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RETIREMENT AND HEALTH

LABOUR MARKET

Inequalities in Expected Retirement Duration and Expected Disability-Free Retirement Duration

Julien Blasco* and Ulysse Lojkine**

Abstract – This paper develops an indicator of disability-free retirement expectancy (DFRE) based on mortality, disability and retirement data, alongside the retirement expectancy (or expected retirement duration – RE) indicator. It then measures inequalities in disability-free retirement expectancy by gender, socio-occupational category and educational attainment, using French data from 2009 to 2019. The social gradient of DFRE, for a given sex, is steeper than that observed for RE: a blue-collar worker can expect fewer years of retirement than a manager of the same sex, and for these years, more years of retirement with disabilities. These inequalities in DFRE are explained by differences in disability and mortality, partly offset among men by the earlier retirement of blue-collar workers and the less educated. Over the period, RE declined as a result of later retirement, whereas, due to a reduction in disability, DFRE rose among women and the more highly educated.

JEL: I14, J14, J26
Keywords: retirement, health, inequalities

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Debates on the sustainability and fairness of the pension system and its parameters often raise questions concerning the division of lifetime between working years and retirement years. For example, according to the explanatory memorandum to the 2003 French pension reform bill, the article 5 aimed "to stabilise the ratio between time spent working and time spent in retirement by 2020 in order to ensure the sustainability of pay-as-you-go schemes and fairness between generations" (Fillon, 2003).

Life expectancy – particularly its evolution over time and its distribution across the population, has been a frequent topic of both public debate and economic and statistical research into the French pension system. Life expectancy has been rising steadily for several decades, both at birth and at age 60[1] (Thélot, 2025). Without reforms aimed at postponing the retirement age, most of these gains in life expectancy would probably have led to an increase in the length of retirement, thereby reducing the ratio of working life to retirement duration. The reforms of 2003, 2010 and 2014 may have helped to stabilise this ratio (Aubert & Rabaté, 2014).

However, this average ratio may conceal major differences between population categories, particularly according to gender and social category. Although women live longer than men, they tend to retire later (Bonnet *et al.*, 2006). Similarly, blue-collar workers – despite retiring earlier on average than people in higher-level occupations – have a lower life expectancy (Blanpain, 2024b). Expected length of retirement – an indicator of the number of years an individual can expect to spend in retirement – can thus be used to measure how these different factors combine. Bonnet *et al.* (2023) show that blue-collar workers, despite an earlier average retirement age, spend less time in retirement than people in higher-level occupations (two years for men). Women spend between three and five more years in retirement than men, depending on their profession and socio-occupational category (SOC).

However, such an indicator treats all retirement years as equal. Yet these years may differ greatly in terms of health and its impacts on daily life, notably due to age-related loss of autonomy. It may therefore be beneficial to use the indicator of disability-free life expectancy (or "healthy life expectancy"), in which the notion of disability refers to lasting limitations in daily activities for health reasons.

Between 2008 and 2023, disability-free life expectancy at birth increased much less than life expectancy for men, and even declined slightly for women (Deroyon, 2024). As a result, the proportion of years lived without disability over the entire lifespan fell on average. By contrast, disability-free life expectancy at age 65, which is more relevant for measuring the conditions under which the population ages, increased faster than life expectancy, for both men and women. This long-term trend can be observed in most developed countries (Robine *et al.*, 2020).

Once again, these general trends mask major differences between population categories, and especially between social categories. It is well established that blue-collar workers, compared with people in higher-level occupations, are subject to a "double disadvantage": not only is their life expectancy lower, they also spend more of those years with disabilities (Cambois *et al.*, 2008; Cambois *et al.*, 2011). Other national studies have highlighted the same phenomenon according to level of educational attainment: people with a secondary or lower level of education live shorter lives, and with more disabilities, than higher-education graduates (Sihvonen *et al.* (1998) show this in Finland, Remund & Cullati (2022) in Switzerland). In a review of international literature, Cambois *et al.* (2020) show that this "double-penalty" phenomenon, although varying in magnitude according to gender and country, is found in all developed countries.

Whether measured at birth or at age 65, disability-free life expectancy can nevertheless give only a partial idea of the time spent in retirement without disability. To measure this, it is necessary to take account of the distribution of years of disability over a person's lifetime, and the age at which people retire. Yet this is strongly correlated with disability.

However, the link between disability, length of career and the age at which pension rights are claimed is not clear-cut: on the one hand, certain mechanisms enable people with disabilities[2] to retire earlier than the rest of the population; on the other hand, people with disabilities generally have less complete careers, forcing them to retire later to reach a given pension level. Retirement itself is also known to have an impact on health (Caroli *et al.*, 2023). Finally, in France, people with severe disabilities claim their pensions

1. Life expectancy at birth rose from 81.9 years for women in 1995 to 85.6 years in 2024 (+3.7 years), and from 73.8 to 80.0 years for men (+6.2 years). Life expectancy at age 60 also increased significantly over this period, by 2.9 years for women and 4.0 years for men in 2023.

2. Early retirement mechanisms exist for long careers, arduous occupations or permanent incapacity for work (see Aubert, 2025). However, in the latter case, the people concerned will benefit from few or no years of disability-free retirement.

later on average than the rest of the population (Aubert, 2020). International comparative analyses also show wide variations between relative lengths of careers with and without disability. In some countries, the proportion of years spent working with disability (in absolute terms or relative to the total number of years lived with disability) is much higher than in others (Lièvre *et al.*, 2008).

Inequalities between generations, sexes and socio-occupational categories in terms of expected length of time spent in retirement (or retirement expectancy, RE hereafter) are thus combined with inequalities in terms of disability, which may, in certain cases, reinforce or compensate for them. It is known that blue-collar workers live fewer years without disability than people in higher-level occupations, and that they also live fewer years in retirement. To what extent do they live fewer years of disability-free retirement? Does their earlier retirement allow them to more years of disability-free retirement?

Several studies have calculated durations of disability-free retirement. For example, based on disability data and career simulations, the Pensions Advisory Council (COR) (2015) and Cazenave-Lacroutz & Godet (2016) produced time projections and a comparison by gender of disability-free retirement durations. Aubert (2020, 2022) moved from this career-simulation-based approach to one modelled on the principle of life expectancy, i.e. a fictitious trajectory constructed from point-in-time data. Aubert (2020) thus calculated the proportion of years spent with disability within the retirement period for women and men, and showed that women spend a larger proportion of their retirement with incapacity because they reach more advanced ages. Aubert (2022) calculated an approximate disability-free retirement expectancy (hereafter referred to as DFRE), using aggregate data on the prevalence of disability and retirement rates by age, requiring a strong assumption that these two variables are uncorrelated.

We have adopted the DFRE approach in this paper, this time using individual data linking disability status with retirement status and breaking down the results by gender and socio-occupational category, and by gender and educational attainment, describing changes over almost ten years. We also introduce a method for decomposing differences in RE or DFRE (whether between two periods, between sexes, or between two socio-occupational or educational categories) in order to explain these disparities by differences in mortality, retirement age and prevalence of disability. In addition to estimating disability-free retirement expectancies for both sexes and for different socio-occupational and educational categories, the analysis of the breakdown of these differences is, to our knowledge, an original contribution of this paper.

We apply these indicators to French data for the 2009–2019 period by combining mortality tables by socio-occupational category and educational attainment published by INSEE with measures of disability and retirement status from the Statistics on Incomes and Living Conditions (SRCV – *Statistiques sur les ressources et les conditions de vie*) survey. For each of two periods (2009–2013 and 2017–2019), we calculate life and retirement expectancy with and without moderate or severe disability, by sex and socio-occupational category, and by sex and qualification.

We mainly highlight the following four phenomena.

1. The “double disadvantage” for members of lower socio-occupational categories: not only do members of lower socio-occupational categories (white-collar workers[3] and blue-collar workers) have a lower retirement expectancy than the members of higher categories (higher-level occupations), the gap is even greater when it comes to disability-free retirement expectancy. This finding updates and extends the “double disadvantage for workers” observed by Cambois *et al.* (2008) and Cambois *et al.* (2011) with regard to life expectancy. We also show that this phenomenon is also observed according to educational attainment, when comparing individuals with upper-secondary qualification level or less (*baccalauréat* or less) and higher-education graduates.

2. While social inequalities in retirement expectancy can be explained by differences in mortality, social inequalities in disability-free retirement expectancy can largely be explained by differences in disability between social groups. In both cases, among men, earlier retirement among members of lower socio-occupational categories plays a partially compensatory role.

3. Between the 2009–2013 and 2017–2019 periods, retirement expectancy fell for both men and women, while disability-free retirement expectancy increased for women, but not for men.

3. *Lower-level socio-occupational category, corresponding to clerical and service workers (*employés *in French), excluding managers.*

4. These average trends conceal disparities according to educational attainment and social category. In particular, between the two periods, the gaps between higher education graduates and people with below-baccalaureate-level qualifications widened both for retirement expectancy and for disability-free retirement expectancy.

In the remainder of this article, we begin by describing the data used and the method used to calculate the indicators and break down the differences between categories (Section 1). We present our results in Section 2 and discuss their interpretation and certain methodological choices in Section 3, before concluding.

1. Data and Method

1.1. Data

We used mortality tables by sex, age, socio-occupational category and qualification that are calculated by INSEE (Blanpain, 2024b) on the basis of the Permanent Demographic Sample, an individual panel produced by INSEE that includes census data. These tables are calculated over periods of several successive years, the most recent being 2000–08, 2009–13, 2017–19 and 2020–22. They are adjusted and smoothed.[4]

We selected two periods: 2009–2013 and 2017–2019.[5] In the rest of the text, when a statement does not explicitly refer to a given period, it refers to the 2017–2019 period. These data cover the whole of France, excluding Mayotte for 2009–2013 and including Mayotte for 2017–2019.

In these mortality tables, we used the "Survival" column. Calculated on the basis of mortality quotients, it gives the probability of survival by age reached during the year: survival at age a, for $a \geq 30$ years, corresponds to the probability of being alive on 1st January of the year of one's a-th birthday for a fictitious cohort of 30-year-olds who experience at each age the mortality probabilities observed during the period under consideration.

We also compiled, by period and sex, an aggregated mortality table for the "*Baccalauréat* or less" qualification category, grouping together all qualification categories except for higher education, calculated as the average, at each age, of the mortality quotients for these sub-categories, weighted according to their respective weight in the population.[6] We used the same procedure to calculate mortality quotients by social category, both sexes combined.

The proportions of retired and non-retired people with or without a moderate or severe disability were estimated from the *Statistiques sur les ressources et les conditions de vie* (SRVC – Statistics on Household Resources and Living Conditions) survey, covering metropolitan France.[7] For synchronic comparisons between categories, we stacked the data corresponding to the mortality table periods described above (2009 to 2013 on the one hand, 2017 to 2019 on the other).

We used the first level of the classification of occupations and socio-occupational categories (*professions et catégories socioprofessionnelles* – PCS), which comprises six categories, known as socio-occupational categories (*catégories socioprofessionnelles* – CS). For reasons of sample size (and because age-specific disability differed little between the sexes), the prevalence of age-specific disability for farmers, tradespeople, shopkeepers and company managers was calculated for all sexes combined. In the body of the article and the figures, we focus on the four most numerous socio-occupational categories (higher-level occupations, intermediate-level occupations, white-collar workers and blue-collar workers).[8]

Qualification levels are taken from question DIP11 of the SRCV (DIP14 before 2013) and grouped as follows to correspond to the INSEE mortality tables by qualification: no qualification; lower secondary (*brevet des collèges*), primary education certificate (*certificat d'études primaires*); CAP (*certificat d'aptitude professionnelle*) and BEP (*brevet d'études professionnelles*) or equivalent vocational qualifications; upper secondary (*baccalauréat*) (general, technological, vocational) or equivalent; higher-level qualification (including paramedical and social at *bac*+2 level, BTS (*brevet de technicien supérieur*), DUT (*diplôme*

4. Blanpain (2024a) provides a detailed explanation of the smoothing methods used – a modified Brass method (Leridon & Toulemon, 1997) up to 2009–2013, followed by the spline method (Camarda, 2012). The same document indicates that for the periods of interest to us, these treatments are minor as the raw data are of good quality, .

5. For earlier periods, data on disability and retirement appear to be unavailable for comparison, due to a break in the series since the 2008 SRVC survey. Finally, we do not take account of the most recent data (2020–2022) as this corresponds to the COVID period, which was exceptional from the perspective of mortality and health.

6. This weight is measured in the SRVC survey within the corresponding age category (see below on the age categories used).

7. We also reproduced our calculations, for the 2017–2019 period, based on the Labour Force Survey; this robustness test is discussed in Appendix 2.

8. In French, respectively cadres, professions intermédiaires, employés *and* ouvriers.

universitaire de technologie), bachelor's and master's degrees).[9]

Over the period in question, this survey covered all ordinary households in metropolitan France. Around 25,000 observations are available each year, including more than 7,000 individuals between the ages of 30 and 100 of each sex. In the end, we had more than 2,000 observations of individuals aged between 30 and 100 for each sex × period × qualification or sex × period × socio-occupational category for people in higher-level occupations, intermediate-level occupations, white-collar workers and blue-collar workers. Furthermore, the scope of the survey excludes collective housing, which tends to underestimate the prevalence of limitations in the oldest age groups, where the proportion of people living in collective housing (nursing homes for dependent elderly people) increases, particularly among people with disabilities in daily life.[10]

Disabilities are measured on a self-reported basis; the question asked of respondents is: "Have you been limited for at least six months, due to a health problem, in the activities that people usually do?" Three answers are possible: "Yes, severely limited", "Yes, limited, but not severely" and "No, not limited at all". The measurement of disability, using this question, is known as GALI, (Global activity limitation index).[11] The same indicator, from the same survey, is used by DREES (Ministerial Statistical Office for the Ministries of Health and Social Affairs) (Deroyon, 2024) to calculate disability-free life expectancy by sex. The GALI indicator is also used in European comparisons, via the EU-SILC (Statistics on Income and Living Conditions) survey system.

Retirement status is also measured on a self-reported basis, by the response to the question on "main work situation" in the SRCV survey, which also considers anyone aged 70 or over as retired.[12]

Adopting the DREES method of calculating disability-free life expectancy, the joint prevalence series of disability status and retired status was constructed by calculating an average per five-year age bracket (30 to 34, ..., 80 to 84) and then a category from 85 to 100 corresponding to the ages at which fewer observations are available.[13] On average, these sex × period × (socio-occupational category or qualification) × age group cells contain several hundred observations. The smallest correspond to male white-collar workers aged 85 and over in 2017–2019 (27 observations) and, for the same period and age group, female higher education graduates (52 observations).

1.2. Method and Indicators

We thus have the following variables for each group G (sex, socio-occupational category and qualification groups) and for each age a, at date t:

– $DF_{G,a,t}$ the proportion of disability-free people (within the meaning of the GALI indicator) among people of age a in group G at time t;

– $Retired_{G,a,t}$ the proportion of retirees;

– $DFR_{G,a,t}$ the proportion of disability-free people among retirees;

– $Survival^{attained}_{G,a,t}$ the proportion of survivors in group G at time t, at attained age a.

From this last variable, we calculate the survival at exact age (probability of being alive on the day of one's a-th birthday for the fictitious cohort mentioned above): $Survival^{exact}_{G,30,t}=1$ and, for subsequent ages, $Survival^{exact}_{G,a,t}=\frac{1}{2}\left(Survival^{attained}_{G,a,t}+Survival^{attained}_{G,a+1,t}\right)$, and then the number of person-years[14] at age a (the number of years lived at age a within the same fictitious cohort):

$$L_{G,a,t}=\frac{1}{2}(Survival^{exact}_{G,a,t}+Survival^{exact}_{G,a+1,t}).$$

Given that all the life expectancies, total or without disability, and retirement durations, total or without disability, in this article are calculated at age 30, we can then define life expectancy at age 30 as:

$$LE_{G,30,t}=\sum_{a=30}^{100}L_{G,a,t}.$$

It should be borne in mind that life expectancy in a given year is a statistical calculation measuring the average lifespan of a fictitious generation

9. *Adding the dimension of educational qualifications to that of socio-occupational category has several advantages: educational attainment is more stable after age 30 and is more easily comparable between countries. See Section 3.2 for a discussion of the differences between these two indicators.*
10. *In 2021, 5% of people aged 65 or over lived outside ordinary housing (Daguet, 2025). In Section 3.1 we discuss the implications of this exclusion and show that it is unlikely to change the conclusions.*
11. *For a more detailed discussion of the benefits and limitations of this indicator, see Cambois* et al. *(2015), who study the effect of variations in the wording of the question, Dauphin & Eideliman (2021) and Louvel & Monirijavid (2024) who study the overlapping of this indicator with other self-reported or administrative disability indicators, and Galenkamp* et al. *(2020) for an international perspective.*
12. *See Appendix 3 for two robustness tests on the effect of the definition of retired status on our results.*
13. *In Appendix 1, as a robustness test, we examine an alternative imputation method based, at advanced ages, on a linear model of disability prevalence as a function of age; the results show little sensitivity to this variant.*
14. *Except for the final age, where we apply the conventional assumption of a constant mortality quotient for all subsequent years, giving:* $L_{G,100,t}=Survival^{exact}_{G,100,t}/m_{G,99,t}$ *where* $m_{G,99,t}$ *denotes the central mortality quotient at age 99 – see Jagger* et al. *(2001), p. 25-26 and Preston* et al. *(2001), p. 48.*

that is subject at each age to the mortality rates observed in that year. It differs from the average lifespan.

The Sullivan (1971) method (see Jagger *et al.*, 2001 for a more recent presentation) enables the combination of age-specific mortality tables and age-specific disability prevalence tables to calculate disability-free life expectancy:

$$DFLE_{G,30,t} = \sum_{a=30}^{100} L_{G,a,t} \times DF_{G,a,t}.$$

This is therefore the expected number of years lived without disability for a fictitious cohort that experiences, at each age, the observed risk of death at that age in the population, and, conditional on its survival, the observed risk of disability at that age. This method of calculating life expectancy and disability-free life expectancy at age 30 is that used by INSEE and DREES. Indeed, the values we measure for each sex and each of our two periods correspond to the average, in each period, of the annual values reported by Deroyon (2024).[15]

This definition of disability-free life expectancy using the Sullivan method is based on disability prevalence data at different ages. An alternative definition that is sometimes used is based on flow data, i.e. data on people entering or leaving disability. The scarcity of this information in surveys explains why Sullivan's method has become standard.[16] If disability flows at each age are stable over time, the two meanings coincide, whereas if disability flows at each age tend to decline over time, then the Sullivan indicator has a higher value than the flow-based indicator.

Years of disability measured in disability-free life expectancy are not necessarily consecutive or confined to the end of life. On the one hand, although the disabilities covered by the question used last at least 6 months, they may nevertheless be temporary. On the other hand, it should be borne in mind that the disability-free life expectancy indicator does not reflect a life trajectory, but the aggregate cumulative number of years lived with or without disability at different ages in a population.

It should be noted that calculating a survival rate or life expectancy by social category requires the combination of mortality data across age groups, without taking account of mobility between social categories over the course of a lifetime: the indicator therefore describes a fictitious cohort that belongs to the same social category throughout its career. This decision to neglect post-age-30 mobility between categories seems legitimate in the case of qualification categories. For the aggregated six-group socio-occupational category, immobility is strong but mobility is very real: between 1998 and 2003, Deauvieau & Dumoulin (2010) observed outward mobility flows of around 10 to 20% for each socio-occupational category. The indicators should therefore be interpreted as describing fictitious trajectories internal to socio-economic categories. This assumption nevertheless appears more legitimate for socio-occupational categories than for finely defined income categories (Lojkine, 2022).

Retirement expectancy. Retirement expectancy is also calculated using the Sullivan method, i.e. the trajectory of an individual is simulated based on the assumption that person-years lived in retirement at age *a* equal the product of person-years calculated from the mortality table and the proportion of retirees observed at that age. We thus obtain a retirement expectancy $RE_{G,30,t}$, as in the literature (e.g. Leinonen *et al.*, 2018):

$$RE_{G,30,t} = \sum_{a=30}^{100} L_{G,a,t} \times Retired_{G,a,t}.$$

This calculation method follows exactly the same principle as working life expectancy (Weber & Loichinger, 2022). Note that this notion of retirement expectancy does not overlap exactly with another common notion, that of period-specific retirement duration, calculated as the difference between life expectancy and the period-specific retirement age. This is because retirement expectancy combines the proportions of retirees observed at each age with mortality quotients, whereas the period-specific retirement age is unaffected by the possibility of death before retirement. Estimating retirement duration as the difference between life expectancy and the period-specific retirement age therefore systematically gives a lower figure than retirement expectancy. The magnitude of this bias, both for expectations and for differences in expectations between categories, is discussed in Online Appendix S1 (link to Online Appendix at the end of the paper).

Disability-free retirement expectancy. Similarly, disability-free retirement expectancy is calculated as $DFRE_{G,30,t}$:

$$DFRE_{G,30,t} = \sum_{a=30}^{100} L_{G,a,t} \times \mathrm{Retired}_{G,a,t} \times DFR_{G,a,t}.$$

15. Some figures match to within 0.1 years. This may be explained by our use of Blanpain's (2024b) survival rates calculated directly over multi-year periods, which do not necessarily coincide exactly with the averages of survival curves obtained from mortality for each year separately.

16. The SRCV survey is a panel survey that could, in theory, enable the estimation of transition probabilities between disability statuses; however, the size of the survey sample does not allow reliable estimates of these transition probabilities to be obtained by socio-occupational category or by qualification.

Once again, the use of the Sullivan method based on prevalence data at a given age leads to a different measurement from one based on the observation of flows, when these flows are not stable over time. In this case, the retirement age rises slightly each year during the period under study, and the DFRE calculated using the Sullivan method is therefore higher than a DFRE calculated using retirement flows.

Decomposition of differences in RE and DFRE. When two groups, such as male blue-collar workers and men in higher-level occupations, have different retirement expectancies, what proportion of this difference is explained by the difference in retirement-age profiles, and what proportion by the difference in mortality quotients at each age? We can address this question in the following manner: if male blue-collar workers had the same mortality quotients as men in higher-level occupations, and differed from them only in their retirement ages, how would their retirement expectancies differ? Or if male blue-collar workers had the same proportion of pensioners as men in higher-level occupations at each age, and differed from them only in mortality quotients, how would their retirement expectancies differ? Formally, this amounts to decomposing the RE gap between two groups G and G' as follows:[17]

$$\begin{aligned} RE_{G,30,t} - RE_{G',30,t} &= \sum_{a=30}^{100}\left(L_{G,a,t}\times Retired_{G,a,t} - L_{G',a,t} \times Retired_{G',a,t}\right) = \sum_{a=30}^{100}\left(L_{G,a,t}-L_{G',a,t}\right)\times Retired_{G,a,t} \\ &+ \sum_{a=30}^{100} L_{G,a,t}\times\left(Retired_{G,a,t}-Retired_{G',a,t}\right) \\ &- \sum_{a=30}^{100}\left(L_{G,a,t}-L_{G',a,t}\right)\times\left(Retired_{G,a,t}-Retired_{G',a,t}\right). \end{aligned}$$

The first term is called the "survival-explained gap", the second the "retirement-explained gap" and the third the "residual" or "interaction" term. The same method naturally applies to give a similar breakdown of a DFRE gap:

$$\begin{aligned} \underbrace{\Delta DFRE_{G-G',30,t}}_{\text{DFRE gap}} &= \underbrace{\sum_{a=30}^{100}\Delta Survival_{G-G',a,t}\times Retired_{G,a,t}\times DFR_{G,a,t}}_{\text{due to mortality}} \\ &+ \underbrace{\sum_{a=30}^{100} Survival_{G,a,t}\times \Delta Retired_{G-G',a,t}\times DFRE_{G,a,t}}_{\text{due to retirement}} \\ &+ \underbrace{\sum_{a=30}^{100} Survival_{G,a,t}\times Retired_{G,a,t}\times \Delta DFR_{G-G',a,t}}_{\text{due to disability}} + \text{Residual}, \end{aligned}$$

where the first term refers to the gap explained by mortality, the second to the gap explained by retirement and the third to the gap explained by health. This method also enables the decomposition of the temporal changes in these indicators for a group G, from date t to date t'.

Confidence intervals. As mentioned above, our mortality data are those of Blanpain (2024b), supplied without confidence intervals.[18] The confidence intervals presented here reflect only error in measuring the prevalences of retirement, disability and disability-free retirement.

As indicated above, these prevalences are calculated as averages within age categories. Let $C = \{[[30, 34]], ..., [[85, 100]]\}$ denote the set of these age categories; we can write (as in Jagger, *et al.*, 2001, p. 27):

$$DFLE_{G,30,t} = \sum_{A\in C}\sum_{a\in A} \mathrm{L}_{G,a,t}\times DF_{G,a,t},$$

whence:

$$V\left(DFLE_{G,30,t}\right) = \sum_{A\in C}\left(\sum_{a\in A} L_{G,a,t}\right)^2 \times V\left(DF_{G,a,t}\right).$$

The variance $V(DF_{G,a,t})$ is calculated under the assumption of independent draws between observations of different individuals (without taking account of the sampling design of the SRCV survey), but allowing for the correlation between any successive observations of the same panel individual in stacked waves of the SRCV survey (clustered standard error).

The same principle is used to calculate the variance of RE and DFRE. The 95% confidence intervals are then calculated on the basis of these variances. While these confidence intervals take account of the error in estimating prevalences due to the number of observations in each age category for a given social group and period, by construction, they ignore two other sources of error: firstly, as indicated above, error due to mortality; secondly, error due to the method used to impute age-group prevalences. On the second point, we refer you to Appendix 1 for a comparison with another imputation method.

2. Results

First, we present the differences in retirement (RE) and disability-free retirement expectancy (DFRE) between social categories for the 2017–2019 period, and then decompose these differences into the respective contributions of differences in mortality, disability and retirement. We then examine the changes between 2009–2013 and 2017–2019 in RE and DFRE by sex, measuring the contributions of their various determinants. Lastly, we

17. We have chosen this decomposition because the indicators produced can be interpreted in the same way. We have thus departed slightly from the decomposition of differences in disability-free life expectancy proposed by Nusselder & Looman (2004, p. 317), which has the advantage of not involving a residual.

18. Blanpain (2024a, p. 33-34) presents 90% confidence intervals for life expectancies by social category, but not for mortality quotients or survival rates.

disaggregate this analysis by presenting changes in the differences between social categories and between educational attainment levels.

2.1. The "Double Disadvantage" for the Working Classes

Over the 2017–2019 period, the retirement expectancy of a 30-year-old woman was 24.3 years, including 12.2 disability-free years and 19.0 years without severe disability. For a man, retirement expectancy was 20.2 years, including 10.8 disability-free years and 16.4 years without severe disability. The advantage of women over men with regard to DFRE (1.4 years) is therefore much smaller than with regard to RE (4.1 years).

