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The effect of unemployment duration on reservation wages: Evidence from Belgium

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Abstract

This paper studies the effect of unemployment duration on asking (reservation) wages. We provide new evidence based on unique longitudinal data on unemployed workers in Belgium, who were surveyed about self-reported reservation wages at the start of the unemployment spell, and after 3 and 6 months of unemployment duration. Our estimates suggest that reservation wages decline with unemployment duration by about 0.4 percent per month, or 5 percent per year, and that cross-sectional estimates are biased upward. We find stronger declines among men and among workers who earned high wages in their previous jobs. The paper discusses these findings in the light of learning models and discusses the implications of falling reservation wages for the debate on the effect of unemployment on wages.

1) Introduction

There is a long-standing debate on how unemployment affects the wage demands of unemployed job seekers (Mortensen, 1970; Feldstein & Poterba, 1984; Addison et al., 2013; Krueger & Mueller, 2016). The reservation wage – also known as the acceptance wage or the minimum asking wage – is the minimal wage level at which a job seeker would accept a job offer. Economic intuition about market clearing and supply and demand pressures predicts that asking prices decline after a period of time if a product fails to sell – both in the labor market and in any other market such as the housing market. However, the phenomenon is perhaps more relevant in the labor market because it is likely to affect unemployment levels and starting wages across the business cycle.

Various job search theories explain why the optimal reservation wage might decline with unemployment duration, either because the value of leisure time and unemployment benefits decline over time (Kasper, 1967; Mortensen, 1986), because job seekers face liquidity constraints (Crosslin & Stevens, 1977), because of human capital depreciation or because of imperfect information and learning about the expected benefits of continued job search (McCall, 1970; Burdett & Vishwanath, 1988). However, little is known about the empirical relation between unemployment duration and reservation wages. Most existing studies rely on cross-sectional data, but identification is not straightforward with such data because OLS estimates are likely to be biased because of the reverse causal pathway from reservation wages to unemployment duration, and because valid instrumental variables are hard to find in this setting. There are some longitudinal studies but almost all these studies rely on yearly data from panel surveys (Addison et al., 2009; Pannenberg, 2010; Addison et al., 2013) in which identification is complicated by the fact that a full year is a long time step in unemployment spells and that it is often difficult to account for transitions between interviews. To the best of our knowledge, there is only one study that relies on high-frequency longitudinal data. This seminal study by Krueger & Mueller (2016) contains the best available evidence – perhaps the only credible available evidence – on the effect of unemployment duration on reservation wages. Krueger and Mueller conclude that “reservation wages decline at a modest rate” and the fact that in their full sample the estimated duration effect was not statistically significant led others to interpret the Krueger and Mueller study as showing that reservation wages are not affected by unemployment duration (Schmieder et al., 2016).

In this study we present original longitudinal data about the asking wages of job seekers in Belgium to identify the effect of unemployment duration on reservation wages and to investigate how this effect varies with job seekers’ characteristics. The job seekers in our data were interviewed three times with three-month intervals. The main findings are that reservation wages decline with unemployment duration by about 0.4 percent per month and we find stronger declines among men and among workers who earned high wages in their previous jobs.

This study aims to contribute to various literatures. First, this study provides evidence on the decline of reservation wages with unemployment duration. The finding itself, which is essentially that asking prices are lowered over time as a product fails to sell, should be less surprising than the fact that this is the first study to provide this evidence. As we argued above, this is largely a result of challenges relating to data and methods. And although our results may appear to contradict the findings of Krueger and Mueller (2016), we argue that they do not: their point estimates are close to ours and we will argue that a decline of the asking wage by 5 percent per year is substantial and not modest. We also contribute to the empirical literature by showing that cross-sectional estimates of the effect of unemployment duration on reservation wages are subject to substantial upward bias because exits from the unemployed population are non-random, and that adding the usual controls only partly eliminates this bias.

Second, understanding the effect of unemployment duration on reservation wages is important because it provides one channel for the effect of the unemployment rate on wages across the business cycle. The findings of this study are in line with the recent literature that shows that the wages of workers starting in new jobs fall with unemployment duration (Schmieder et al., 2016) and that starting wages are much more cyclical than wages in the overall labor market that are usually characterized by wage stickiness (Pissarides, 2009; Haefke et al., 2013). Even in the presence of downward wage rigidity within jobs and within employment spells, the reservation wage channel opens up a causal pathway from unemployment to the wage bill because the mean unemployment duration is countercyclical. Our results also imply a reinterpretation of earlier findings reported in the literature on starting wages. Schmieder et al. (2016) assume that asking wages are constant over the unemployment spell (they rely on Krueger and Mueller (2016)) and therefor interpret the 0.8 percent per unemployment month fall in starting wages as a duration effect resulting from stigma, human capital depreciation, etcetera. Our estimates suggest that the duration effect on starting wages is smaller than what these authors report.

Third, we report empirical evidence on the heterogeneity of the duration effect on reservation wages with respect to gender and previous wage levels. Inspired by the learning model of Burdett and Vishwanath (1988) and assuming wage dispersion and job seekers who do not know the distribution of their wage offer function, we argue that job seekers infer the wage offer distribution using information about their previous wages (anchoring or reference-dependence) so that workers with high previous wages experience stronger declines in reservation wages as they receive signals about their wage offers during search. We speculate that heterogeneity of the duration effect by gender could result from either gender differences in risk-aversion and self-confidence or gender differences in the relation between unemployment duration and the disutility of being unemployed. We will show that the empirical evidence is more in line with the latter than the former explanation.

