effects is straightforward.

6 When defining the period dummies as $\boldsymbol{g}_t$, this implies that the following two conditions hold: $\sum_t g_t = 0$ and $\sum_t g_t\, t = 0$.

Table 3 LPM estimation results for one-year re-employment probability (1999–2008)

	Baseline specification men	Extended specification men	Baseline specification women	Extended specification women
Birth cohort				
<1940	Reference	Reference	Reference	Reference
1940–1944	0.02*** (3.8)	0.02** (2.4)	0.02** (2.2)	0.01 (1.2)
1945–1949	0.08*** (12.8)	0.04*** (4.8)	0.08*** (8.0)	0.03** (2.5)
1950–1954	0.11*** (15.4)	0.06*** (6.7)	0.10*** (10.0)	0.04*** (3.0)
1955–1959	0.10*** (13.4)	0.05*** (5.5)	0.11*** (9.8)	0.04*** (2.8)
1960–1964	0.10*** (12.8)	0.05*** (5.3)	0.09*** (8.1)	0.02 (1.2)
1965–1969	0.10*** (12.3)	0.05*** (4.8)	0.08*** (6.7)	−0.01 (3.80)
1970–1974	0.11*** (12.8)	0.05*** (5.2)	0.10*** (8.4)	0.00 (0.1)
1975–1979	0.14*** (16.4)	0.09*** (8.4)	0.15*** (12.7)	0.04** (2.4)
1980–1984	0.17*** (19.1)	0.12*** (10.7)	0.19*** (15.6)	0.06*** (3.6)
>1984	0.18*** (19.1)	0.13*** (10.9)	0.19*** (14.9)	0.04*** (2.6)
Age category				
15–24	Reference	Reference	Reference	Reference
25–29	−0.01*** (5.2)	−0.01*** (3.1)	−0.02*** (7.3)	0.01*** (2.7)
30–34	−0.06*** (17.7)	−0.05*** (13.6)	−0.09*** (24.1)	−0.04*** (9.8)
35–39	−0.10*** (16.7)	−0.09*** (20.5)	−0.10*** (24.7)	−0.05*** (9.6)
40–44	−0.14*** (31.6)	−0.12*** (23.2)	−0.11*** (22.4)	−0.05*** (8.9)
45–49	−0.19*** (37.3)	−0.16*** (26.6)	−0.16*** (27.3)	−0.09*** (13.4)
50–54	−0.26*** (45.9)	−0.23*** (32.4)	−0.26*** (39.1)	−0.18*** (22.1)
55–59	−0.42*** (67.5)	−0.37*** (47.4)	−0.41*** (55.5)	−0.32*** (33.9)
60–64	−0.47*** (66.4)	−0.43*** (49.5)	−0.47*** (54.3)	−0.41*** (37.6)
Year of UI inflow				
1999	0.06*** (41.6)	0.05*** (29.8)	0.04*** (20.4)	0.01*** (5.4)
2000	0.03*** (19.7)	0.01*** (7.7)	0.04*** (18.2)	0.02*** (7.2)
2001	0.00 (0.5)	0.01*** (7.3)	0.03*** (13.7)	0.05*** (19.9)
2002	−0.04*** (27.3)	−0.02*** (14.1)	−0.02*** (10.2)	0.01*** (2.8)
2003	−0.08*** (56.6)	−0.07*** (44.8)	−0.07*** (41.2)	−0.05*** (24.9)
2004	−0.05*** (36.3)	−0.06*** (39.0)	−0.07*** (42.1)	−0.08*** (44.5)
2005	−0.01*** (7.4)	−0.01*** (9.1)	−0.04*** (21.4)	−0.04*** (23.8)
2006	0.04*** (23.3)	0.04*** (25.7)	0.01*** (6.9)	0.01*** (8.4)
2007	0.06*** (38.9)	0.06*** (36.9)	0.06*** (34.7)	0.05*** (29.9)
2008	0.00 (1.3)	−0.01*** (4.4)	0.03*** (19.8)	−0.02*** (13.2)
Search exemption	–	−0.05*** (14.5)	–	−0.03*** (5.2)
Potential benefit duration	–	−0.0022*** (40.7)	–	−0.0038*** (18.7)
Idem, squared	–	1.45×10^{-6}***(19.8)	–	5.88×10^{-7}*** (20.9)