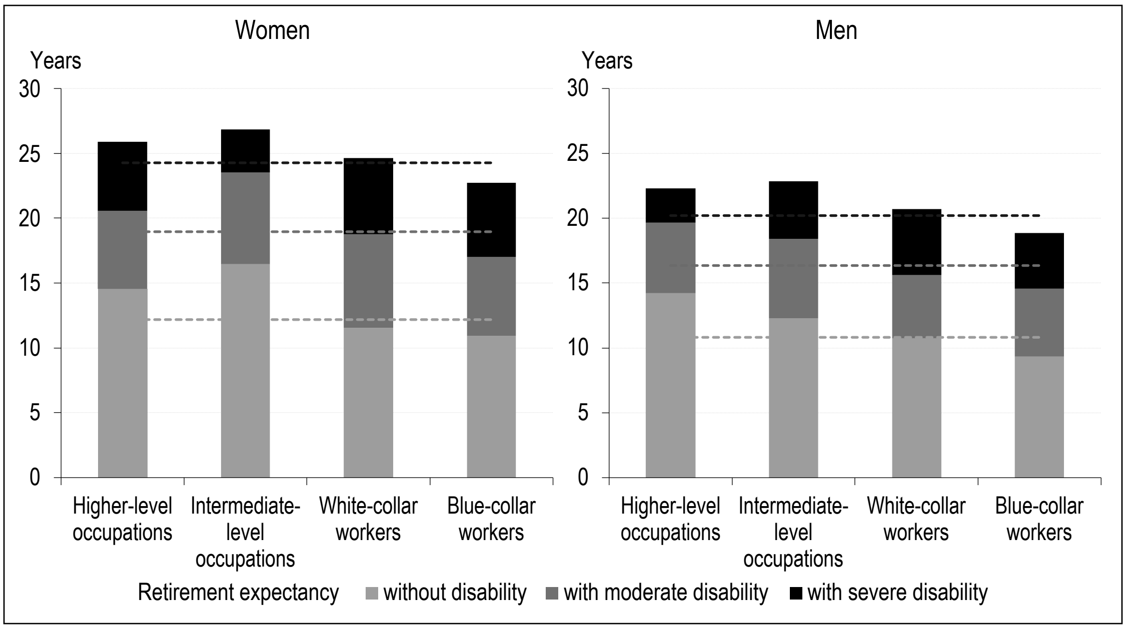
Members of lower socio-occupational categories, whether defined as white-collar workers or blue-collar workers, or by having lower-than-upper-secondary (*baccalauréat* or less) educational attainment, have a lower RE than members of higher-level categories, defined as people in higher-level occupations or higher education graduates. A male blue-collar worker can therefore expect 18.9 years of retirement: 3.5 fewer years than a man in a higher-level occupation (Figure I).

They also spend fewer years in disability-free retirement: the disability-free retirement expectancy of male blue-collar workers is 9.3 years, i.e. 4.9 fewer years than men in higher-level occupations. This means that the DFRE gap exceeds the RE gap by 1.4 years. Similarly, a man with no qualifications can expect 4.9 fewer retirement years than a male higher education graduate, but 5.7 fewer disability-free retirement years (Figure II). Therefore, earlier retirement, on average, among blue-collar workers does not compensate either for their lower life expectancy or for their greater exposure to disability. Of the 10.4 fewer disability-free years over a person's lifetime (from age 30), just over half are "spent" during retirement.

The same applies to women. The gap between a female white-collar worker and a woman from a higher-level occupation is 1.3 years in terms of RE, but 3.0 years in terms of DFRE (for female blue-collar workers versus female higher-level occupations, it rises more modestly from 3.2 to 3.6 years between the two indicators, and the difference is no longer significant). Likewise, the gap between a woman with no qualifications and a female higher education graduate is 5.2 years for RE but 6.6 years for DFRE.

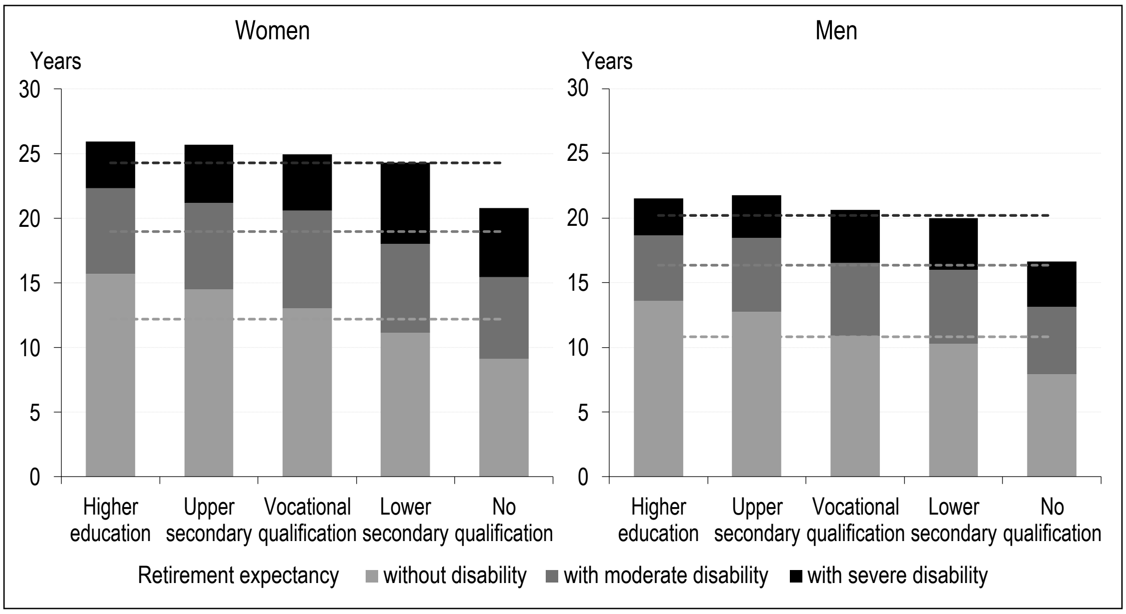
The expression "double disadvantage for the working classes" is inspired by the article by Cambois *et al.* (2008), entitled "*La 'double peine' des ouvriers*" ("A double disadvantage for blue collar workers"), which used data from 2003 to describe a similar phenomenon in terms of life expectancy: the gap between socio-occupational categories in disability-free life expectancy

Figure I – **Retirement expectancy, with and without disability, by socio-occupational category and gender**



Note: The dotted lines represent, from bottom to top, the average retirement expectancy without disability, without severe disability and without total disability, by gender.
Source: SRCV, 2017–2019, and Mortality tables by socio-occupational category, 2017–2019, INSEE.

Figure II – **Retirement expectancy, with and without disability, by qualification and by gender**



Note: The dotted lines represent, from bottom to top, the average retirement expectancy without disability, without severe disability and without total disability, by gender.
Source: SRCV, 2017–2019, and Mortality tables by qualification, 2017–2019, INSEE.

was greater than the gap in life expectancy. We show that the phenomenon persisted in 2009–2013 and 2017–2019, and although the authors' data are not strictly comparable to ours, we observe relatively stable gaps over time (see Online Appendix S2, Figure S2-I). We also show that the phenomenon is observed across educational attainment levels (Online Appendix S2, Figure S2-II). This qualification-based approach enables the detection of larger gaps in life expectancy and length of retirement – the gap between a person with no qualifications and a higher education graduate is greater than the gap between a blue-collar worker and someone in a higher-level occupation.

The gaps in DFRE are also greater than those in RE when comparing higher-level occupations with white-collar workers or higher education graduates with holders of a vocational training qualification (*brevet*, CAP or equivalent). The same applies to severe disability.

Overall, for both men and women, the overall social gradient of disability-free retirement expectancy is steeper than that of retirement expectancy (see Online Appendix S2, Figures S2-III and S2-IV). The only exception concerns women in intermediate-level occupations, whose total and disability-free retirement expectancy is higher than that of female higher-level occupations.

When calculating the same indicators by socio-occupational category, but for both sexes combined, the "double disadvantage" is still observed for blue-collar workers (see Appendix 4, Figure A4-I): their gap relative to higher-level occupations is 3.4 years for RE, and 4.7 years for DFRE. The situation differs for the other categories, since for both sexes combined, the retirement expectancy for white-collar workers and intermediate-level occupations is higher than that for higher-level occupations. This is due to a sex composition effect: women, who live longer on average, are over-represented in these two categories, and under-represented among higher-level occupations. When the social category is measured by educational attainment for both sexes combined, these composition effects are less pronounced and the social gradient for RE is clearly visible, as is the steeper gradient for DFRE (see Appendix 4, Figure A4-II).

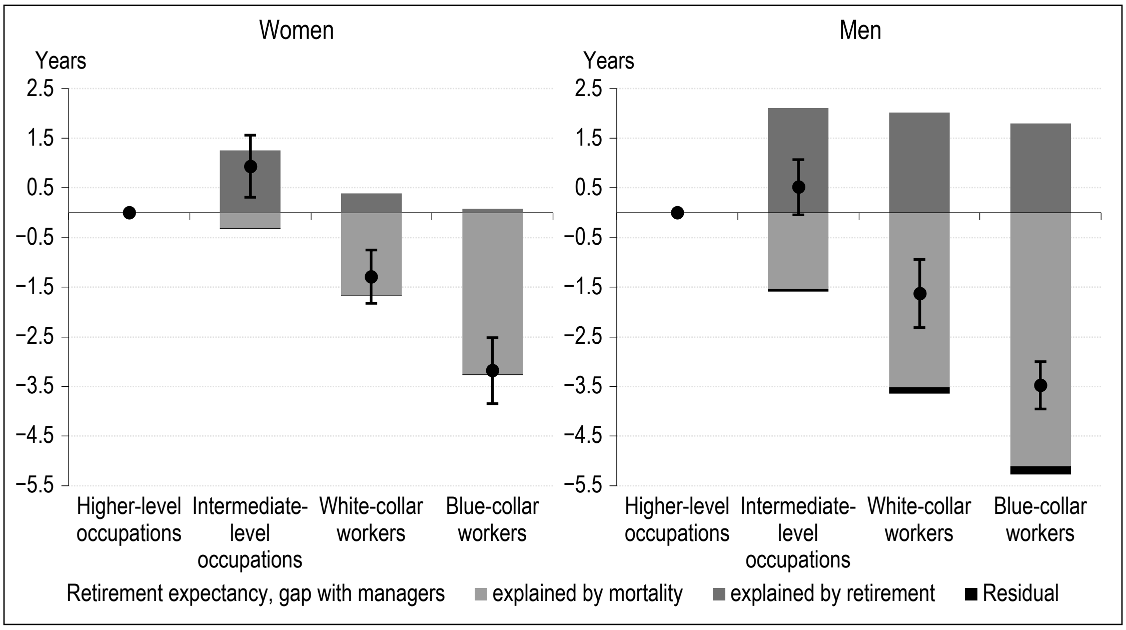
2.2. Decomposition of Differences

The method proposed in Section 1.2 decomposes the gaps between categories into the respective contributions of differences in mortality, disability and retirement. Figures III and IV show that the differences in retirement expectancy between socio-occupational categories and between educational attainment categories are mainly driven by differences in mortality. Differences in retirement age play a partially compensating role, especially for men. For example, if male blue-collar workers retired at the same age as men in higher-level occupations, and differed from them only in mortality, their retirement expectancy gap would increase from 3.5 to 5.1 years.

It is this same compensatory effect of retirement age that explains why, despite higher mortality quotients at each age, both men and women in intermediate-level occupations have a higher retirement expectancy than both men and women in higher-level occupations. For these categories, the earlier retirement age more than compensates for their lower life expectancy.[19]

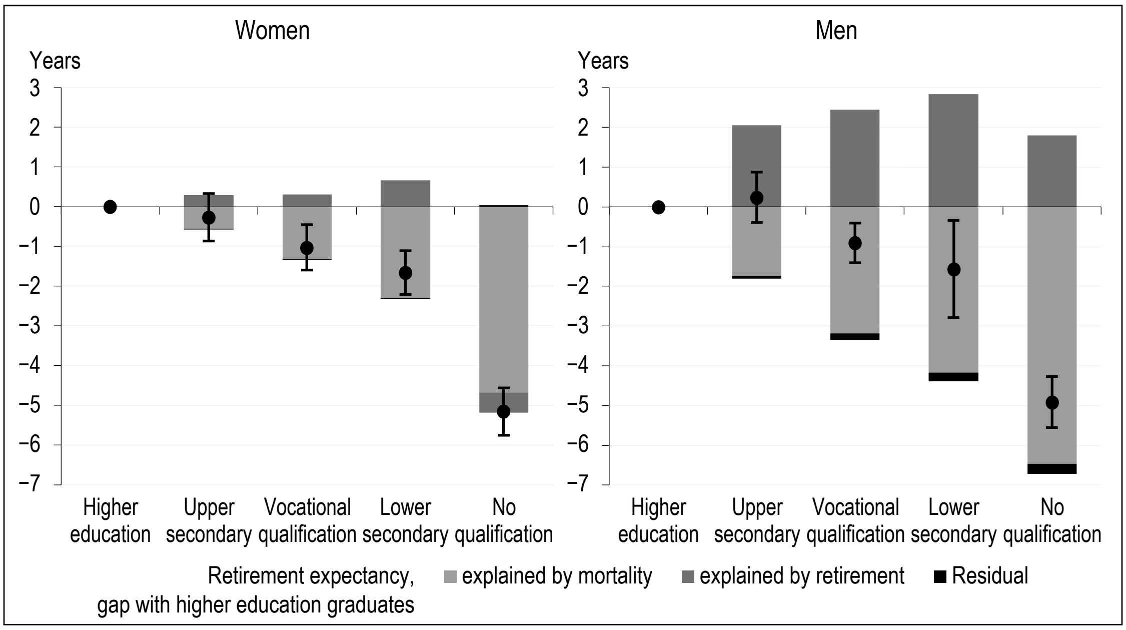
19. Aubert (2024) also shows that, in general, remaining life expectancy at full-rate pension age is higher for those who reach full-rate pension at a younger age. This undermines the idea that the reforms of the last thirty years, which have made it possible to reach the full rate earlier, have only benefited the groups with the lowest life expectancy.

Figure III – **Breakdown of the difference in retirement expectancy between socio-occupational categories**



Source: SRCV, 2017–2019, and Mortality tables by socio-occupational category, 2017–2019, INSEE.

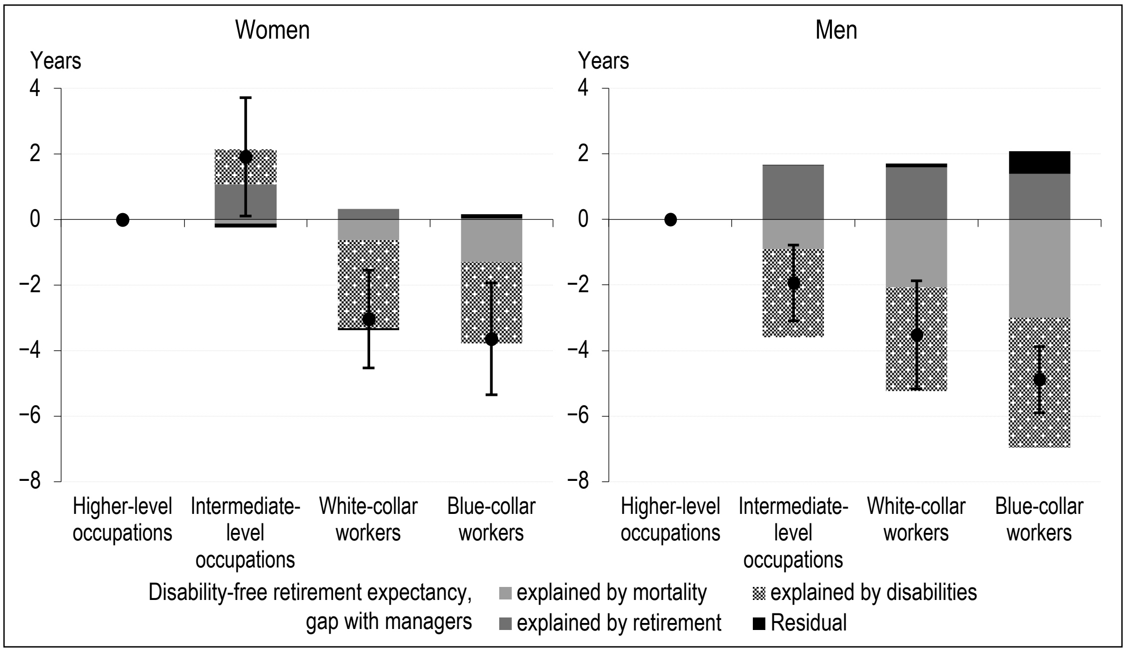
Figure IV – **Breakdown of the difference in retirement expectancy between qualification levels**



Source: SRCV, 2017–2019, and Mortality tables by qualification, 2017–2019, INSEE.

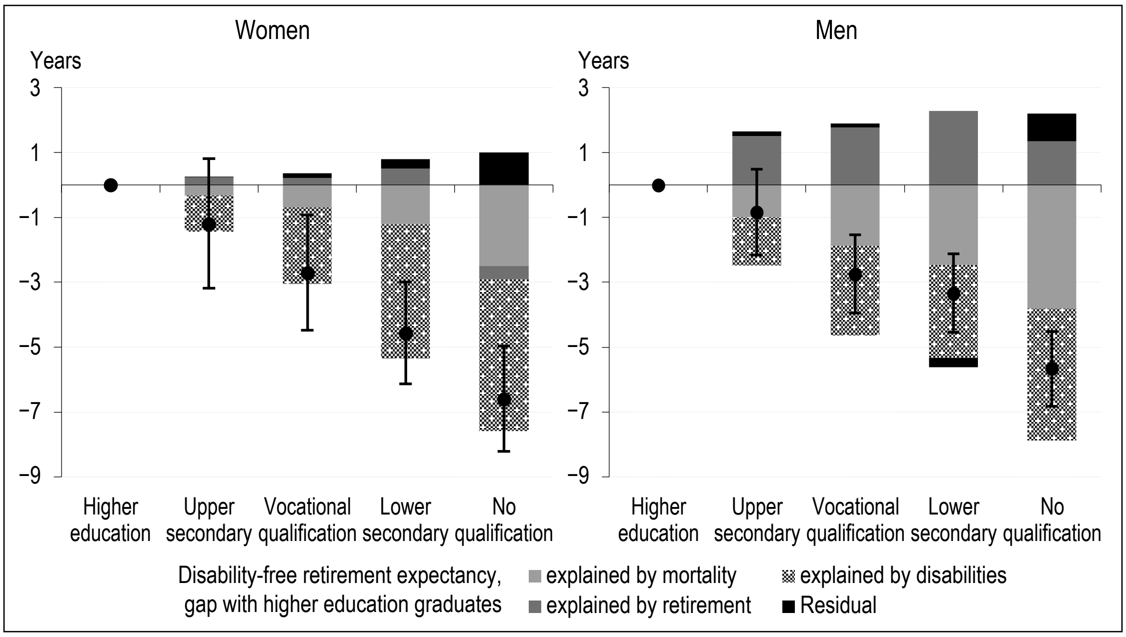
Figures V and VI reproduce the same exercise for disability-free retirement expectancy gaps between socio-occupational categories and between educational attainment categories, respectively. Most of these gaps can be explained by differences in disability, and to a lesser extent by differences in mortality; retirement age once again plays a partially compensating role for men. For example, the disability-free retirement expectancy gap of −3.5 years between male blue-collar workers and men in higher-level occupations breaks down as follows: −3.2 years are explained by differences in disability, −2.1 years are explained by differences in mortality, and a compensation of +1.6 years is explained by differences in retirement age.

Figure V – **Breakdown of the difference in retirement expectancy without disability between socio-occupational categories**



Source: SRCV, 2017–2019, and Mortality tables by socio-occupational category, 2017–2019, INSEE.

Figure VI – **Breakdown of the difference in retirement expectancy without disability between qualification levels**



Source: SRCV, 2017–2019, and Mortality tables by qualification, 2017–2019, INSEE.

For women, the compensatory effect of retirement age between socio-occupational categories is once again much weaker, and the differences are mainly explained by differences in disability. For example, between female white-collar workers and female higher-level occupations, the disability-free retirement expectancy gap is −3.0 years, broken down into −2.7 years due to disability and −0.6 years due to mortality, with a compensating effect of +0.3 years due to retirement age.

Finally, it should be noted that the compensatory role of retirement age also affects the gap between sexes. Women have 4.1 more years of RE and 1.4 more years of DFRE than men. The difference in retirement ages reduces the first gap by 1.0 year, and the second by 0.7 years (see Figure A4-III in the Appendix). This can be explained by the later retirement of women, due to more frequent career interruptions which delay the accumulation of the years of pensionable service required to obtain the full-rate pension (see, for example, Bonnet *et al.*, 2006).

2.3. Aggregate RE and DFRE Trends

How do the indicators studied here change between our two periods: 2009–2013 and 2017–2019? It is well established that during the 2010s, retirement was taken later and later due to later career start dates and the implementation of several pension reforms. In 2010, the Woerth reform introduced a phased shift in the statutory retirement age from 60 to 62 over the 2011–2016 period, and an increase in the age of cancellation of the early-retirement penalty from 65 to 67 from 2016 onwards. The Touraine reform of 2014 scheduled an increase in the contribution period required to obtain the full-rate pension, previously set at 41.5 years, at a rate of one additional quarter every three years. The early-retirement scheme for long careers has been modified several times.[20]

In fact, a decline in retirement expectancy, which was more pronounced for women (−0.4 years) than for men (−0.2 years, not significant), was observed between our two periods. This change in retirement duration had been anticipated (Aubert & Rabaté, 2014, p. 79) and has already been confirmed, notably by the French Pensions Advisory Council (*Conseil d'Orientation des Retraites* – COR), whose annual report tracks a retirement duration indicator defined as the difference between life expectancy and the period-specific retirement age (*Conseil d'Orientation des Retraites*, 2024, p. 123).[21] This is therefore a known result that we confirm using the RE indicator, which takes account of deaths before retirement age and which, for this reason (see Online Appendix S1), does not perfectly coincide with retirement duration indicators such as those used by the COR.

The decomposition method described in Section 1.2 can be used to quantify the two forces at work over this period: the negative contribution of later retirements is only partly offset by the positive contribution (greater for men) of the reduction in mortality. Thus, if the retirement rate at each age had remained constant over time, the fall in mortality would have generated an increase of 1.0 year in the RE of men over the period, instead of the observed decline (Figure VII).

Disability-free retirement expectancy (Figure VIII) increases for women (+0.6 years) and decreases for men (−0.3 years, not significant). For both sexes, retirement age makes a negative contribution; the positive net change for women is explained by a favourable change in survival at disability-free retirement ages.

2.4. Widening Gaps by Educational Attainment Level

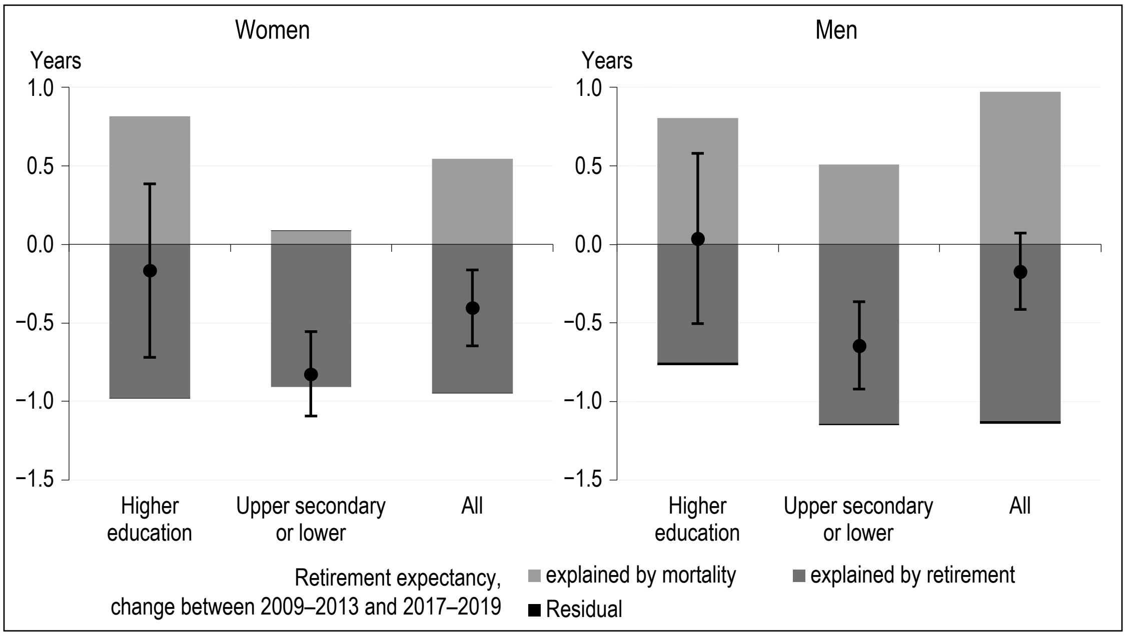
Figures VII and VIII show that the trend in RE and DFRE is not uniform by level of educational attainment: higher education graduates benefit from more favourable trends in RE and DFRE than the average, while the other qualification categories, grouped in the "*Baccalauréat* (upper secondary) or less" category, experience less favourable trends. In particular, the decline in RE in this category is at least twice the average, both for women (−0.8 years) and men (−0.6 years).

The gap between higher education graduates and those in the "*Baccalauréat* or less" category increased for RE between the two periods, by 0.7 years for both men and women. For DFRE, the increase in the gap between these educational attainment categories – of 1.2 years for men and 1.3 years for women – is not significant.

In short, as Figure IX shows, between 2009–2013 and 2017–2019, for both men and women, the gradient by educational attainment level widened, in terms of total and disability-free retirement expectancy, but also in terms of total and disability-free life expectancy.

20. The conditions were first tightened in 2009, and then extended in 2012, thereby mitigating the effect of the increase in the statutory retirement age.
21. These studies also show that this decline can be seen as a reversal of the gains in retirement duration achieved during the previous decade. Since the scope of our study begins on the eve of the 2010 pension reform, we cannot observe this phenomenon here.

Figure VII – **Change in retirement expectancy between 2009–2013 and 2017–2019 by qualification and its breakdown**



Source: SRCV, 2009–2013 and 2017–2019, and Mortality tables by qualification, 2009–2013 and 2017–2019, INSEE.

Figure VIII – **Change in disability-free retirement expectancy between 2009–2013 and 2017–2019 by qualification and its breakdown**

Source: SRCV, 2009–2013 and 2017–2019, and Mortality tables by qualification, 2009–2013 and 2017–2019, INSEE.

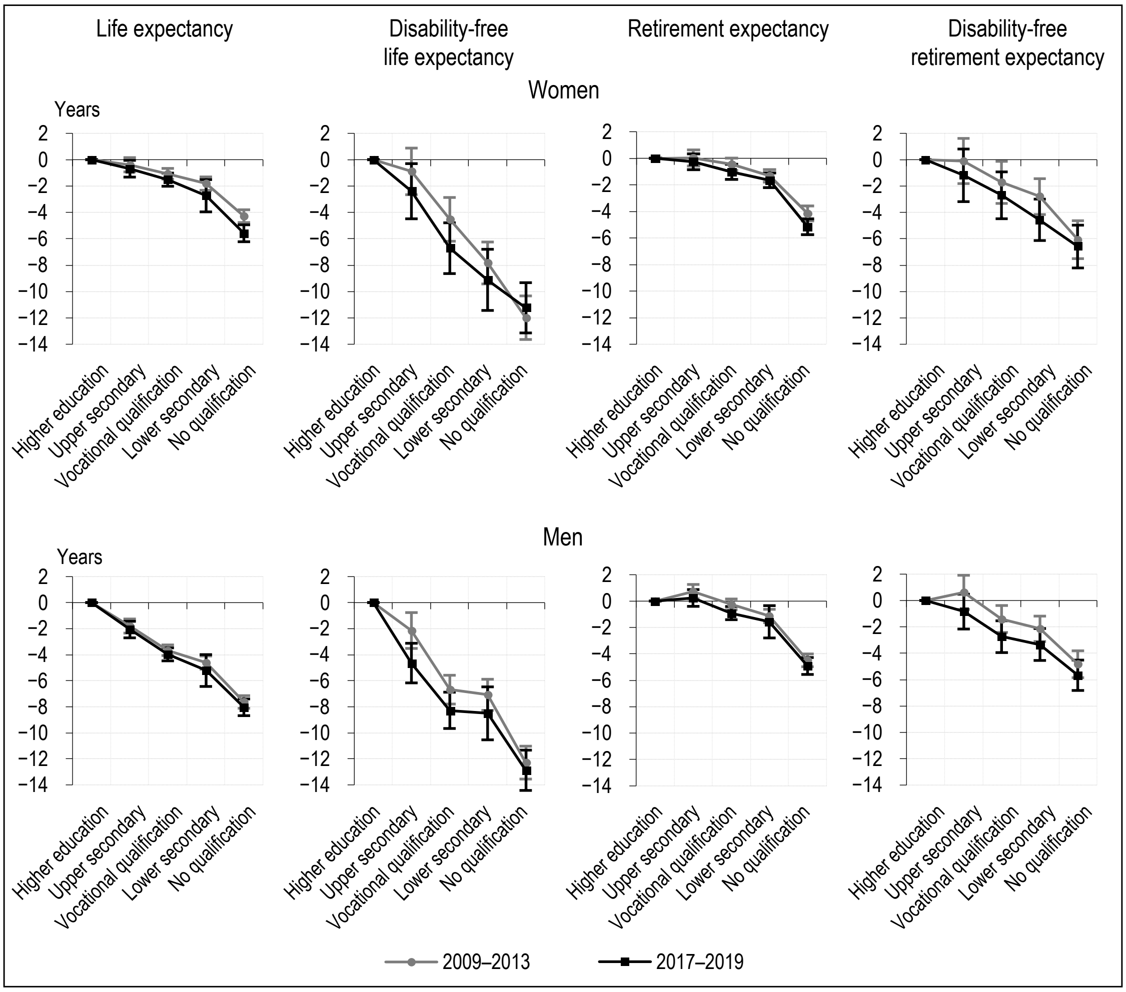
Changes by socio-occupational category leave less room for clear-cut conclusions. The results are presented in Appendix 4 (see Figures A4-V and A4-VI): the gradient in retirement expectancy by socio-occupational category more or less stable for both sexes; for disability-free retirement expectancy, it decreases for women but increases for men.