The remainder of the article is structured as follows. Section 2 presents theories about the effect of unemployment duration on reservation wages in a framework of nonstationary job search models. Section 3 discusses the problem of identification of duration effects and reviews the empirical literature. Section 4 introduces the data and design of our study. Section 5 presents the results and section 6 concludes. More details about the data and additional results are included in various online appendices.

2) Theory

The various theories about the effect of unemployment duration on reservation wages can best be understood within the framework of job search models. Search models assume that job seekers have imperfect information in the sense that job offers arrive sequentially and that job seekers do not know which vacancies and which wages will be offered in the near future. The job seeker must decide to either accept or decline a job offer in order to maximize her discounted lifetime income. If she decides not to accept a job offer now and to continue searching, then there are benefits in terms of a higher expected wage offer in the future but also search costs in terms of acquiring information about additional vacancies and the opportunity cost of the time taken up during search, which is a function of foregone earnings and unemployment benefits. A rational job seeker will accept an offer if its associated future income stream is at least as big as the expected income stream would be in the case of continued search.

Assume that job offers arrive randomly following a Poisson process characterized by an arrival rate λ. Job offers are random draws from a wage offer distribution $H(w)$. The time that a job seeker is searching instead of working has a value b, which is a function of the value of leisure time, the disutility of being unemployed, unemployment benefits and direct costs related to job search. Job seekers know b, λ and $H(w)$ but they do not know when job offers will arrive and what will be the offered wage, and there is no recall of previously rejected job offers. It is assumed that job seekers are risk-neutral and maximize the expected present value of lifetime income over an infinite time horizon using a constant discount rate r.

The job search model is called stationary if λ, $H(w)$ and b are constant over time and do not depend on the realizations of the offer times and the wage offers (van den Berg, 1990). In a stationary model the job seeker's expected lifetime income is independent of time or unemployment duration and the optimal strategy for the job seeker is to choose a reservation wage RW that equates the marginal cost to the marginal benefits of continued job search. In continuous time, the optimal strategy is to adopt a decision rule in which any wage offer above this reservation wage is accepted (Mortensen, 1986):

$$RW = b + \frac{\lambda}{r}\int_{RW}^{\infty}(w - RW).H(w).dw \qquad (1)$$

If $\lambda(t)$, $H(w,t)$ or $b(t)$ are functions of time (elapsed unemployment duration), then the search model is non-stationary and the optimal reservation wage $RW(t)$ becomes a function of time. Van den Berg (1990) showed that the optimal reservation wage is a function that satisfies the differential equation

$$\frac{dRW}{dt} = r.RW(t) - r.b(t) - \lambda(t)\int_{RW(t)}^{\infty}\big(w - RW(t)\big).dH(w,t) \qquad (2)$$

and that $RW(t)$ decreases over time if either $b(t)$ is decreasing, if $\lambda(t)$ is decreasing or if the mean of the wage offer function is decreasing.

Theories about the effect of unemployment duration on reservation wages all require some explanation about how unemployment duration affects $b(t)$, $\lambda(t)$ or $H(w,t)$. A first set of theories hypothesizes channels relating to changes in the value $b(t)$ of time spent on job search. Kasper (1967) suggested that reservation wages decline because the marginal utility of leisure decreases with unemployment duration. Likewise, if the disutility of being unemployed increases with unemployment duration, then the reservation wage should fall. Also, if unemployment benefit payments are reduced as a sanction for insufficient job search intensity (Abbring et al, 2005), or if benefit levels decrease over time, then reservation wages can be expected to decline because unemployment benefits lower the opportunity cost of search. If unemployment benefits are a discontinuous function of time, for example when unemployment benefits are limited in duration, then the reservation wage should decrease in a gradual way because the optimal strategy for job seekers is to reduce their reservation wage as the time limit approaches. A related theory assumes that unemployed workers face liquidity constraints (Mortensen, 1986; Crosslin & Stevens, 1977). If it is assumed that job seekers must self-finance the costs of search using their previous savings, then the period open for job search is limited to the moment when these savings are depleted and the optimal reservation wage will decline with search tenure as this time limit approaches. Such liquidity constraints would not appear if credit markets would function properly, but Mortensen (1986) argues that unemployed workers face difficulties in borrowing in official credit markets. Crosslin and Stevens (1977, p. 1298) extend this theory of liquidity constraints by considering both the amount of "financial and *psychic* resources available to the individual jobseeker". If psychic resources are used in the search process and if these limit the period open for search, then the optimal reservation wage will again decrease with search tenure.

