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AGE AND GENERATIONS

Editorial

Économie et Statistique becomes *Economie et Statistique / Economics and Statistics*

This latest issue marks a new milestone in the development of *Économie et Statistique*, which now becomes *Economie et Statistique / Economics and Statistics,* under the sign of continuity and internationalisation, as reflected in the new name.

Articles will continue to appear in French in the print edition of the journal and in its online edition, which can be accessed on the Insee website, free of charge, on the day of publication. The novelty is that the online articles will also appear in English at the same time.

Published by Insee, the journal will continue to present articles on economic and social phenomena, drawing on data from official statistics and other sources, together with articles on statistical and econometric methodology of interest to applied research. The journal will also continue to issues alternating between articles on a variety of topics, and special issues focusing on a particular theme or major official statistical surveys. As part of its effort to reconcile scientific excellence with readability, the journal continues to target a broad and diverse readership which includes students, researchers, methodologists, producers and users of statistical data and political and economic decision-makers.

It intends to go on informing the economic and social debate in France. In today's globalised world, many economic and social phenomena concern advanced economies beyond the borders of France; the same economic and social policy issues often recur, even if the responses may differ at the international, European and national levels. While many surveys are designed within a European (Eurostat) or international (OECD) context, and official statistics are governed by European regulations, *Economie et Statistique / Economics and Statistics* now seeks to fuel thinking and shed light on these phenomena through national and foreign contributions written from an international perspective.

These questions and phenomena are the subject of forthcoming issues. They include: the financial crisis and its transmission to the real economy ten years later; technical progress and uneven regional development; housing inequality; and the Big Data revolution and the construction of official statistics by European national statistical institutes. The present issue on "Age and generations" looks at differences between public and private sector pensions and intergenerational comparisons of living standards, drawing extensively on data from Insee's Household Expenditure surveys (*enquêtes Budget de famille*). Professor Sir Richard Blundell of University College London, renowned for his work on consumption and savings behaviour, on labour markets, and his contributions to panel data econometrics, has done us the honour of writing the foreword to this issue. Didier Blanchet, the former Editor-in-Chief and a specialist in pensions and demographic issues, returns to the subject of intergenerational equity in the introduction to this issue, a project to which he made a substantial contribution.

A great deal of work has gone into opening up the journal and helping it find its place in the demanding academic world. Founded in 1969, nearly half a century ago, the journal was initially designed to publish the work of Insee's statisticians and researchers. As of the late 1980s, it began to publish a growing number of articles by authors from outside the Institute. Pierre Morin, Editor-in-Chief from 1993 to 2011, accentuated the policy of opening up the journal and gradually aligned its way of working with that of other scientific journals. Articles undergo thorough review procedures and, in 2003, a scientific council was set up to review the previous year's publications and help define the editorial strategy. Then Didier Blanchet, who directed the journal from 2011 to 2016, set up an editorial board in 2013 to assist the scientific council in selecting articles. The position that the journal now occupies was thus acquired through a constant effort to support authors, article by article. The editorial team has every intention of continuing this effort.

For more than a quarter of a century, the priorities of the journal have been to open up to a broad range of topics and disciplines rooted in the worlds of official statistics and research. These priorities are confirmed by this latest initiative to open up the journal further still, by attracting authors, including reviewers, and addressing readers outside the French-speaking world, and thus provide *Economie et Statistique / Economics and Statistics* with the means to gain international recognition. □

Laurence Bloch, Editor-in-Chief

Foreword
On the importance of taking a life-cyle view in understanding generational issues

Professor Sir Richard Blundell
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Accurate measurement of the standard of living and of generational comparisons is of growing importance in the policy debate. Are baby-boomers better off than generations born later? How should we compare the standard of living across families of different size? How do individual earnings relate to family income? Are public pensions more generous that private sector pensions?

Snapshot measures of transfers, taxes and pension entitlements can be very misleading and typically do not reflect the reality of either the level or distribution of standard of living. What is required is careful and detailed empirical work on all sources of incomes and outlays for families and for the individuals that make up family units. It is particularly important to take a life-cycle view of incomes, consumption and the standard of living, especially in understanding generational issues.

My own research[1] has stressed the value of combining micro-data measures of earnings, income and consumption for different birth-cohorts to learn how individuals and families 'insure' themselves against adverse labour market shocks across their life-cycle. From this analysis we can examine the trade-off between social insurance and self-insurance[2]. A key aspect of recent trends in the standard of living, especially among the poor, has been the growing inequality in the labour market earnings of men[3]. For men in the UK, low hourly wages and low hours of work increasingly go together. Over the last two decades the growth in tax-credits and female earnings have offset this trend and, for the vast majority of the population, total net household income inequality has been much more stable. But for how long? And what about top incomes? As in other developed economies, the top 1% have been very different. Their share of net total household income increased dramatically.

I would like to emphasise the importance of careful empirical work on detailed individual data with comprehensive treatment of taxes and transfers, placed in a life-cycle setting, for studying changes in the standard of living and for generational comparisons. This is precisely the aim of the papers in this volume. It is great to see this new issue of *Economie et Statistique / Economics and Statistics* address these key issues in research on economic statistics.

1. *Income Dynamics and Life-Cycle Inequality: Mechanisms and Controversies (2014),* Economic Journal*, 124(576), 289–318.*
2. *Labor Income Dynamics and the Insurance from Taxes, Transfers, and the Family. Joint with Michael Graber, and Magne Mogstad (2015),* Journal of Public Economics*, 127, 58–73.*
3. *Two decades of income inequality in Britain: the role of wages, household earnings and redistribution. Joint with Chris Belfield, Jonathan Cribb, Andrew Hood, and Robert Joyce (2017),* IFS Working Paper *W17/01 (forthcoming* Economica*).*

Age and generations: a general introduction

Didier Blanchet *

This issue of the journal brings together five contributions devoted to comparing standards of living depending on age and generation: methodological contributions relating to equivalence scales and to the econometrics of pseudo-panels; the initial results for France of National Transfer Accounts (NTA) that break down National Accounts aggregates on the basis of age; and comparisons of pension entitlements between public and private sector employees. We return to four of the questions they raise. The first is the issue of separating age, period, and cohort effects: how it is conducted should depend on the question asked. We then advocate a plural approach to intergenerational inequalities, consisting in looking at them from several complementary angles: for example, by referring not only to monetary income, but also to health, and access to education and employment, or housing. We continue by examining the concept of “lifecycle deficit”, which is calculated by the NTA, and is the gap between what a generation consumes and what it produces through its labour throughout its existence. We discuss how it ties in with the broader issue of sustainability, which is the prospective part of the issue of intergenerational fairness. A minimalistic criterion of intergenerational fairness could be that each generation should be watchful to ensure that the next ones enjoy living conditions at least as good as it did. Finally, we comment on the various possible avenues for comparing pension entitlements in the public and private sectors: the difficulty of measuring contribution effort is an argument in favour of an overall approach combining direct salary and all of the pension entitlements.

JEL codes: D31, E01, H55.

Keywords: age, generations, sustainability, pensions.

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Translated from: « Âges et générations : une introduction générale ».

This issue of *Economie et Statistique* is inaugurating a new version for the journal, which now becomes *Economie et Statistique / Economics and Statistics*, with all of its articles in the electronic edition being systematically bilingual, in French and in English. By keeping the whole of its traditional format in French, the journal wants to continue to enlighten the national economic debate on the basis of work done inside or outside the public statistics system. Simultaneously, publication in English will give that work broader international visibility. The new editing team should be thanked and congratulated for taking the initiative of producing this new version.

As in the past, the research work will be included either in mixed issues or in special theme issues. This issue is part-way between the two types. It is not strictly a special issue systematically covering a single topic. However, the studies it presents have a common denominator that justifies them being grouped together, namely the question of measuring and comparing standards of living depending on age or generation.

Two articles are of the methodological type. The one by **Henri Martin** is devoted to evaluating equivalence scales, a recurrent topic for the journal (Bloch & Glaude, 1983; Glaude & Moutardier, 1991; Hourriez & Olier, 1998) and that is an essential stage in evaluating the standards of living of households whose demographic structure evolves from one phase of the lifecycle to another. It implements one of the possible approaches to the question, namely the approach based on subjective perceptions of standard of living, rather than on indirect indicators such as the shares devoted to food spending or to adult-specific spending, such indicators being more objective but also conventional and probably outdated. The work is interesting because it implements that method on the latest edition of the French Household Expenditure Survey (*enquête Budget de famille*), and because it shows how the results are sensitive to the choice of specification that is made, which would suggest the method needs to be used with precaution: what it gives is a range of possibilities and the work of comparing the standards of living of the households should take that uncertainty into account. The other methodological contribution is the one by **Marine Guillerm**, who offers an instructive presentation of the pseudo-panels method and of some of its recent technical developments, with an application to the relationships between age, generation, and possession of wealth. The article highlights well the relationships between pseudo-panels and true panels, and the way in which the former can constitute interesting alternatives to the latter.