3. Discussion and Limitations

3.1. Disabilities in Non-Ordinary Housing

This analysis is based on data from surveys traditionally conducted in ordinary housing, i.e. excluding institutions, hostels and communities. In 2021, 5% of people aged 65 or over lived in

Figure IX – **Difference with higher education graduates in terms of life expectancy and retirement expectancy, total and without disability, in 2009–2013 and 2017–2019**



Source: SRCV, 2009–2013 and 2017–2019, and Mortality tables by qualification, 2009–2013 and 2017–2019, INSEE.

non-ordinary housing,[22] 97% of them in nursing homes for dependent elderly people or in retirement homes (Daguet, 2025). Yet in 2019, 85% of people living in nursing homes for dependent elderly people had lost autonomy within the meaning of the Gerontological Autonomy, Iso-Resource Groups (AGGIR – *Autonomie Gérontologique Groupes Iso-Ressources*) scale (Balavoine, 2022). Although this scale does not coincide precisely with the GALI indicator, it can be estimated that a very large majority of people in residential care for the elderly, and therefore of seniors in non-ordinary housing, are living with disabilities.

The use of survey data for ordinary housing therefore tends to underestimate the proportion of people with disabilities at older ages. Given the social differences in residence in nursing homes for dependent elderly people (former blue-collar workers are known to be more numerous, Roy (2023)), this may affect the measurement of social inequalities in disability-free retirement expectancy.

To assess the impact of this phenomenon on our results, we used census data to measure the rate of residence in non-ordinary housing by narrow age band, sex and educational attainment level[23] (Figure X). This enabled us to adjust our distribution of disability rates by age and category, making the most unfavourable assumption that all people in non-ordinary housing live with disabilities after age 70.

The impact on DFRE estimates varies from −0.2 years to −1.1 years depending on sex and educational attainment category. It is stronger for women, who reach more advanced ages and

22. This proportion rises rapidly at the oldest ages: it is 19% for those aged 85 or over, and 42% for those aged 95 or over.

23. The last profession declared is not available in our data. That is why we were unable to conduct this analysis by profession and socio-occupational category (PSOC), because most people are classified in the "Retired"category after a certain age. For the proportion of people living in non-ordinary housing by social group, see Daguet (2025).

spend more of their time in institutions.[24] For the same reasons, it is also stronger for higher education graduates, but the gap with the other educational attainment categories is smaller than that between the sexes.

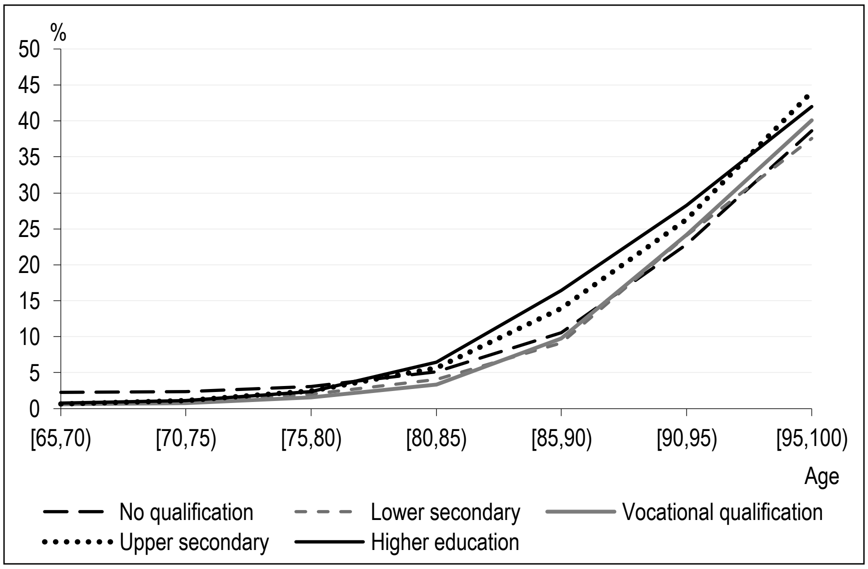
Consequently, the main results are not greatly altered by this variation: the differences in DFRE between male higher education graduates and men with no qualifications fall from 5.7 years to 5.5 years, and from 6.6 years to 5.9 years for women (Figure XI).

3.2. Inequalities by Socio-Occupational Category or by Educational Attainment Level

In this article, we present inequalities in disability-free retirement expectancy by socio-occupational category and highest qualification obtained. Although the analyses for these two dimensions are similar, they are not equivalent.

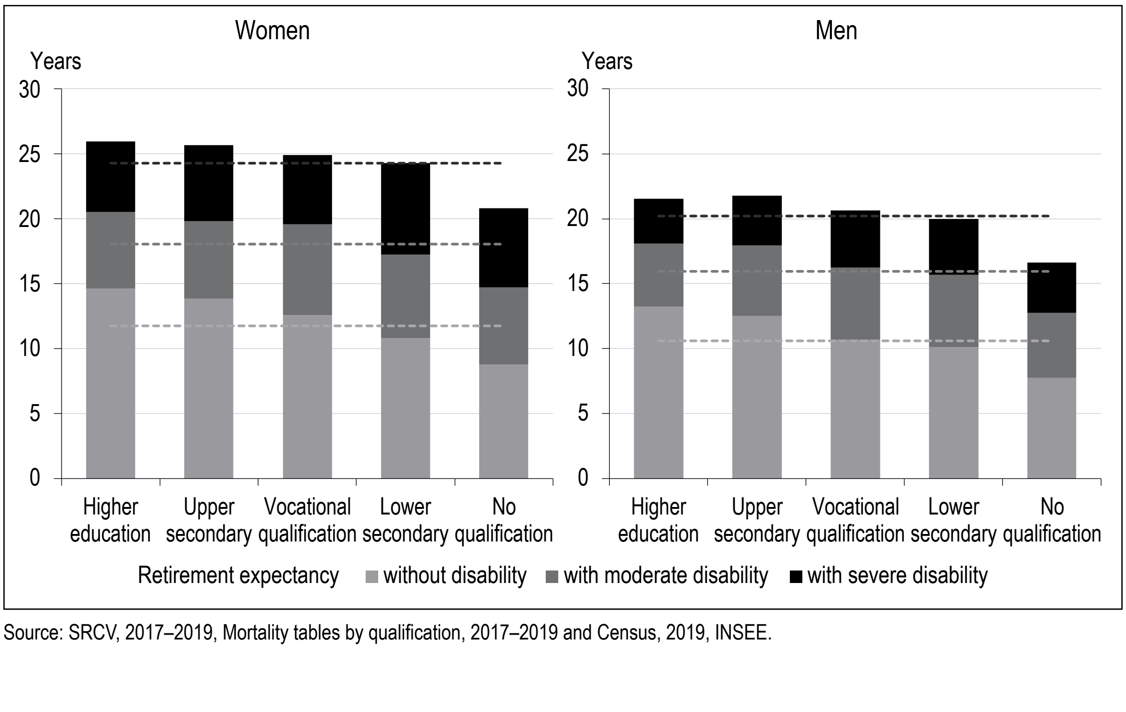
24. In fact, women are largely predominant in residential establishments for the elderly (72.8% in 2019, Balavoine, 2022).

Figure X – **Proportion of people living outside ordinary housing by age for different levels of qualification**



Source: Census, 2019, INSEE.

Figure XI – **Retirement expectancy, with and without disability, by qualification and gender, including people living in non-ordinary housing and those with disabilities**

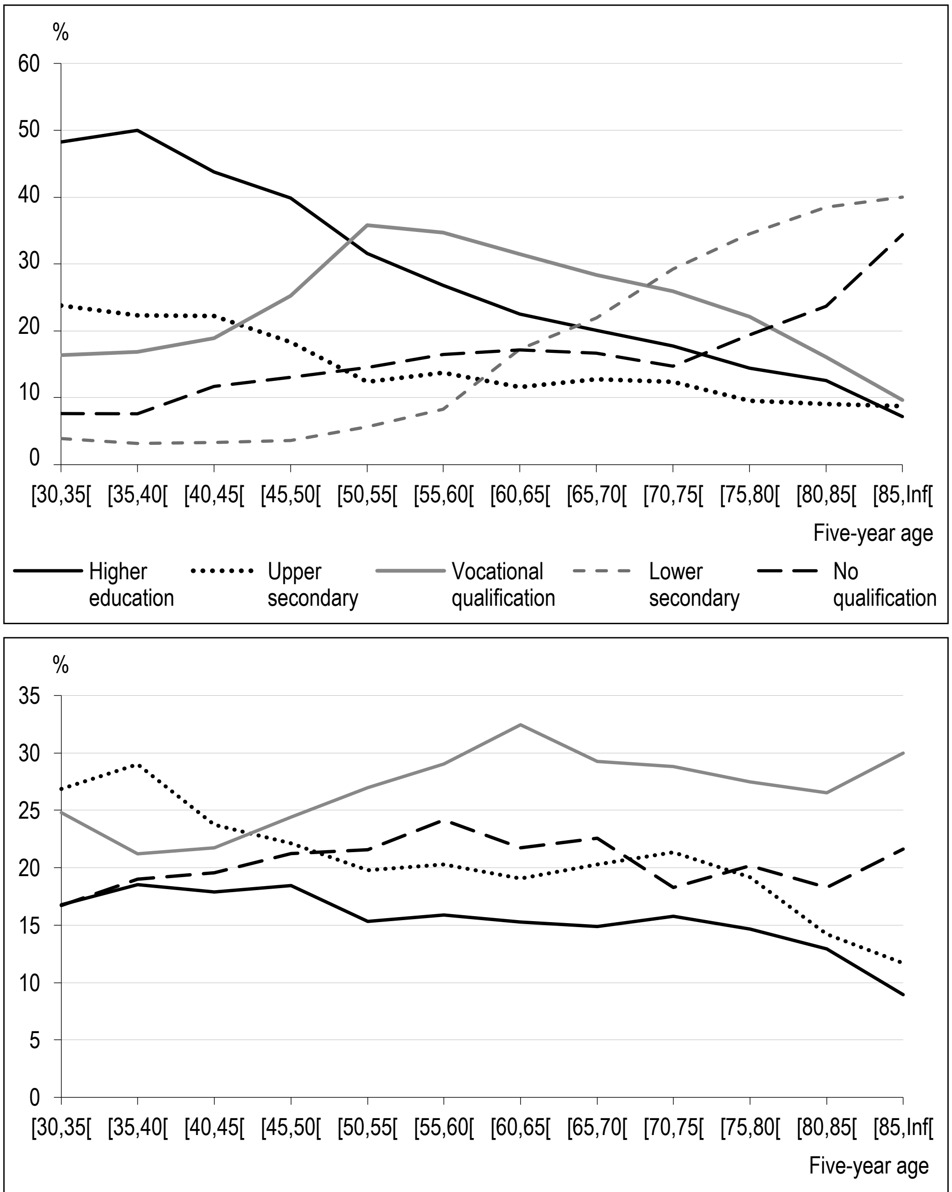


Source: SRCV, 2017–2019, Mortality tables by qualification, 2017–2019 and Census, 2019, INSEE.

Educational attainment and socio-occupational category are both relatively correlated with a person's socio-economic situation and are therefore correlated with life expectancy and disability. However, they do not necessarily have the same impact on people's health: Geyer *et al.* (2006) show, for example, that educational attainment is more closely correlated with certain pathologies than socio-occupational category, and conversely less so for other pathologies.

Analysis by socio-occupational category enables the extension of the results already established in France for disability-free life expectancy (Cambois *et al.*, 2020) and retirement expectancy (Bonnet *et al.*, 2023). In the present case, however, it has the disadvantage of being restricted to the four categories of higher-level occupations, intermediate-level occupations, white-collar workers and blue-collar workers, for reasons of sample size (there are too few farmers and tradespeople/shopkeepers/company managers for analysis). Furthermore, socio-occupational category is less relevant for distinguishing women in the oldest cohorts, a significant proportion of whom have not engaged in professional activity (this applies

Figure XII – **Distribution of the population in metropolitan France by age, socio-occupational category and qualification**



Source: SRCV 2019, INSEE.

to 10% of women aged 65 or over and 15% of women aged 80 or over in 2019).

The advantage of educational attainment is that qualifications can be obtained at an early age and concern the entire population. Furthermore, this dimension is more easily comparable between countries: it was, for example, the variable used by Mäki *et al.* (2013) to compare social inequalities in disability-free life expectancy in eight European Union countries. By contrast, socio-occupational category requires non-trivial reprocessing before it can be converted into internationally comparable categories. In a study of mortality rates according to socio-economic category, Mackenbach *et al.* (2008) were able to compare 19 countries on the basis of educational attainment, but only nine on the basis of socio-occupational category.

At the individual level, educational attainment is a more stable measure over time than socio-occupational category, which may change over the course of a career. However, the general rise in the level of educational attainment in the population means that the proportion of higher education graduates is substantially lower at the oldest ages (for the oldest cohorts): they therefore represent a much more specific category (Figure XII). Similarly, the proportion of people with no more than the *brevet des collèges* (lower-secondary qualification) is much higher for these ages. This phenomenon also applies to socio-occupational categories: higher-level occupations and intermediate-level occupations are much less common in older cohorts.

Among those aged 60 or over, who are of interest to us in calculating our indicators, blue-collar workers represent a share of the population close to that of people with no qualifications (21% and 19% respectively). However, the age distribution is not the same: the proportion of people with no qualifications is much higher over the age of 80, while the proportion of blue-collar workers is relatively stable. This problem also arises, to a lesser extent, for socio-occupational categories, because the proportion of higher-level and intermediate-level occupations is lower at older ages: they therefore represent a more specific population than for the ages at the start of retirement. It is therefore more difficult to interpret retirement expectancy in these situations, where the category evolves across generations.

* *
*

By combining INSEE mortality tables with questions on retirement and disability in the SRCV (Statistics on Resources and Living Conditions) survey, we calculated expected retirement expectancy with and without disability in France over several years, breaking it down by gender, socio-occupational category and educational attainment. We highlight a "double disadvantage" for members of lower socio-occupational categories relative to higher categories, whether in terms of educational attainment or socio-occupational category: they have a lower retirement expectancy and can expect to spend a greater proportion of their lives with disabilities.

Differences in retirement expectancy between socio-occupational categories or between educational attainment levels are mainly explained by differences in mortality, while differences in disability-free retirement expectancy are mainly explained by differences in terms of disability. Among men, members of higher socio-occupational categories retire later, which helps to reduce but not eliminate these gradients.

The earlier retirement of blue-collar workers gives them more disability-free retirement time than higher-level occupations, but this is more than offset by their higher mortality and greater exposure to disability. It can also be seen that women, who benefit from significantly longer retirement expectancy than men thanks to their greater life expectancy, spend a larger proportion of their retirement with disability. Women therefore have a much smaller advantage over men in terms of disability-free retirement expectancy than in terms of retirement expectancy overall.

Between 2009–2013 and 2017–2019, we show that later retirement reduced the retirement expectancy of both men and women despite lower mortality, but that improvements in disability reduction led to an increase in disability-free retirement expectancy. However, these aggregate trends conceal disparities, particularly according to educational attainment. For both men and women, the gradient by educational attainment for life expectancy and retirement duration, both total and disability-free, became steeper between 2009–2013 and 2017–2019.

Analyses based on the retirement expectancy indicator, its evolution and its distribution across different population categories can therefore be enriched by taking account of the years lived with disability. At the conclusion of this research, it appears appropriate to propose the incorporation of the total and disability-free retirement expectancy indicators into the toolkit of indicators available to the public institutions responsible for monitoring the retirement system in France, in order to monitor their development and distribution and help inform debates on the inter- and intra-generational equity of the retirement system. ☐

Link to Online Appendix:
www.insee.fr/en/statistiques/fichier/9008738/ES549_Blasco-Lojkine_Online-Appendix.pdf

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APPENDIX 1

IMPUTING PREVALENCES AT ADVANCED AGES: AVERAGE BY AGE CATEGORY OR LINEAR SMOOTHING?

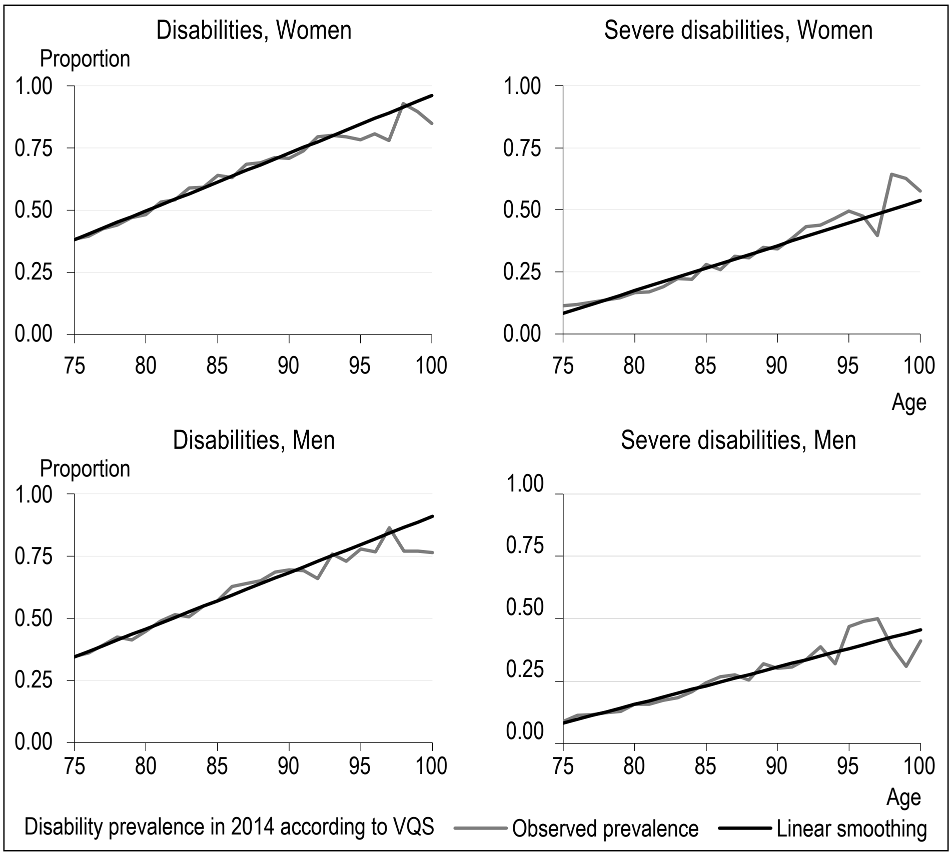
As explained in Section 2.2, and in accordance with the DREES (Ministerial Statistical Office for the Ministries of Health and Social Affairs) method (Deroyon, 2024) for calculating disability-free life expectancy, we imputed retirement status crossed with disabilities by five-year age categories from age 30 to 84, and then within a broad category of ages 85 and over.

This imputation method is debatable, particularly for prevalence rates of disability. Imputation by age category may therefore appear to be a strong assumption, whose relevance we verify in this appendix by carrying out a robustness test, consisting in estimating the prevalences of retirement status cross-classified with disability using a five-year moving average up to age 74 (which differs little from imputation by five-year categories), and then by linear interpolation between the ages of 75 and 100.

The realism of this alternative imputation method is suggested by the VQS (Daily Life and Health) survey, a large-scale survey conducted by DREES, which does not include information on occupation and qualifications, but does enable the estimation of disability prevalences by age and sex up to advanced ages. Based on the 2014 wave of this survey, which includes 121,000 observations of individuals aged between 75 and 100, Figure A1-I shows the closeness of the raw averages by detailed age to a linear model. We cannot verify whether this linear relationship between age and disability prevalence applies within each social category, but Figure A1-II suggests that it also applied within the largest of our categories, female white-collar workers, in 2009–2013.

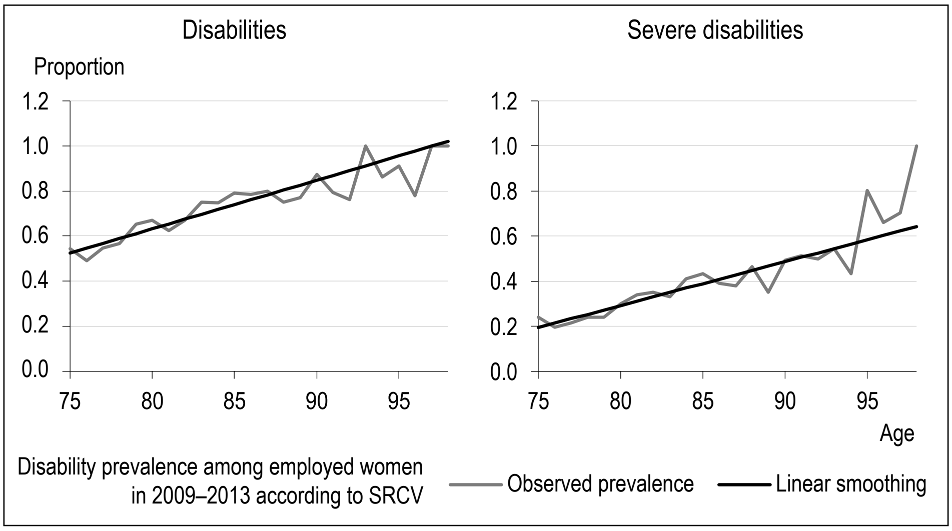
A comparison of the two imputation methods shows that the method has little effect on our results. Taking the example of DFRE by occupation for men in 2017–2019, the largest difference between the two methods proves to be for higher-level occupations, with DFRE that is 0.2 years lower when estimated using the linear smoothing method. Changes between periods, aggregated by sex or broken down by social category, are likewise not significantly affected.

Figure A1-I – **Comparison between the prevalence of disability estimated directly as the average prevalence by detailed age and by a linear model for all 75-100 year olds**



Source: Survey *Vie quotidienne et santé* 2014, DREES.

Figure A1-II – **Comparison between the prevalence of disability estimated directly as the average prevalence by detailed age and by a linear model for employees aged 75 to 100**



Source: SRCV, 2019–2013, INSEE.

APPENDIX 2

WHICH SOURCE: SRCV OR THE LABOUR FORCE SURVEY?

In this paper, we use the *Statistiques sur les ressources et les conditions de vie* (SRVC – Statistics on Household Resources and Living Conditions) survey as the basis for imputing the cross-classified prevalences of retirement and disability at different ages by socio-occupational category, educational attainment, sex and period. The Labour Force Survey (*Enquête Emploi*) is another survey containing the required variables. Our decision to use the SRCV survey was motivated by two reasons. First, this survey is used by DREES for the annual publication of the aggregate disability-free life expectancy indicator, and by Eurostat for the publication of the same indicator, harmonised with the other European Union countries (the SRCV survey is part of the EU-SILC system). Second, the question on limitations in daily life has been asked in the same manner without a break in series since 2008, enabling the production of our various indicators for the 2009–2013 and 2017–2019 periods, for which mortality tables according to social category by Blainpain (2024b) are available, whereas the question has only been asked since 2013 in the Labour Force Survey.

For the 2017–2019 period, however, it is expedient to compare our measurements with the same indicators calculated from the Labour Force Survey, which benefits from a larger sample and whose measurement of the retirement age is closer to the administrative reference sources. The change of survey affects both retirement prevalences and disability prevalences, and therefore both the retirement expectancy (RE) and disability-free retirement expectancy (DFRE), in different ways. Thus, by changing from the SRCV to the Labour-Force Survey, RE decreases (from 20.2 to 18.9 years for men and from 24.3 to 23.4 years for women) and DFRE increases (from 10.8 to 11.5 years for men and from 12.2 to 13 years for women). This means that at a given age, respondents to the Labour Force Survey report being disability-free more often and retired less often.[25]

The change of source also affects the gaps between categories, only slightly among men but more markedly among women. For example, when moving from the SRCV to the Employment survey, the RE gap between a female blue-collar worker and a woman in a higher-level occupation shrinks from 3.3 to 2.6 years, whereas the DFRE gap increases from 3.6 to 5.0 years.

However, these variations do not affect our qualitative results for the 2017–2019 period. On the contrary: the “double disadvantage” phenomenon is still observed using the Labour Force Survey, and becomes even more pronounced among women.

In the future, upcoming official statistical surveys will provide an even more accurate measurement of inequalities in disability by socio-occupational category, notably through the inclusion of the GALI question in the 2025 annual Census survey onwards and the conduct of the Health and Territories survey in 2025 and 2026.

25. The first point had previously been made by Aubert (2020, p. 5).

APPENDIX 3

MEASURING RETIREMENT STATUS

As explained in Section 1.1, we measure whether or not a person is retired on the basis of the response to the SRCV survey question on main employment status, corrected by considering anyone aged 70 or over as retired. In this appendix, as a robustness test, we explore two other conventions: defining retired status by the fact of receiving a retirement pension, and defining it on the basis of main labour-market status, but without the requirement of considering anyone aged 70 or over as retired.

A3.1. Measured Based on Receipt of a Retirement Pension

In this first robustness test, we define retired status on the basis of the RETRAITES_I variable in the SRCV survey, which provides the amount of retirement and pension income received in the previous year, obtained by matching the survey with tax data. The fact of this being the previous year's income introduces a time lag relative to the respondent's age, albeit of less than one year, since the survey is conducted annually between January and April. In this test, as in our main exercise, we consider anyone aged 70 or over as retired.

This measure leads to higher estimated retirement expectancies (0.8 years for men and 1.3 years for women). This may be due to situations in which employment and retirement are combined, where a pension is received by someone who is not retired in terms of their main labour market situation. This gap is unevenly distributed across social categories and therefore affects the measurement of the RE and DFRE gradient. For example, the RE gap between a blue-collar worker and a member of a higher-level socio-occupational category narrows from 3.5 to 3.0 years for men and from 3.0 to 1.8 years for women.

However, our qualitative conclusions are robust for this specification. In particular, the "double disadvantage" effect is still observed; the trends between RE and DFRE periods are similar, and gaps between educational attainment categories also widen.

A3.2. Measured Based on Uncorrected Labour Market Status

The second robustness test consists in of using the response to the question on main labour market status without correcting for people aged 70 or over.

This variation does not significantly affect our results for men, whether aggregated by socio-occupational category or by educational attainment, in 2017–2019 as in 2009–2013, indicating that throughout these two periods, few men aged 70 had an administrative status other than retired.