Table 3 continued

	Baseline specification men	Extended specification men	Baseline specification women	Extended specification women
Idem, third polynomial	–	-1.39×10^{-10}***(17.2)	–	-1.70×10^{-11}***(21.1)
Controls	–	✓	–	✓
Observations	793,196	793,196	583,329	583,329
R^2	0.10	0.13	0.08	0.11

Absolute t-values are in parentheses. *,**,*** indicate significance at the 10 %/5 %/1 % level. ‘Search exemption’ is a dummy variable that equals one for older workers who were not required to actively search for a job, i.e. individuals aged 57.5 or older who entered UI before January 1, 2004. PBD is measured in months. Remaining controls include personal characteristics (ethnicity, household position, urbanity of the living area), characteristics of the last job (duration of last job, type of employment, reason of dismissal and industry sector of last job) and UI size in hours

Starting with the baseline specification (1), the parameter estimates of birth cohorts show that the one-year re-employment probability of men significantly and substantially improves among subsequent birth cohorts, holding the effects of age and calendar time constant. The largest improvements in the one-year re-employment probability are observed among the oldest cohorts. Compared to the base group of men born before 1940, men born in 1945–1949 have an 8 % point larger probability of finding re-employment within one year after UI inflow. Since the majority of these men qualifies as older workers, it is safe to say that the one-year re-employment probability has increased for new generations of older men. For cohorts born later than 1950, the growth in the one-year re-employment probability flattens out, but increases again for men born after 1974. The effects of age are larger than those of birth cohorts. According to our parameter estimates, the one-year re-employment significantly decreases with age, most notably when men reach the age of 55. Compared to men aged 50–54, the one-year re-employment probability of men aged 55–59 is 16 % point lower.

The period dummies, which capture both changes in the business cycle and policy reforms (in the baseline specification), also have a substantial and significant impact on the one-year re-employment probability. The parameter estimates of the year dummies reveal that men who entered UI in 1999 and 2007 had the largest probability of finding re-employment within one year. Men who entered UI in 2003 were less successful: their one-year re-employment probability was 14 % point lower than in 1999 and 2007.

As the second column of Table 3 shows, the extended baseline specification improves the explanatory power of our model by adding personal characteristics, characteristics of the last job and UI entitlement characteristics as additional controls. Controlling for these variables causes the birth cohort effect to reduce in size by up to 5 % point at maximum. In addition, the one-year re-employment probability of the oldest birth cohorts changes favorably compared to subsequent birth cohorts. The age effects only get somewhat smaller, suggesting that differences in personal characteristics, characteristics of the last job, UI entitlement characteristics and UI eligibility criteria are confined to birth cohorts and not age groups.

The parameter estimates of the extended baseline specification that are of special interest to us are those of ‘search exemption’ and PBD since they measure the effect of recent UI reforms on the re-employment rate of older workers. The reported coefficient of 'search exemption' shows that eligible men (i.e. men aged 57.5 years or older who entered UI before January 1, 2004) have a 5 % point lower one-year re-employment probability than non-eligible men of equal age. The estimated effect is similar to the one found by Bloemen et al. (2013). Table 3 also displays the three polynomial coefficient values for the PBD length. Given these coefficient estimates, the implied parameter effect of the reduction in the PBD from 60 to 38 months is a 5 % point higher one-year re-employment probability. This UI reform, which came into effect as of 2006, predominantly affects the job finding rate of older men since a PBD of 60 months requires an extensive labor market history. Given the average tenure of workers who are older than 55, the overall impact of the PBD reduction amounts to 3 % point.

The last two columns of Table 3 present the parameter estimates of the (extended) baseline specification for women. According to our estimates, the effect of birth cohort on the one-year re-employment probability of women is comparable to those for men (ceteris paribus). Moreover, instead of showing continuous growth, the birth cohort effects are virtually stable for women born between 1950 and 1975. A possible explanation is that, in this period of evolving social norms toward paid female employment, women with lower productivity rates (than earlier cohorts) entered the labor market (Euwals et al. 2011). This may have compensated the (positive) effects of increasing education levels of women.

In the extended baseline specification, the female birth cohort effects decrease substantially, suggesting that the controls explain a large part of the variation in the one-year re-employment probability among subsequent cohorts of women (more so than for men). As in the case of men, the one-year re-employment probability of the oldest birth cohorts now starts to look more favorable. There is also a substantial reduction in the size of the age effects. One explanation may be that women with children are more picky in accepting job offers (e.g. because they want to work part-time or have lower desired traveling distances). So once we account for children in the household, the one-year re-employment probability of these mothers becomes larger.