Very much related to that article is the one by **Hyppolite d'Albis** and **Ikpidi Badji** that poses the question of how to separate age effect, generation effect, and period effect for the income and consumption of the households observed from 1979 to 2011. Their work is part of an international project to compile National Transfer Accounts (NTA), that project also being represented in this issue with the article by d'**Hippolyte d'Albis**, **Carole Bonnet**, **Julien Navaux**, **Jacques Pelletan** and **François-Charles Wolff.** The aim of such transfer accounts is to go as far as possible into disaggregating the results of national accounts according to age, on the basis of all of the microeconomic data that so permit (see the website of the National Transfer Accounts Project[1] and United Nations, 2013). We should salute the ambition of this work and emphasise its concern for mapping

1. http://www.ntaccounts.org/web/nta/show

as well as possible with the data from the national accounts. National accounts do not have all the answers, but they offer the advantage of providing a consistent accounting framework within which to put the results from sources enabling finer analyses to be made. That is what is done here. Such is also the spirit of Distributional National Accounts (DINA[2]) that come from the work of Atkinson, Piketty and Saez on high incomes (Atkinson et al., 2011), gradually broadened to all income and wealth distributions. That is also the spirit of the efforts made at the OECD for compiling household accounts that are disaggregated both into age and into socio-professional category (Fesseau & Van de Ven, 2014), using work conducted and continued at Insee since the early 2000s (Accardo et al., 2009). For NTA, the focus is on breaking down by age exclusively, which is a whole subject on its own and manifestly an issue in a context of population ageing and of adapting the transfer systems to accommodate that ageing. The focus on the sole dimension of age is offset by the project offering international comparison and attempting to propose series that are as long as possible.

For its part, the article with which this issue opens, by **Patrick Aubert** and **Corentin Plouhinec**, looks at a highly debated aspect of the fairness of the French intergenerational transfer system, namely comparing pension entitlements in the public and private sectors. It is accompanied by comments by **Antoine Bozio**.

This preface is not going to propose systematic discussion of all of the points covered by the articles. It will limit itself to providing some perspective for four of the questions that they raise. The first perspective concerns the question of implementing pseudo-panel methods and more specifically the issue of identifying the effects of age, of period, and of cohort. The subject might appear technical, but there are important issues at stake when the idea is to decide to what extent successive generations are "privileged". We would emphasise that it is important not to approach the question from an econometric angle only: it is necessary firstly to think about the exact nature of what we are seeking to measure. The second perspective consists in advocating a plural approach to intergenerational comparisons, consisting in looking at them from several complementary angles. The third perspective consists in linking up the issue of "lifecycle deficit" as calculated by the NTA and the question of sustainability, which is quite simply the prospective part of the issue of intergenerational equality. Finally, we return to the question of intergenerational comparison of pension calculation rules, by continuing Bozio's discussion about it. How can we compare two pension schemes whose principles are very different, and how can we go beyond the eternal debate about their relative generosity?

Age, period, and cohort effects: how and why should we distinguish between them?

Separating age effect, period effect, and cohort effect is a problem with which we are systematically confronted whenever we have data by age over a long period. This issue is older than the emergence of the term "pseudo-panel". It is a traditional subject for demographers who, it might be said, did pseudo-panel work for a long time without knowing it. The demographic indices that are in most

2. *http://wid.world/wid-world/*

widespread use are period-based indices constructed by aggregating data that are broken down into age, but it is often the generation effects that we are seeking to read or to anticipate behind the period effects: what will be completed fertility levels for successive generations, and what can period life expectancies tell us about longevity changes according to year of birth? For a long time, the demographic approach to this problem remained descriptive and non-econometric, using graphical representations of the type to be found in the articles by Guillerm and by d'Albis and Badji, graphical representations of the age effects for periods or successive generations that have also been widely used for analysing variables such as employment rates or profiles of salaries according to age.

The article by Guillerm reminds us that the term "pseudo-panel" has another origin. It was in the 1970s that panels of microeconomic data became progressively accessible, those panels monitoring elementary units of observation over time, those units being either households or enterprises. What was and still is expected of such data is that they should help to solve a fundamental problem of econometric inference on cross-sectional data, the estimation biases resulting from the non-observed heterogeneity of the units analysed, when it is correlated with the variables that we are seeking to explain. Having repeated cross-section data makes it possible to neutralise this heterogeneity by consenting to assume it is constant over time. But these "real" or "true" panel data are not always available, and many statistical sources are still in the form of repeated independent cross-sections, without individual monitoring. And, even when the same individuals are monitored from one wave to another, a problem we come up against is attrition, which can be selective and also correlated with the phenomenon of interest. Such attrition often leads to limiting ourselves to short panels, which do not lend themselves well to analysing phenomena whose development all through the lifecycle of the units we want to analyse.

It was in response to these various questions that came the idea of seeing whether grouping together cross-sectional data into homogeneous cells that are monitored over time might be an interesting alternative, preserving most of what is contributed by true panels while also addressing some of their limitations (Deaton, 1985). The term "pseudo-panel" thus includes the old descriptive practice of arranging into cohorts the successive values of data measured according to age, but also augments it with an econometric problem to be solved, namely using that data for explanatory analysis, with the same expectations as what was expected from true panels. This dual aspect (descriptive and econometric) appears clearly in the articles by Guillerm and by d'Albis et Badji, which use both traditional graphical visualisations distinguishing age effects at given periods or for given generations, and econometric modelling of the phenomena of interest, namely wealth in the former case, and income and consumption in the latter.

Regardless of whether the angle of approach is descriptive or econometric, separating the age, period, and cohort effects poses the same problem. In the descriptive approach, its most usual expression is the fact that the apparent effect of age is not the same depending on whether it is looked at from a cross-sectional angle or from a longitudinal angle, those angles neutralising respectively the role of generation, and the role of period. The econometric approach comes up against the same difficulty and reformulates it in terms of identifiability. The three effects of age, period, and cohort of the APC (Age-Period-Cohort) models are linear and identifiable only up to one linear term because age is equal to the difference between the current date and the year of birth.

With the problem having been reformulated from this econometric angle, we are naturally led to seek econometric answers to it, i.e. the choice of the identifying constraints that make it possible to remove this indeterminacy. The risk to be avoided is the risk of approaching the issue only from the angle of econometric technique, while losing sight of the basic question, namely knowing what exactly we want to measure. The econometric strategy to be implemented depends on what we are seeking to estimate. The study by d'Albis and Badji is a good illustration of this problem. Depending on the case, it is possible to choose deliberately to "load" the generation effect rather than the period effect by purging it of any time trend – this is what the Deaton and Paxson (1994) method does – or *vice versa*.

Knowing which of the two options to choose really does depend on the question asked. An example of when it would seem abnormal to remove any trend from the period effect is the case of contributions to productivity. Admittedly, productivity includes a generational component – the rise in the level of initial training of the successive generations – but for the most part, it is period phenomenon: innovations take place at each period, and, at least up to certain point, they benefit simultaneously and cumulatively all of the generations working at that period. It is difficult to imagine representing that component by a variable devoid of any trend, and which would alternate between periods of growth or, conversely, of recession. In such a case, we need to find other ways of solving the problem of identifiability, e.g. by estimating the generation effect through the observable impact of the level of education of each generation.

Conversely, "loading" the generation effect to a maximum extent is fully justified if the aim is to know how the same progress in productivity benefits the standards of living of the successive generations. Even if the rise in standard of living between generations were due only to period effects, without owing anything to the specific characteristics of the successive generations, it would still remain that the cumulative result of all of these period effects would indeed enable each generation to be richer than the preceding one, and that is the message that we want to emphasise. In which case, this time trend must be found at generation level. This is what d'Albis and Badji do: they use the APC method to highlight that, over all of the cohorts born from 1901 to 1979, no generation has been disadvantaged compared with its elders. More precisely, the standard of living of the baby boomers is higher than that of the generations born pre-war, and lower than that of the generations that come after them.