On the contrary, this measure reduces the RE of women (23.6 instead of 24.3 years), due to the significant proportion of women aged 70 or over who define themselves as housewives. This gap varies according to socio-professional category, altering the RE gap between categories: a difference of 2.8 years instead of 3.2 years between female blue-collar workers and women in higher-level occupations. However, the "double disadvantage" remains for members of lower socio-occupational categories.

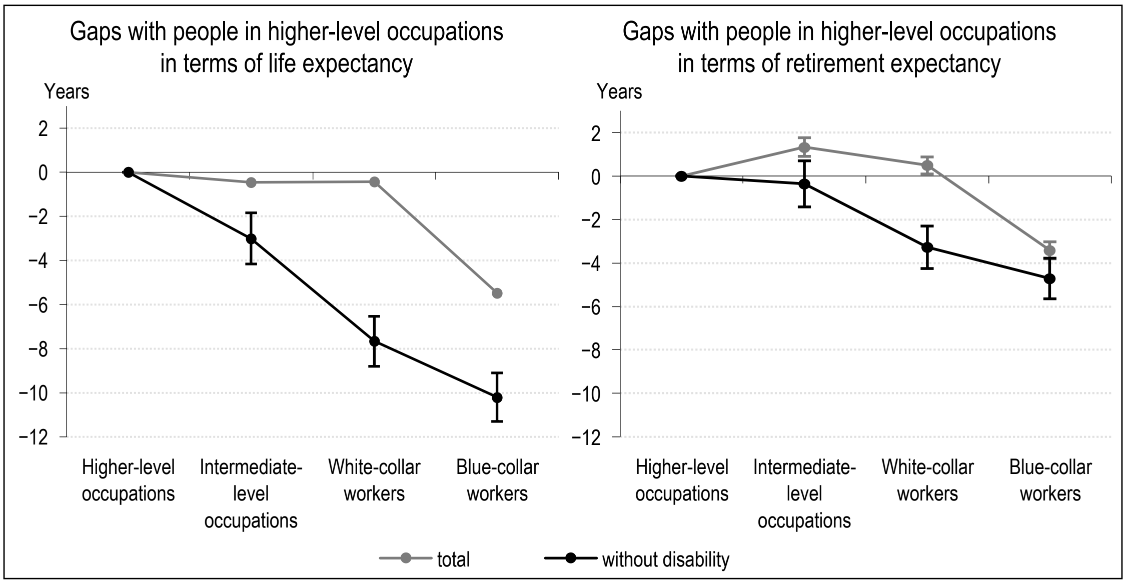
More significantly, this variant affects the overall pattern of changes. Whereas according to our main measure, RE decreased between the two periods for both sexes (–0.4 years for women), according to the raw measure, it increased for women (+0.2 years, not significant). This significant contrast can be explained by a change: the proportion of women aged between 70 and 79 not reporting retired status was 8% in 2009–2013, and 3% in 2017–2019. The vast majority of these women reported being housewives. Therefore, during the same period when women were retiring later and later, they were increasingly likely to report their status as retired after age 70 – certainly because cohorts reaching those ages were women who had accumulated more years of pensionable service on the labour market.

The assumption adopted in the main text of this paper is to consider housewives aged 70 or over as retired, as they seem to be closer to this status than to working. This enables the opposing roles of changes in retirement ages and mortality to be highlighted. However, it would not necessarily be less legitimate to drop this assumption, which would introduce a third force, the extension of pension-system coverage: the stabilisation or even the extension of retirement expectancy, which is an average, would then be likely to result from the combination of lower mortality for most women, later retirement for those covered by the pension system, and a reduction in the proportion of a cohort not benefiting directly from the pension system.

APPENDIX 4

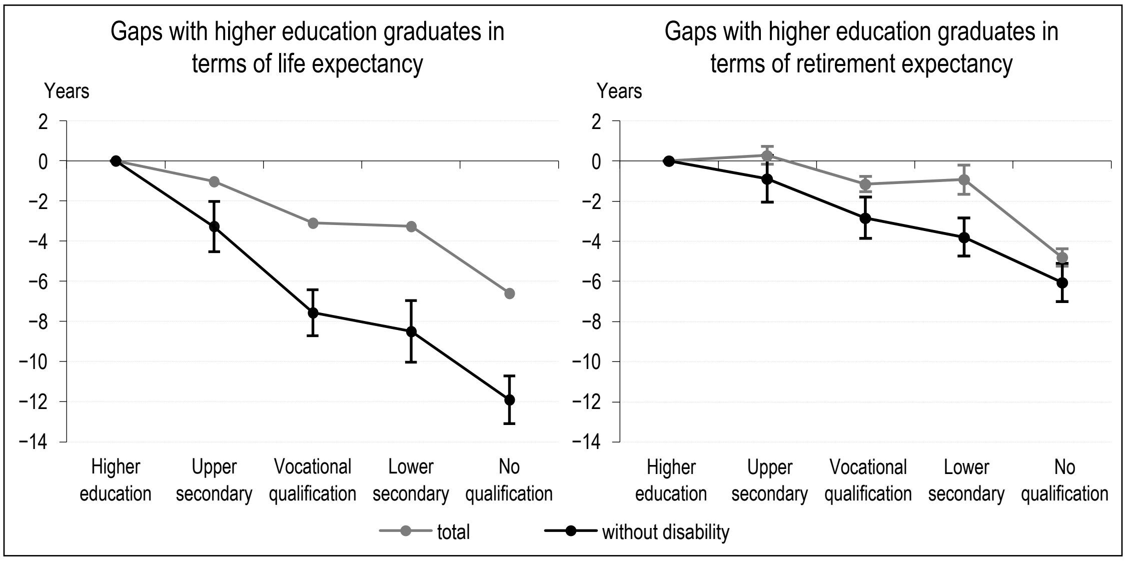
ADDITIONAL FIGURES

Figure A4-I – **Differences in life expectancy and retirement expectancy, total and without disability, between socio-occupational categories**



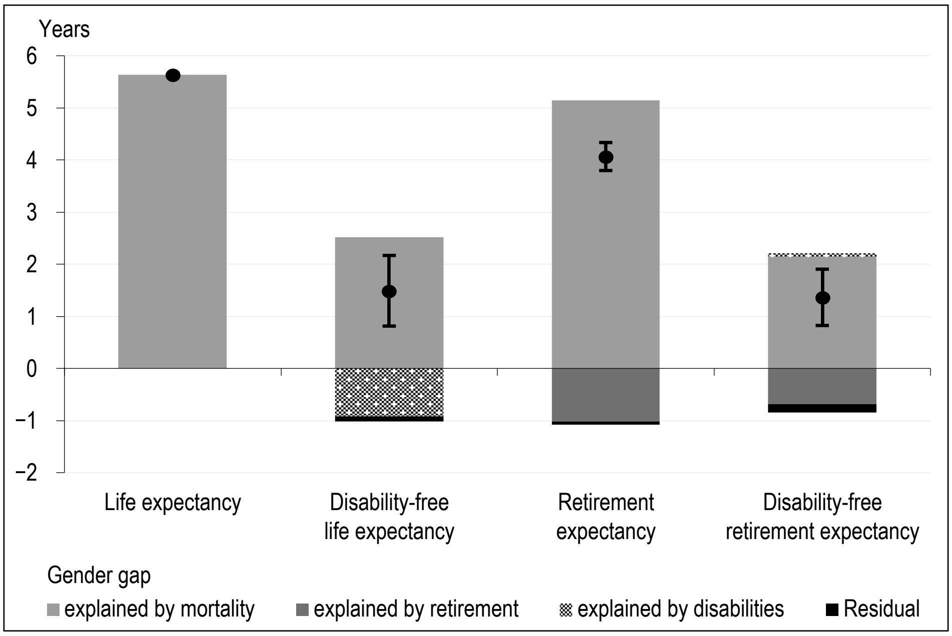
Source: SRCV, 2017–2019 and Mortality tables by socio-occupational category, 2017–2019, INSEE.

Figure A4-II – **Differences in life expectancy and retirement expectancy, total and without disability, between qualification levels**



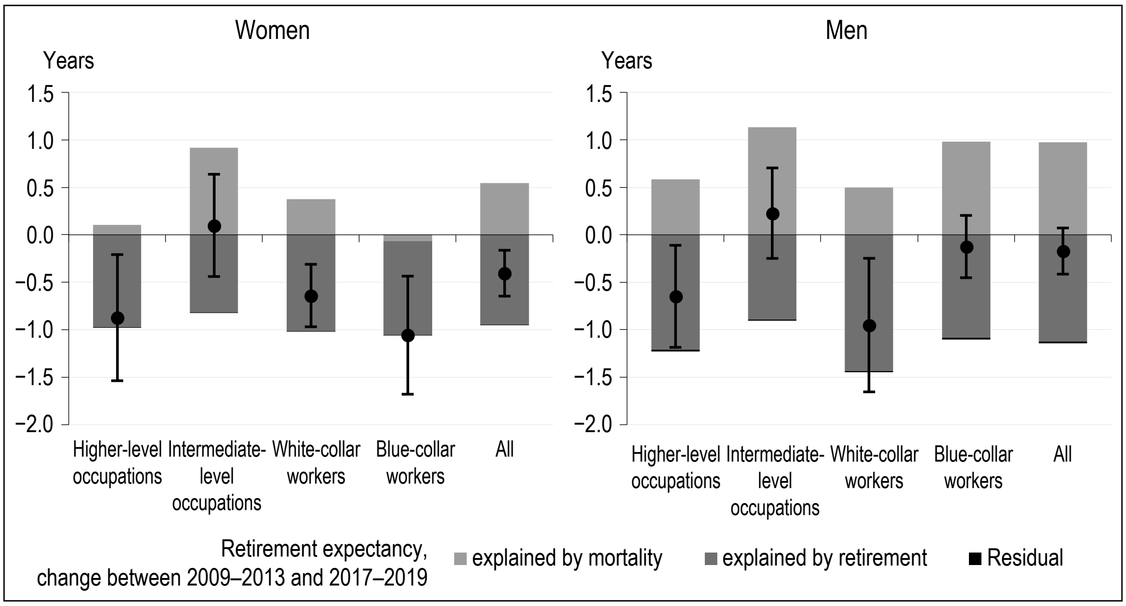
Source : SRCV, 2017–2019 and Mortality tables by qualification, 2017–2019, INSEE.

Figure A4-III – **Gender gap in life expectancy and retirement expectancy, with and without disability, and its breakdown**



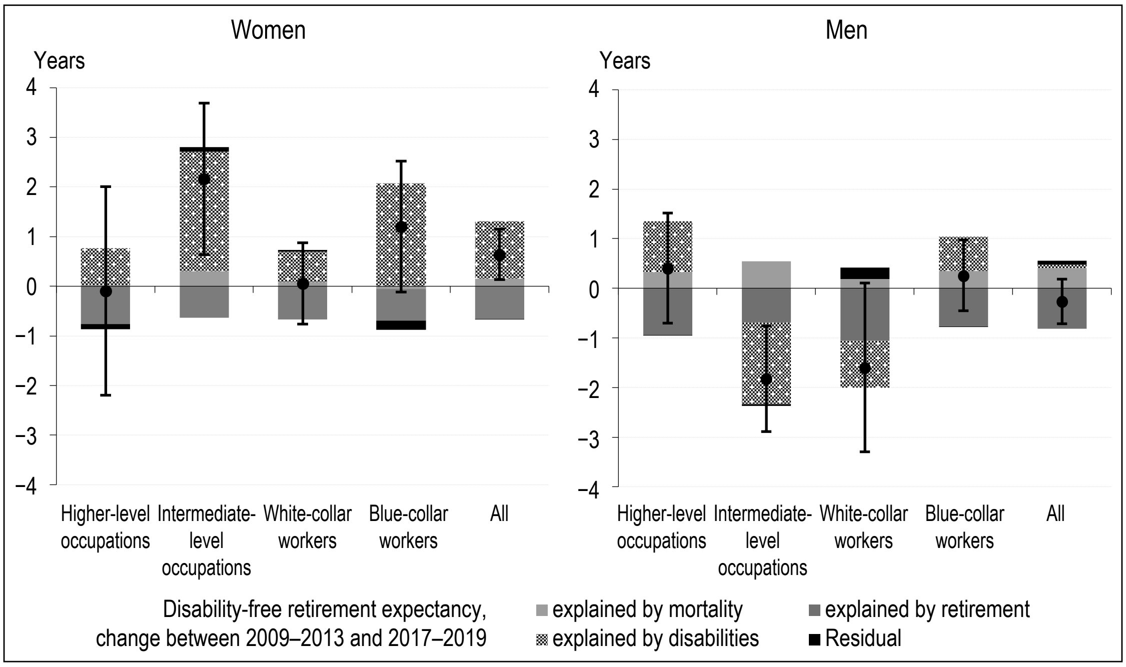
Source: SRCV, 2017–2019 and Mortality tables by qualification, 2017–2019, INSEE.

Figure A4-IV – **For each socio-occupational category, change in retirement expectancy from 2009–2013 to 2017–2019 and breakdown of this change**



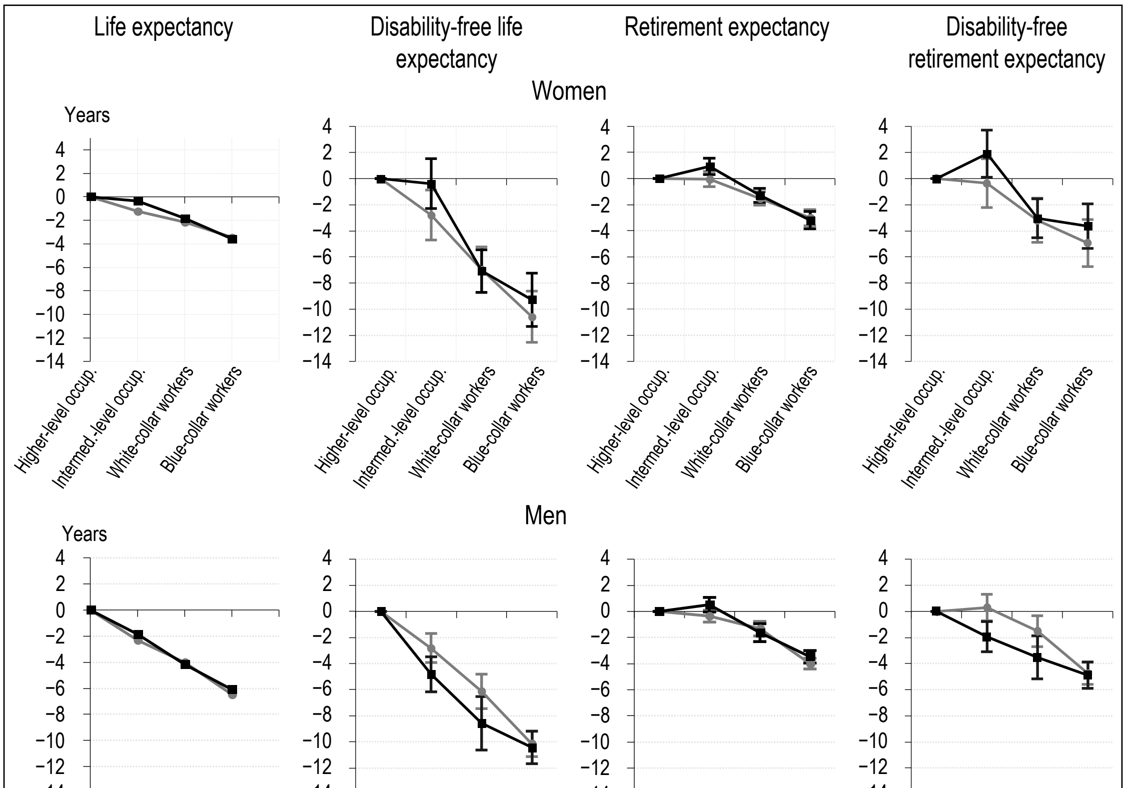
Source: SRCV, 2009–2013 and 2017–2019, and Mortality Tables by socio-occupational category, 2009–2013 and 2017–2019, INSEE.

Figure A4-V – **Change in disability-free retirement expectancy between 2009–2013 and 2017–2019 and its breakdown**



Source: SRCV, 2009–2013 and 2017–2019, and Mortality tables by qualification, 2009–2013 and 2017–2019, INSEE.

Figure A4-VI – **Differences in life expectancy and total disability-free retirement expectancy between socio-occupational categories, in 2009–2013 and 2017–2019**



Source: SRCV, 2009–2013 and 2017–2019, and Mortality tables by socio-occupational category, 2009–2013 and 2017–2019, INSEE.

Three Approaches to Labour Market Discrimination Against Workers of North African Origin

Émilie Arnoult*

Abstract – Whether assessed through correspondence tests, self reported experiences or differences in unemployment rates among individuals with identical characteristics, discrimination in the labour market against individuals of North African origin is substantial and is observed across all levels of education. This article draws on the results of a correspondence study conducted by the IPP and ISM-Corum in 2019, data from the Trajectories and Origins 2 survey (*Trajectoires et Origines* 2), and the 2019 and 2020 French Labour Force Surveys to propose complementary approaches in addressing discrimination. It shows that, whilst the extent of the differences observed between individuals of North African origin and those with no migrant background may vary depending on the approach, these differences can be interpreted in light of the specific characteristics of each approach: the assumptions on which they are based, their scope, or their analytical framework. Comparing these approaches thus provides a better understanding of the various mechanisms through which discrimination operates in the labour market and helps to grasp the multiple forms it can take.

JEL: J7, J15, J64, C52, C93
Keywords: discrimination, unemployment, gender, ethnic origin

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In 2019–2020, nearly one in five people reported having personally experienced discrimination on grounds prohibited by law (Lê *et al.*, 2022). Discrimination is defined as any distinction, exclusion or restriction that undermines equal treatment between individuals. French law specifies the prohibited criteria (such as ethnic origin, gender, religion, disability and age, amongst others) as well as the areas (employment, housing, access to services, education, etc.) in which such practices are banned and punishable. They constitute an offence that is severely punished under the Criminal Code, as they cause harm to those who suffer them and have negative effects on society. For several decades, the social sciences have been addressing this issue and working to understand the mechanisms underlying the existence of such a phenomenon. The literature shows that, although discrimination is a criminal offence and various tools are in place to combat it, the problem persists.

Among individuals who reported having experienced discrimination, the workplace is the setting most frequently cited, both by respondents to the Trajectories and Origins (*Trajectoires et Origines* TeO) surveys (INSEE, 2023) and by those surveyed in the Security Experiences and Perceptions survey (*Vécu et ressenti en matière de sécurité*) (SSMSI, 2022). Ethnic origin is, moreover, a factor frequently cited by victims of discrimination. Across all sectors, 41% of victims report that the discrimination they experienced was at least partly based on their ethnic origin, whether real or perceived (SSMSI, 2022). When the effect of other characteristics is offset, ethnic origin is the main factor explaining the likelihood of reporting having experienced discrimination, both in 2008–2009 and in 2019–2020 (Lê *et al.*, 2022). In employment, whilst ethnic origin is not the criterion most frequently cited by victims themselves, it is cited by those who report having witnessed such discrimination (Défenseur des droits, 2023).

Discrimination can have multiple consequences for those who experience it. On the one hand, it can undermine their confidence, both in themselves and in institutions. This is reflected, for example, in the low number of complaints filed by victims, who most often explain their decision not to take legal action by saying that it would have been *"pointless"* (Défenseur des droits, 2023; Lê *et al.*, 2022). It can also damage their health or increase their sense of insecurity. Among those who have experienced discrimination or discriminatory violence on the grounds of their ethnic origin, skin colour or religion, 56% report that the discriminatory behaviour they experienced caused psychological harm, as well as an increased sense of insecurity in their neighbourhood or village (SSMSI, 2022).

Given the scale and persistence of this trend, this article aims to offer three perspectives on the discriminatory process at work in the labour market against people of North African origin.[1] We propose to compare the results of three different approaches, thereby accounting for the limitations inherent in each. In a recent literature review on the methodological challenges of measuring discrimination, Neumark (2018) highlighted the benefits of combining results. He also noted that the results of correspondence tests could be more widely used in the design of public policies to combat discrimination.

The first approach involves identifying instances of discrimination faced by people perceived as North African during the recruitment process by employers, using an experimental method that replicates a potentially discriminatory situation. It is based on a series of correspondence tests conducted between December 2019 and April 2021 by ISM-Corum and the IPP[2] under the aegis of DARES.[3] The second approach looks at individuals' perceives experiences on discrimination in job search process. It is based on self-reported experiences of unequal or discriminatory treatment in the TeO2 survey, conducted by INSEE and INED in 2019 and 2020. The third approach identifies "residual" discrimination by comparing the unemployment risk among people with no direct migrant background to that of people of North African origin, using INSEE's 2019 and 2020 Labour Force Surveys. This approach does not measure hiring discrimination, but focuses on its potential consequences in terms of unemployment.

In the following section, we examine in greater detail the strengths and limitations of each of these approaches and discuss how they complement one another. The second section presents the data and methodology. The results are presented and discussed in the third section.

1. The decision to study discrimination against people from this region is based not only on the findings in the literature, but also on the data we use in this article.
2. *Respectively Inter Service Migrants*—Centre d'Observation et de Recherche sur l'Urbain et ses Mutations *(Inter Service Migrants—Centre for Observation and Research on Urban Areas and Their Transformations) and* Institut des Politiques Publiques *(Institute for Public Policies).*
3. Direction de l'animation de la recherche, des études et des statistiques *(Directorate for Research, Studies and Statistics).*

1. Literature Review

In 1957, Gary Becker analysed why economic agents prefer interacting with certain individuals rather than others. He defined taste-based discrimination as resulting from a disutility associated with interacting with people perceived as different. When recruiters have strong prejudices against a social group to which they do not belong, they may consider that the interaction costs involved in recruiting someone from that group would be too high, leading them to favour applicants who are similar to themselves. Following on from this work, Phelps (1972) and Arrow (1973) highlighted the existence of a statistical bias that arises from imperfect information about individuals' membership in minority social groups. Recruiters tend to attribute to applicants the average characteristics of the group to which they believe they belong. If they consider that these groups, on average, possess less productive characteristics or that the quality of the services they provide is lower, their risk aversion leads them to favour applicants from the majority group. Beyond the numerous theoretical extensions that have developed around these two lines of thought,[4] determining which inequalities result from discrimination can prove to be a complex exercise (Delattre *et al.*, 2013). The empirical approaches that have emerged each have their own advantages and limitations.

1.1. An Experimental Approach to Discrimination

Correspondence studies were developed to detect the existence of discrimination by isolating differences in treatment between two groups based solely on a criterion prohibited by law. These field experiments rely on a protocol designed to account for observed differences between population groups that might be linked to differences in characteristics[5] or to applicant self-selection.

In the case of hiring discrimination, the process is straightforward: fictitious applications are created and submitted in response to a large number of job vacancies, then the responses received for each applicant profile are analysed and the discrepancies measured. This approach is based on the assumption that recruiters' biases may lead them to treat applications differently due to applicants' perceived membership in distinct social groups. A statistically significant average difference in treatment between applicant profiles indicates the existence of discrimination between these groups. Whilst conducting tests on companies seeking to recruit may raise ethical questions, given that reviewing applications represents a cost for them, this method places them in a real-world recruitment context (Du Parquet & Petit, 2019; Bertrand & Duflo, 2017). Recruiters are unaware that they are being tested and do not adjust their behaviour (King *et al.*, 2013). The observed differences reflect solely the judgements based on the group to which they assume the applicants belong.

This initial approach appear to be the most suitable for directly measuring discrimination: the clarity and simplicity of the method facilitate an understanding of the results particularly as it does not rely on complex econometric models (Delattre *et al.*, 2013). The main difficulty lies in developing a rigorous protocol to ensure the quality of the measurement obtained; a recent literature review conducted by Neumark (2018) details the methodological challenges faced by researchers.

The scope of the study is defined in advance. The occupations are chosen carefully, in line with the hypotheses that guided the implementation of the experiment. They are also selected to limit the risk of detection, which could influence recruiters' behaviour. To this end, the selected occupations must attract a sufficiently large number of applications for those sent by the researcher to go unnoticed.

Once the number of fictitious applications to be sent per job vacancy has been determined, particular attention is paid to their design to ensure that they are of comparable quality: educational qualifications, places of study, professional experience, personal characteristics, language proficiency, etc., are carefully selected. Only one indication that does not affect applicants' productivity is included in the CVs to suggest that they belong to a specific social group, through first and last names, for example. The content of CVs and cover letters is often exchanged between applicants to ensure that any observed differences in treatment do not reflect preferences for specific applicant content (Du Parquet & Petit, 2019).

Correspondence tests assume that there is no self-selection bias. The procedure for submitting applications is designed so that job-seeking

4. For more details on developments and the links between taste-based discrimination and statistical discrimination, see Rathelot & Safi (2022).

5. There are many factors influencing the productivity of applicants and the criteria used by recruiters. The choice of the applicant to be hired is most often based on experience and skills, but also on the motivation, flexibility or availability of the person being recruited (Bergeat & Rémy, 2017), and these criteria vary depending on the recruitment context and the roles involved (De Larquier & Marchal, 2020).

behaviour does not vary according to applicant profiles (Du Parquet & Petit, 2019), ensuring that the measurement of discrimination reflects only the difference in treatment by recruiters. To ensure that the measurement is as representative as possible of the real labour market, the fictitious applications are, in most cases, sent in response to published job vacancies (Adamovic, 2020).

Due to the constraints imposed by the experimental protocol, correspondence studies can be criticised for the limited scope of their results. Firstly, they can only measure potential discrimination, that which would be observed if applicants behaved as defined in the experimental protocol, rather than actual discrimination in the labour market (Heckman, 1998). Secondly, the results reflect only those recruitments that result from the publication of an online job advertisement.[6] Finally, correspondence tests are restricted to a specific number of occupations, a given geographical area and a specific period, calling into question the generalisability of their results (Fougère *et al.*, 2011; Delattre *et al.*, 2013). Given these limitations, Neumark (2018) highlights the value of comparing correspondence tests with other approaches to strengthen the validity of the results.

1.2. An Approach Based on Self-Reported Experiences of Discrimination

Studying the experiences of people who have been victims of discrimination constitutes a "subjective" approach. It involves asking people about their perception of their own situation. It therefore focuses on how inequalities and discrimination are internalised by those who experience them. These experiences can indeed lead to an anticipation of discrimination (Delattre *et al.*, 2013) and alter the behaviour of people who are victims, for example in job search process. Awareness of the risk of unfair treatment in the labour market can lead to self-selection behaviours. While correspondence studies are criticised for not reflecting actual job-seeking behaviour in the labour market (Heckman, 1998), self-reported experiences of discrimination are more representative of it. Depending on their profile, applicants may apply to different job vacancies and may not encounter the same recruiters.

Despite the precautions taken in collecting and analysing these data, the reliability of results from this approach is often questioned due to their subjectivity. According to Carcillo & Valfort (2020) or Delattre *et al.* (2013), asking people about their own experiences may just as easily underestimate as overestimate discrimination. On the one hand, being regularly confronted with unfair or discriminatory treatment could lead respondents to underestimate the frequency of such incidents. Conversely, they might perceive certain inequalities as discrimination when in fact they are related to factors other than their membership in a particular group.

Special attention is therefore paid to the collection of this type of data: it requires the development of specific surveys, such as the Trajectories and Origins surveys in France (Beauchemin *et al.*, 2023). Safi & Simon (2014) distinguish between self-reported discrimination, where respondents answer a direct question about their experiences of discrimination, and situational discrimination, where they consider that they have been treated unfavourably because of their personal characteristics in specific situations, places or during particular interactions. Their analyses reveal, on the one hand, a strong correlation between indicators of self-reported and situational discrimination, consistent with the idea that self-reported experiences are indeed grounded in concrete facts. On the other hand, they show that self-reported experiences are consistently lower than situational experiences, suggesting an underestimation of self-reported discrimination rather than a process of victimisation.