Finally, we find that the search requirement exemption for women aged 57.5 and older before 2004 reduced their one-year re-employment probability with 3 % point, which is roughly half the effect on the one-year re-employment probability of eligible men. Possibly, we are dealing with a more selective group of motivated women whose search behavior is not affected much by the introduction of search requirements. The effect of the other UI reform on the one-year re-employment probability, however, is somewhat larger: a 22-month reduction in PBD results in an 8 % point larger one-year re-employment probability. At the same time, the overall impact of the reduction is limited, as the labor market histories of women are generally shorter than for men.

3.3 Birth cohort and policy effects

As we have argued earlier in Section 2, the re-employment probabilities of older workers have increased at a faster pace than those of younger workers in the period under investigation. This is also reflected in our model outcomes: both birth cohort effects and policy effects have lessened the gap between older and younger workers. To assess the relative size of these two explanations, we use our model outcomes to decompose the growth in the one-year re-employment probability of older men and women (aged 55–65) between 1999 and 2008. In this time period, the increase in the one-year re-employment probability of men amounted to 5 % point. Acknowledging that the year 2008 had a negative impact on the one-year re-employment probability of 6 % point compared to 1999, these men actually realized an increase in the re-employment probability of 11 % point. Of this increase, approximately 4 % point is due to birth cohort effects.[7] Policy effects are responsible for the remaining increase: 5 % point is due to the search obligation, and 3 % point was due to the reduction of the PBD.[8] Hence, policy effects were the main drivers of the growth in the one-year re-employment probability of older men in the period 1999–2008.

For women, the 6 % point increase that was realized between 1999 and 2008 becomes an increase of 9 % point if one takes account of the (negative) period effects. Here, the birth cohort effect, the introduction of search requirements and the PBD reduction were each responsible for a 3 % point increase in the one-year re-employment probability.[9] So, similar to men, the increase in the one-year re-employment probability of older women was predominantly driven by policy changes.

To conclude, we find that birth cohort effects have stimulated job finding rates among both men and women in our period of observation, also at old age. Although our extended model has shown that personal characteristics, characteristics of the last job and policy reforms explain part of the birth cohort effects, our conclusion remains the same. Birth cohort effects explain part of this growth, but policy changes seem to have been of greater importance.

3.4 Sensitivity analyses

So far, our analysis assumes period effects to be common among age and cohort groups. Accordingly, changes in re-employment rates of older worker are essentially identified from two sources: changes in policy variables that are targeted at older workers and (changes in) cohort effects. It may however be that unobserved time trends—like other targeted policies or compositional changes that are unobserved in our data—are relevant too. Particularly in light of the rapid increase in employment rates of older workers, changes in the composition of older workers that enter into UI may have been important.

To test for the robustness of our outcomes to the common trends assumption, we re-estimate our model in a more flexible way, namely by including linear time trends for the oldest age categories in our sample. Table 4 displays the outcomes of a model

[7] This percentage is calculated by taking the difference between the average effect of birth cohort on the one-year re-employment probability of the cohorts that represented older workers in 1999 and the cohorts that represented older workers in 2008.

[8] For the effect of the PBD reduction, we take into account that only half of all older men were entitled to more than 38 months of PBD in the period under investigation.

[9] Only 40 % of all older women were entitled to more than 38 months of PBD, resulting in an effect of 3 % point of the PBD reduction on the increase of the one-year re-employment probability.

Table 4 LPM sensitivity analyses for re-employment probability for men (1999–2008): Flexible time trend for older workers and sustainable re-employment as outcome measure