Loading the generation effect to a maximum extent is also what we would do if we had full longitudinal data that we merely needed to sum over all of the lifecycles, dispensing with the estimation of an APC model: this is how we proceed in the demographic field when we wait for the life cycles to be observed completely before we say what the real developments of the phenomenon of interest over the lifecycles are. This reminds us that the purpose of APC models is not necessarily to identify period effects and generation effects *per se*. They might be seen merely as calculation intermediaries making it possible to give messages about what will become of the generations without waiting for the ends of their lives. But that means going over to a forecasting approach, for which there is no miracle solution: a forecast requires assumptions. The crucial assumption of the APC approach is that age effects are stable. That assumption is necessary if we want to be able to give messages about the overall lifecycles of generations for which we observe only the ends or the beginnings of their lives, merely on the basis of the very incomplete information we have about them. The assumption of the age effects being stable can only be an approximation. This applies particularly for the effect on income on becoming retired. The improvement in pension entitlements up until the middle of the 1980s rather distorted the age profile of income in favour of retirees, and a reverse movement is expected, ultimately, under the effect of the reforms put in place since the second half of the 1980s, with, in particular, the change to price indexation for the main parameters for calculating pension entitlements. All this would urge us to look behind econometrics. For properly answering the question of the standard-of-living prospects for the younger generations, the APC model is merely an indicative tool that cannot replace more in-depth projection exercises such as the ones regularly conducted for pensions.

Comparing successive generations: we also need to vary the points of view

To continue on this subject, we should point out the limitations – and also the advantages – of another approach to this APC problem, namely the Age-Period-Cohort-Detrended (APCD) models approach that is mentioned briefly at the end of the article by d'Albis and Badji, taken from Chauvel (2013). That decomposition method puts on an equal footing both of the polarised solutions consisting in transferring all of the trend effects either onto the period effect or onto the generation effect, by making both of the two effects stationary. It might be said that the idea is to evaluate the period and generation effects only insofar as they deviate from the general trend. Its limitation can easily be seen: if there is an upward trend in the objective standard of living over time, it seems really difficult to ignore that in comparing successive generations. Nevertheless, this double correction of the trend effect can play an interesting part in attempting to reconcile this message of growth in the standard of living with the perceptions of the relative intergenerational situations. Here, we return to a classic theme of comparisons of well-being or happiness over time, namely the paradox suggested by Easterlin in 1974 whereby an improvement in objective standard of living as measured by the indicators of the national accounts is not found in the evolution of subjective (self-reported) well-being because such well-being is evaluated by the interested parties in terms of difference relative to their aspirations. In such a case, it is the accelerations or slowdowns in growth that translate into variation in perceived well-being. The messages provided by the APC model and by the APCD model are then complementary, one for reporting an objective reality, and the other for reporting the way it is perceived.

Along the same lines, we can point out another way of introducing this concept of relative perception in analysing intergenerational inequalities. For each generation, it is possible, at each age, to look at the way it is situated relative to the other age groups at the same period (Legris & Lollivier, 1996; Blanchet & Monfort, 2002). Let us imagine a general growth trend that is beneficial for everyone but with a specific generation who, at each period, manages to enjoy a relatively bigger slice of the instantaneous cake at each age: for example, if it has enjoyed a generous family policy when it was young, if it has not had to bear upward transfers that are too high while it was working, and if, on retirement, it benefits from transfers that have not yet been reduced too much. It might well be that this generation will not ultimately have a standard of living that is higher than the standards of living of the succeeding generations. However, the fact that it has given the impression of doing better at each period of its existence is a point that deserves to be reported. We can note that a phenomenon of this type may appear from the article by d'Albis et al. concerning salaries by age. The employee earnings increase from one period and from one generation to the next. But it can also be observed that the generation of 1954, aged 35 in 1979, was already at the mode of the distribution of salaries by age in 1979, and was at that mode again 10 years and then 21 years later, in 1989 and in 2000, a little as if, at each date, it had managed to access the highest paid jobs of the time. This type of phenomenon is doubtless worth looking at a little closer, and, at the very least, it shows the utility of varying the points of view.

Varying the points of view can also consist in multiplying the number of dimensions of well-being or happiness that are used to make comparisons between generations. It is possible, in particular, to refer to Clerc et al. (2011) who use the dimensions of monetary income, health, and access to education, to employment and to housing. By trying to disaggregate the aggregates of national accounts on the basis of age, it can be said that the NTA, like Distributional Accounts, or accounts by social category meet one of the recommendations that had been made by the Stigliz-Sen-Fitoussi report for going beyond the limitations of national accounts: the recommendation to go beyond the average (Stigliz et al., 2009). For the NTA, this expression should be understood as "going beyond the instantaneous averages", which, *de facto*, are not necessarily representative of the experiences of all of the generations concerned throughout their lifecycles. But another aspect of its recommendations is ignored, the one for also going beyond an "all-monetary" logic. A possible explanation for the difference between the message of d'Albis and Badji and the perceived intergenerational inequality might, for example, lie in individuals weighting the consequences of difficulties of access to employment more heavily than as mere monetary consequences, which seems to be a classic result in the literature on determinants of subjective well-being.

Lifecycle deficit and sustainability: how are they related?

The article by d'Albis et al. can also be looked at from another aspect of the Stiglitz message, the one concerning measuring sustainability. This issue of intergenerational equality and the issue of sustainability are really very closely related. Actually, it is quite difficult to agree on exactly what the concept of intergenerational equality covers, but when we ask ourselves the question from a prospective angle, i.e. equality with respect to future generations, there is quite a simple minimalistic criterion that consists in saying that every generation should be watchful to give the next generations the assurance of living conditions that are at least equal to those that it was able to enjoy. It is from this perspective that we can question the concept of "lifecycle deficit", which is one of the main indicators of the NTA and that gives the article its title. The idea is to compute the difference or gap between what one generation consumes and what it produces through its labour throughout its existence. In the first phases of its existence, each generation consumes only, and then it becomes productive, and its production outstrips its consumption during its working adulthood, at the end of which its production returns to zero and it becomes a pure consumer again. Naturally, this observation is clearly nothing new, and the contribution that the article makes is rather to quantify this phenomenon, and above all to examine how it evolves over time, be it under the effect of behavioural changes or indeed, at aggregate level, because of the variation in the relative weights of the age brackets that is induced by population ageing. The question is whether or not generations tend to consume an increasingly large fraction of what they produce through their labour, thereby reducing accordingly what they pass on to the next generations. Can this go as far as to a situation of overconsumption in which generations consume more than they produce over their lifecycle?

This aspect of the NTA Project descends directly from an earlier initiative recalled by the authors, namely the attempts made by Kotlikoff et al. to compile accounts by generation. However, there are two important differences. The first is that that approach focused on the issue of fiscal transfers, i.e. on comparing what each generation contributed to and cost the public finances, with the idea of being able to say "who pays for whom" in the game of intergenerational redistribution. The second is the rather militant nature of the approach: speaking out against undue reception of public resources by certain generations, namely those who benefited from the expansion of the Welfare State while leaving some of its financial burden to the next generations (Kotlikoff, 1992). As d'Albis et al. recall, fifteen years ago the journal *Économie et Prévision* had devoted a special issue to discussing that approach (Malgrange & Masson, 2002). The NTA, for their part, take into account the sum of what the generations produce, earn in labour income, and consume – generations can leave, at the same time, both a large public debt and also considerable private assets, it is the result of the two that is important. The NTA also do that in a more detached spirit. The message of the article is intended to be moderate, even though the message is that the deficit has grown. In particular, the gap widened from 1979 to 1989. Expressed in consumption points, we went from a surplus of 6.2% in 1979 to a deficit of 15.3% in 1989, and that deficit has remained roughly stable since then.

Even though it is superior to accounting that is limited to public transfers, the messages drawn from the indicator can nevertheless call for some precaution, but that can be in two opposite directions. In the indicator, there are some details that lead to the problem of sustainability being overestimated and others that tend to lead it to be underestimated.

On the overestimation side: as it is defined, the "lifecycle deficit" compares consumption and production, which is quantified by the inflows of labour income throughout the whole career path. We might be tempted to understand that one generation penalises the next ones whenever it consumes more than it has produced directly through its labour. That would be to forget the part played by capital income.