Drawing also on data from the TeO survey, Meurs (2017) highlighted the close link between situational discrimination and the objective position of individuals in the labour market. The author showed that people who are unemployed, even though one would expect them to be in employment given their personal characteristics, also reported being the most likely to experience unfavourable treatment during recruitment, for a given profile. The approach to discrimination based on self-reported experiences appears to be consistent with the approach based on measuring inequalities of position in the labour market.

1.3. An Approach Based on Inequalities in the Unemployment Risk

The models on which this third approach is based were originally developed to study wage discrimination (Oaxaca, 1973; Blinder, 1973). They involve estimating wage equations based

6. The choice of the recruited applicant may stem from other channels, such as personal or professional connections or labour market intermediaries (Adamovic, 2020; Bergeat & Rémy, 2017), and the discrimination risk may vary depending on the recruitment channel (Rémy & Valat, 2025). The work of Challe et al. *(2020) shows, however, that the measurement of discrimination against people of presumed North African origin is not sensitive to the recruitment channel tested.*

on a large set of observable characteristics, and then measuring the difference between actual wages and the average wage for workers with identical characteristics. However, these methods have been criticised for relying on the concept of a "reference group" and for assuming that the effect of the explanatory variables in this reference group would be the same as that observed in the potentially discriminated group in the absence of discrimination (Popli, 2022). There have since been methodological developments to overcome this limitation, notably by Di Nardo *et al.* (1996). We have taken into account these developments in this article. The method developed by Di Nardo *et al.* is valued for being intuitive and easy to implement, for not relying on the concept of a reference group, and for being better suited to the study of dichotomous variables (Popli, 2022), such as the risk of being unemployed rather than in employment.

While the experimental approach using correspondence tests proves to be the most relevant for detecting hiring discrimination, it only covers the applicant selection step and does not allow for addressing the disparity in circumstances between two populations, particularly regarding employment, wages, etc. (Neumark, 2018). The same applies when discrimination is examined through self-reported experiences: experiencing unfair or discriminatory treatment does not necessarily lead to unemployment. Approaching the issue through the study of inequalities in the unemployment risk is therefore complementary: it involves studying the consequences of inequalities or discrimination identified by other approaches rather than seeking to detect their existence (Popli, 2022). It makes it possible to formulate hypotheses about the sources of these inequalities (Delattre *et al.*, 2013) and to study their consequences on the structure of the labour market.

Certain precautions are required when interpreting the results produced. This approach does not allow for the identification of differences in unemployment risk that are solely attributable to discrimination (Chappe & Eberhard, 2021; Carcillo & Valfort, 2020; Jugnot, 2019). The residual approach merely confirms that the persistent differences in unemployment risk are not related to the variables included in the model. While applying such methods to large administrative databases offers a better representation of the labour force (Fougère *et al.,* 2011), the control variables available in these datasets are limited. Athari *et al.* (2019) explain that some of the differences observed between the population with no migrant background and that of immigrant origin could be attributed to factors not reflected in administrative data, such as proficiency in French or difficulties in having qualifications obtained abroad recognised. In addition, labour market participation patterns and the choice to apply for certain job vacancies rather than others may also influence the unemployment risk.

Despite the difficulty in identifying unexplained differences in unemployment risk as discrimination, this latter approach to discrimination is complementary and produces results that are generally consistent with other approaches. Neumark (2018), for example, confirms that hiring discrimination detected in correspondence tests is more pronounced among population groups experiencing the greatest average wage inequalities or unemployment risks.

2. Data and Methodology

The three approaches to discrimination discussed in this article require the use of different datasets. Particular attention is paid to ensuring comparability of the field of study across the three approaches.

2.1. Correspondence Tests

2.1.1. Experimental Protocol and Data Collection

A correspondence study was conducted by the IPP and ISM-Corum under the aegis of DARES between December 2019 and April 2021.[7] The aim was to obtain an "overall" measurement of discrimination based on the applicants' gender and perceived origin (French or North African).[8] Twelve occupations were selected (Table 1) for the variety of their skill levels (low-skilled or skilled, with or without managerial responsibilities) and required qualifications (ranging from CAP (*Certificat d'Aptitude Professionnelle*, Vocational Training Certificate) to Master's degree), level of labour market tension (low or high, measured using the 2019 Labour Needs survey—*enquête Besoins en main-d'oeuvre*) and the extent of female representation (based on the share of women and men working in these occupations in the 2015 Annual Declarations of Social Data (DADS, *Déclarations administratives de données sociales*). The websites of Pôle Emploi, followed by Indeed, APEC, L'hôtellerie-restauration, Cadremploi and Cadreo

7. *Data collection was suspended between 15 March and 1 July 2020. The applicant callback rate decreased overall between the period before 15 March and that after 1 July, but the differences in response rates by ethnic origin are of comparable magnitude (Arnoult* et al.*, 2021).*
8. *See Breda* et al. *(2022) for more details on the experimental protocol.*

were daily consulted. Careful consideration was given to the selection of job vacancies tested to ensure coverage of the mainland France. Certain job vacancies were systematically excluded to minimise the risk of detection: those from an establishment that had already been tested, those advertised via employment agency, and those not recruiting online. The number of tests carried out per day and per occupation was limited so that data collection was spread across the entire calendar year. In total, 2,400 job offers were tested.

For each occupation, two levels of experience were tested in order to assess whether the level of discrimination depends on experience. For low-skilled and skilled roles without managerial responsibilities, applications with limited experience (4 to 6 years' professional experience) were submitted for half of the job vacancies tested, and applications with extensive experience (14 to 16 years' experience) for the other half. Experienced applications were sent to half of the vacancies for skilled roles involving managerial responsibilities, and highly experienced applications (29 to 31 years' experience) to the other half.

For each occupation and level of experience, four fictitious applications (female/male, with perceived French/North African origins) were created, each containing a CV and a cover letter, for a total of 104 applications (13 occupations × 2 levels of experience × 4 profiles). Their content (degree title and place of graduation, professional experience, skills, etc.) was designed to ensure they were of similar quality. Despite these precautions, and to ensure that the preference for an applicant was not the result of an imbalance between the content of the fictitious applications, these were systematically exchanged between the four applications. They varied according to two criteria, gender and ethnicity, which are suggested by the surnames and first names attached to them. Gaddis (2017; 2019) emphasises the importance of the choice of applicants' identities, which constitutes the only information relating to gender and ethnic origin available to recruiters. Whilst surnames and first names may suggest a link to migration, they can also convey information about social background, which may bias the measurement of discrimination related to other criteria (Crabtree & Chykina, 2018; Delattre *et al.*, 2013). To avoid this, a list of the most common first names by gender, age group[9] and social background (based on the father's socio-occupational category: executive or intermediate professional, clerical worker or

9. The most common first names were selected for each age group of applicants, based on their level of experience (inexperienced, experienced, highly experienced).

Table 1 – **Occupations included in the field experiment**

Occupations	Qualification	Level of experience tested	Level of qualification	Gender balance	Tension	Number of tests
Administrative employee	Low level of qualification	Inexperienced, experienced	CAP to Baccalauréat(3)	Predominantly female	Low	240
Electrical fitter/order picker(1)				Predominantly male	Low	240
Sales assistant				Mixed	Low	240
Kitchen hand				Predominantly male	High	240
Recruitment officer / HR manager(2)	Qualified with no managerial responsibilities		2 to 5 years of tertiary education	Predominantly female	Low	240
IT sales engineer				Predominantly male	Low	240
Management controller				Mixed	Low	240
IT developer				Predominantly male	High	240
Production engineer	Qualified with managerial responsibilities	Experienced, highly experienced		Predominantly male	Low	240
Store manager				Mixed	Low	120
Restaurant manager				Predominantly male	Low	120

(1) The role of order picker, with the same characteristics, was introduced in addition to that of electrical fitter from July 2020.
(2) The number of job vacancies for the role of HR manager soon proved to be insufficient; the four tests that had already been carried out were added to those for recruitment officers.
(3) High School Diploma.
Source: DARES/IPP/ISM correspondence study (2019–2021).

manual worker) among people with no migrant background (neither immigrants nor descendants of immigrants) and those of North African origin was compiled using data from the TeO1 survey. A survey was then conducted among 453 people to identify the 24 first names (2 genders × 2 geographical origins × 3 age groups × 2 social background categories) for which gender and ethnic origin were most often correctly attributed. These first names were then randomly assigned to each application before submission. The surnames selected were common names occurring throughout mainland France.

2.1.2. Estimating the Effect of Ethnic Origin on the Probability of Receiving a Callback

Comparing the callback rates for applicants whose first and last names suggest a North African background with those whose first and last names suggest a French background provides an estimate of discrimination based on ethnic background. To back up the results, an analysis with all other factors being equal was conducted, based on the following *logit* model:

$$\pi_{ij} = \log\left(\frac{p_{ij}}{1-p_{ij}}\right) = \alpha + \beta_1 orig_i + \beta_2 fem_i + \epsilon X_j + \tau E_{ij} + \mu_{ij}$$

where π_{ij} is the log odds obtained from the estimation of a logistic model, measuring the likelihood that the establishment j responded positively to the application i. $orig_i$ is the variable of interest, which is equal to 1 if the applicant is perceived to be of North African origin and 0 if they are perceived to be of French origin. fem_i is an indicator that is equal to 1 if the applicant is a woman and 0 otherwise. The vector X_j includes the contract type (open-ended or fixed-term), the business sector, the size of the establishment and an indicator of whether it is located in the Paris region, as well as two contextual variables: the share of immigrants and the unemployment rate in the local labour market area. E_{ij} is a set of variables relating to the submission of applications: the CV submitted from among the four created for the given occupation and level of experience, the order of application submission, and the quarter in which the application was submitted.[10] μ_{ij} is an error term.

2.2. Experience of Discrimination

2.2.1. The TeO2 Survey

The Trajectories and Origins 2 survey (TeO2 —*Trajectoires et Origines*), conducted by INED and INSEE between July 2019 and November 2020 surveyed people aged 18 to 59 living in standard housing in mainland France, ensuring adequate representation by geographical area of origin for immigrants and descendants of immigrants, as well as those from the French overseas territories. It aimed to study the influence of migratory origins on living conditions and social trajectories. The sample consisted of approximately 27,200 respondents.

In this article, the analysis of experiences of hiring discrimination is based on a contextual approach, as defined by Safi & Simon (2014). It is identified by the response to the question *"In the last five years, have you been unfairly refused a job?"*, asked of all those who had looked for a job at least once in the last five years, regardless of their employment status at the time of the survey.

In this context, a person's ethnic origin is defined by administrative criteria: their country of birth and nationality at birth, and those of their parents. People of North African origin are those who are immigrants—born abroad to foreign parents and residing in France, whether or not they have since been granted French nationality—or descendants of immigrants—born in France to one or two immigrant parents—from Algeria, Morocco or Tunisia. People with no direct migrant background are those who are neither immigrants nor descendants of immigrants. This approach to ethnic origin differs from that based on first names and surnames in correspondence tests. Grandchildren of immigrants may indeed have first names and surnames suggesting a migrant background in the first approach, whereas they would be considered to have no direct migrant background in the second. To our knowledge, there is no research available to estimate the extent of the difference between identifying ethnic origin using administrative data and using first names and surnames, nor the potential implications for the study of discrimination (an overestimation or underestimation of the trend), as first names may not always reflect a person's connection to migration, and the transmission of surnames may extend beyond two generations. It is therefore conceivable that some of the differences observed between this approach and the experimental approach may result from this, if surnames and first names suggest a migrant background that differs from that identified by administrative data. To study the experience of self-reported discrimination, we adopt an

10. The indicator of the quarter in which the application was submitted allows for any bias related to the data collection period to be taken into account.

approach centred on social groups rather than on the stated grounds for discrimination. Although ethnic origin is the reason most frequently cited by people of North African origin who report having been unfairly denied a job in the last five years, it is mentioned by fewer than half of them (45%). Analysing differences in exposure to the discrimination risk between social groups thus partly overcomes the internalisation mechanisms that may influence how people interpret and categorise their experiences of unfair treatment, particularly the tendency to attribute them to one criterion of discrimination rather than another. This approach also facilitates comparison with the results obtained from the residual approach presented in Section 1.3, which examines unexplained differences in unemployment risk between these same social groups.

The scope of the analysis has been adapted to align as closely as possible with that defined by the field experiment protocol. It includes individuals who have completed their initial education, who have sought employment in the last five years, regardless of their employment status at the time of the survey, and who hold a qualification ranging from CAP to five years of tertiary education (Master's degree level) in France. Potential experience is measured as the difference between the date the highest qualification was obtained and the survey date. The least experienced individuals (less than four years of potential experience) are excluded from the sample, as are the oldest (aged over 55 and/or with more than 31 years of potential experience). This sample consists of 6,549 respondents, of whom 908 are from North Africa and 1,887 have no direct migrant background.

2.2.2. Estimating the Effect of Ethnic Origin on the Experience of Unfair Job Rejection

The individual likelihood of reporting an unfair job rejection is estimated using the following *logit* model:

$$\pi_i = \log\left(\frac{p_i}{1-p_i}\right) = \beta_0 + \beta_1 orig_i + \epsilon X_{j(i)} + \tau E_i + \mu_i$$

π_i represents the log odds that the individual i (living in local employment area j) reports having been unfairly refused a job in the last five years; p_i is the likelihood they will report such an event; $orig_i$ is an indicator equal to 1 if they are immigrant or descendant of immigrants from North Africa, and 0 otherwise. E_i is a vector of individual characteristics including: gender, educational qualifications, potential experience and its square, employment status and contract type, family status, the spouse's employment status, number of children, administrative recognition of a disability, and the quarter during which the survey was made. $X_{j(i)}$ contains two variables specific to the individual's local employment area: an indicator of residence in a QPV (priority neighbourhood) and an indicator of residence in the Paris region, as well as two contextual variables measured from the 2019 population census data: the share of immigrants and the unemployment rate in the local employment area. μ_i is an error term.

2.3. Inequalities in the Unemployment Risk

2.3.1. Labour Force Surveys

The 2019 and 2020 French Labour Force Surveys (*enquêtes Emploi*) are used to examine differences in unemployment rates in mainland France according to migrant background. These surveys are conducted quarterly by INSEE and provide information on people's employment status, their connection to migration, and a range of socio-demographic characteristics. As in the TeO2 survey, ethnic origin is addressed using an administrative approach that distinguishes between people of North African origin, whether they are immigrants or second-generation descendants of immigrants, and people with no direct migrant background, who are neither immigrants nor descendants of immigrants.

To ensure comparability with the results of the previous two approaches, unemployment rates are calculated for the population aged 18 to 55 who have completed their initial education, hold a CAP to five years of tertiary education (Master's degree level) in France, and have 4 to 31 years of potential work experience. Unlike the data from the TeO2 survey, the Labour Force Surveys do not provide information on where the most recent qualification was obtained. However, we exclude from the study those who arrived in France after obtaining their highest qualification. Finally, by definition, the analysis is restricted to employed and unemployed individuals. The final sample comprises 149,051 observations, of which 7,332 are from North Africa.

2.3.2. Analysis of Differences in Unemployment Rates by Ethnic Origin

We apply the decomposition method proposed by Di Nardo *et al.* (1996). This involves calculating the difference between the unemployment rate of individuals of North African origin (u_{magh}) and that of individuals with no direct migrant background (u_{SAM}), and then distinguishing within this difference between the "explained" portion, attributable to observable differences in characteristics between the two population groups, and the "unexplained" portion.

As a first step, we estimate the likelihood of membership in the group of immigrants and descendants of immigrants from North Africa rather than in the population with no migrant background, based on individual characteristics:

$$P\left(orig_i = 1|X\right) = \theta\left(X\beta\right),$$

where $orig_i = 1$ is whether the individual is immigrant or descendant of immigrants from North Africa, and $orig_i = 0$ is whether they belong to the population with no direct migrant background. X represents the same individual characteristics as those taken into account in the approach to discrimination based on experiences (Section 1.2): gender, potential experience and its square, educational qualification, marital status—single/in a relationship, the spouse's employment status, the presence of children in the household, including children under six years of age, official recognition of a disability or reduced activity capacity, residence in a neighbourhood identified as a priority area by urban policy, residence in the Paris region, and the quarter of the survey; the unemployment rate and the share of immigrants within the local employment area.

The individual re-weighting factor ($\psi_i\left(X\right)$) is then measured using the conditional probabilities estimated in the first step:

$$\psi_i\left(X\right) = \frac{P(orig_i = 0 \mid X)}{P(orig_i = 1 \mid X)} \times \frac{N_{orig_i=1}}{N_{orig_i=0}},$$

where N is the size of the populations concerned. The counterfactual u'_{SAM} unemployment rate is the rate that would be observed among the population with no migrant background if it had the same observable characteristics as the population of North African origin. It is obtained by applying the individual re-weighting factor $\psi_i\left(X\right)$ to the individual weight π_i of people with no migrant background, i.e.:

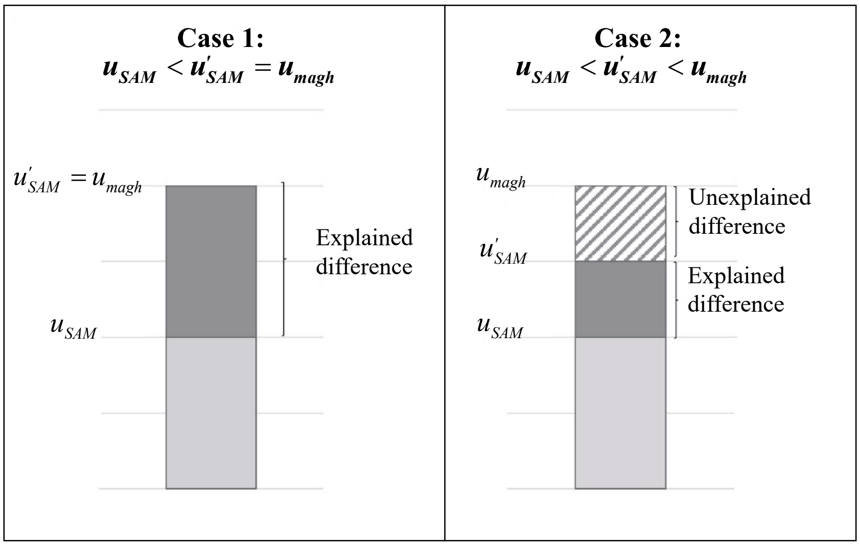
$$u'_{SAM} = \frac{\sum_{i:\,orig_i=0} \pi_i \psi_i\left(X\right) U_i}{\sum_{i:\,orig_i=0} \pi_i \psi_i\left(X\right)}$$

where $U_i = 1$ if the individual i is unemployed and 0 otherwise.

The difference in observed unemployment rates ($u_{magh} - u_{SAM}$) can be broken down into an "explained" component due to differences in observable characteristics ($u'_{SAM} - u_{SAM}$) and an "unexplained" component ($u_{magh} - u'_{SAM}$), see figure.

If the unemployment rate for people with no migrant background is equal to the unemployment rate for people of North African origin with the same characteristics (case 1), the entire observed difference between the two groups can be explained by differences in characteristics. However, if the unemployment rate for people of North African origin remains higher than that which people with no direct migrant background would have if they shared the same observable characteristics as the former (case 2), part of the difference in unemployment risks remains unexplained. This difference could, in whole or in part (depending on the relevance of the control variables used), stem from the discrimination faced by people of North African origin in the labour market.

Figure – **Breakdown of differences in unemployment rates using the Di Nardo method**



Source: author.

3. Results

3.1. 32% Lower Callback Rate for Applications Suggesting North African Origin

A "callback" refers to any contact from the employer, including a request for a phone call, further information regarding the application, or an invitation to an interview. Of the 9,600 applications submitted, 2,693 resulted in a callback, representing 28% of the total (Table 2). 19% received an explicit rejection while 53% received no response. Among applications where the first and last names suggest North African origin, 22.8% were contacted by recruiters, compared with 33.3% of those where the identity suggests French origin. The callback rate for the North African group is thus 32% lower than that of the French group: this figure is in line with the average callback rate differences measured in other correspondence studies conducted in France on the same groups (see, for example, Duguet *et al.*, 2010; Berson, 2013; Petit *et al.*, 2013; Pierné, 2013; Foroni *et al.*, 2016; Edo *et al.*, 2019; Chareyron *et al.*, 2023).

Ethnic differences in the treatment of job applications are slightly greater among men, who are more likely than women of the same origin to receive no response. For instance, the response rate for men whose name suggests North African origin is 34% lower than that for men perceived as French, compared with a 29% lower rate for women. However, this difference between female and male is not statistically significant. Studies based on French data do not provide a clear picture of the interaction between gender and ethnic origin, and the lack of statistical significance in the observed difference may reflect the diversity of the occupations included in the study. For the roles of administrative assistant, secretary and accountant, Edo *et al.* (2019) showed, for example, that discrimination against people of North African origin is as prevalent among women as it is among men. Conversely, differences based on ethnic origin were more apparent among women for the role of IT specialist requiring a Master's degree (Petit *et al.*, 2013), while they were higher among men for the role of cashier (Berson, 2013).

All else being equal, these differences in treatment are confirmed (Table 3): the effect of ethnic origin on the likelihood of receiving a callback from the recruiter remains similar once the characteristics of the establishments and the experimental protocol are included in the analysis. The discrimination risk based on ethnic origin among male applicants remains slightly higher than among female applicants, but this small difference is not statistically significant.

3.2. Hiring Discrimination Reported Twice as Often Among People of North African Origin

When asked about their experience of being unfairly refused a job in the last five years, immigrants or descendants of immigrants from North Africa are twice as likely to report this as

Table 2 – **Responses received by applicants by gender and perceived origin**

	All applicants		Female applicants		Male applicants		Total
	French	North African	French	North African	French	North African	
Negative response	819	990	421	492	398	498	1,809
No response	2,382	2,716	1,177	1,336	1,205	1,380	5,098
Positive response	1,599	1,094	802	572	797	522	2,693
Callback rate (%)	33.3	22.8	33.4	23.8	33.2	21.8	28.1
Callback rate ratio	0.685 (0.021)		0.713 (0.026)		0.657 (0.027)		
Callback rate odds ratio	0.591*** (0.021)		0.623*** (0.028)		0.559*** (0.027)		

Note: The callback rate corresponds to the share of applicants who received a positive response. The callback rate ratio is the ratio between the callback rate for applications of perceived North African origin and that for applications of perceived French origin. Standard deviations are shown in parentheses.
Reading note: Across all applications, the callback rate for applicants whose identity signals a North African origin is 32% lower than that for those signalling a French origin (22.8/33.3 = 0.68).
Field: 9,600 applications (2,400 job vacancies x 4 applications), mainland France.
Source: DARES/IPP/ISM correspondence study (2019–2021).

Table 3 – **Likelihood of receiving a callback following an application (logistic regression)**

	All applicants		Male applicants		Female applicants	
	(1)	(2)	(1)	(2)	(1)	(2)
Presumed North African origin	0.591*** (0.021)	0.587*** (0.021)	0.559*** (0.027)	0.556*** (0.027)	0.623*** (0.028)	0.620*** (0.028)
Female	1.059* (0.034)	1.062* (0.034)	-	-	-	-
Control variables		X		X		X
Number of observations	9,600	9,600	4,800	4,800	4,800	4,800
Pseudo R^2	0.012	0.024	0.014	0.027	0.009	0.024

Note: The likelihood of an application receiving a callback is estimated using logistic regression without control variables (Model 1) and with control variables (Model 2). The table shows the odds ratios. These may differ significantly from 1 at the 1% (***), 5% (**) or 10% (*) significance levels. Standard deviations are shown in parentheses.
Reading note: All other things being equal, the likelihood of being contacted by a recruiter rather than not being contacted is 41.3% lower among applications from candidates of perceived North African origin than those of perceived French origin (column (2): 1–0.587 = 0.413).
Field: 9,600 applications (2,400 job vacancies x 4 applications), mainland France.
Source: DARES/IPP/ISM correspondence study (2019–2021).

those with no direct migrant background (32% compared with 16%, Table 4). This difference does not vary significantly by gender: men of North African origin are 2.1 times more likely than those with no direct migrant background to report that they have been unfairly refused a job in the last five years; this ratio is 1.9 among women. This finding is consistent with those of several previous studies (Défenseur des droits, 2023; Lê *et al.*, 2022; SSMSI, 2022), which put forward the hypothesis of a correlation between perceived discrimination and that measured objectively using correspondence tests.

Once socio-demographic differences have been controlled for, differences in the likelihood of reporting unequal treatment in job search persist (Table A1 in the Appendix). Individuals of North African origin are relatively more likely to report having experienced discrimination in the last five years than those with no direct migrant background. In line with the results from the experimental approach, discrimination against individuals of North African origin does not vary by gender. Meurs (2017) had shown, based on the first edition of the TeO survey, that reports of discriminatory situations in job seeking were more frequent among men than among North African women. Whilst there does not appear to be a difference between women and men of North African origin in our study, this may be due to its more limited scope of analysis. Indeed, Brinbaum *et al.* (2016) explained that the difficulties faced by men who are descendants of North African immigrants when entering the labour market were mainly due to their early exit from the education system with no qualification. People without qualifications are, however, excluded from our scope of analysis. Furthermore, Lê *et al.* (2022) showed that self-reported discrimination changed between the two rounds of the TeO surveys, with, in particular, an increase in reports of sexist discrimination, which may have reduced the gap observed between women and men of North African origin.

Table 4 – **Experiences of unfair job rejection according to gender and migrant background**

	No direct migrant background (%)	North African origin (%)	Ratio	Odds ratio
All	16.0	31.9	2.004 (0.129)	2.726 (0.585)
Males	17.3	35.7	2.068 (0.189)	2.886 (0.828)
Females	14.8	28.1	1.899 (0.174)	2.465 (0.813)

Note: Standard deviations are shown in parentheses.
Reading note: People of North African origin are 2.004 times more likely than those with no direct migrant background to report that they have been unfairly denied a job (31.9% versus 16.0%).
Field: Individuals aged 18 to 55 with 4 to 31 years of potential work experience, who have obtained qualifications in France ranging from CAP to Master's degree level, have completed their initial education and training, and live in standard accommodation in mainland France.
Source: INED-INSEE, Trajectories and Origins 2 survey (2019–2020).

3.3. Higher Unemployment Risk Among Immigrants and Descendants of Immigrants From North Africa

Discrimination in access to the labour market is associated with a higher unemployment risk for individuals of North African origin (Table 5). In 2019–2020, their unemployment rate was nearly three times higher than that of individuals who are neither immigrants nor descendants of immigrants (14.1% compared with 5.2%). This difference can only be partially explained by differences in observed characteristics: if individuals without direct migrant background had the same characteristics (in terms of potential experience, qualifications, etc., see Section 2.3.2) as those of North African origin, the difference in unemployment rates between the two groups would be only 5.6 percentage points, compared with 8.9 points in reality. Thus, when the differences in average characteristics between the two groups are controlled for, 63% of the observed unemployment rate gap remains unexplained.

Although our analysis excludes from its scope those without qualifications, foreign nationals who obtained their qualifications abroad, the oldest and least experienced workers, who are likely to face the greatest inequalities in treatment in the labour market, the unexplained difference in unemployment remains of the same order of magnitude as that estimated by Athari *et al.* (2019) for the entire working population. Difficulties in mastering the language or in having qualifications obtained abroad recognised, cited by these authors as a possible explanation, do not appear sufficient to account for the increased unemployment risk among individuals of North African origin.

Unlike the two previous approaches, which show no significant differences by gender, the unexplained share of unemployment by origin is much higher among men (79%) than among women (48%), a finding already highlighted by Athari *et al.* (2019).

Applied to the population as a whole, the three approaches to discrimination tend to yield similar results, highlighting the disadvantage faced by individuals of North African origin in accessing the labour market. This finding reinforces the conclusions of the existing literature, which highlights both the close link between differences in unemployment risk and perceived discrimination (Meurs, 2017), and the correlation between the results of correspondence tests and unexplained unemployment risk (Neumark, 2018).