Another set of theories hypothesizes channels related to changes in the expected income benefits of job search. One reason is the depreciation of human capital. If general human capital – the type that is transferrable across firms – depreciates over the unemployment spell, then productivity and reemployment wages decline with unemployment duration (Ljungqvist & Sargent, 1998). If job seekers are aware of the implied downward shift of the mean of the wage offer distribution then the optimal asking wage falls. A related theory assumes asymmetric information and regards unemployment duration as a signal of low productivity. Vishwanath (1989) proposes a search model with a wage offer function that depends on the information the firm has about the worker: if it is assumed that job seekers are aware of such stigma effects associated with unemployment duration, then the reservation wage falls in this model. Another theory is based on learning by job seekers about the expected benefits of search (McCall, 1970; Burdett & Vishwanath, 1988). Whereas standard search models assume that the distribution of the wage offer function is known to job seekers, Burdett and Vishwanath (1988) assume that job seekers do not know the wage offer distribution corresponding to their skills and that job seekers infer the distribution using the signaling information in the received job offers by adjusting their reservation wage upward if they receive high wage offers and downward if they receive low wage offers. Since – by assumption – job seekers who receive a higher wage offer than their reservation wage, accept the offer and are selected out of the pool of unemployed, the population of remaining job seekers is composed of job seekers who received low wage offers relative to their reservation wages and who reduce their reservation wages, so that individual reservation wages decline with unemployment duration in the population of job seekers.

Assume that at each time t when an offer is received, the worker guesses the mean of the wage offer function μ_t as a weighted average of the previous guess $\mu_{t'}$ and the new wage offer w_t:

$$\mu_t = (1 - \gamma_n).\mu_{t'} + \gamma_n.w_t \quad (3)$$

where the optimal weight $0 \leq \gamma_n \leq 1$ is a decreasing function of n because $\mu_{t'}$ becomes more reliable as n increases (see Burdett and Vishwanath, 1988). This learning model can be extended to inferences about the job offer arrival rate λ. Assume that workers do not know the arrival rate and that they use information about the interarrival times between wage offers to infer the arrival rate λ_t at each time t when a new offer is received. Since $1/\lambda$ is the expected waiting time until the next wage offer, learning about the arrival rate can be thought of as an adjustment process $1/\lambda_t = (1 - \gamma_n).1/\lambda_{t'} + \gamma_n.\Delta t$, where $\lambda_{t'}$ is the previous guess of the arrival rate and $\Delta t = t - t'$ is the interarrival time between the two last received wage offers. Job seekers who initially set the arrival rate too low, are more likely to exit from unemployment so that the reservation wage can be expected to decline in the population of remaining job seekers. The arrival rate also depends on search intensity, which could change with unemployment duration if, for example, job seekers become more discouraged over time.

These theories relate to the causal pathway from unemployment duration in the direction of reservation wages. However, search theory also proposes a basic channel for reverse causality. A higher reservation wage increases the expected duration of unemployment because it reduces the probability that the wage attached to a randomly drawn job offer exceeds the reservation wage during the search process (Mortensen, 1970). This issue of reverse causality has been the main obstacle to identify these effects in empirical work.

3) Identification and the empirical literature

There is a large empirical literature on the determinants of reservation wages. Various studies show that women set lower reservation wages than men at the start of an unemployment spell, even after controlling for their previous job (Caliendo et al., 2017; Le Barbanchon et al., 2021). Constant et al. (2017) find that second-generation migrants set higher reservation wages than first-generation migrants, who they argue still have their home country's wages in mind when forming beliefs about the wage offer distribution. Koenig et al. (2016) study how reservation wages change over the business cycle and find that they are less cyclical than expected, which they argue can be explained by reference-dependence on previous wages. DellaVigna and Paserman (2005) find that the effects of impatience and time preferences on reservation wages are essentially zero, and Le Barbanchon et al. (2019) find an almost zero effect of unemployment benefit duration on the level of the reservation wage at the start of the unemployment spell.

Identifying the effect of unemployment duration on reservation wages is not straightforward. Cross-sectional data on reservation wages and unemployment durations are likely to lead to biased estimates of the duration effect. One source of bias is the non-random selection into work, discontinued search and survey non-response. For example, if more motivated job seekers set lower reservation wages and are more likely to find a job, then the estimated duration effect would be biased upward. Reverse causality also produces bias. Job seekers who choose lower reservation wages experience shorter unemployment spells because a wage offer is more likely to exceed their reservation wage, which leads to upward bias in estimates of the duration effect that rely on cross-sectional data.

Better estimates of the causal impact of unemployment duration on reservation wages can be obtained using panel data in which both unemployment duration and reservation wages are surveyed at different moments in time. A fixed effects estimate only exploits information about changes in reservation wages within individuals who remain unemployed, so it is not susceptible to the potential selection bias resulting from reverse causality. Moreover, this effectively controls for all individual fixed effects so that the potentially confounding role of unobservables is limited. The fixed effects estimator is consistent if the unobservable characteristics of job seekers are constant over the unemployment duration period under study. This is the identifying assumption of the estimates presented in this study. Although fixed effects estimates eliminate most sources of bias, within-individual changes in characteristics over the duration of unemployment may still confound the relation. For example, a decline of the motivation of job seekers with unemployment duration would produce upward bias if the more motivated set lower reservation wages and if motivation is unobserved (although such a channel via motivation could be considered to be part of the duration effect of interest).