Firstly, under steady-state conditions, we can have situations that are entirely sustainable and in which each generation consumes more than the income from its labour because it is also possible for a fraction of the capital income to be consumed too without calling sustainability into question. This applies whenever the trend is for the rate of return on the capital to be greater than the rate of growth of the economy – the famous relationship "r>g" put forward in the work by Piketty

(2013). That inequality enables each generation to consume all of the income from its labour, and a fraction of the income from the capital, and nevertheless to allow the stock of capital to grow at a rate greater than or equal to g, which is a condition sufficient for that growth to be sustainable.

Secondly, outside steady-state conditions, labour income does not represent a stable fraction of the sum of what the working population produces: the sharing of the value added is deformed and that can distort the message of a comparison between consumption and salaries alone. Perhaps that plays a part in explaining what we observe between 1979 and 1989. The starting point of 1979 is a situation in which, following the first oil crisis, the sharing of the value added was deformed considerably in favour of the salaries, in a way that the policies of the 1980s successfully sought to reabsorb. The initial surplus might be due to this atypical sharing of the value added, and its subsequent resorption would thus have been due to a phenomenon of returning to normal.

It is thus worthwhile to look at the other terms of the accounting equations presented in the article. But fully exploring this idea of taking capital into account could also lead to messages that are less optimistic about the issue of sustainability. If we follow what the Stiglitz report says about it, the concept of the capital that each generation passes on to the next ones should be broadened to include many other dimensions in addition to those that are monitored by national accounts (Blanchet et al., 2009; Antonin et al., 2011). Two main candidates for such broadening of the concept of capital are intangible capital and environmental capital. The message about sustainability can find itself reversed again: a generation might have consumed more than the sum of the salaries over the lifecycle – negative message – but less than the sum of what it has produced overall during its lifecycle, once the earnings from productive capital have been included – hence a positive message – and at the same time leave to the next generations a smaller amount of capital than it had inherited, if the stock of productive capital in the usual sense of the term being maintained or increased is more than counterbalanced by what is taken from the natural assets.

Naturally, in the current state of progress of the NTA, we cannot complain that this broadened vision of the concept of capital was not approached from the outset. A major step has already been taken by broadening Kotlikoff's initial approach to beyond merely accounting for taxes and transfers. And, since they are mapped on national accounts, the NTA are necessarily limited by them: no accounting for natural assets, and intangible asset accounting that is in its infancy. Similarly, some might see it as restrictive to limit the vision of production to market production only: describing non-working retirees purely as consumers naturally ignores their home production. A section of the NTA project that is not presented in this file also aims to take this home production into account, from a gender accounting perspective (d'Albis et al., 2017). There are thus many avenues open, and we should not hesitate to approach them by going beyond the highly normed framework of the core of the national accounts system.

A final remark can be added to that: the authors emphasises the dynamic nature of their approach, i.e. the fact that they have managed to construct accounts in relatively long series, over 35 years, which, so far, only applies to a minority of countries taking part in the project. That is indeed an advantage, but the observation window nevertheless remains too short to reconstruct genuine generational histories. This is another point about which the choice of the term "lifecycle deficit" might call for a caveat, as the authors also admit. These lifecycles are only pseudo-lifecycles here. To return to using demographer's vocabulary, the deficits we compute are transversal, i.e. those that a fictitious generation would have, knowing throughout its life the consumption and production conditions by age of the current period. To put things another way, it can be said that the usual concept of aggregate savings rate is also a cross-sectional concept that we might want to transform into a longitudinal one: when the instantaneous rate of savings falls, the behaviour of exactly which generation(s) is the cause of that fall? Answering this question would be advantageous, but, once again, it has a largely prospective dimension and we would certainly not recommend trying to use the same type of APC approach as in the article by d'Albis et Badji: we really find it difficult to see how gaps between consumption and production that are observed either at the very beginning or at the very end of the lifecycle could be used as a basis for econometrically estimating differences in overall deficits over the lifecycles of the generations in question. The only solution is to push backcasting further for the earlier generations, and push forecasting also further for the more recent ones. Indeed, that is what Kotlikoff-style generational accounts were led to do. The NTA should be seen as providing only a fraction of what is required for full generational balance sheets, namely the fraction that covers observing the past, to be supplemented by projection exercises of the same type as those that are constructed for studying pensions. This offers us an opportunity to link up with the first of the articles in this issue, the one by Aubert and Plouhinec.

Comparing pension entitlements: what indicators should be preferred?

An important dimension of these intergenerational transfers are the ones related to pensions spending, about which the journal has already written abundantly, including very recently from this intergenerational comparison angle (Dubois & Marino, 2015). Here, the issue raised is more intragenerational, namely comparing the "generosities" of pension entitlements between the public and private sectors, an eminently divisive subject in the French public debate. Naturally, we cannot make do with naïve comparisons, such as merely putting the average levels of pension of civil servants and private-sector employees side-by-side, as we still often see done. There is no sense in making this type of comparison because the populations have very different average levels of qualification. There would be sense in making this comparison only if France had chosen a Beveridge-type pension system aiming to allocate the same levels of pension to all retirees, regardless of their qualifications and regardless of their past jobs and salaries. This principle is not the one on which the French pensions system was built, in which pension and past salaries are closely linked. The comparison should be made at identical salary levels, which is done here by simulating application of the rules of one or the other of the systems to standard cases whose career profiles have been set. This way of doing things does not completely exhaust the debate, as shown by the discussion by Bozio, but it does show that it is not possible to have a simple and unequivocal position on comparing the two types of rules: one or other of the two systems is the more favourable depending on the standard cases examined.

However, the comparison remains limited to one indicator, namely the replacement rate. Here too, the question arises of diversifying the points of view, and also of the possibility combining these points of view by using a single synthetic index. A first dimension that is lacking from the analysis is the time for which the pension is paid, and that depends both on the age at which the pension starts being drawn and also on life expectancy. Taking it into account would not pose any particular technical problem, and it is possible to consider combining level of pension and length of payment in the form of an aggregate indicator of present or discounted pension entitlements, which is what the literature often refers to as a wealth equivalent of pension entitlements. But that is still not enough: higher or lower overall entitlements depending on employment categories or grades are not necessarily synonymous with inequality if they are in exchange for past contribution efforts that are more or less large. That is what indicators of return of the pensions system or of return on contributions sought to verify.

The latter type of indicator calls for a few comments to be made. It is quite often excluded from the French debate on pensions due to it having a connotation that is overly "funded pensions" and because of the argument that, in any case, equalising these rates of return on contributions cannot constitute a target for equality, either intragenerationally or intergenerationally.

These two arguments can be refuted. As regards the former, admittedly it is true that calculating return indicators can lead to messages that are apparently unfavourable to "pay-as-you-go" pension financing. Under steady-state conditions, the average return that pay-as-you-go can guarantee to the beneficiaries insured under the scheme is equal to the rate of growth g of the economy, whereas the average return of funded-schemes is, in principle the interest rate r. Funded schemes therefore appear to offer higher performance than pay-as-you-go ones whenever the above mentioned condition "r>g" is satisfied. But we do not necessarily need to banish such a comparison, because it does not necessarily convey the message that going over to funded schemes would be beneficial for all. Firstly, there is the fact that such a transition would penalise the transition generations because it is not possible to go back from pay-as-you-go to funded without having to bear a period of double contribution, or without giving up at least some of the entitlements acquired by the generations who have already reached retirement. Secondly, even under steady-state conditions, the relationship r>g may be valid only as a trend. The advantage of funded schemes then comes at the price of the pensions paid being more sensitive to economic uncertainties, and the crisis of 2008 with the sudden devaluation of assets is there to remind us of that. Finally, regardless of whether or not g is close to r, the mere fact that it remains positive provides an argument that is useful in defence of pay-as-you-go: it makes it possible to invalidate the widespread theory that it is synonymous with the young generations contributing "at a loss", that line of speech doubtless also weighing heavily on the "perceived" intergenerational inequality that we mentioned above.

As for the idea that equalisation of rates of return is not a standard for equality or fairness, that idea is naturally quite true. Social justice entirely legitimises having rates of return that are higher for the least privileged categories – it is the very principle of redistribution – or indeed for categories exposed to particular constraints. But that in no way precludes looking at those rates of return, quite the opposite in fact. Far from opposing redistributive logic, calculating the rate of return on contributions can constitute one of the means of managing such redistribution, by making it possible to check that it is acting in the right direction, with a rate-of-return gradient sloping the other way to the primary resources gradient.