However, the approach based on inequalities in the unemployment risk yields more pronounced differences among men than among women, whereas the other two approaches do not reveal any significant difference in discrimination by ethnic origin between men and women. The lower level of participation in labour market among North African women could reduce their relative risk of unexplained unemployment. Indeed, although there are few differences in participation rates between men of North African origin and those with no direct migrant background, the participation rate of North

Table 5 – **Unexplained component of unemployment risk among people of North African origin compared with those with no direct migrant background**

	All	Females	Males
Unemployment rate among immigrants and descendants of immigrants from North Africa (1)	14.1	15.2	13.1
Unemployment rate among individuals with no migrant background (2)	5.2	5.3	5.1
Counterfactual unemployment rate among individuals with no migrant background (3)	8.5	10.4	6.8
Observed gross difference (4) = (1) – (2)	8.9	9.9	8.0
Explained difference (5) = (3) – (2)	3.3	5.1	1.7
Unexplained difference (6) = (4) – (5)	5.6	4.8	6.3
Unexplained component of the unemployment rates difference (7) = (6)/(4) (%)	63.0	48.0	79.0

Note: Differences in unemployment rates are measured using the decomposition method of Di Nardo *et al.* (1996).
Reading note: The unemployment rate is 14.1% among individuals of North African origin, compared with 5.2% among those with no direct migrant background. If the latter group had, on average, the same observable characteristics as the former, their unemployment rate would be 8.5%. Thus, 63% of this difference is unexplained.
Field: Individuals aged under 55 with 4 to 31 years of potential work experience, who have obtained qualifications in France ranging from CAP to Master's degree level, have completed their initial education and training, and live in standard accommodation in mainland France.
Source: INSEE, Labour Force surveys (2019–2020).

African women is significantly lower than that of women with no direct migrant background.[11] Given the discrimination they face in the labour market, they may withdraw from it in contexts where men of the same origin are more likely to face unemployment or precarious employment.

3.4. An Analysis by Educational Level and Gender Using the Three Approaches

When recruiters are looking to fill a post requiring a qualification ranging from a CAP to a *Baccalauréat* (High School Diploma) level, they are more likely to contact men perceived as French (35.6%) than those perceived as North African (17.9%, Table 6). For this type of position, differences in callback rates by origin are more pronounced among male than female applicants: the callback rate for men perceived as North African is 50% lower than that for men perceived as French, compared with a 40% lower rate among women.

The difference in treatment is smaller for applicants to jobs requiring a higher education qualification (two to five years of tertiary education) than for applicants to jobs requiring a secondary school qualification. Both female and male perceived as North African receive, on average, 23% fewer callbacks than those of perceived French origin.

When recruiting for higher qualification levels, recruiters make less distinction between applicants based on their perceived origin; the signal conveyed by the qualification appears to reduce the role of prejudice in selecting applicants with similar profiles. This reduced discrimination risk when the qualification is higher was previously found in a meta-analysis summarising the results of 97 correspondence studies (Quillian *et al.*, 2019), which revealed a statistically significant reduction in discrimination based on origin for positions requiring a higher education qualification compared with a Baccalauréat level. Although the available data do not allow confirmation, the authors suggest that the lower discrimination risk may reflect the greater information available to recruiters when hiring university graduates, which leaves less room for bias and stereotypes than in the case of applicants with only secondary education. Nevertheless, the negative effect of ethnic origin on the likelihood of receiving a callback remains statistically significant for hiring requiring a higher education qualification (Table A2 in the Appendix).

A relatively higher share of individuals of North African origin report unfair job rejection when they hold a higher education qualification rather than a secondary school qualification (Table 7). Among secondary school graduates, those of North African origin are 1.7 times more likely to report unfair job rejection than those with no direct migrant background. This ratio is higher, at 2.4, for those with two to five years of post-secondary education, particularly for men (2.6) and, to a lesser extent, for women (2.1). These results remain valid when all other factors are held constant (Table A3 in the Appendix).

Finally, when discrimination is examined in terms of differences in unemployment rates, modest differences are observed among men depending on their educational level, with a lower unexplained share of unemployment among those with the highest qualifications (Table 8). For women, the unexplained difference in unemployment rates, based on observable characteristics, goes in the opposite direction:

11. *For more details, see the thematic information sheet on "Labour market participation", INSEE (2023).*

Table 6 – **Callback rates by origin, gender and qualification level**

	Callback rate for individuals of North African origin	Callback rate for individuals of French origin	Callback rate ratio
CAP to Baccalauréat			
All applicants	18.1	33.0	0.548 (0.024)
Female applicants	18.3	30.3	0.604 (0.034)
Male applicants	17.9	35.6	0.503 (0.029)
2 to 5 years of tertiary education			
All applicants	25.9	33.6	0.771 (0.032)
Female applicants	27.5	35.5	0.775 (0.044)
Male applicants	24.3	31.6	0.769 (0.040)

Note: See Table 2.
Source: DARES/IPP/ISM correspondence study (2019–2021).

Table 7 – **Experiences of unfair job rejection by gender, migrant background and educational level**

	No direct migrant background (%)	North African origin (%)	Ratio
CAP to Baccalauréat			
All applicants	17.4	29.5	1.695 (0.113)
Female applicants	15.8	26.6	1.686 (0.162)
Male applicants	19.3	32.3	1.674 (0.157)
2 to 5 years of tertiary education			
All applicants	14.3	34.1	2.385 (0.149)
Female applicants	13.8	28.8	2.091 (0.201)
Male applicants	15.0	39.7	2.643 (0.237)

Note and field: See Table 4.
Reading note: Among individuals with qualifications ranging from the CAP to the Baccalauréat, those of North African origin are 1.7 times more likely than those with no direct migrant background to report that they have been unfairly denied a job (29.5% versus 17.4%).
Source: INED-INSEE, Trajectories and Origins 2 survey (2019–2020).

Table 8 – **Unexplained component of unemployment risk among people of North African origin compared with those with no direct migrant background, according to educational level**

	CAP to Baccalauréat			2 to 5 years of tertiary education		
	All	Females	Males	All	Females	Males
Unemployment rate among immigrants and descendants of immigrants from North Africa (1)	19.6	21.5	18.2	8.5	9.1	7.9
Unemployment rate among individuals with no migrant background (2)	7.2	8.1	6.6	3.4	3.3	3.5
Counterfactual unemployment rate among individuals with no migrant background (3)	12.1	16.3	8.8	5.3	5.7	4.8
Observed gross difference (4) = (1) – (2)	12.4	13.5	11.6	5.2	5.9	4.3
Explained difference (5) = (3) – (2)	4.8	8.3	2.3	1.9	2.5	1.2
Unexplained difference (6) = (4) – (5)	7.5	5.2	9.3	3.3	3.4	3.1
Unexplained component of the unemployment rates difference (7) = (6)/(4) (%)	61	39	80	63	58	72

Note and field: See Table 5.
Source: INSEE, Labour Force surveys (2019–2020).

it is much greater for those with higher education qualifications than for those with secondary education. Among women with secondary school qualifications, 39% of the gap in unemployment risk cannot be explained by differences in observable characteristics between North African women and those with no migrant background, compared with 58% among women with higher education qualifications.

* *
*

Several factors could explain the differences observed depending on how discrimination in the labour market is approached. Firstly, the experimental protocol defined for conducting correspondence tests allows for a more precise measurement of differences in treatment for a given profile. The applications are identical in all respects, whereas in practice, degree specialisations or the reputation of the awarding institutions may vary by gender and origin. Thus, the lower discrimination risk for higher qualification levels in the correspondence study approach, which is not observed in the other two approaches, may be due to the fact that the latter two also reflect other sources of inequality, including those related to educational backgrounds.

The difference between the approaches could also be due to job-seeking behaviour, which can vary significantly between applicants. Valat

(2016), for example, shows that women from migrant backgrounds are less likely to draw on their personal networks when looking for work than men. They appear to find employment relatively more often through labour market intermediaries (De Larquier & Rieucau, 2015). However, recruitment for managerial positions most often takes place through personal connections or labour market intermediaries outside the public employment service (Bergeat & Rémy, 2019). As the correspondence tests only cover recruitment conducted via the publication of job vacancies, they are not fully representative of the difficulties faced by women of North African origin in accessing skilled employment.

Finally, downward occupational mobility could also be a factor. For a given profile, individuals with an immigrant background are more likely to hold a job below their skill level when they have a higher education qualification (INSEE, 2023). Beauchemin *et al.* (2022) showed that the returns on education were lower for the descendants of non-European immigrants, among whom the share of those with a higher education qualification working in intermediate or managerial roles is significantly lower. Thus, experiences of unfair job rejection may not be reflected in an increased unemployment risk, particularly among male immigrants or descendants of immigrants with higher education. Women with the lowest qualifications, often constrained to jobs below their skill level, may be more likely to exit the labour market, mechanically reducing their risk of unexplained unemployment compared with men, who are more exposed to it.

A comparative approach to discrimination highlights results that no single method, considered in isolation, can reveal. The three approaches used in this article generate complementary findings, and their comparison contributes to a more nuanced understanding of the complexity of discrimination. Differences in results, where they exist, are partly due to the specific characteristics of each method, including their assumptions, scope and analytical frameworks. By combining these three approaches, we can overcome the limitations inherent in each and better capture the diverse forms that labour market inequalities can take across migrant backgrounds. ☐

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Table A1 – **Likelihood of reporting unfair job rejection in the last five years (logistic regression)**

	All		Males		Females	
Migration background (*Ref.:* No direct migrant background)						
Of North African origin	2.502***	(0.399)	2.643***	(0.645)	2.574***	(0.535)
Of Sub-Saharan African origin	2.563***	(0.508)	1.672*	(0.500)	3.950***	(1.058)
From another EU27 country	1.082	(0.196)	1.107	(0.324)	0.994	(0.240)
From another country	1.408*	(0.260)	1.210	(0.320)	1.601*	(0.411)
Female (*Ref.:* male)	0.869	(0.145)	-		-	
Higher education level (*Ref.:* CAP to Bac)	1.150	(0.225)	1.08	(0.299)	1.185	(0.331)
Potential experience	0.967	(0.049)	1.020	(1.020)	0.893*	(0.057)
Potential experience squared	1.002	(0.001)	1.000	(0.002)	1.004**	(0.002)
Employment status (*Ref.:* Open-ended contract - CDI)						
Fixed-term contract - CDD	1.123	(0.297)	0.929	(0.354)	1.212	(0.423)
Temporary worker	3.432**	(1.460)	4.419**	(2.295)	0.649	(0.356)
Apprentice, student	0.803	(0.369)	0.510	(0.252)	1.381	(0.903)
Unemployed	1.871**	(0.452)	1.703	(0.592)	2.003*	(0.629)
Has never worked	1.390	(0.640)	1.572	(0.948)	1.334	(0.820)
Other not in employment	1.850*	(0.568)	2.408	(1.106)	1.480	(0.555)
Number of observations	4,889		2,226		2,663	

Note: The likelihood of reporting unfair refusal of employment is estimated using logistic regression, first for all applications, then for female applicants, and finally for male applicants. Odds ratios differ significantly from 1 at the 1% (***), 5% (**) or 10% (*) significance levels. Standard deviations are shown in parentheses.
Field: Individuals aged under 55 with 4 to 31 years of potential work experience, who have obtained qualifications in France ranging from CAP to Master's degree level, have completed their initial education and training, and live in standard accommodation in mainland France.
Source: INED-INSEE, Trajectories and Origins 2 survey (2019–2020).

Table A2 – **Likelihood of receiving a callback following an application according to educational level (logistic regression)**

	CAP to Baccalauréat			Higher education level		
	All	Males	Females	All	Males	Females
Presumed North African origin	0.434*** (0.028)	0.378*** (0.032)	0.500*** (0.042)	0.688*** (0.029)	0.690*** (0.041)	0.685*** (0.039)
Female	0.873** (0.050)	-	-	1.192*** (0.047)	-	-
Control variables	X	X	X	X	X	X
Number of observations	3,840	1,920	1,920	5,760	2,880	2,880
Pseudo R^2	0.065	0.073	0.066	0.020	0.022	0.019

Note: Estimation of a logistic model; the odds ratios are presented here, with their standard deviations shown in parentheses. The control variables are the type of contract, the business sector, the size of the establishment, an indicator of location in Île-de-France, the proportion of immigrants and the unemployment rate in the employment area, the CV template used, the order in which applications were sent, and the quarter in which they were sent.
Source: DARES/IPP/ISM correspondence study (2019–2021).

Table A3 – **Likelihood of reporting unfair job rejection (odds ratios)**

	CAP to Baccalauréat			2 to 5 years of tertiary eduation		
	All	Males	Females	All	Males	Females
Migration background (*Ref.:* No direct migrant background)						
Of North African origin	1.697* (0.375)	1.454 (0.443)	2.189** (0.756)	3.759*** (0.894)	5.800*** (2.264)	2.897** (0.939)
Of Sub-Saharan African origin	1.450 (0.457)	0.795 (0.401)	2.600** (1.175)	3.827*** (1.043)	3.222** (1.252)	5.102*** (1.998)
From another EU27 country	0.969 (0.254)	0.794 (0.311)	1.03 (0.377)	1.104 (0.295)	1.259 (0.533)	0.803 (0.293)
From another country	1.206 (0.301)	1.008 (0.400)	1.436 (0.501)	1.530 (0.401)	1.180 (0.462)	1.740 (0.592)
Female (*Ref.:* male)	0.820 (0.187)	-	-	0.960 (0.240)	-	-
Potential experience	0.994 (0.069)	1.06 (0.111)	0.937 (0.086)	0.891 (0.068)	0.979 (0.117)	0.815** (0.073)
Potential experience squared	1.000 (0.002)	0.999 (0.003)	1.002 (0.002)	1.005** (0.002)	1.001 (0.003)	1.008*** (0.003)
Employment status (*Ref.:* Open-ended contract - CDI)						
Fixed-term contract - CDD	0.827 (0.285)	0.973 (0.490)	0.615 (0.264)	1.482 (0.594)	0.940 (0.605)	2.005 (0.961)
Temporary worker	2.906** (1.419)	3.526** (2.067)	0.486 (0.384)	6.082** (4.361)	15.097*** (13.261)	0.348 (0.253)
Apprentice, student	0.986 (0.540)	0.761 (0.427)	1.367 (1.215)	0.202** (0.154)	0.227 (0.289)	0.300 (0.236)
Unemployed	1.574 (0.465)	1.553 (0.621)	1.565 (0.612)	2.754*** (1.047)	3.101* (1.837)	2.456* (1.135)
Has never worked	1.375 (0.711)	2.932 (1.919)	1.045 (0.709)	1.406 (0.850)	1.000 (0.000)	3.351 (2.477)
Other not in employment	1.983 (0.844)	4.594** (2.789)	1.259 (0.725)	1.913 (0.844)	1.785 (1.161)	1.789 (0.899)
Number of observations	2,375	1,176	1,199	2,514	1,041	1,464

Note: Estimation of a logistic model. Standard deviations are shown in parentheses. The following are also included as control variables: presence of children, spouse's employment status, recognition of a long-term illness, the proportion of immigrants in the employment area and the local unemployment rate, as well as indicators for the quarter during which the survey was made, priority neighbourhood and region.
Field: Individuals under 55 with 4 to 31 years of potential work experience, who have obtained qualifications in France ranging from CAP to Master's degree level, have completed their initial education and training, and live in standard accommodation in mainland France.
Source: INED-INSEE, Trajectories and Origins 2 survey (2019–2020).

Protecting Jobs, Preserving Efficiency: Insights from European Short-Time Work Schemes

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Abstract – Short-time work (STW) programmes have been central to European labour-market policy during the Great Recession and COVID-19. This survey integrates theory, cross-country institutional design, and microeconometric evidence to assess whether and how STW succeeds in stabilising employment, preserving firm-specific human capital, and mitigating employees' loss of earnings. Comparative analysis of Belgium, France, Germany, and Italy shows that targeting temporary shocks, combining monitoring with financial incentives for beneficiary firms, and limiting programme duration are crucial to maximise benefits. By contrast, untargeted or prolonged STW not only generates deadweight losses but also delays necessary labour reallocation. Well-designed programmes support firms and workers efficiently while preserving labour-market adjustment.

JEL: E24, J22, J23, J63, J65
Keywords: Short-time work, labour hoarding, employment, firm survival, unemployment insurance

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Short-time work (STW) schemes subsidise temporary reductions in working hours to help firms adjust to transitory—generally unexpected—shocks without resorting to layoffs.[1] In contrast to standard unemployment insurance, STW preserves existing job matches and limits the destruction of firm-specific human capital. By stabilising employment and incomes, STW reduces inflows into unemployment, alleviates fiscal pressure on unemployment insurance systems, and contributes to aggregate demand stabilisation through support for household consumption (Dengler *et al.*, 2025). These schemes may also have distributive effects, particularly during large shocks that disproportionately affect lower-wage workers (Christl *et al.*, 2022).

STW gained prominence during the Great Recession and reached unprecedented levels during the COVID-19 crisis. In April 2020, roughly 15% of employees in Germany, 30% in Italy, 35% in France, and 30% in Belgium were enrolled in STW schemes (Hijzen & Salvatori, 2022). The sudden collapse in labour demand created a high risk of widespread layoffs and bankruptcies, prompting governments to mobilise and expand STW policies on a massive scale.

However, this rapid expansion revived long-standing concerns regarding programme design. Broad eligibility rules, weak monitoring, and limited financial responsibility for firms can generate moral hazard, encouraging overstated output losses, excessive reductions in hours, or enrolment of workers without genuine hour reductions. These behaviours can produce fiscal costs and distort labour reallocation by maintaining non-viable jobs for too long. While the precise challenges of coping with these distortive firm behaviours vary across contexts, this issue has received limited attention in previous surveys. In particular, even if the objective of STW programmes designed for economic reasons is not to protect against expected/seasonal or structural shocks, this is an unintended consequence in many settings, which enlarges the fiscal cost and compromises effectiveness. We return to these questions in the rest of the article, drawing on recent empirical evidence.

This article provides a systematic analytical review of STW as a labour-market policy instrument in Europe, focusing on four countries: Belgium, France, Germany, and Italy. This group is particularly relevant for several reasons. First, they have long-standing and institutionalised STW schemes repeatedly mobilised in response to cyclical downturns. Second, they share relatively stringent employment protection legislation that restricts external flexibility, while hours and wage-setting rigidities heighten the importance of internal adjustment tools such as STW. Third, they experienced large yet heterogeneous expansions of STW during both the Great Recession and the COVID-19 crisis, offering a rich comparative framework for examining the effects of the institutional framework on the adoption and effectiveness of such measures.

Our contribution is threefold. First, we present a conceptual framework for understanding how STW operates within unemployment insurance systems in countries characterised by strong employment protection and labour-market rigidities. We highlight the expected benefits of STW—employment preservation, income stabilisation, and support for aggregate demand—and the potential inefficiencies arising from information asymmetries and moral hazard, which can generate fiscal costs and distort labour reallocation.

Second, we adopt a comparative institutional perspective to document the evolution of key design features—eligibility rules, conditionality, working-time arrangements, and co-financing mechanisms—during non-recession periods, the Great Recession, and the COVID-19 crisis in Belgium, France, Germany, and Italy. This perspective allows us to examine how differences in institutional architecture shape programme utilisation, targeting, and effectiveness, and to provide a systematic view of the policy trade-offs in high-employment-protection environments, both during aggregate economic downturns and in periods of lower programme use.

Third, we connect theoretical mechanisms, institutional design, and empirical findings by analysing three policy instruments that have received comparatively less attention in existing surveys: monitoring, co-financing mechanisms, including co-payments and experience-rating premia, and limits on the maximum duration of support. We further discuss how efforts to strengthen programme targeting through tighter

1. These programmes are also referred to as short-time compensation schemes. In several European countries, such as Belgium, France, and the Netherlands, STW schemes are also known as "temporary unemployment" or "partial unemployment". They should not be confused with temporary layoffs, which are common in the United States and Canada and allow employers to temporarily suspend the employment relationship. In principle, temporary layoff policies enable dismissed workers to be rehired by their original employers (Feldstein, 1976). Temporary layoffs differ from STW schemes because STW maintains the employment relationship continuously. STW programmes should also be distinguished from furlough schemes, such as the one used in the United Kingdom, which allow for a full suspension of work. While some STW schemes may exceptionally permit a complete suspension of activity, they differ from furloughs in their ability to accommodate partial reductions in working time while keeping the employment relationship intact.

eligibility criteria and enhanced monitoring may compromise the timely delivery of assistance and increase administrative burdens. Finally, we highlight that employer co-financing and duration limits can discourage misuse and limit labour hoarding and reallocation inefficiencies while preserving the programme's ability to respond to temporary shocks. Drawing on recent empirical contributions, we show how these instruments can shape programme utilisation, improve targeting, and mitigate labour allocative distortions.

The remainder of the article is organised as follows. Section 1 develops a conceptual framework that synthesises existing research on STW, clarifying how these schemes operate within unemployment insurance systems, support employment and income stabilisation, and may generate fiscal and allocative inefficiencies arising from information asymmetries and moral hazard. Section 2 provides a comparative analysis of the institutional architecture and policy developments of STW schemes in Belgium, France, Germany, and Italy, documenting core design features during stable economic periods and tracing their evolution during the Great Recession and the COVID-19 crisis. Section 3 reviews empirical evidence on the effects of STW on firms and workers, drawing primarily on microeconometric evaluations from a broader set of European countries. Section 4 distils lessons for future policy design, with particular attention to targeting mechanisms, through monitoring and co-financing, and to the role of binding limits on maximum programme duration, then we conclude.

1. Conceptual Framework of STW

This section develops a conceptual framework to understand STW schemes in European labour markets. We first explain why STW is a necessary component of social insurance systems characterised by strict employment protection legislation and rigid wage and working-time agreements. Second, we describe the insurance value of STW for firms, workers, and society. Third, we discuss the main trade-offs involved in STW schemes, highlighting how the programme's insurance function may generate inefficiencies, including fiscal externalities that result in deadweight costs and distortions in labour reallocation.

1.1. Rationale for STW Inclusion in the Social Insurance System

This subsection illustrates why STW is a necessary component of European social insurance systems, given institutional and financial constraints on labour adjustments during unexpected temporary shocks.

Several arguments explain why layoffs may become excessive relative to the social optimum during temporary economic shocks (Cahuc & Carcillo, 2011). The following paragraphs show how STW helps limit these excessive layoffs, that is, job destructions that are inefficient because they exceed the socially optimal level. First, when faced with a temporary economic downturn, firms encounter a trade-off between optimising their long-term labour force needs—by retaining more workers than strictly necessary during the crisis—and avoiding short-term cash-flow shortages caused by ongoing labour costs. Short-term financial constraints, such as limited cash flow and the worsening of financial market imperfections during economic shocks, hinder firms' ability to engage in labour hoarding, leading to increased job destruction (Giroud & Mueller, 2017; Melcangi, 2024) and potentially lower firm survival during downturns.

In principle, firms could also adjust labour costs by reducing working hours or lowering wages. However, in many European countries, labour market institutions (LMIs) and collective agreements restrict the scope for such adjustments. Pay cuts may harm employee morale or trigger voluntary quits (Bertheau *et al.*, 2025; Davis & Krolikowski, 2025), while reductions in working hours are often constrained by rigid agreements on hours and employment conditions (Jäger *et al.*, 2024; Villanueva & Adamopoulou, 2022). Germany stands as an exception, where flexible wage bargaining between firms and workers can lead to mutually agreed-upon downward wage adjustments following negative shocks (Brinkmann *et al.*, 2024; Jäger *et al.*, 2022). More generally, collective agreements negotiated at the national or sectoral level tend to limit firm-worker negotiations, reducing the capacity to preserve job-match surpluses and allocate labour efficiently during downturns (Jäger *et al.*, 2023; Acemoglu, 1995; Hall & Lazear, 1984; Giupponi & Landais, 2023).

Another argument concerns the limited incentives for firms to internalise the broader social costs of layoffs, such as increased government spending, lower tax revenues, and long-term unemployment scarring (Boeri & Cahuc, 2023). Standard unemployment insurance programmes typically do not fully align firms' private incentives with these social costs. An experience rating system can address this by linking firms' unemployment insurance contributions to their layoff histories, thus incentivising them to limit excessive dismissals (Blanchard & Tirole, 2008; Cahuc & Malherbet, 2004; Feldstein, 1976).[2] In theory, full experience rating, where firms fully cover the costs of unemployment insurance benefits paid to a firm's former workers, would eliminate excessive dismissals. However, during downturns, financial constraints may make full experience rating suboptimal, as it could worsen firms' financial distress, increasing insolvency risk, and reducing hiring (Cahuc & Carcillo, 2011; Johnston, 2021). These risks arise because the experience rating system would tax firms concurrently with their layoff decisions. Given these limitations, unemployment insurance alone may not be sufficient to curb layoffs, underscoring STW's role in offering firms greater flexibility.

In contrast, strict employment protection laws (EPL) can deter excessive layoffs by restricting and discouraging dismissals. However, maintaining stringent EPL requirements during temporary economic shocks exacerbates the financial strain of firms affected by these shocks (Boeri & Cahuc, 2023). In such contexts, STW provides the necessary flexibility to manage labour costs without the negative impacts of rigid EPL.

Taken together, these institutional and financial limitations reinforce the need for a flexible adjustment mechanism, such as STW, to stabilise employment during transitory shocks.

1.2. Insurance Value of STW During Temporary Economic Shocks

STW schemes aim to preserve employment relationships, providing insurance for both firms and workers during temporary economic shocks. Unlike unemployment insurance, which compensates workers in the event of involuntary job loss, such as a layoff, STW aims to prevent layoffs altogether, sparing both parties from associated costs.

For firms, layoffs result in the loss of firm-specific human capital, severance costs, and the need to rehire and train workers once the economy recovers. For workers, STW provides financial support via government subsidies for hours not worked, while being able to keep their earnings for hours worked, and shields them from long-term earnings losses and re-employment difficulties (Schmieder *et al.*, 2023). Beyond financial effects, layoffs can negatively impact psychological well-being and economic security, with such consequences often intensifying during recessions (Drydakis, 2021; Rodriguez Conde *et al.*, 2025).

These mechanisms suggest that STW can influence *firm-level outcomes* (employment, survival, productivity, and profitability) and *worker-level outcomes* (income trajectories and career stability). Section 3 examines empirical evidence on these expected effects, assessing whether different programme designs deliver the predicted benefits while considering potential trade-offs and inefficiencies (see Section 1.3).

At the aggregate level, STW generates positive effects through three channels: *(i)* a reduced fiscal burden relative to the unemployment insurance alternative, *(ii)* stabilisation of consumption and demand, and *(iii)* spillover effects for non-users. First, by limiting layoffs, STW reduces unemployment insurance claims, easing the fiscal burden from unemployment insurance—a link explored in Section 1.3. Second, by mitigating the risk of job loss and associated income reductions, STW stabilises demand during recessions (Jaravel, 2022). Faced with unemployment risk and borrowing constraints, workers may cut consumption, triggering a contractionary spiral. STW dampens this effect by reducing the incentive for precautionary savings (Dengler *et al.*, 2025). Finally, STW also benefits firms that do not directly make use of it. Within the theoretical framework developed by Cahuc (2024), the mere existence of STW within the social insurance system leads firms to internalise, in their expected profit calculations, the possibility of resorting to it in the event of a future economic shock. This anticipation mechanism thereby reduces the expected risk of job loss for employees of these firms, even in the absence of any actual use of the scheme. STW also mitigates "rat race" dynamics and congestion externalities on the labour market, which arise when job seekers compete for a limited number of vacancies during economic downturns. This mechanism operates through a reduction in the influx of displaced workers into unemployment, which prevents a deterioration in re-employment rates for non-users (Montenegro & Hijzen, 2024).

Overall, these mechanisms show that STW provides multi-level insurance: stabilising firms' labour forces, safeguarding workers' incomes and careers, and mitigating macroeconomic and labour market disruptions.