Given the lack of good quality longitudinal data, the empirical literature has struggled to identify the effect of unemployment duration on reservation wages. Most studies use cross-sectional data and the earlier ones rely on OLS regressions (Kasper, 1967; Barnes 1975; Lynch 1983). Crosslin and Stevens (1977) adopt an instrumental variables approach, but their instruments for unemployment duration (gender, race and unemployment benefits) are likely to be directly related to reservation wages (Jones, 1988). Almost all longitudinal studies rely on yearly data from panel surveys (Addison et al., 2009; Pannenberg, 2010; Addison et al., 2013) in which identification is hard because respondents in those samples are not all recent job seekers at the time of the survey. Moreover, a full year is a long time step in unemployment spells and it is usually hard in these data to account for respondents who re-enter into unemployment. For example, using data from the European Community Household Panel, Addison et al. (2009) find a significant effect of duration on reservation wages in only out of the thirteen countries in the study – and this effect is positive, which is hard to explain.

The best available evidence is in Krueger and Mueller (2016), who use high-frequency longitudinal data (weekly data for a period of 24 consecutive weeks) on unemployed workers in New Jersey. Their point estimates of the effect of unemployment duration range from 0.05 to 0.14 percent per week of unemployment, or 2.4 percent to 7.3 percent per year of unemployment, and they conclude that "reservation wages decline at a modest rate over the spell of unemployment". In the full sample the estimated effect is statistically insignificant, but Krueger and Mueller report significant effects in subsamples of older workers, workers with a lot of savings and workers looking for a full-time job. A possible drawback of their high-frequency data is that asking job seekers about their reservation wages every week could lead to measurement error or consistency bias if respondents try to be self-consistent. It is also worth noting that the response rates in the study were low, although the authors attempt to adjust for this problem.

4) Data and design

We collected survey data among short-term unemployed job seekers in Flanders, the Dutch speaking region of Belgium. The data are unique in the sense that the same respondents were questioned about their reservation wage up to three times at various moments in time. Reservation wages were measured by asking respondents how high the net monthly wage needs to be so that they would accept an offer for the kind of job they are looking for. The reservation wages of respondents who indicated that they were looking for a part-time job, were converted to fulltime equivalent wages (see online appendix A for more details on this conversion). Many of the variables used in this study are based on self-reported information provided by the job seekers in the surveys, such as the reservation wage, the wage in their previous job and the number of application letters they recently sent out. The other variables, including demographics and information on participation in counseling and coaching by the Flemish Public Employment Services agency (PES), were added from administrative sources.

We started with a random sample of 6,000 individuals who registered as unemployed in October 2011 with the PES (this is a requirement to receive unemployment benefits) and who had had a paid job before they became unemployed. These participants were contacted and asked to answer an online or paper questionnaire (wave 1). The respondents were contacted again for a follow-up questionnaire after three months (wave 2) and after six months (wave 3). In the empirical analyses we measure unemployment duration as the time elapsed since the time of the survey in wave 1. This measure underestimates the time since registration as unemployed with the PES because participants filled in our survey a few weeks after registration and the measure contains some measurement error because not all job seekers registered with the PES and responded to the survey at exactly the same moment, but this error is small compared to the time horizon of our study. Measuring the start of the unemployment spell accurately is also difficult because the moment when job seekers register as unemployed with the PES differs somewhat from the moment they reportedly started searching for a job when job seekers are asked in a survey. This self-reported search duration is not necessarily a better measure of unemployment duration because job seekers could have been searching for a job while they were still in their previous job. Our survey asked job seekers if they are currently actively searching for a job and how many weeks they have done so. Participants who say they are not actively searching are excluded from the analysis and we also exclude job seekers who state that they had been actively searching for more than 20 weeks. This issue about the timing of the start of the unemployment spell should not be a source of bias because our main results rely on first differences and the time span between waves is the same for every respondent.

Table 1 presents an overview of the number of observations in our analyses across the three waves and the various selection criteria we imposed on the sample. The total number of respondents to the surveys were 1,744 in wave 1 (implying a response rate of 29%), 1,138 in wave 2 (65% of the participants in wave 1) and 963 in wave 3 (85% of the participants in wave 2). These response rates are substantially higher than in the survey used in the study of Krueger and Mueller (2016) that was discussed earlier. A closer inspection of our data indicates that the attrition across the waves is not random in the sense that younger, low-skilled and immigrant job seekers are more likely to drop out of the sample, but – as explained earlier – this should not be a source of bias in fixed effects estimates. More important is the drop-out resulting from exits to work and discouragement. 30 percent of the job seekers had found a job by wave 2 and 40 percent by wave 3. We excluded job seekers who state in our survey that they are currently not actively searching for a job. Most of these respondents indicate that they are either sick, pregnant or in a training program. We further restrict the sample by only including job seekers who state in wave 1 that they started looking for a job less than 20 weeks ago. Finally, we excluded a small number of observations where information about the reservation wage is missing and some outliers by dropping respondents who gave extreme reservation wages of either below 500 euro or above 5000 euro. This leaves us with a sample size of 1,024 respondents in wave 1, 381 respondents are observed in both waves 1 and 2, and 217 respondents are observed in all three waves.