Ideally, it is this approach that we would like to be able to apply to public-private comparison. But, here we come up against an apparently insurmountable difficulty, namely how to measure the contribution effort objectively. Even in the private sector, it is not patently clear how to measure it. The usual approach is to identify it by the rate of contribution, and more exactly by the overall rate of contribution combining the employee and the employer contributions. What needs to be identified is the share of the direct salary that the employees deprive themselves of to fund their future pension. The assumption made is that the overall cost of labour is set exogenously independently of the legally required sharing between employer and employee contributions. This is particularly true if the overall cost of labour is imposed on the economy by the state of international competition. If employees want to keep their jobs at a given overall labour cost, they are obliged to accept that the sum of the two contributions is deducted from their net salary, and it is thus them who ultimately fund their pension entitlements. But the assumption cannot be totally true. And this calculation ignores the fact that, in the private sector, a large fraction of the pension entitlements is funded by non-contributory levies, to which this reasoning does not apply: in 2013, the employer and employee contributions covered only 72% of the total spending for the pension schemes (*Conseil d'orientation des retraites*, 2015).

The problems are even more difficult to overcome in the public sector. We have an apparent rate of contribution for the employee only, the public employer contribution taking the form of a balancing subsidy. Can we consider that that subsidy is ultimately also paid by the employee, i.e. the idea that, if the French State did not have to pay the pensions of its former civil servants, it would give back to its employees all of the resulting savings made? Perhaps that would be true in a world of permeability and of total competition between the labour markets of the public-sector employees and of the private-sector employees, but the assumption is a strong one, of course.

This makes it particularly interesting to look at an alternative avenue mentioned in Antoine Bozio's discussion and making it possible to bypass the issue of contribution effort, at the price of broadening out to an overall comparison including both salaries and pensions. The approach would be to consider that, in an employee's instantaneous pay, there are two components, namely a direct net salary, immediately pocketed, and a deferred salary, constituted by pension entitlements. What can then be aimed for is an overall comparison of the sums of direct net salaries and deferred ones, both discounted all over the lifecycle. As a result, we broaden the issue to include more than the pension entitlements, but is that not the real problem on which we should focus in making such comparisons between the public and the private sector? If replacement rates that are higher for certain categories or grades in the civil service are accepted in compensation for a lower direct net salary and/or for specific constraints, this type of indicator will give us the correct message regarding equality between the two categories of population in terms of present overall entitlements as discounted over the lifecycle. There would be inequality only if one or the other of the categories totalled direct salary and deferred salary that were both higher, for identical qualifications and job characteristics. Here, there is an area to be explored that brings together analysis of pensions and analysis of wage inequalities.

That said, another way out of the suspicion of unequal treatment between the two categories of the population would be to work more to have the rules of the two types of schemes converge. The bottom line is that, if the rules and the modes of funding were totally identical, the issue of comparing the two pensions systems would become a non-subject, and comparing the two categories could boil down to comparing salaries alone, leaving the salary alone to compensate for the specific constraints of the various jobs. But that is another story, and would doubtless take a very long time to achieve. □

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Differences between public and private sector pensions: an analysis based on career profile simulations

Patrick Aubert * and Corentin Plouhinec **

Following the alignment of the rules for private and public sector pension schemes, which began with the reform of the French pension system in 2003, there remain a number of differences. These include structural variations between the two schemes, the definition of the reference salary (salary over the best 25 years in the private sector or salary excluding bonuses over the final six months in the public sector).

We simulate the application of the two types of rules to several standard civil service careers. The impact on the replacement rate is not homogeneous: for the generation born in 1955 preparing to retire in 2017, applying private sector rules would be more beneficial for a standard category B civil servant, but less beneficial for a teacher, and slightly less beneficial for an "A+" category manager. This is the result of the interplay of the factors that determine pension amounts with each type of rule: the proportion of bonuses in the total remuneration for civil service schemes (the higher this proportion, the lower the pension amount as a proportion of the final pay), the level and slope of the wage trajectory for private schemes (the more the slope is ascending and the greater the proportion of pay over the social security ceiling, the lower the pension as a proportion of the final pay).

A change from one sector to another during a career can have a significant and varied impact on the replacement rate. It often leads to a lower replacement rate than would be achieved by remaining employed in either the public or the private sector throughout a career (for identical net salaries at all ages), but there are some configurations where a change of sector leads to a higher replacement rate: for example, the case of a category A+ civil service manager whose career finishes with around ten years in the private sector.

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Reminder:

The opinions and analyses in this article are those of the author(s) and do not necessarily reflect their institution's or Insee's views.

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As in many OECD countries[1], the French pension system is characterised both by the range of mandatory schemes and a diversity of rules for acquiring pension benefits and calculating pensions. In addition to the basic general scheme (Cnav), the French system has some basic occupational pension schemes (agricultural employees, craftsmen and merchants), special employee schemes (civil servants and some private sector employees[2]), and self-employed schemes (self-employed professions, self-employed agricultural workers, etc.).

Such a diversity of rules generates much debate about the equity between schemes – as evidenced by simply reading through the parliamentary debate proceedings around the last reform of the pension system, and observing the frequent references to the situation of special schemes and civil servants. This led the legislator to explicitly mention this issue among the general objectives and principles of the pension system, stating that "*individuals covered by social security shall be treated equitably with regard to the pension period and amount, regardless of their scheme"* (Paragraph II of Article L111-2-1 of the French Social Security Code). Furthermore, monitoring disparities between schemes was underlined as one of the specific missions of the new Committee for Pensions Monitoring, since the Act which establishes this committee states that it will be required to "*examine the situation of the pension system, with regard in particular, [...] to comparative pension benefits under the various pension schemes*" (Article 4 of Act no. 2014-40 of 20 January 2014). This issue is also regularly assessed by the Pensions Advisory Council (COR, 2009; 2014; 2015b; 2016a and b) and the *Cour des Comptes* (French National Audit Office) (2003; 2016).

While questions around equity are raised for all special schemes, given their importance in the French pension system, the debate often focuses on comparison between the general scheme, which covers most private sector employees, and civil service schemes[3]. This article also focuses on these schemes.

Beyond the obvious differences in rules and structure between schemes, which are primarily a product of history, the question of the equity or potential inequity of treatment between civil servants and private sector employees[4] is particularly complex – not least because it raises the question of the equity standards to consider, which are not set out in law. In any case, there would be little sense in limiting the question to the similarity or uniformity of rules, because identical rules applied to different groups of people do not always ensure equity, while diversity of rules does not necessarily result in pension inequality. Employment structures, career profiles and pay vary considerably depending on a person's career – in whole or in part – as a civil servant or private sector employee.

These differences in employment structure between the private and public sectors significantly complicate straightforward descriptive statistical comparison between sectors. While mean pension amounts are higher for former civil servants – a mean of €2,520 per month at the end of 2014 for former Central Government civilian public servants, €1,840 per month for former local authority and public hospital workers, and €1,770 per month for former private sector employees, at the end of a full career affiliated to a single plan (Drees, 2016, p.44) – the differences are explained first and foremost by the fact that, on average, the public sector workforce has higher qualifications. We therefore cannot use these differences in their current state to assess whether or not the rules for public sector pension schemes are more "generous". Equally, comparisons between replacement rates (that is, the ratio of total pension amount at retirement to the final salary at individual level) can be deceptive, although the impact of structural effects on this indicator is probably lower than for the pension amount. Although the most recently available data show that replacement rates between the private and public sectors are fairly close to one another - the median replacement rate following a full career is slightly lower for individuals who finish their career in

1. *For example, Germany, Belgium, Spain and even Japan have a specific plan for civil servants, with some specific rules – however in Spain and Japan, this plan has recently been closed to new members. In other countries (e.g. Canada, the USA, the Netherlands, the United Kingdom or Sweden), private pension funds exist alongside the public system which is the same for everyone. These vary between employers, and therefore differ between public sector and private sector employers (COR General Secretariat, 2014a).*

2. *Schemes organised for certain professions (miners, sea fishermen, solicitor clerks and employees, electric and gas company employees, etc.) or operated in some companies (SNCF, RATP, Banque de France, Opéra de Paris, Paris Chamber of Commerce, etc.).*

3. *At the end of 2014, the general scheme represented 12.9 million pensioners, around 82% of all pensioners on French schemes (employees and self-employed workers from the public and private sectors). The various civil service schemes represented 2.8 million people, around 18% of the total (some of these pensioners were also under the general scheme) (Drees, 2016, p. 9). Other special employee schemes accounted for just over 600,000 pensioners, i.e. around 4% of the total.*

4. *This article uses the term "private sector" schemes to refer to those under the general scheme and the Agirc and Arrco supplementary schemes. This is something of a simplification, because some public sector employees are also covered by the general scheme, while some private sector employees are not.*

the public sector than for those who finish their career in the private sector (73.9% and 75.2% respectively) for people born in 1946 (Senghor, 2015, p.5) - this similarity does not demonstrate equal treatment under identical characteristics. Given that the pension system performs vertical redistribution, which means that the replacement rate generally decreases with the end of career salary level, we might have expected a bigger difference in the median replacement rate between former civil servants and private sector employees, given that, on average, they have higher qualifications, and therefore salaries.