2. *Experience rating applies within the unemployment insurance in the United States. Unemployment benefits are financed by firms that pay individualised rates that reflect the actual unemployment insurance costs imposed on the system by the firm's layoffs. While these taxes are computed based on the layoff history of a firm, experience rating is paid concurrently (Guo & Johnston, 2021). Empirical evidence by Duggan* et al. *(2023) demonstrates that U.S. firms laid off fewer workers in response to labour-demand shocks during the Great Recession when facing higher layoff penalties.*

1.3. Towards Efficient STW Insurance: Key Trade-Offs

STW schemes provide rapid and flexible support to firms and workers during temporary shocks, stabilising employment and protecting workers' income. However, these benefits come with fundamental trade-offs rooted in information asymmetries between firms and governments. The latter ones cannot perfectly identify which firms genuinely face large, temporary shocks and would otherwise resort to layoffs without incurring slow and costly verification procedures (Burdett & Wright, 1989; Jaravel, 2022). Moreover, public authorities often struggle to distinguish temporary, unexpected shocks from predictable seasonal (Cahuc & Nevoux, 2018) or structural fluctuations (Mosley & Kruppe, 1996), leading to inefficiencies when firms outside the programme's intended target are nevertheless protected. These informational frictions lie at the core of potential inefficiencies, which we discuss in the following section.

1.3.1. Information Asymmetry and Moral Hazard

The information asymmetry generates moral hazard: firms may strategically manipulate eligibility criteria or overuse STW benefits by excessively reducing working hours (Burdett & Wright, 1989; Jaravel, 2022), even when financially stable or growing. This mirrors the moral hazard problem observed in standard unemployment insurance, but here it is firms' behaviour, rather than workers', that is affected.

Firms may also benefit from "cross-subsidization," receiving support despite not facing genuinely temporary shocks. For instance, firms with seasonal fluctuations or structural problems may enrol workers or reduce hours excessively, thereby benefiting disproportionately from STW without facing a genuine temporary and unexpected decline in activity (Boeri & Cahuc, 2023; Cahuc & Nevoux, 2018). Some countries explicitly limit such misuse (i.e., Switzerland) or establish separate schemes to prevent it.

Misreporting hours of work or enrolling workers without actual reductions further illustrates these behaviours. In Germany, Bossler *et al.* (2023) provide evidence of such practices: employees on STW often worked more than reported and, in some cases, did not experience any reduction in working time during the COVID-19 crisis. In practice, high administrative verification costs and imperfect government information amplify the moral hazard problem.

1.3.2. Fiscal Externalities and Deadweight Costs

Moral hazard generates fiscal externalities: some public resources are allocated to firms that would not have laid off their workers even without the STW programme, creating deadweight fiscal costs (Burdett & Wright, 1989; Cahuc, 2019; Van Audenrode, 1994). In other words, public funds are spent without producing additional social benefits.

Deadweight fiscal costs may also arise when STW preserves structurally non-viable jobs that would be destroyed in the medium term regardless of programme take-up. Moreover, cross-subsidisation can increase the overall fiscal cost, as firms that do not face genuine temporary shocks but rather predictable seasonal downturns may benefit disproportionately (Cahuc & Nevoux, 2018).

Empirical illustrations of these fiscal externalities include the following. First, during the COVID-19 crisis in France, Lapeyre (2023) documents misreporting of firms' eligibility status to access more generous STW benefits. Second, simulations from a dynamic search-and-matching model by Albertini *et al.* (2022) indicate that a 10 percent increase in employer contributions to the most generous STW programme used at the peak of the crisis in France would have reduced demand for the scheme without harming aggregate employment, highlighting that overly generous subsidies induced excessive hour reductions and generated substantial fiscal externalities, thereby increasing the overall cost of the programme. Third, Bossler *et al.* (2023) document free-riding in Germany during the COVID-19 pandemic, with firms misreporting hours and generating potential deadweight fiscal costs.

Few studies have analysed the net fiscal cost of STW relative to traditional unemployment insurance in the absence of STW. During the Great Recession in Italy, each euro spent on STW cost society €1.38, lower than unemployment insurance alone (€1.5–2.5 per marginal euro) (Giupponi & Landais, 2023). In Switzerland, STW's fiscal benefits outweighed its costs, meaning that the programme effectively paid for itself (Kopp & Siegenthaler, 2021). This cross-country variation underscores the importance of institutional design. Consistent with this evidence, Stiepelmann (2026) shows, using a calibrated search-and-matching model, that a system combining unemployment insurance with an optimally designed STW scheme is fiscally more efficient than one relying on unemployment insurance alone. Indeed, STW limits job separations which, under a pure unemployment insurance regime, generate negative fiscal externalities by imposing a collective cost on public finances.

1.3.3. Labour Market Reallocation Distortions

Beyond fiscal externalities, STW can also affect labour reallocation, a key driver of aggregate productivity (Cahuc *et al.*, 2014; Cahuc, 2019). Just as moral hazard and cross-subsidies create deadweight fiscal costs by allocating resources to firms that do not need support, inefficiencies in labour allocation arise when STW preserves jobs that would naturally have been reallocated in the absence of the programme.

By stabilising employment, STW may maintain temporarily low-productive but viable jobs, allowing firms to ride out shocks without resorting to layoffs (Boeri & Bruecker, 2011; Cooper *et al.*, 2017; Hijzen & Venn, 2011). However, supporting structurally non-viable jobs can slow the transition of labour to more productive firms, generating allocative inefficiencies (Mosley & Kruppe, 1996; Giupponi & Landais, 2023; Cahuc, 2024). The extent of these distortions depends on both the duration and the nature of the shock.

In the case of persistent recessions, reallocation inefficiencies may be amplified, as STW subsidises jobs that would otherwise have disappeared, crowding out vacancies and slowing labour mobility towards higher-productivity jobs. As a result, STW can increase hiring costs for more productive firms and for new entrants, tightening the labour market (Cooper *et al.*, 2017; Giupponi & Landais, 2023; Jaravel, 2022). Importantly, reallocation inefficiencies may also arise in short-lived and sector-specific shocks, as STW can crowd out the voluntary labour hoarding that firms would otherwise undertake, given that incentives to hoard labour endogenously are stronger during brief downturns in the absence of STW (Diaz *et al.*, 2025). This mechanism may amplify adverse effects on sectoral reallocation and output losses. Moreover, STW may impede reallocation even more during short recessions, as the discouraging effect of the scheme on workers' job search in declining sectors is stronger when aggregate uncertainty is lower, further reducing incentives for low-surplus workers to move towards expanding sectors. Finally, the COVID-19 crisis introduced additional complexity due to heterogeneous sectoral impacts and accelerated structural shifts, such as the rise of remote work and gig employment (Barrero *et al.*, 2021; Boeri & Cahuc, 2023). In this context, STW may have preserved temporarily viable jobs while hindering the reallocation of structurally non-viable ones, with potential implications for aggregate productivity (Mohimont *et al.*, 2024; Meriküll & Paulus, 2024). However, empirical evidence on whether STW has significantly impeded reallocation remains limited and is still emerging.

In this way, the labour reallocation inefficiencies complement the fiscal deadweight costs discussed in Section 1.3.2: both represent indirect social costs of STW arising from moral hazard and programme design, highlighting the trade-off between providing insurance and maintaining STW efficiency.

1.4. Asymmetric Contract Protection

In addition to fiscal externalities and labour reallocation distortions, STW programmes can cause inefficiencies through unequal protection across different types of contracts. By design, STW often favours open-ended contracts with high replacement costs, leaving fixed-term or low-cost workers underserved (OECD, 2021; Cahuc & Carcillo, 2011; Cahuc, 2019). During large temporary shocks, firms typically adjust by not renewing fixed-term contracts and/or by laying off workers with weaker attachment to the firm. In this context, STW reinforces incentives to place workers on open-ended contracts into the scheme, as these workers face higher dismissal costs and stronger employment protection.

This asymmetry can exacerbate labour market segmentation and limit the movement of workers across contract types, hindering transitions from fixed-term to permanent positions (Hijzen & Venn, 2011). Asymmetric contract protection represents a form of allocative inefficiency: certain workers benefit disproportionately from insurance while others remain exposed, potentially slowing overall labour market adjustment and reducing aggregate productivity.

Evidence from Italy confirms that STW predominantly protected open-ended contracts during the Great Recession (Giupponi & Landais, 2023), although other studies find no clear reduction in fixed-term contract use (Kopp & Siegenthaler, 2021). Recognising these asymmetries is important for understanding the full scope of trade-offs in STW design and for ensuring that the programme balances insurance provision with efficiency across the entire labour market.

1.5. From Theory to Practice and Evidence

The conceptual framework developed in Section 1 highlights why STW emerges as a key component of social insurance systems characterised by strict employment protection, limited internal flexibility in hours and wages, and financially constrained firms facing temporary shocks. When firms cannot easily adjust labour costs without resorting to layoffs (Section 1.1), STW provides a mechanism for smoothing temporary fluctuations in labour demand while preserving valuable job matches. As Section 1.2 discusses, this insurance tool benefits firms, workers, and the broader economy by stabilising employment, protecting incomes, and reducing public expenditure relative to unemployment insurance.

Yet Section 1.3 also makes clear that these benefits come with fundamental trade-offs. Information asymmetry between firms and policymakers creates scope for moral hazard behaviours, including misreporting and cross-subsidisation, generating fiscal externalities and deadweight losses. Similarly, by preventing separations, STW may distort labour reallocation, particularly when shocks are persistent or heterogeneous, and may strengthen existing labour market segmentation by disproportionately protecting open-ended contracts. These inefficiencies arise directly from the same design features that make STW valuable during temporary shocks, underscoring the challenge of calibrating the programme to offer insurance while preserving incentives for efficient adjustment.

Translating these theoretical insights into practice requires examining how countries design and implement their STW systems. Institutional arrangements—such as eligibility rules, monitoring procedures, employer co-financing, benefit generosity, and maximum duration—not only determine the extent to which firms use the scheme but also shape the severity of moral hazard, the size of fiscal externalities, and the potential distortions to labour reallocation. For instance, stricter targeting mechanisms may reduce deadweight costs but risk excluding financially constrained firms that truly need support, whereas more generous and broadly accessible schemes may stabilise employment at the cost of efficiency.

2. STW Design and Policy Changes During Major Shocks

It is essential to understand how STW schemes are designed in order to interpret variations in their effectiveness from one country to another. This section examines the institutional design and policy evolution of STW programmes in four countries, Belgium, France, Germany, and Italy, in response to major economic shocks, namely the Great Recession and the COVID-19 pandemic. We first outline the core design features of STW schemes—eligibility rules, conditionality requirements, working-time arrangements, and co-financing mechanisms—and discuss how they support STW's aim to stabilise employment during temporary downturns. We then document how these features evolved across the two crises, emphasising both common adjustments and crisis-specific innovations. Understanding the structure and adaptability of these schemes is essential for interpreting cross-country differences in programme effectiveness and the inherent trade-offs between employment protection, fiscal costs, and labour reallocation.

2.1. Design Features of STW Programmes

Across countries, STW programmes generally share similar design features. These are eligibility criteria for workers and firms, follow-up conditions upon enrolment, varying regimes of working time reduction, and different co-financing levels by participating firms and workers (Cahuc & Carcillo, 2011; Giupponi *et al.*, 2022). These features define how STW operates under stable business cycle conditions, prior to any crisis-specific adjustments. They are designed to address firm- or sector-specific idiosyncratic shocks. We summarise the main design features of STW linked to economic downturns for Belgium, France, Germany, and Italy in the pre-COVID-19 period (see Table).

Eligibility criteria are a set of rules for firms and workers to qualify for and gain access to STW. Workers, whether on open-ended or fixed-term contracts, may be required to meet minimum past employment requirements in some countries (i.e., akin to eligibility for unemployment insurance). Firms must generally demonstrate a temporary decline in revenues or demand to qualify, driven by cyclical economic downturns or, in a few countries, by anticipated seasonal fluctuations. We notice that the German STW system is the only one in which eligibility for firms is tied to specific thresholds of working-time reduction. By contrast, the Belgian STW system features lax eligibility criteria for firms.

The take-up of STW policies typically rises during aggregate economic crises. However, in some countries, STW can also be activated to cope with predictable seasonal fluctuations in sales or profits linked to the economic activity of specific industries (Belgium and France) or structural shocks (Italy). By contrast, Italy has a distinct STW scheme to address predictable seasonal declines in revenues, whereas Germany has two distinct STW schemes to tackle, respectively, anticipated seasonal fluctuations and structural downturns.[3]

Conditionality requirements are designed to maintain employment relationships while preserving workers' employability. Firms are expected to retain jobs during the STW period, and workers may be required to participate in training or upskilling programmes. These measures ensure that firm-worker matches are preserved during temporary shocks, facilitating smooth reintegration once business conditions recover.

Working-time arrangements under STW are operated via the maximum potential duration of STW support and the modality of hours reduction, whether full or partial suspension. These two features are most often linked: full suspensions of activity generally correspond to lower maximum duration of STW support. Setting a maximum duration reinforces the temporary nature of STW support and its core function: subsidising labour hoarding, which is itself a transitory phenomenon. Firms retain workers during downturns in anticipation that the job match will become profitable again once business conditions improve (Cahuc *et al.*, 2014). However, prolonged usage may restrain the reallocation of workers to more productive sectors and firms, in particular in times of economic recovery, when business conditions improve. Therefore, restricting the programme's duration may thus curtail potential adverse effects of STW on labour reallocation, which is among the social costs of STW discussed in Section 1.3. We observe that Belgium and Italy are the countries with the longer maximum potential duration of STW support.

Co-financing mechanisms allocate costs among workers, firms, and the government. Workers generally receive compensation for lost hours at statutory replacement rates often equal to or higher than standard unemployment insurance, often subject to statutory caps. Firms may contribute via co-payments, experience-rated contributions, or sectoral agreements.[4] These mechanisms aim to reduce overuse and moral hazard. We learn that Italy presents the highest statutory replacement rate for workers, whereas Belgium features the lowest one. Belgium and Italy are the sole countries in which an experience rating system for employers applies.

2.2. Design and Policy Changes Across Crises in Belgium, France, Germany, and Italy

A notable feature of STW programmes is their built-in flexibility, allowing adjustments to eligibility rules for programme access, working-time arrangements, conditionality rules, and co-financing structures in response to major aggregate fluctuations in demand.

For instance, easing STW access when major crises hit aligns with STW's objective of delivering timely support to firms and workers during economic downturns. Its streamlined procedures for application, approval, and payment enable a swift reduction of labour costs for firms while ensuring that workers receive prompt income compensation. The speed of STW implementation is therefore crucial to achieving its stabilising objectives and preventing unnecessary job losses. Another example of policy flexibility is that employers' co-financing generally decreases during acute downturns and gradually rises as the economy recovers. As discussed in Section 1.2, the rationale is that, during unanticipated aggregate shocks, firms typically face greater financial market frictions and more constrained access to credit, necessitating stronger government support (Cahuc & Carcillo, 2011). By contrast, reinstating financial incentives as soon as the economy recovers creates incentives for firms to exit prolonged, and potentially no longer warranted, use of the programme.

Belgium, France, Germany, and Italy—whose STW schemes have been embedded in their social insurance systems since the first half of the twentieth century (Cahuc, 2024)—benefited from mature administrative infrastructures that enabled swift policy modifications during both crises. This institutional maturity stands in contrast to countries that introduced STW programmes only in response to the Great Recession.

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3. The Belgian and French STW systems allow for take-up linked to predicted seasonal fluctuations in addition to cyclical downturns (Table). By contrast, the German and Italian STW systems operate distinct schemes for seasonal shocks—namely Baugewerbetarif *in Germany and* Cassa Integrazione Guadagni Ordinaria *in Italy—separate from the instruments used for cyclical downturns. Additionally, Germany operates a scheme to tackle structural firm downturns (*Transferkurzarbeitergeld*).*

4. Co-payments generally rise with reduced hours and wages and are paid concurrently with STW use. By contrast, experience rating, like the system for unemployment insurance in the US (Guo & Johnston, 2021) or disability insurance in Europe, imposes deferred, progressive premiums based on past STW uptake. We discuss the experience rating system in STW in Section 3.

Table – **Core features of STW programmes in Belgium, France, Germany, and Italy (stable business cycle conditions before COVID-19)**

	Belgium	France	Germany	Italy
Programme	*Chômage temporaire*	*Activité partielle*	*Kurzarbeit*	CIGS
Eligibility	Initially designed for blue-collar workers; firms self-declare a temporary decline in activity, including seasonal fluctuations. White-collar workers were progressively included after 2008, but eligibility requires sectoral or firm-level agreements and verifiable evidence of significant economic difficulties, making access more stringent than for blue-collar workers.	Firms must demonstrate a temporary slowdown in activity, including seasonal fluctuations or structural needs (reorganisation or restructuring). Access is often linked to sectoral or firm-level agreements.	Firms must show a temporary reduction in working hours affecting a minimum share of employees (pre-crisis: 30% of workforce, ≥30% hours reduction).	Firms experiencing economic shocks, company crises, restructuring, or liquidity problems, subject to firm-size thresholds (≥15 FTE) and sectoral eligibility.
Replacement rate	65% of gross daily wage (capped).	70% of the reference wage (floor at the minimum wage).	67% of net wage (capped), with higher rates for employees with children.	80% of wage (capped).
Co-financing	Experience rating for blue-collar workers based on STW intensity (110 days per worker per year), with delayed payments to avoid liquidity pressures; additional social security contributions and optional sectoral or firm-level top-ups.	Government contribution of €7.74 per hour not worked for firms with fewer than 250 employees and €7.23 for firms with 250 employees or more; employers cover the remainder to reach the 70% replacement rate. Since 2013, employer cost is zero for gross hourly wages below €11.	Employers pay social security contributions on unworked hours.	Employer top-ups, experience rating at the firm level after 9 months of STW use, and social security contributions.
Working-time arrangements	Partial or full reduction in days per week; maximum duration of 3–12 months for partial reduction and 4 weeks for full reduction. Firms must resume activity for one week before reusing STW during the same year.	Partial or full reduction in hours for a maximum of 6 months per year.	Partial or full reduction in hours; maximum duration of 6 continuous months, with no yearly limit.	Partial or full reduction in hours; maximum duration of 36 months over a 5-year period.

Sources: OECD (2020); Cahuc & Carcillo (2011); Giupponi *et al.* (2022).

We summarise the main design features and policy changes in Belgium, France, Germany, and Italy during the Great Recession (2009) and the COVID-19 crisis (March 2020–June 2021), highlighting both shared patterns and country-specific innovations (see Table S1 in Online Appendix—link to the Online Appendix at the end of the article). The following sections examine each core design feature—eligibility criteria, conditionality requirements, working-time arrangements, and co-financing mechanisms—and describe how they were modified in response to these crises, linking the discussion to the mechanisms highlighted in Section 1.

2.2.1. Eligibility Criteria

During the Great Recession, eligibility was generally limited to firms demonstrating temporary economic need and, in some cases, structural needs, such as planned reorganisation (Italy). Firms in Italy faced additional restrictions based on size and specific industry/sectors. Germany's *Kurzarbeit* STW programme extended access to firms with at least 10% of the workforce experiencing a minimum 10% drop in working time relative to a pre-crisis reference period. Eligibility was thus enlarged as the pre-crisis thresholds corresponded to 30% of the workforce. As of 2009, France launched the *Activité Partielle de Longue Durée* (APLD) for firms to cope with shocks expected to be longer than the general short-term nature of STW. Workers' eligibility was typically tied to contract type (temporary or permanent) and working-time arrangement (full-time or part-time) and prior entitlement to unemployment insurance. In Belgium, white-collar workers gained temporary eligibility in 2009, with permanent eligibility established in 2012.

During COVID-19, access was significantly eased through crisis-specific regimes: Belgium implemented the Force Majeure Corona scheme, France reinstalled the APLD as of July 2020, and Italy introduced CIG-Covid. Eligibility was virtually extended to almost all firms of all sizes, and in some cases proof of economic need was no longer required. Retroactive applications were allowed in France and Italy. In Germany, the threshold for workforce reduction to access STW was lowered from 30% to 10%, as during the Great Recession. For workers, coverage was extended to temporary agency workers. These eligibility expansions directly reflect the insurance role of STW highlighted in Section 1, aiming to preserve firm–worker matches and maintain employment even under widespread shocks, while also reducing labour market segmentation by covering previously excluded workers.

Unlike conventional STW, which addresses temporary shocks, the APLD in France targeted firms facing more persistent but non-terminal difficulties and offered exceptionally long support: up to 24 months within a 36-month period (Calavrezo & Walkowiak, 2025). Access was conditional on prior collective bargaining at the firm, establishment, or sectoral level and required explicit commitments to job retention and worker training.[5] In contrast, the conventional STW scheme involved far fewer obligations and no mandatory agreements, highlighting the APLD's more conditional and negotiated approach to employment support.

2.2.2. Conditionality Requirements

During the Great Recession, firms were typically required to maintain employment during and shortly after STW participation, sometimes submitting recovery plans.

During COVID-19, as uncertainty was rising and the shock was becoming more persistent, conditionality increasingly emphasised training and skill development. In France, APLD participants had to provide firm-specific training during non-working hours, reimbursed by the government. Optional firm-specific training was also introduced in Germany and Belgium, focusing on skills directly relevant to employees' current positions rather than general transferable skills. These measures reinforce the mechanisms discussed in Section 1, ensuring that temporarily idle workers retain relevant skills and supporting the preservation of viable jobs.

2.2.3. Working-Time Arrangements

During the Great Recession, STW typically allowed partial or total reductions in working hours, with country-specific maximum durations (e.g., up to 12 months in Germany, six months in France, and 36 months over five years in Italy). In France, the yearly number of subsidised hours per employee rose to 800, or 1,000 hours for sectors most severely affected by the Great Recession, such as textiles and automobile manufacturing.

The **COVID-19** shock prompted substantial extensions: Belgium's Force Majeure Corona scheme had no defined maximum; France's *Activité Partielle* allowed a maximum of STW hours per worker equal to 1,607 per year, well above the extension implemented during the Great Recession; Germany extended Kurzarbeit from 12 to 24 months. As noted in Section 1.3.2, prolonged STW support can, in theory, risk protecting structurally non-viable jobs, potentially hindering labour reallocation and productivity recovery once economic conditions improve.

2.2.4. Co-Financing Mechanisms

During the Great Recession, the financing was shared among workers, firms, and the government. Workers received partial compensation for lost hours, subject to statutory caps, while firms contributed via direct payments, sectoral top-ups, or experience-rated contributions, discouraging overuse, and aligning incentives with employment preservation. For instance, in France, the government subsidy amounted to €3.84 per hour per employee for establishments employing up to 250 workers and €3.33 for those with more than 250 employees. Therefore, the firm was responsible for covering the difference between the benefits paid to the worker and the flat government subsidy.

During COVID-19, employer contributions were temporarily reduced or eliminated in most countries to ease liquidity constraints. In Belgium and Italy, firms faced zero costs under

5. Social dialogue played a key role in France, enhancing transparency and mitigating workers' job insecurity associated with prolonged reductions in hours.

emergency schemes. In France, the government financed compensation up to 4.5 times the minimum wage, with employers covering any excess to maintain a 70% replacement rate. A key distinction was introduced in June 2020 between sectors particularly affected by COVID-19 and those less affected, formalised through the classification of industries into protected and non-protected sectors at the 5-digit level. While all employers remained responsible for compensation above 4.5 times the minimum wage, firms in non-protected sectors were required to co-finance 10% of STW costs below this ceiling, whereas protected-sector firms were exempt from this increase (Lapeyre, 2023).[6] In Germany, firms' payments of social security contributions for unworked hours were reduced by 50%. Across countries, workers' replacement rates increased: Belgium to 70% of gross wages, France up to 84% depending on earnings level, Germany to 70% from the fourth month and 80% from the seventh month (higher for workers with children), and France APLD at 70% of gross wage. These adjustments provided timely income support while ensuring continued employment.

2.2.5. Key Differences Between Crises

During the Great Recession, these four countries were able to easily expand their existing STW programmes that were already part of strict employment protection systems. They had the highest take-up rates in Europe and kept unemployment rates low, which is often seen as a sign of success. Nonetheless, their institutional frameworks exhibited significant differences. First, employer cost-sharing varied markedly. Belgium and Germany reduced employers' costs to zero, whereas France had some contributions, which may help explain France's comparatively low take-up and limited intensity of use. Second, targeting features diverged. Germany set a quantitative threshold—at least 10% of workers having their hours cut by at least 10%—to weed out companies that were only facing mild shocks. Italy limited eligibility to businesses with more than 15 full-time equivalents and left out some industries, which made the distribution of support more selective. Belgium and France, on the other hand, used broad administrative assessments of economic hardship. Third, the adjustment margins available to countries varied. In Germany, decentralised negotiations over pay and working hours provided firms with an alternative mechanism for labour hoarding, reducing reliance on STW and influencing the implementation of Kurzarbeit. Overall, these institutional contrasts shaped take-up patterns and the employment-stabilising role of STW, reflecting differences in policy design.

The COVID-19 crisis triggered similar design adjustments as those observed during the Great Recession, with two notable differences, reflecting the exceptional nature of this latter crisis. First, whereas during the former crisis the relaxation of eligibility criteria still excluded certain firms from access, such as those with more than 15 full-time equivalents in Italy, eligibility during the COVID-19 crisis was extended to virtually all firms in most countries. This difference reflects the changing nature of firms' constraints: during the Great Recession, STW primarily addressed liquidity shortages that limited firms' ability to hoard labour, while during COVID-19, mandatory business closures and complete halts in working time added a non-financial dimension to labour market disruptions. As a result, a broader spectrum of firms, beyond financially constrained small enterprises, became eligible for support.

Secondly, as governments sought to phase out the exceptionally generous support provided in the early stages of the pandemic, countries such as France introduced distinctions between firms in more affected and less affected sectors. These categorisations, based on evolving health and economic indicators, aimed to dynamically recalibrate the scope of support—concentrating resources on genuinely distressed, pandemic-hit sectors while excluding firms no longer in need. This approach sought to reduce negative reallocation inefficiencies and minimise deadweight losses, as discussed in Section 1.3.

3. STW in Practice: Do Designs Deliver the Predicted Effects?

This section reviews key findings on the effects of STW on firms and workers and examines whether different programme designs achieve the outcomes predicted by theory. We focus on recent causal microeconometric evaluations, distinguishing evidence by country and crisis episode (Great Recession vs. COVID-19). These evaluations rely on quasi-experimental designs to account for non-random take-up (i.e., selection bias) and ensure credible causal inference.

6. Below the cap, employers were now required to contribute proportionally, covering 10% of the worker's gross hourly wage, while the government financed the remaining 60%. As under the previous arrangement, for wages above the 4.5 minimum wage cap, the employer bore the full cost of STW compensation, ensuring continuity of the 70% replacement rate.

From the firm perspective, STW enables labour hoarding, supporting survival, particularly for liquidity-constrained firms, and helps preserve human capital, which can influence future investment in both human and physical capital. Accordingly, STW may affect employment, survival, profitability, and productivity over both short- and medium-term horizons.

From the worker perspective, STW provides insurance against *(i)* income loss from reduced working hours and *(ii)* unemployment scarring, which can have persistent effects on earnings trajectories and employment probabilities. The policy thus shapes individual labour-market outcomes in both the short and medium term.

Early cross-country studies on the Great Recession suggest that STW stabilised aggregate employment during temporary downturns (Boeri & Bruecker, 2011; Cahuc & Carcillo, 2011; Hijzen & Venn, 2011; Hijzen & Martin, 2013; Brey & Hertweck, 2020). While these analyses partially account for selection bias, they cannot fully separate country-specific features or heterogeneous crisis effects, which may jointly affect STW take-up and employment.

In contrast, causal microeconometric studies exploit firm-level variation within a fixed institutional setting to estimate causal impacts. Because programme take-up is non-random, naïve comparisons risk conflating STW effects with pre-existing firm characteristics. Credible counterfactuals are therefore essential. Although randomised experiments are rare, quasi-experimental approaches enable robust estimation at both firm and worker levels. Nevertheless, early evaluations of the effects of STW during the Great Recession reached mixed conclusions, with results heavily influenced by the identification strategies employed (Bellmann & Gerner, 2011; Calavrezo *et al.*, 2010).