Table 1. Number of observations across waves

Number of observations	Wave 1	Wave 2	Wave 3
Total respondents	1,744	1,138	963
Unemployed in wave *t*	1,744	802	574
Actively searching in wave *t* and earlier waves	1,265	499	295
Self-reported search < 20 weeks in wave 1	1,090	413	242
Non-missing RW in wave *t*	1,037	392	230
Non-outlying RW in wave *t*	1,024	392	230
Non-missing RW in wave *t* and earlier waves	1,024	381	217

The main analysis estimates changes over time of the reservation wage (first differences) and individual fixed effects regression models $rw_{it} = \alpha_i + \beta d_t + \epsilon_{it}$, where rw_{it} is the reservation wage of job seeker i in wave t, d_t is the unemployment duration in months and α_i control for individual fixed effects. The first difference estimates are based on the 381 and 217 job seekers in waves 2 and 3 that remain unemployed over this period. The estimates identify the effect of unemployment duration on reservation wages in the population of job seekers that remain unemployed between the measurement periods in the survey and under the assumption that there are no time-variant unobservables that are correlated with reservation wages. We will show that the first differences and fixed effects estimates are very similar, we will compare these estimates to OLS estimates to study the amount of bias in cross-sectional studies and we will explore the heterogeneity of the effect with respect to job seeker characteristics by adding interaction terms to the model. Although the panel nature of the data and the design allow us to estimate the effect of unemployment duration on reservation wages controlling for individual fixed effects, there are certain limitations to the analysis presented below. First, the fixed effects estimator has a price in terms of statistical power. Standard errors of fixed effects estimates tend to be large because they rely only on within-individual variations in the data. This increases the risk of not finding an effect that is significantly different from zero, while the true effect is different from zero (type II error). Second, the use of fixed effects estimators is even more troublesome in the case of unemployment duration, because the many transitions of job seekers to employment and discouragement rapidly limit the sample sizes and further inflate the standard errors involved. Third, the data used relate to an unemployment duration of six months with measurements at zero, three and six months. One implication is that the unemployed who find a job in less than three months are in fact excluded from the population we study. The estimated effects in this study concern the population of unemployed who remain in unemployment for at least three to six months. We cannot exclude the possibility that the duration effect is different for job seekers who remain unemployed for a short duration so that they leave the sample before the second measurement period (the duration effect is probably greater in this group because job seekers who reduce their reservation wage more rapidly are more likely to accept a job offer and exit from the sample). Another implication is that we have no data on longer-term effects, so it is unclear whether the estimates remain valid outside the sample time frame.

Table 2 shows descriptive statistics for the main variables in the sample at wave 1. The mean reservation wage is 1580 euro. Immigrants are defined as persons whose current or previous nationality was of a state outside of the EU or EFTA. About 9 percent of the sample at wave 1 are immigrants, which may seem implausible low but it should be kept in mind that we do not study the population of job seekers but the population of new job seekers who had a paid job before they became unemployed. 26 percent worked part-time in their previous job. Educational levels are measured in 4 categories. The mean previous monthly wage was

1591 euro net. 6% had had a meeting with a job coach from the Public Employment Service (PES). On average, job seekers had sent out 12 applications for job vacancies in the last three months. 75 percent were receiving unemployment benefits at the moment of survey 1 (the non-receivers are mostly people who became unemployed very recently). The average level of benefits in this sample is 671 euro (the level is set to zero for those who have no benefits).

Table 2. Descriptive statistics

	count	mean	sd
Reservation wage (euro)	1024	1579.72	437.94
Male (%)	1024	.42	.49
Immigrant (%)	1024	.09	.28
Age (years)	1024	37.12	10.52
Part time (%)	1017	.26	.44
Degree: No secondary (%)	1024	.30	.46
Degree: Secondary (%)	1024	.44	.50
Degree: BA (%)	1024	.16	.37
Degree: MA (%)	1024	.11	.31
Previous wage (euro)	917	1591.46	489.34
PES job coaching (%)	1024	.06	.23
Applications (number sent)	982	12.20	16.46
UB receiver (%)	923	.75	.43
UB level (euro)	923	671.40	468.73
Self-reported search duration (weeks)	1024	8.39	5.68

Note: Sample includes all job seekers at wave 1.

5) Results

a) The effect of unemployment duration on reservation wages

Table 3 describes the relation between reservation wages and unemployment duration in our sample. Panel A presents the mean reservation wages in each wave in the pooled cross-section, i.e. the sample that combines the job seekers in all three waves. The mean reservation wage slightly increases from 1580 euro in wave 1 to 1590 euro in wave 3, which is an increase of 1.2% or 0.2% per month on average. The pooled cross-sectional estimate is equivalent to a standard cross-sectional estimate in a sample of job seekers with varying lengths of unemployment duration under the weak assumption that separation and hiring rates are constant over time. So panel A presents a good estimate of the effect of unemployment duration on reservation wages that we would expect to see in a standard cross-sectional study. The large drop of the sample size with duration in panel A, particularly between waves 1 and 2, is due to the large job finding rates of newly unemployed. This large drop demonstrates the fact that populations of unemployed at different durations of unemployment are hard to compare and severely risk introducing selection bias.