	Baseline specification	Baseline specification; flexible trend older unemployed	Extended specification; flexible trend older unemployed	Baseline specification; sustainable one-year re-emp. prob.
Birth cohort				
<1940	Reference	Reference	Reference	Reference
1940–1944	0.02*** (3.8)	0.01 (1.6)	0.01* (1.8)	0.01* (1.8)
1945–1949	0.08*** (12.8)	0.06*** (6.7)	0.03*** (3.4)	0.02** (2.2)
1950–1954	0.11*** (15.4)	0.08*** (8.2)	0.05*** (5.0)	−0.01 (0.9)
1955–1959	0.10*** (13.4)	0.07*** (7.2)	0.04*** (4.2)	−0.03*** (3.4)
1960–1964	0.10*** (12.8)	0.07*** (7.0)	0.04*** (4.1)	−0.05*** (4.7)
1965–1969	0.10*** (12.3)	0.07*** (6.8)	0.04*** (3.8)	−0.05*** (4.8)
1970–1974	0.11*** (12.8)	0.08*** (7.3)	0.05*** (4.2)	−0.04*** (3.6)
1975–1979	0.14*** (16.4)	0.11*** (10.4)	0.08*** (7.3)	0.00 (0.1)
1980–1984	0.17*** (19.1)	0.14*** (12.8)	0.12*** (9.5)	0.04*** (3.0)
>1984	0.18*** (19.1)	0.15*** (13.3)	0.13*** (9.9)	0.09*** (6.2)
Age category				
15–24	Reference	Reference	Reference	Reference
25–29	−0.01*** (5.2)	−0.01** (5.2)	−0.01*** (3.0)	0.00 (0.2)
30–34	−0.06*** (17.7)	−0.06*** (17.7)	−0.05*** (13.5)	−0.03 (7.0)
35–39	−0.10*** (16.7)	−0.10*** (26.8)	−0.09*** (20.3)	−0.08** (12.8)
40–44	−0.14*** (31.6)	−0.14*** (31.6)	−0.12*** (23.0)	−0.12*** (16.8)
45–49	−0.19*** (37.3)	−0.19*** (37.3)	−0.16*** (26.4)	−0.18*** (21.6)
50–54	−0.26*** (45.9)	−0.26*** (45.9)	−0.22*** (32.2)	−0.27*** (27.6)
55–59	−0.42*** (67.5)	−0.45*** (52.2)	−0.37*** (40.1)	−0.41*** (38.7)
60–64	−0.47*** (66.4)	−0.51*** (47.3)	−0.44*** (39.2)	−0.45*** (39.2)
Year of UI inflow				
1999	0.06*** (41.6)	0.06*** (41.4)	0.05*** (29.7)	0.02*** (13.7)
2000	0.03*** (19.7)	0.03*** (19.6)	0.01*** (7.6)	−0.01*** (7.5)
2001	0.00 (0.5)	0.00 (0.5)	0.01*** (7.3)	0.02*** (10.7)
2002	−0.04*** (27.3)	−0.04*** (27.2)	−0.02*** (14.0)	−0.01*** (6.5)
2003	−0.08*** (56.6)	−0.08*** (56.3)	−0.07*** (44.6)	−0.04*** (29.0)
2004	−0.05*** (36.3)	−0.05*** (36.1)	−0.06*** (38.8)	−0.01*** (9.9)
2005	−0.01*** (7.4)	−0.01*** (7.3)	−0.01*** (9.1)	0.04*** (31.4)
2006	0.04*** (23.3)	0.04*** (23.3)	0.04*** (25.7)	–
2007	0.06*** (38.9)	0.06*** (38.8)	0.06*** (36.9)	–
2008	0.00 (1.3)	0.00 (1.1)	−0.01*** (4.5)	–
Age specific time trend older workers				
55–64 Year-1998	–	0.0039***(4.8)	0.0009 (1.15)	–
Search exemption	–	–	−0.05*** (14.2)	−0.03*** (9.7)

Table 4 continued

	Baseline specification	Baseline specification; flexible trend older unemployed	Extended specification; flexible trend older unemployed	Baseline specification; sustainable one-year re-emp. prob.
Potential benefit duration	–	–	−0.0022*** (40.3)	−0.0021*** (37.2)
Idem, squared	–	–	1.45×10^{-6}***(19.7)	1.36×10^{-6}***(17.4)
Idem, third polynomial	–	–	-1.38×10^{-10}***(17.1)	-1.29×10^{-10}***(15.0)
Controls	–	–	✓	✓
Observations	793,196	793,196	793,196	576,286
R^2	0.10	0.10	0.13	0.12

Absolute t-values are in parentheses. *,**,*** indicate significance at the 10 %/5 %/1 % level

specification that adds a specific time trend for workers of 55–65 years of age. We also include the outcomes of the baseline model—without age specific time trends and with controls—as a reference point. We focus on providing sensitivity analyses for men, arguing that those obtained for women were alike.

The second column of Table 4 makes apparent that age specific time trends for the older workers are substantial and significant. Without controlling for individual and policy effects, we observe a positive time trend for workers who are older than 55 years of age, which accumulates to about 4 % point in the time period under investigation. However, the trend coefficient becomes insignificant if we include individual controls and the PBD and the search exemption as explanatories (see the third column of Table 4). This suggests that the differences in time trends can largely be explained by the observed policy changes. Moreover, the coefficient estimates of the policy variables are robust with respect to inclusion of time trends. Finally, note that the coefficient values of the cohort and age effects are close to those for the baseline model with all controls and policy variables as explanatories.