The following subsections first examine firm-level outcomes, distinguishing short-term effects—observed during STW take-up—from medium-term effects, and exploring heterogeneity by firm characteristics and exposure to shocks. The discussion then turns to worker-level outcomes.

3.1. Effects of STW on the Economic Behaviour of Firms

3.1.1. Employment Effects

STW influences employment along two main margins. On the intensive margin, firms respond to temporary demand and/or revenues shocks by adjusting labour demand at the intensive margin, reducing hours per worker. On the extensive margin, STW can preserve jobs that might otherwise be destroyed, generating immediate short-term effects, particularly in firms facing severe demand shocks (Cahuc *et al.*, 2021), thereby stabilising overall employment. However, deadweight losses may arise if jobs that would have survived without intervention are subsidised, which are likely concentrated in firms facing mild demand shocks.

Over the medium term, employment gains are likely to persist, as firms avoid the costs of rehiring and hours per worker gradually recover, given that the scheme is intended to be temporary. However, when short-term gains dissipate, STW may delay the exit of non-viable jobs, potentially hindering long-term employment growth in both treated and untreated firms and slowing labour reallocation (a summary of these expected effects is presented in Table S2 in Online Appendix).

Some micro-econometric evaluations for the Great Recession (see Table S3 in Online Appendix for a summary) provide robust empirical evidence consistent with these mechanisms. These studies rely on credible identification strategies that exploit exogenous variation generated by institutional features of STW programmes using large administrative datasets. Taken together, these papers highlight a common pattern: STW reduces hours per worker, while its employment effects depend critically on the severity of firms' demand shocks, programme targeting, and how persistent the aggregate downturn was.

In France, Cahuc *et al.* (2021)[7] use administrative discretion in the approval of STW applications as an instrumental variable (IV) for STW take-up among single-establishment firms. For the average firm, STW induces a sharp decline in hours per worker—28 percent in the short term—but no employment gains, except among firms facing large adverse shocks. In those firms, hour reductions translate into substantial increases in both headcount and total hours, consistent with the view that STW provides insurance precisely when jobs are truly at risk. In particular, firms with a predicted growth of hours of work below the median—estimated based on pre-crisis firm characteristics and used by the authors as a proxy for large negative demand shocks—that took up STW in

7. The effect sizes reported in the text are drawn from Cahuc et al. *(2021). A revised version appears as Cahuc* et al. *(forthcoming). The main conclusions are unchanged, though some estimates differ slightly across versions.*

2009 experienced, in the short run (i.e., during the first year of take-up), a 15 percent decline in hours per worker, accompanied by increases of 42 and 34 percent in employment and total hours worked, respectively. Yet, much of the French STW use occurred in firms experiencing mild shocks, generating deadweight losses, or what the authors refer to as large windfall effects. The low costs employers face when using STW—particularly for protecting low-wage workers, who are largely concentrated in sectors that make recurrent use of the programme—may further help explain these windfall effects (Cahuc & Nevoux, 2018).[8]

In Belgium, Bermúdez *et al.* (2025) exploit variation in firm-level STW take-up stemming from differential eligibility rules for blue- and white-collar workers (see Table). Using grouped administrative data for firms with 5 to 50 employees, they examine the effects of STW on employment and wages. Consistent with the French evidence, STW primarily reduces hours per worker without increasing average employment or wages. However, the effects vary sharply across sectors. In manufacturing—the sector most exposed to the crisis—one additional FTE of STW per firm corresponds to a 2.3 percent reduction in hours per worker and a 5 percent increase in headcount employment. These short-term gains disappear after one year, and no positive effects emerge in less affected sectors, implying substantial deadweight fiscal costs when programme access is broad. This evidence further suggests that STW may have delayed the exit of non-viable jobs that entered the programme at the onset of the crisis.

In Italy, Giupponi & Landais (2023) exploit STW eligibility rules related to firm size and sectors among firms with 15–25 FTEs (see Table). STW induces a 40 percent reduction in hours per worker and a 45 percent increase in headcount employment in the short term, although these effects fade after the first year of take-up. Their analysis also reveals strong heterogeneity across firm characteristics: firms facing liquidity constraints benefit the most, as STW eases short-term financial pressures and allows employment to rebound once activity resumes (Giroud & Mueller, 2017; Melcangi, 2024). Interestingly, although most STW users were low-productivity firms prior to the crisis, medium-term employment gains per hour reduced are concentrated among high-productivity firms. By contrast, very low-productivity firms exhibit no such gains in the medium term—even though they reduce hours per worker more extensively—consistent with Biancardi *et al.* (2022). These findings underscore the dual role of the programme: it can stabilise viable matches but may be less effective—or even distortive—when it ends up protecting a substantial number of non-viable jobs, as in Belgium at the onset of the crisis.

Kopp & Siegenthaler (2021) analyse the impact of STW during the Great Recession in Switzerland using an event study difference-in-difference (DID) design. They compare changes in the outcomes of successful and unsuccessful applicants, examining both average and dynamic effects around the event of application to the programme. They document sizeable short- and long-term employment increases (9–17% up to 4.5 years after application). Switzerland's strict monitoring and selective access—largely restricted to firms facing temporary cyclical shocks—are likely drivers explaining the programme's effectiveness and limited deadweight costs, especially given the country's short-lived recession.

Evidence for the immediate aftermath of the Great Recession in Germany (Brinkmann *et al.*, 2024) suggests that, in an institutional environment characterised by substantial internal wage flexibility, extensions of STW duration limits generated deadweight losses by preserving jobs that would have survived through wage bargaining alone, albeit at the cost of lower wages per employee. The authors exploit a policy-induced regression discontinuity design arising from the 2012 extension of the maximum STW duration. Comparing firms just above and below the eligibility threshold, they find no employment effects, as firms instead adjust wages downward when the potential duration of STW benefits tightens.

The Great Recession experience shows that STW consistently reduces hours per worker, but employment gains arise primarily in firms facing severe, temporary shocks or liquidity constraints. Programmes with strict targeting and monitoring, as in Switzerland, generate positive employment effects, whereas broadly accessible schemes risk substantial deadweight losses. Moreover, these effects tend to be short-lived, particularly when firms subsequently undertake significant structural adjustments.

The COVID-19 pandemic provides a distinct context: an unprecedented and abrupt collapse in economic activity, accompanied by substantial

8. In practice, although co-financing requirements were in place in France during the Great Recession, the cost to employers of enrolling low-wage workers in STW was substantially lower than for high-wage workers, which in turn induced higher take-up among low-wage workers.

STW expansions, simplified eligibility rules, and large-scale programmes. The existing empirical evidence (see Table S3 in Online Appendix) remains preliminary, and credible causal identification is particularly challenging given the near-universal nature of both the shock and the policy response.

Moreover, evidence from the COVID-19 period extends beyond traditional STW programmes to include furlough schemes implemented in several countries to fully suspend working time. A number of studies (see Table S3 in Online Appendix) document that these furlough schemes generated substantial short-term employment protection without raising major concerns about their overall effectiveness (Bennedsen *et al.*, 2023; Beņkovskis *et al.*, 2023; Montenegro & Hijzen, 2024; Meriküll & Paulus, 2023). Other contributions—although not always based on micro-econometric identification—highlight that large programme expansions entailed significant deadweight losses due to firms' behavioural responses, thereby increasing the risk of inefficient allocation of support as the pandemic evolved (Albertini *et al.*, 2022; Brinkmann *et al.*, 2024; Kagerl, 2024; Dengler *et al.*, 2025).

Bennedsen *et al.* (2023) evaluate Denmark's furlough programme using a selection-on-observables strategy that exploits managers' pre-policy predictions of layoffs and furlough use under a no-policy scenario—collected before programme implementation—as counterfactuals. Combining these survey data with administrative records, they find that firms using furloughs reduced layoffs by about 24 percentage points relative to comparable non-treated firms, which would otherwise have dismissed workers during the first year of the pandemic.

Beņkovskis *et al.* (2023) study the Latvian furlough scheme using a local-projection DID design, documenting positive effects on headcount employment both in the short term and persistently through the end of 2021 for first-wave users relative to pre-pandemic levels and to non-users. Montenegro & Hijzen (2024), using an RDD that exploits variation in programme generosity in Spain, show that each furloughed worker corresponded to roughly one job saved, indicating the absence of short-term deadweight effects. Meriküll & Paulus (2023) evaluate Estonia's COVID-era STW programme through propensity score matching, comparing treated firms with non-users experiencing similar turnover declines. They find that treated firms reduced layoffs relative to their matched counterfactuals, with approximately 20 percent of supported jobs preserved in the short term.

Early micro-econometric evidence from the COVID-19 period therefore indicates that pandemic-era job-retention schemes were highly effective in preventing layoffs, particularly in sectors unable to operate during lockdowns. One key feature stands out: most of the furlough programmes examined above were introduced for the first time during COVID-19, at a moment when an exceptionally large number of jobs were at risk. The absence of prior experience with these schemes, combined with the severity of the shock driven by mandated closures, likely limited their use by firms facing predictable seasonal fluctuations rather than genuine, unexpected demand shocks.

Another set of studies underscores important inefficiencies created by very broad or universal eligibility, high replacement rates, and—in some cases—slow adjustment of programme parameters as the pandemic evolved. These features raised concerns about sizeable deadweight losses (Albertini *et al.*, 2022; Kagerl, 2024) and about the risk of inefficient labour reallocation with potential aggregate productivity consequences (Dengler *et al.*, 2025).

Albertini *et al.* (2022) find that, in France, the STW programme during the pandemic successfully prevented job destruction but also excessively reduced working hours, thereby protecting jobs that would have survived without support. Kagerl (2024) evaluates the German STW programme during the COVID-19 crisis using an event-study combined with coarsened exact matching. The study shows that firms using STW increased employment by 3.5 percent in the short term but experienced no medium-term effects, indicating only temporary stabilisation driven by the protection of a large number of low-productivity job matches. These results also suggest that the net effects on employment remain limited due to significant windfall effects associated with widespread use of the scheme.

Dengler *et al.* (2025) quantify the aggregate effects of Germany's programme using a New Keynesian model with search-and-matching frictions and incomplete financial markets. They highlight several insights. First, although STW resulted in overall productivity losses, these are moderate relative to the sizeable employment benefits during aggregate shocks. Second, when monetary policy is constrained by the zero interest rate lower bound, STW further reduces workers' income risk, dampens precautionary savings, and stabilises aggregate demand by preserving employment. Finally, discretionary relaxations of eligibility criteria or increases in generosity become progressively less effective as they protect more firms that would have not fired even in the absence of the programme, yielding limited additional employment gains.

Overall, the COVID-19 experience illustrates that STW can serve as a powerful emergency stabiliser for employment, particularly when a large share of sectors face severe operational constraints as at the peak of the crisis. However, the effectiveness of such programmes—especially once sectors begin to reopen—depends heavily on programme design and timely adjustments that balance generosity with targeting, thereby limiting support to jobs that are either viable and genuinely at risk of destruction.

3.1.2. Effects Beyond Employment

Beyond employment, STW can shape a wide range of firm outcomes, influencing both short-term adjustment margins and longer-term performance. By subsidising labour hoarding during temporary revenue declines and preserving existing job matches, STW may affect firm survival probability, as well as productivity and investment decisions in the short and medium term.

Giupponi & Landais (2023) find that, in Italy during the Great Recession, STW increased firm survival by 10 percentage points one year after treatment, with stronger effects for liquidity-constrained firms. However, they detect no significant survival effects for less productive firms—defined based on pre-crisis productivity—consistent with the absence of employment effects for this group. Similarly, Kopp & Siegenthaler (2021) show that Swiss firms using STW during the 2008–2009 financial crisis were 5.5–9 percentage points more likely to remain operational 4.5 years later. Evidence on survival during the COVID-19 period remains limited, but Beņkovskis *et al.* (2023) demonstrate that reduced firm exits account for an important share of STW's positive employment effects in Latvia during this episode.

Beyond survival, STW can influence productivity and profitability. By preserving employment, the programme reduces turnover costs and protects firm- and industry-specific human capital. These mechanisms have implications for productivity as well as for future investment decisions and can translate into higher profits when demand recovers. Nonetheless, Giupponi & Landais (2023) and Biancardi *et al.* (2022) find that, although STW boosts short-term employment in Italy during the Great Recession, it temporarily lowers per-employee productivity (proxied by value added per worker) due to fewer hours worked, with no effect on productivity per hour in the short or medium term. Giupponi & Landais (2023) also find no significant impact on physical capital investment in either the short or medium term.

Moreover, Biancardi *et al.* (2022) further examine the effects of STW on firm performance in Italy during the Great Recession, emphasising the role of unionisation. Although STW leads to a modest short-term decline in profits—reflecting that the fall in per capita productivity exceeds the reduction in labour costs—these effects dissipate after STW uptake. Importantly, the degree of unionisation shapes these responses. In highly unionised firms, STW leads to stronger reductions in hours worked per employee but has only limited effects on labour costs, as unions negotiate to prevent wage reductions (e.g., via wage top-up payments) while promoting work-sharing, with no adverse effects on profitability. In contrast, weakly unionised firms, which exhibit smaller reductions in hours worked per employee, may achieve substantial labour-cost reductions through wage negotiations. These reductions are accompanied by larger short-term declines in productivity and profits than those observed in highly unionised firms.

Overall, the evidence shows that STW can improve firm survival and stabilise employment during downturns, particularly for liquidity-constrained firms. Its effects on productivity and profitability are more mixed: reductions in hours tend to lower per capita productivity without affecting hourly productivity in the short term, and there is little evidence of medium-term efficiency gains or increased investment. Institutional features such as unionisation further shape these outcomes by influencing wage and cost adjustments.

3.2. Effects of STW on Worker Outcomes

STW protects workers in two main ways: by providing income insurance for reduced hours and by mitigating the scarring effects of unemployment, which can have persistent consequences for future earnings and employment prospects. These mechanisms generate both short-term and medium-to-long-term impacts on worker outcomes.

In the short term, STW stabilises employment by preventing layoffs, allowing workers to maintain their jobs and labour market attachment. Over the medium term, retained employment can translate into faster earnings recovery and lower risk of long-term unemployment. Broad, untargeted programmes may, however, offer limited gains if they subsidise low-productive jobs that are unviable in the medium term, delaying layoffs.

Assessing the dynamic effects of STW requires comparing STW-insured employees to different counterfactual groups, i.e., the scenarios workers would have faced in the absence of the programme. Recent studies primarily adopt two counterfactuals: *(i)* unemployed individuals and *(ii)* employees similar to STW participants but not covered by the programme.

Some studies from the Great Recession period suggest that STW positively impacts earnings and employment prospects but highlight varying short- and medium-term effects. Tilly & Niedermayer (2016), who calibrate a job search and matching model assuming non-negotiable contracts in terms of hours and wages, find that STW participation in Germany causes a sharp but temporary decline in earnings, albeit less severe than for unemployed individuals. They find no medium-term earnings penalties for STW participants, unlike unemployed workers, due to STW's ability to preserve employment relationships and its higher replacement rates compared to unemployment insurance systems.

Giupponi & Landais (2023) find that, in the short term, workers in STW have the same employment probability one year after programme participation as similar workers employed in non-eligible firms. However, this probability declines over time for STW users. Three years after treatment, the employment probabilities of STW workers converge towards those of similar workers in non-eligible firms who were laid off. Likewise, earnings per employee in STW decline in the short term due to reduced hours and, in the medium term, become comparable to the earnings of similar laid-off workers in non-eligible firms. This suggests that in Italy, STW had no significant medium-term effect on employment prospects, consistent with its lack of impact on firm headcount employment during the Great Recession.

Overall, STW appears to provide only short-term insurance, particularly in Italy during the Great Recession, where the economic shock was more persistent than in other countries and a larger fraction of structurally unviable jobs was treated by the programme. In prolonged downturns, firms eventually resort to layoffs, reducing workers' earnings over time. Moreover, Giupponi & Landais (2023) show that in high-productivity firms (defined prior to the crisis), medium-term earnings of STW participants are significantly higher than those of laid-off workers in comparable firms. This suggests that STW improves medium-term employment prospects only for workers in high-productivity firms.

During the COVID-19 period, Brinkmann *et al.* (2024) leverage exogenous worker-level variation in STW eligibility within firms, exploiting the statutory retirement age cutoff, beyond which individuals become ineligible. They focus on workers in firms that took up STW in April 2020 but had not done so in the previous three months. Using an RDD approach, they compare workers who reached the statutory retirement age up to three months before or in April 2020 (never-eligible workers) with workers in the same firm who reached the statutory retirement age up to three months after April 2020 (eligible workers). They find no difference in the probability of remaining employed at the same firm between eligible and ineligible older workers. In this case, STW is directed toward workers with high baseline retention probabilities. Although specific to older workers—who face higher layoff costs—this finding nevertheless highlights potential deadweight losses from subsidising workers unlikely to be laid off.

Taken together, existing studies provide only a partial view of how STW shapes workers' longer-term employment trajectories. While its short-term stabilising effects are well established, evidence on medium- and long-term outcomes remains limited and highly context-dependent. Most worker-level analyses derive from environments in which STW sustained a substantial number of low-productivity or otherwise non-viable firms; much less is known about systems that more effectively screen such firms out. A fuller understanding of these dynamic impacts—across different types of firms and workers—remains an important direction for future research.

3.3. Main Insights From the Empirical Evaluation of STW

The empirical literature shows that STW delivers its intended short-term employment protection, but its overall effectiveness depends critically on whether support reaches firms at genuine risk of destroying jobs. The largest gains arise in sectors facing severe temporary shocks and among liquidity-constrained firms, where STW reduces hours while stabilising headcount and improving survival. When access is broad, however, programmes frequently subsidise firms that would not have laid off workers, generating deadweight effects, insuring predictable or seasonal fluctuations, and in some cases sustaining structurally weak, low-productivity jobs.

For workers, STW offers immediate insurance by preventing layoffs and limiting income losses. Medium-term outcomes, however, depend on firms' underlying viability: workers in productive, resilient firms experience smoother earnings and employment recovery, whereas those in structurally weak firms see few sustained gains, mirroring the limited medium-term effects observed at the firm level.

These issues were especially visible during the Great Recession, when positive effects were concentrated in severely affected firms, while broad eligibility and low employer costs often increased windfall use. Medium-term outcomes remain highly heterogeneous and sensitive to programme design, the nature of the shock, and firms' underlying fundamentals. During the COVID-19 crisis—although evidence is still emerging—employment gains appear more widespread, particularly in settings where job-retention schemes were introduced for the first time. These programmes protected a large share of jobs during mandatory shutdowns, generating sizeable short-term employment gains. However, in countries with pre-existing STW systems, the exceptional generosity and duration of COVID-era schemes heightened fiscal costs and raised concerns about delaying necessary reallocation.

Taken together, the evidence shows that STW is most effective when shocks are temporary, firms are fundamentally viable, and support is well-targeted and time-limited. Yet this evidence is drawn largely from periods of aggregate downturn. There is little causal evidence on STW effectiveness during more stable periods, primarily because low take-up severely limits identification.

Overall, the findings underscore that the effectiveness of STW critically depends on careful targeting, appropriate duration, and adequate co-financing, ensuring that temporary shocks are mitigated without unnecessarily subsidising firms or jobs that would have survived independently. These insights motivate the next section, which draws lessons for refining STW design to optimise insurance provision while minimising distortions.

4. Lessons for Future STW Design

The experience of recent crises, and especially the COVID-19 pandemic, has revived interest in how STW schemes can be designed to provide rapid stabilisation (Section 1.1) while minimising inefficiencies and fiscal costs (Section 1.3). Despite the central role of STW in countries' social-insurance responses—often exceeding the take-up of unemployment insurance—evidence on how specific design features influence take-up and performance remains limited, particularly compared with the unemployment insurance literature (Schmieder & von Wachter, 2016; Spinnewijn, 2020). This section proposes two core sets of design levers: employer co-financing, such as co-insurance and an experience rating premium, and setting a maximum duration of STW support.

In this section, we first emphasise that an efficient STW design would hinge on screening firms experiencing genuine, temporary shocks to ensure that support is properly targeted. We then discuss how tightening eligibility criteria to strengthen programme targeting and the monitoring of those may compromise the timely delivery of STW assistance and can be costly. Finally, we highlight how employer co-financing and limits on the maximum duration of support can curb misuse and labour-hoarding reallocation inefficiencies while preserving STW responsiveness.

4.1. Targeting and Monitoring

Effective STW requires screening firms to ensure support reaches those experiencing genuine, temporary shocks. Section 3.1 shows that persistent employment and survival gains mainly occur in liquidity-constrained firms or those hit by severe, temporary demand shocks (Cahuc *et al.*, 2021; Giupponi & Landais, 2023; Bermúdez-Barrezueta *et al.*, 2025). Monitoring reduces the risk of subsidising jobs that would have survived (Cahuc *et al.*, 2021) or supporting structurally weak firms (Giupponi & Landais, 2023). Ex-ante criteria typically rely on revenues or demand indicators, while ex-post checks verify actual reductions in working time.

Extensive screening, however, is costly and can slow access, undermining STW's rapid stabilisation role (Boeri & Cahuc, 2023). While stricter eligibility improves targeting, real-time verification is administratively demanding. Empirical evidence suggests that employer co-financing and duration limits can curb misuse more efficiently, at lower administrative cost (Sections 4.2–4.3).

4.2. Co-Financing

Building on the fiscal externalities highlighted in Section 1.3.2, financial tools such as co-payments and experience rating are designed to internalise part of STW's cost and discourage misuse along both the intensive (hours per worker) and extensive (number of workers) margins. In Belgium, the experience rating system applies at the individual worker level, making employer contributions depend on past STW usage per worker per year rather than only on aggregate firm-level take-up. This design has two advantages. First, unlike experience rating within unemployment insurance, tax rates are not only linked to a firm's past STW take-up history but are also deferred and paid in the following calendar year. This design reduces potential liquidity pressures on firms. Secondly, the Belgian experience rating system ensures that firms are accountable for repeated or intensive use while maintaining a fair distribution of reductions across employees. Relative to unemployment insurance, STW promotes a more even distribution of working-time reductions across the workforce rather than concentrating job losses among a few workers (Abraham & Houseman, 2014; Giupponi *et al.*, 2022). In Belgium, the experience rating system, adjusted according to the use of STW per employee, aims to strengthen this redistributive role, which was one of the objectives behind the introduction of this scheme.

When costs are borne directly by employers and rise with STW take-up, excessive use is naturally deterred. By contrast, pooled contributions via sectoral funds or general social security payments weaken this link, reducing incentives for efficient programme use. Because strict ex-ante screening can delay support, financial levies provide a cost-effective alternative. By penalising prolonged or excessive uptake, experience rating systems discourage reliance on STW for structural and seasonal anticipated shocks rather than temporary unexpected downturns.

To avoid straining firms' liquidity, payments can be deferred and smoothed over time (e.g., similar to the structure of the U.S. experience rating system in unemployment insurance; Guo & Johnston, 2021). Levy-based incentives also have distributional implications: labour hoarding primarily benefits workers with high replacement costs, potentially leaving lower-wage or fixed-term workers underserved (OECD, 2021). STW design should consider this to avoid reinforcing labour-market segmentation.

Causal evidence, though limited due to data constraints and endogenous policy design, supports the effectiveness of these measures. Further research will be needed to better understand the causal effect of financial incentives on the use of STW.[9]

4.3. Maximum Duration Limits: Maintaining Focus on Temporary Shocks

STW programmes ought to be limited to short-term temporary economic shocks. Long-term or structural use can hinder the survival of viable businesses, slow labour reallocation, and reduce productivity gains. Research demonstrates that while widespread, untargeted programmes significantly postpone layoffs, and that persistent employment and survival effects primarily occur in businesses experiencing severe but transient shocks (Cahuc *et al.*, 2021; Giupponi & Landais, 2023).

Without the need for extra steps like mandatory registration with employment services, maximum duration limits combined with financial incentives can discourage abuse and lessen administrative burden. These restrictions aid in maintaining STW's status as a stabilising instrument as opposed to a structural support system. However, these restrictions are not without trade-offs: by limiting the duration of use of the scheme, they may lead some employees to move to new firms, which conflicts with the very objective of STW, namely preserving firm-specific human capital and maintaining a strong link between the employee and their employer.

Consistent with this principle, training under STW should focus on firm-specific skills to preserve human capital within the enterprise. General training that facilitates worker mobility outside the firm may be inappropriate within STW, as it encourages structural adjustment rather than temporary stabilisation. General training may become relevant only if the negative shock persists and transitions into a structural problem, beyond the scope of STW.

Overall, STW programmes designed to address temporary problems—complemented by appropriate financial incentives and maximum duration limits—can provide rapid stabilisation, protect workers, and minimise distortions

9. Recent studies have drawn on natural experiments exploiting kinks in employers' co-financing schedules—notably the changes in the Belgian experience rating system (Cockx et al.*, 2026) and the changes in employers' contributions made in France during the COVID-19 period (Lapeyre, 2023). These studies draw respectively on detailed Belgian and French administrative data (at the establishment and employee level) and analyse the impact of financial incentives on the use of STW, both at the intensive and extensive margins.*

(Cahuc *et al.*, 2021; Giupponi & Landais, 2023; Bennedsen *et al.*, 2023; Montenegro & Hijzen, 2024). It is also important to ensure that co-financing requirements and maximum duration limits can be adjusted throughout the crisis period—with regard to co-financing, by reducing it to zero during the downturn, before gradually restoring it as the economy recovers.

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*

This survey has examined STW programmes as a key labour-market policy instrument in European economic downturns. First, we elucidated the mechanisms by which STW offers employment insurance and temporary income support by creating an integrated conceptual framework, emphasising the trade-offs between stabilisation, fiscal costs, and allocative efficiency. This framework assisted in determining how programme design features, including financial incentives, duration restrictions, and targeting, influence the impacts of STW and mediate the trade-off between efficiency and insurance.

Second, comparative analyses from Belgium, France, Germany, and Italy during the Great Recession and COVID-19 crisis show that institutional design has a significant impact on programme uptake and efficacy. Given that persistent employment, survival, and productivity gains are concentrated among liquidity-constrained firms or those facing large temporary declines in demand, STW is thus most effective when subsidies are targeted toward these firms. Broad or untargeted programmes, on the other hand, typically result in deadweight losses without long-term benefits and postpone necessary labour reallocation. As seen during COVID-19, rapid STW expansions highlight the trade-off between accurate targeting and prompt support.

Thirdly, well-designed STW reduces unemployment scarring, protects earnings, and maintains jobs, according to microeconometric studies at the firm and worker levels. If financial incentives, such as experience rating or co-payments, are carefully calibrated to prevent penalising firms facing exogenous shocks, they can prevent excessive use and internalise fiscal costs. The temporary nature of STW is reinforced by maximum duration limits, which also prohibit structural use that might impede reallocation.

Finally, our survey points to remaining research gaps. Medium- and long-term impacts on productivity, profitability, and workers' earnings trajectories remain underexplored, as do distributional effects across contract types, age groups, and sectors. Further work is needed to understand the optimal balance between targeting through monitoring and co-financing mechanisms, as well as the interaction of STW with the complementary labour hoarding strategies and working-time arrangements.

Overall, this analysis shows that STW can minimise distortions and financial costs while offering prompt, targeted stabilisation. By systematically connecting theory, cross-country institutional design, and empirical evidence, it also offers practical insights for improving STW schemes so that they remain a reliable and effective tool for managing transient economic shocks in Europe. ☐

Link to the Online Appendix:
www.insee.fr/en/statistiques/fichier/9008748/ES549_Bermudez-et-al_Online-Appendix.pdf

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- Un fichier anonymisé du manuscrit complet (texte, illustrations, bibliographie, éventuelles annexes) indiquant en première page uniquement le titre, le résumé, les codes JEL et les mots-clés.

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