Panel B presents the mean reservation wages in a subsample which only contains job seekers who were still unemployed at wave 2, so that the sample means at waves 1 and 2 are now calculated using the same group of 381 respondents. In this group the mean reservation wage decreases from 1613 euro in wave 1 to 1582 euro in wave 2, which corresponds to an average decrease of 0.6% per month. Panel C further restricts the sample only to those 217 job seekers who remained unemployed in all three waves. The mean reservation wage in this subsample decreases from 1649 euro in wave 1 to 1604 euro in wave 3, which represents an average decline 0.45% per month (the rate of the decline is about 0.35% in the first three months and it accelerates to 0.55% in months 3 to 6). The results are similar if the reservation wage ratio, which is defined as the reservation wage divided by the previous wage, is used as a measure instead of the reservation wage itself (online appendix B presents the reservation wage ratios in a way that is directly comparable to Feldstein & Poterba (1984) and Krueger & Mueller (2016)).

Table 3 also reveals how reservation wages are correlated with selection into the sample. In wave 1 the mean reservation wage in the full sample is 1580 euro, while it increases to 1613 euro (panel B) and 1649 euro (panel C) as the sample is restricted to only those who were still unemployed at waves 2 and 3 respectively. This implies that in wave 1 the reservation wage of those who eventually go on to remain unemployed until after wave 3 is 4.4% above the average reservation wage in the full sample. This finding is consistent with job search theory which predicts that a lower reservation wage, all else equal, increases the probability of finding a job. The probability of selection into the sample by remaining unemployed is not just correlated with the initial level of the reservation wage, but also with the rate of decline of the reservation wage across waves. Compare the reduction of the reservation wage in the first three months between panels B and C. Among those who remain unemployed for at least three months the reservation wage declines by 0.6% per month on average in the first three months, whereas this rate is only 0.35% among those who remain unemployed for at least six months. This suggests that job seekers who lower their reservation wage less rapidly, are more likely to remain unemployed and to be included in the sample.

Table 3. Mean reservation wages by unemployment duration (in euro)

	A. Pooled cross-section		*B. Longitudinal (sample at wave 2)*		*C. Longitudinal (sample at wave 3)*	
Wave	N	Mean (SE)	N	Mean (SE)	N	Mean (SE)
1 (month 0)	1024	1579.7 (13.7)	381	1612.6 (23.8)	217	1649.3 (34.9)
2 (month 3)	392	1582.6 (22.1)	381	1582.3 (22.6)	217	1631.8 (32.2)
3 (month 6)	247	1589.9 (27.0)	-	-	217	1604.9 (28.7)

Table 4 presents regression estimates of the effect of unemployment duration on reservation wages in different specifications. In the linear models (1-3) the dependent variable is the reservation wage (in euro), whereas in the log-linear models (4-6) the dependent variable is the natural logarithm of the reservation wage so that the estimated effects can be interpreted in relative terms. Models (1) and (4) present estimates based on simple models, which essentially reproduce the earlier results from panel A in table 3: cross-sectional OLS estimates suggest that the mean reservation wage increases by 0.1% per month. The coefficient is positive, but it is small and it does not significantly differ from zero. Models (2) and (4) produce somewhat better estimates by controlling for observables. These estimates would suggest that reservation wages decline by about 0.2% per month (not significant). The coefficient estimates of the controls provide further insight into the determinants of reservation wages. Although these estimates could be affected by omitted variables, they suggest that the reservation wages of men are 3.1% above those of women, after controlling for their previous wage level and the other included variables. Older job seekers, part-time workers and the higher educated also ask higher wages and the estimated elasticity with respect to previous wages is 0.437, i.e. a 1 percent increase of the previous wage increases the reservation wage by approximately .437 percent.

Models (3) and (6) present the main results. These fixed effects models exploit only within-individual variation in reservation wages to estimate the effect of unemployment duration on reservation wages. We find that reservation wages decline with unemployment duration by about 0.4% per month. These fixed effects estimates are more negative than the OLS estimates (although controlling for the usual observables in a cross-sectional study appears to reduce bias by more than half). The statistical significance of the FE estimates does not depend on the fact that the sample includes respondents who were only inquired once: restricting the sample to respondents whose reservation wages were observed at least two times, produces the same results. To test whether the rate of decline of the reservation wage is constant, we also estimated a specification that includes a quadratic term for unemployment duration (results not shown): the point estimates suggest that the rate of decline decreases with a minimum reservation wage reached at 5 to 7 months of unemployment duration, but the coefficients are not statistically significant (which could be related to limitations in the sample size and the limited variation we can exploit in our fixed effects specification).

Table 4. The effect of unemployment duration on reservation wages

	Linear models			Log-linear models		
	(1)	(2)	(3)	(4)	(5)	(6)
	OLS	OLS	FE	OLS	OLS	FE
Months	1.547	-3.439	-7.356*	.001	-.002	-.004**
	(4.183)	(3.634)	(3.764)	(.002)	(.002)	(.002)
Male		52.051**			.031**	
		(25.129)			(.013)	
Immigrant		-1.510			.008	
		(30.775)			(.018)	
Age		5.259***			.003***	
		(1.156)			(.001)	
Part time		115.664***			.062***	
		(30.509)			(.017)	
Degree: secondary		-34.499			-.019	
		(24.384)			(.015)	
Degree: BA		41.053			.029	
		(36.195)			(.020)	
Degree: MA		170.478***			.079***	
		(51.920)			(.026)	
Prev. wage (log)		782.177***			.437***	
		(70.504)			(.036)	
PES: job coach.		-18.334			-.000	
		(37.334)			(.021)	
Applications (log)		11.335			.004	
		(10.590)			(.006)	
UB receiver		20.805			.014	
		(24.779)			(.014)	
UB level (log)		75.262**			.050***	
		(30.472)			(.018)	
Observations	1663	1361	1663	1663	1361	1663
Groups			1040			1040