Next to the use of flexible time trends, we also check for the robustness of our results by using an alternative re-employment measures (Table 4; last column).[10] In particular, we use the *sustainable* one-year re-employment probability as a dependent variable. This variable is defined as the probability that an unemployed individual finds re-employment within one year after UI inflow *and* does not claim UI benefits in the two years thereafter. The parameter estimates show that the difference in the one-year re-employment probability between the oldest and youngest birth cohorts of men becomes smaller, or even reverses for some cohorts. This suggests that younger worker cohorts find jobs more often than older workers, but are also more likely to lose these new jobs quickly. It is also noteworthy that this difference has become less marked relative to men born between 1945 and 1964, while these middle aged men are not known for being employed in flexible jobs. A likely explanation is that men

10 We also tested for the robustness of our estimation results by adopting a Probit specification as a functional form. This yielded results that were very similar to the LPM specification.

from the oldest birth cohorts had access to other social insurance schemes (e.g. early retirement, DI and sickness benefits), which kept them out of UI, even after a layoff. In the time period under investigation, the pathways into early retirement and DI were tightened for the oldest workers.

4 Conclusions

Our analysis has shown that policy changes for older unemployed workers have been the dominant drivers of the increased re-employment probability of older workers in the Netherlands that is observed from 1999 to 2008. Specifically, the introduction of search requirements for unemployed workers aged 57.5 and older in 2004 has increased the one-year re-employment probability of men with 5 % point, while the reduction in the PBD has caused the one-year re-employment probability of eligible men to increase with 3 % point (on average for this group). These two effects largely explain the increased re-employment probability of older unemployed workers. Birth cohort effects were also important, though to a lesser extent, amounting to a 3–4 % point increase in the re-employment probability of older workers.

When taking a future perspective, it should be noted that the room for further specific activation policies for older workers seems limited. In particular, the two policy measures that were studied—i.e. the search obligations and the reduction of PBD—probably were the most effective measures that could have been taken. Now that benefit entitlements and search obligations of older and younger workers have become more alike, there is a stronger need for generic policies to increase re-employment rates. This also means that policies should be focused on e.g. improving the education level of workers at younger ages, which in turn contributes to the birth cohort effects. For an optimal policy setting, this means more insight is needed on the importance of various determinants of birth cohort effects in explaining the job prospects of workers, with cultural settings, educational attainment and health conditions as likely candidates. This provides an interesting avenue for future research.

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5 Appendix

See Appendix (Table 5).

Table 5 Overview of all variables

Variable	Type	Explanation	Reference category
Birth cohort	Categorical	Ten dummy variables for five-year birth cohorts	<1940
Age	Categorical	Eight dummy variables for five-year age category	15–24
Year	Categorical	Nine dummy variables for the year of UI inflow	1999
Search exemption	Categorical	Dummy variable for search requirement exemption	No exemption
Ethnicity	Categorical	Three dummy variables for immigrants, individuals born in the Netherlands with one foreign parent and individuals born in the Netherlands with two foreign parents	Native
Household position	Categorical	Three dummy variables for individuals in a partnership without kids, individuals in a partnership with kids and single parents	Single
Urbanity of the living area	Categorical	Four dummy variables for the degree of urbanization of the living area	Very strong urbanization
UI size per week	Categorical	Five dummy variables for the category of UI size hours	1–12 h
PBD	Continuous	PBD in months	N/A
PBD^2	Continuous	PBD in months squared	N/A
PBD^3	Continuous	PBD in months to the third power	N/A
Duration of last job	Categorical	Nine dummy variables for the category of duration of last job in months	1–3 months
Type of employment	Categorical	Two dummy variables for part-time employment or call worker	Full-time employment
Reason of dismissal	Categorical	Five dummy variables for disturbed employer-employee relationship, long-lasting or frequent absenteeism due to illness, bankruptcy of employer, large scale layoffs and no dismissal	Dismissal for economic reasons
Industry sector of last job	Categorical	Seventeen dummy variables for the industry sector of the last job	Agriculture, hunting and forestry
Gross personal yearly income	Continuous	Log of the gross personal yearly income in the year of UI inflow	N/A

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