Without going into a normative discussion of the definition of equity, this article seeks to explain the differences in pensions between civil servants and private sector employees by illustrating the effect of the rules on pension amounts *for a given wage trajectory*, based on several standard careers, and by detailing the various mechanisms involved. First we outline the main differences between the schemes, and then in the second part, we present the results of our simulations which involve applying current private pension schemes rules to various standard civil service careers. This standard case approach is useful in that it neutralises career characteristics and can therefore be used to isolate and detail the effects of the rules for calculating pensions for the standard careers selected.

The differences between private and public schemes

The issue of coverage

First, it should be noted that contrasting the "public sector" and the "private sector" is not as simple as it seems when it comes to the analysis of pensions.

Pension schemes do not have exactly the same coverage as jobs: some public sector employees are affiliated to the general scheme (contractual public sector employees, and some civil servants) and conversely, some civil servants are seconded to the private sector (DGAFP, 2014a, p. 231 and 389).

Furthermore, the two "blocks" are not homogeneous in terms of employer practices and remuneration policies (Daussin-Benichou et al., 2014). This heterogeneity is particularly prominent in the private sector, especially between corporations, small and medium-sized enterprises and very small enterprises. But it is also present in the civil service, e.g. between the State, local authorities and public hospitals.

Last but not least, neither are the two sets of "civil service schemes" and "private employee schemes" fully homogeneous in terms of pension rules. For civil servants, the rules regarding the minimum age at which pensions become payable, for example, are different for military personnel, civil servants in professions classified as arduous or dangerous[5] and "sedentary" civil servants, the latter having the same minimum age as "private sector" employees. At the same time, the rules are not fully uniform for individuals under private sector schemes. Pension rules are only identical for the part of careers subsequent to 1999. Before this date, at which the Arrco supplementary plan was introduced, pension benefits acquired vary for identical salary levels, depending on the specific rules for each supplementary pension fund. Even after 1999, contribution rates to Arrco are not fully homogeneous, because some sectors still have a contribution rate above the contractual rate. The similarity of rules can also only be considered when the supplementary social security offered by some companies (additional pension schemes, "in-house" retirement indemnities and early pensions) is not taken into account.

We also need to remember that individuals can change plans during their career. A substantial proportion of former civil servants actually have multiple pensions, since part of their career has been spent in the private sector, and they therefore also have private sector employee schemes (Aubert et al., 2012).

Differences in rules

Apart from this issue of coverage, the main difference between "private" and "public" schemes is their respective structures. Private schemes are built in stages and include a basic annuities plan (the general scheme), supplementary points

5. Categories referred to as "active" (firemen, municipal police officers, nurses, healthcare assistants, etc.), "super-active" (national police officers, prison officers, etc.) or "insalubrious" (sewage workers). These are professions which generally have no private sector equivalent. As of 31 December 2012, these categories accounted for 160,000 State employees (around 12% of total numbers), 500,000 public hospital employees (around 60% of all civil servants – an estimation that takes into account the fact that on 1 December 2010, half of nurses chose to be recategorised as category A, and are therefore no longer under the "active" category) and 55,000 local authority employees (around 5% to 10 % of total numbers) (DGAFP, 2014, pages 124-127).

schemes (Arrco and Agirc), and any additional professional schemes, with procedures which can vary a great deal (these schemes are not mandatory and therefore only apply to companies and branches that have decided to implement them). Another difference is associated with the level of annual pay: the proportion of remuneration below the social security ceiling (€38,616 annually in 2016) is covered by the basic plan, the supplementary Arrco plan and any additional company plan, while the proportion of remuneration over this ceiling is only covered by the supplementary schemes (Arrco or Agirc, depending on whether the individual has management status or not), and any other additional schemes.

However, the public sector schemes (the State civil service scheme, CNRACL for local authority and public hospital employees, FSPOEIE for government-employed manual workers), offer annuities and are integrated schemes, i.e. a single scheme fulfils the role of all three stages in the private sector scheme at the same time[6]. The pension rate used under these schemes is therefore higher: for a full career, it is 75% of the reference salary under the civil service plan, as opposed to 50% under the general scheme. Moreover, an additional scheme (the RAFP), operating a points and fully funded system, was created in 2005, but the new plan cannot be considered an exact replica of supplementary private sector employee plans for civil servants, because it applies to a remuneration basis that is totally dissociated from that of the integrated schemes.

In order to fully understand the differences between schemes, we first need to restate the formulae for calculating pensions. These can be expressed, under annuity schemes, as follows:

Pension = pension rate x *prorata coefficient* x *reference salary.*

For points schemes, the formula is as follows:

Pension = early retirement reduction factor x *number of points* x *point value.*

The pension rate for basic schemes and the early retirement reduction factor for supplementary schemes, express the modulation of the pension amount depending on the retirement age and the length of contribution under the basic schemes, via a reduced pension for retiring early or extra pension for retiring late in line with a reference rate. It is therefore determined by the age at which the pension becomes payable (the minimum age at which individuals can retire), the required length of contribution for the full rate (the minimum length required in order to avoid a reduced pension for retiring early) and the age at which the reduction is cancelled out. The prorata coefficient for annuity plans expresses the prorata calculation of the pension amount as a function of the length of contribution under the scheme. It is therefore determined by the reference period for a full career, which defines the length required for a prorata calculation of 100%, and by the methods for calculating the period for which contributions have been made under the scheme. This is higher than the length of employment periods, as it also includes periods of involuntary inactivity (unemployment, sickness, etc.) which are treated as paid up, and additional entitlements (credited for the individual's children). Finally, the reference salary under annuity schemes depends in whole or in part on the gross wages received over the individual's career. It therefore does not depend on the contribution rates that have been applied to these wages, whereas the number of points acquired under points schemes does depend on the contributions paid.

Recent pension reforms since 2003 have aligned some of these parameters between public and private schemes (COR, 2015b, p. 5-6). Since the 2003 reform, the rules have been the same for the length of contribution required for the full rate (they were also the same before the 1993 reform), for the reference period of the prorata coefficient denominator (since 2008, this period has been identical to the period required for the full rate, but they differed between 1993 and 2008) and for the common law legal minimum retirement age (which has always been the same for both public and private schemes – the only differences being exceptions granted to some categories). Procedures for yearly pension increases have also been identical between the integrated civil service schemes and the general scheme since 2004.

For other parameters, the differences between schemes are gradually being reduced, but the process of convergence has been spread over a longer period, and has therefore not yet been achieved. The age at which the reduction is cancelled out (2003 reform) and employee contribution rates (2010 reform) will not be fully aligned until 2020.

6. This article does not cover the additional stage provided by personal pension savings schemes (PERP, PREFON, COREM, etc.), for which individuals are solely responsible.

However, some differences remain: the definition of the reference salary for calculating pensions (salary below the social security ceiling for the best 25 years under the general scheme, and final six months excluding bonuses for civil servants) and employer contribution rates (see below); the measurement of the length of contribution (calendar period for civil servants, period based on an annual salary income threshold for private sector employees); the opportunities for early retirement and additional entitlements credited for specific categories (military personnel and "active" category civil servants); additional quarters for children (2 years per child for mothers employed in the private sector, as opposed to one year – under certain conditions – or 6 months for civil servants, depending on whether the child was born before or after 2004); pension increases for large families (a 10% increase in pension for parents of at least three children, regardless of the number of children under private schemes, but increasing beyond the third child under civil service schemes); minimum pension amounts (the minimum guaranteed level is higher in the public sector than in the general scheme); the returns on supplementary or additional schemes[7]; or the eligibility conditions and calculation methods for reversionary pensions (SG-COR, 2014b; COR, 2015b; *Cour des Comptes*, 2016).