* p<.10; ** p<.05; *** p<.01. Significance is based on clustered (OLS) and cluster-robust (FE) standard errors in parentheses. Constant terms are included in all models.

b) Heterogeneity of the duration effect

The effect of unemployment duration on reservation wages differs across individuals and groups if the terms in equation (2) depend on the personal characteristics of job seekers. We argue that the main heterogeneity stems from the fact that job seekers "anchor" their reservation wage to the wage level in their previous job, that they do not know the distribution of the wage offer function and that reservation wages fall with unemployment duration as they adjust the mean of the wage offer function. It is well-known from the empirical literature, as well as from theories about imperfect competition in the labor market (Manning, 2003), that the wages of equally productive workers vary substantially (wage dispersion). If workers are unsure about the distribution of the wage offer function, it makes sense for a job seeker to use the information about the previous wage level w^p to infer the wage offer function: $\mu_0 = f(w^p)$ with $f' > 0$. Similarly, job seekers who happened to have been employed at a low wage previous employer, are likely to set a lower reservation wage and will exit from unemployment sooner than similar job seekers who happened to have worked at a high wage employer. If similar workers receive similar wage offers w_t (on average), then equation (3) predicts stronger declines of the mean of the wage offer distribution (and the reservation wage) for workers with high previous wages.

We also hypothesize that the effect of unemployment duration on reservation wages is stronger among men than among women. The literature suggests that men are more likely to be overconfident, to be risk-loving and to initiate negotiations (Bertrand, 2010; Azmat & Petrongolo, 2014). Such traits put men at an advantage over women in securing promotions or higher wages. But the same traits may lead to longer unemployment spells for men. Overconfident job seekers are more likely to overestimate the mean of their wage offer distribution and job seekers who are risk-loving are more likely to choose a higher reservation wage because the random drawing of wage offers and the choice of an reservation wage in job search models essentially entails a trade-off between the certainty of finding a job soon and the pay-off of securing a higher wage (Dubra, 2004; Pannenberg, 2010). Their confidence and risk profiles might lead men to set higher reservation wages than women. As men and women are confronted with similar actual wage offers, we would expect to see stronger adjustments among male job seekers. There are other possible explanations for a stronger response among men, which do not assume male overconfidence. The disutility of being unemployed could increase at a faster rate among men if they, more than women, become increasingly uneasy with being unemployed – for example because being unemployed does not fit with the breadwinner role expectations that continue to be associated with men. Unemployment may also carry more of a stigma for men then for women because women are more often expected to experience fragmented careers (Mooi-Reci & Ganzeboom, 2015). In these types of explanations, we would expect stronger declines in reservation wages among men, but no higher asking wages at the start of the unemployment period.

In order to test whether the effect of unemployment duration on reservation wages is heterogeneous across job seekers, table 5 presents additional regression models in which interaction terms were added between unemployment duration and a series of dummy variables indicating job seeker characteristics. All these models control for individual fixed effects and so the base model (1) in table 5 is model (6) in the previous table 4. The results show significant interactions with gender and previous wages: males and job seekers with high (i.e. above the median) previous wages experience sharper reductions in reservation wages with unemployment duration. The size of the estimated coefficients suggests, for example, that the mean reservation wage of male job seekers declines at a monthly rate that is .8 percentage points faster than that of female job seekers. The other interactions, including those relating to the number of applications and coaching by the public employment services agency, are not significant. The significance tests are t-tests on each of the estimated coefficients based on cluster-robust standard errors, but it could be argued – as one reviewer pointed out – that the model in Table 5 amounts to the testing of multiple hypotheses which increases the risk of Type I errors. The estimates would not be significant if Bonferroni-type corrections were made to the p-values. On the other hand, it should be kept in mind that not rejecting a null hypothesis does not imply that there is no effect, especially not in fixed effects models where standard errors are inflated because only intra-individual variation is exploited. Some of the non-significant effects are also worth mentioning because their point estimates are substantial: receiving a high level of unemployment benefits (i.e. above the median) and job coaching from the public employment service appear to counteract the effect of unemployment duration on reservation wages.

Table 5. Interactions in the effect of unemployment duration on reservation wages

	Coef.	(SE)
Months (M)	-.004**	(.002)
M. x Male	-.008*	(.004)
M. x Immigrant	.004	(.006)
M. x Age>=50	-.005	(.005)
M. x Part-time	-.006	(.005)
M. x Degree>=BA	-.000	(.005)
M. x High wage	-.010**	(.004)
M. x Job coach.	.007	(.005)
M. x Many applic.	.002	(.004)
M. x High UB	.007	(.004)
Observations	1361	
Groups	822	

* p<.10; ** p<.05; *** p<.01. The dependent variable is the log reservation wage. All variables are centered around their mean. The model controls for individual fixed effects. Significance is based on cluster-robust standard errors.

c) Heterogeneity by gender

Table 5 showed that the asking wages of men decline at a faster rate than those of women. This finding that men respond more strongly than women to unemployment duration, is confirmed by an additional analysis of the extent to which job seekers perceive it is “easy to find a job”: a regression controlling for individual fixed effects shows that both men and women believe it is harder to find a job as unemployment duration increases, but the rate of change for men is almost double that of women (see online appendix C for these results).