The reference salary calculation can give the impression of being more beneficial for public sector schemes because for ascending wage trajectory profiles, the mean of the final 6 months is always higher than that of the best 25 years. However, this "advantage" is balanced out by the fact that the reference salary for civil service pensions is calculated only on part of their earnings. These earnings break down into a "main" part (the basic index-related salary, which depends on the civil servant's index, and therefore primarily on his or her grade and length of service) and an "accessory" part (bonuses[8], indemnities associated with residence, mobility or overtime, family salary supplement, etc.). However, only the main earnings are taken into account for calculating the pension amount under the integrated civil service plan. Their replacement rate, i.e. the ratio of the first pension to the final total salary, is therefore primarily determined by the proportion of bonuses, and therefore decreases as this proportion rises.

In 2012, "accessory" earnings accounted on average for one fifth to one quarter of civil servants' total pay (DGAFP, 2014a, p. 160 and 186). Between the generation born in 1940 and the one born in 1955, this accessory part, observed at the end of the career, changed relatively little for teachers (whether they are category A or B) and category C staff in the active category (prison officers, etc.), but has significantly and regularly increased for other civil service categories, by +5 to +10 percentage points between the 1940 and 1955 generations (DGAFP, 2014b). It should be noted that an increase in the index-related salary can take place at the very end of the career, which leads to a higher pension amount, sometimes referred to as a "*coup de chapeau*". A Drees statistical study seems to show that this phenomenon is not, however, widespread. For example, between the 5 years before the final year, and the final year of a career, the index of civil servants only increased an average of 4.3% for the generation born in 1942. This increase exceeds 10% only for less than one civil servant in ten (Chantel & Collin, 2014).

The difficulty of estimating contribution levels

Contribution rates differ between private sector employees and civil servants, but also between civilian public servants, military personnel, local authority, and public hospital workers. Analysing them presents a significant issue for comparing schemes.

A simple comparison of mandatory contribution rates (employee contribution + employer contribution) reveals very significant differences: in 2015, in comparison with a non-manager private sector employee, the rate was 14 points higher for a local authority or public hospital worker, and 57 points higher for a civilian public servant (Figure I).

However, this kind of comparison is virtually irrelevant, since the bases from which

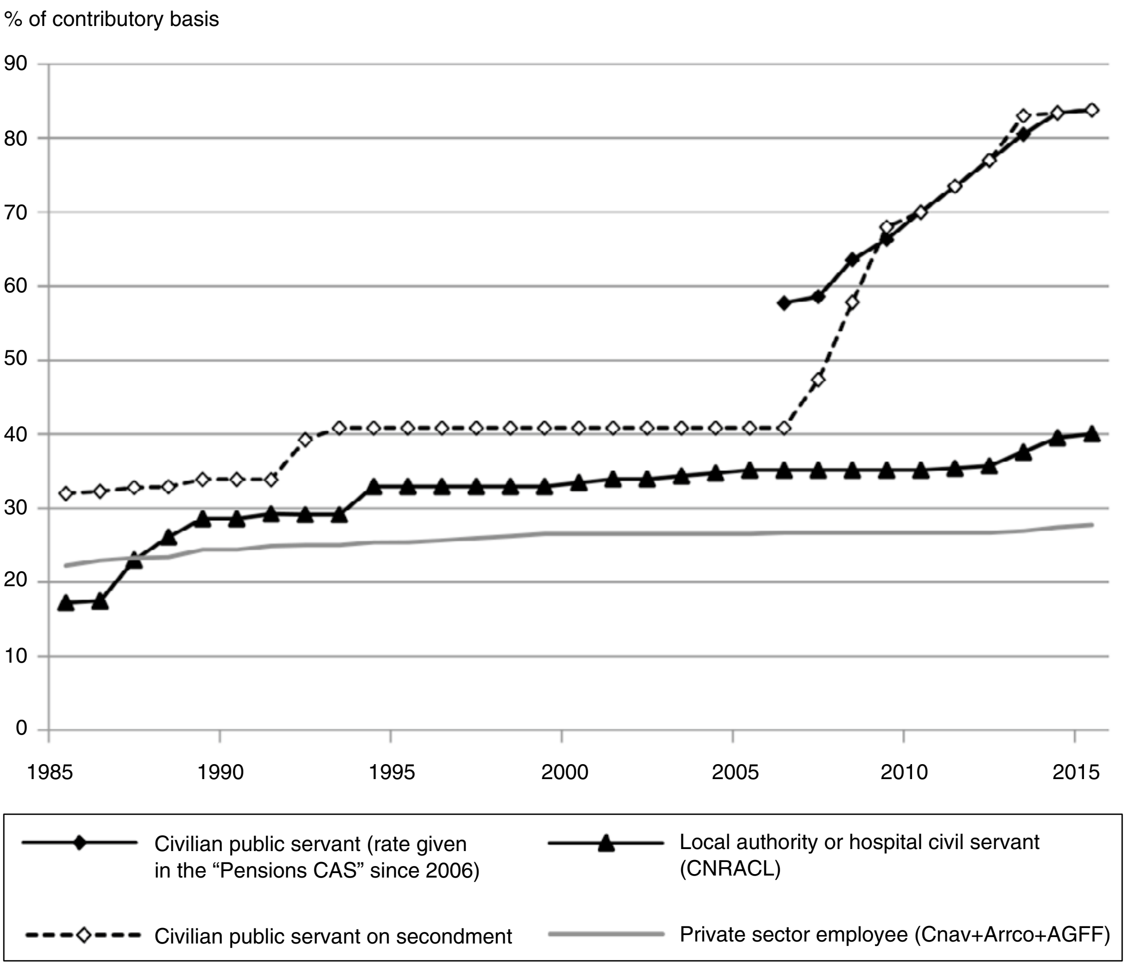
7. The "instantaneous return" is the amount that an individual would obtain in return for an effective Euro contribution if the individual took his or her pension benefit immediately after purchasing it. Under a points plan, it is defined as the ratio between the point value and the point purchase value, multiplied by any call rate. In 2015, the instantaneous yield for Agirc and Arrco was 6.56 % for retirement at the full rate or, taking into account specific contributions that do not generate pension benefits (AGFF contributions, and for managers, CET), 5.21% for non-managers and 5.03% for manager wages below the social security ceiling. For RAFP, this rate is 3.90% for retirement at 62 and 4.76% for retirement at 67.

8. Bonuses in the civil service refer to a permanent component of total earnings; they are not the same thing as the "bonuses" paid occasionally by some employers in the private sector.

contributions are calculated are different, and only represent part of the total earnings. If we consider contributions using a more comparable basis, i.e. the overall earnings including employer contributions, the differences in contribution rates appear to be significantly reduced (in 2013, 15.5% for private sector employees, as opposed to 23.5% for public hospital and local authority civil servants, 35.9% for civilian public servants and 42.2% for military personnel). However, even with a harmonised base, the comparison of contribution rates needs to be interpreted with caution, due to differences in the structure of plan funding – public schemes are funded almost exclusively by social security contributions, while the general scheme receives other sources of funding (SG-COR, 2014b; COR, 2015b and 2016a, p.102-104).

More fundamentally, contributions only give a partial view of employee contribution levels (see Online supplement C2). Some people may accept a lower salary in a sector in return for pension rules that they consider more generous. The lower salary accepted may then be seen as a kind of contribution to pension funding, which needs to be taken into account. So pension comparisons, if we want to be able to consider the contribution levels for various systems, need to take into account salary differentials between sectors, all other things being equal. This makes the analysis extremely complex, because some components cannot be observed, in particular the actual productivity of employees. Analysis ends up being extremely theoretical, since it has to rely strongly on conventional assumptions. It is never really conclusive.

Figure I
Pension contribution rates (employee + employer contribution) since 1985



"Pensions" CAS = "Pensions" special purposes account.
Note: Pension contributions are based on the index-related salary for civil servants, and earnings below the social security ceiling (i.e. segment 1) for private sector employees (non-managers). See Online supplement C1.
Reading note: In 2015, employee and employer pension contributions for local authority or hospital civil servants represented 40% of their gross index-related salary.
Source: legislation.

For this reason, the second part of this article will focus solely on pension amounts, and more specifically on these amounts as a ratio of the final salary – i.e. replacement rates. Since we cannot determine what career and pay each civil servant whose wage trajectory is observed would have had in the private sector, the effect of the pension rules is illustrated by reasoning on the basis of a *given wage trajectory*, i.e. assuming that the wages paid at each age are identical in both sectors.