The question rises as to what explains this gender difference in the effects of unemployment duration. Although we have no conclusive evidence, our data do shed some light on the underlying process. Table 6 compares the reservation wages of men and women at the start of the unemployment period by means of a Oaxaca decomposition (which relies on estimated OLS regression coefficients, involving the usual assumptions). The mean reservation wage of women (1482 euro) is about 10 percent lower than that of men (1641 euro). However, around two thirds of this gender gap in asking wages (58%+8%) can be attributed to gender differences in previous wages and age. Moreover, gender differences in psychological attributes explain very little of the gender gap in asking wages: 4% of the raw gender gap in reservation wages or about 7 euro, i.e. about 0.4% of the reservation wage. The psychological attributes include measures for self-esteem, risk-taking and the extent to which job seekers believe they are good at finding a job (see online appendix D for details). This evidence suggests that the gender gap in reservation wages at the start of the unemployment period is small and that it is not driven by male overconfidence.

Table 6. Decomposition of the gender difference in reservation wages

	Coeff.	Pct.
Overall decomposition:		
Reservation wage: men	1640.69***	.
Reservation wage: women	1482.04***	.
Raw difference	158.65***	100
explained	113.01***	71
unexplained	45.64*	29
Detailed decomposition:		
Previous wage	92.62***	58
Age	13.33**	8
Psychological attributes	7.06	4
Sum of the contributions:	113.01	71
Observations: men	330	
Observations: women	283	

* p<.05; ** p<.01; *** p<.001. Sample of full-time workers at wave 1. The psychological attributes contain measures for self-esteem and risk taking in career decisions.

6) Discussion and conclusions

Our longitudinal data on unemployed workers in Belgium who were interviewed at three-month intervals, show that reservation wages decline by about 0.4 percent per month of unemployment duration – or about 5 percent per year. The results confirm the long-held belief that cross-sectional estimates of the effect of unemployment duration on reservation wages are subject to substantial upward bias because exits from the unemployed population are non-random, and we show that adding the usual controls only partly eliminates this bias. The main difference between this study and earlier longitudinal studies, has to do with the frequency of the data collection. Most earlier studies use yearly data from panel surveys, which produce results that are mixed and in some cases hard to explain – perhaps because this restricts the population to job seekers who remain unemployed for at least one year. One seminal study by Krueger and Mueller (2016) uses weekly data and finds that reservation wages decline at a modest rate. Our estimates are in the middle of the range of their point estimates so the remaining question is the difficult judgment of whether a 5 percent per annum decline in reservation wages is modest or substantial. In our view, a 5 percent annual change is substantial – certainly when compared to changes in other income measures. Between 1950 and 1973, incomes in Western Europe grew at (close to but) less than 5 percent per annum (Crafts & Toniolo, 1996) and yet this is widely agreed to be the 'golden age of European capitalism', 'les trentes glorieuses' and a period of 'exceptionally high growth'. In Greece, the European country hardest hit in the aftermath of the 2008 economic crisis, real wages fell by 2.4 percent per annum between 2007 and 2015 (OECD, 2016) – a drop which again is widely considered to be substantial.

Falling reservation wages imply an additional channel in the effects of unemployment and business cycles on wages. Although empirical evidence on the relation between unemployment and wages shows mixed results (Abraham and Haltiwanger, 1995), studies that focus on newly hired workers suggest that starting wages are much more cyclical (Pissarides, 2009; Haefke et al., 2013). The increased average duration of unemployment over which reservation wages fall is an additional channel through which starting wages can be affected, although other channels, such as productivity fluctuations and reservation wage adjustments resulting from changes in the job offer arrival rate, may be more important. The finding of falling reservation wages may also require a reinterpretation of the observed effects of unemployment duration on starting wages by Schmieder et al. (2016), who assume that reservation wages are constant over the unemployment spell and therefor conclude that the 0.8 percent per unemployment month fall in starting wages does not result from changing reservation wages but from stigma effects and human capital depreciation. Falling reservation wages imply that the duration effect on starting wages may be overestimated. However, part of the decline in reservation wage could be the result of rational job seekers taking these scarring effects – of human capital depreciation and stigma associated with unemployment duration – into account by adjusting their expected wage offer distribution.

We find stronger declines of the reservation wage among men, which cannot be explained by male overconfidence. Instead, the results are compatible with a theory of greater stigma effects against male workers or faster increases in the disutility of being unemployed among men – perhaps because men become uneasy with being unemployed at a faster rate than women because being unemployed does not fit with the breadwinner role expectations that continue to be associated with men. Our finding of stronger declines among workers who earned high wages in their previous jobs, is compatible with a theory of wage dispersion in previous jobs, anchoring of reservation wages to these previous wage levels and an adjustment process of reservation wages to received wage offers during unemployment. Two additional factors showed substantial point estimates, although these were not statistically significant: receiving a high level of unemployment benefits and job coaching from the public employment service appear to counteract the effect of unemployment duration on reservation wages. Further research is needed to study how country differences in such institutions relate to reservation wage dynamics.

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