Disparities in pension amounts for a number of standard careers

Our analysis will involve performing simulations alternating between public sector and private sector rules on a number of standard wage trajectories, based on those developed and frequently used by the Pensions Advisory Council (COR) for its analyses.

The COR has established eight standard careers, four of which are affiliated to the general scheme only, and four of which are affiliated to the civil service scheme only. The simulations will be performed on three of these standard civil service careers. Applying rules from civil service schemes to the careers of private sector employees would present the difficulty of needing first to impute, purely by convention, a breakdown of their wages in terms of basic salary and bonuses. However, it is easy to simulate the application of private sector schemes on wage trajectories for standard cases of civil servants, since we only need to know their total earnings. In practice, these simulations were performed using the CALIPER tool developed by Drees for calculating pension amounts (see Online supplement C3). It is also possible to simulate the application of these rules to just part of a career of a given length, in order to demonstrate the impact of being under both the public and private sector employee schemes over the course of a career.

The standard case approach cannot, and does not intend to, give an overview of the effects of systematically applying the Cnav, Arrco and Agirc rules to all civil servants. It aims to use the example of a few careers under one or more pension schemes (public and/or private) in order to set out in detail the mechanisms involved, and demonstrate the sensitivity of results to some modelling assumptions. A broader perspective would require performing simulations on a representative sample of this population, in order to take into account the weight of each standard career. This article therefore offers a supplementary contribution, and must be read in association with other existing analyses based on representative data, which we will refer to at the end of the article.

The standard career profiles considered

In practice, standard cases correspond to individuals who have worked a full career without interruption, in various civil service categories: a category B sedentary civil servant, whose total end of career earnings include a bonus of around 20% (standard case 5)[9]; a teacher with end of career earnings with a low bonus of around 10% (standard case 6); finally an A+ category manager with end of career earnings with a high bonus of around 33% (standard case 7). The results presented here therefore cover only sedentary categories of civil servants, for whom the rules in terms of age at which pensions become payable and length of contribution required are identical for private employees from the generation born in 1948.

The approach used to build the standard cases was somewhere between a purely theoretical approach which involves selecting individual standard situations by convention, and a purely statistical approach which involves extracting from a sample of observed data a number of real careers that are "representative" of all the others (SG-COR, 2013; COR, 2015a, pages 142-148). More specifically, it is based on statistical analyses of real individual situations to deduce a certain number of realistic career characteristics, in order to produce some stylised standard cases that are simpler than real situations, but are not defined in a completely *ad hoc* way.

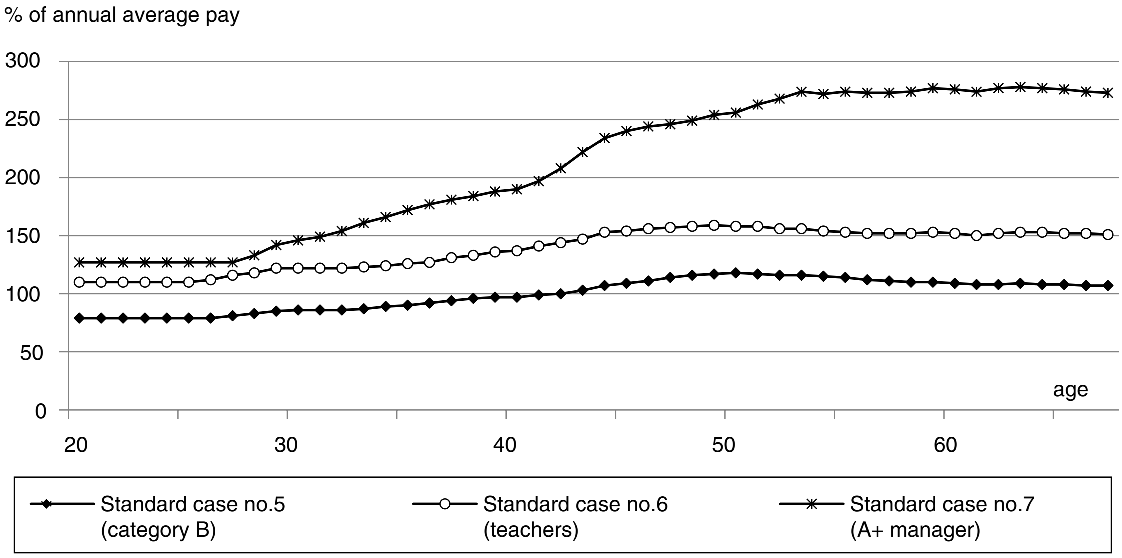
In practice, standard careers are developed on the basis of a statistical analysis conducted by the DGAFP using the Insee civil servant panel (Flachère & Schreiber, 2013). This analysis involved defining, for each standard case, corresponding categories of individuals ("empirical counterparts"), then, for these categories, estimating a wage and bonus proportion profile at each age using the mean values observed for a generation that has completed or virtually completed its career (in this case, the 1950 generation, observed until 2006). The empirical counterpart

9. The term "bonus" is used incorrectly here to refer to all earnings over and above the salary index (including indemnities, overtime, etc.)

for the category B sedentary civil servant standard case generally covers administrative secretaries, inspectors, clerks and higher technicians (excluding, however, category B primary school teachers and police officers). For teachers, it covers accredited or certified teachers and for A+ category staff, magistrates, police commissioners, central administration and local executive managers, engineers, civil administrators, etc.

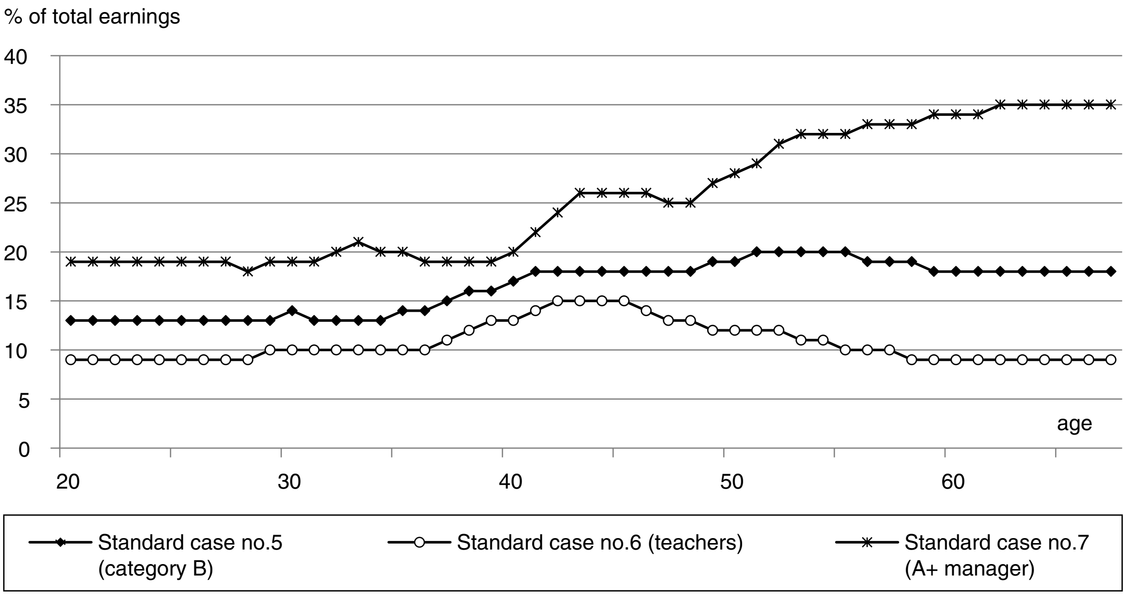
For other generations than the one born in 1950, the *relative* index-related salaries (expressed as a proportion of the annual average pay per capita) and the proportion of bonuses at each age are assumed to be constant and equal to the value observed for the 1950 generation (Figures II et III). This is a conventional assumption that does not take into account actual past changes in the civil service remuneration policy, and in

Figure II
Total earnings as a proportion of annual average pay per capita for the civil servants standard cases



Reading note: salary at 60 for the teacher standard case (no.6) is equal to 152% of the average pay per capita.
Source: COR (2015a) pp. 146-148.

Figure III
Proportion of bonuses in the total earnings for the civil servants standard cases



Reading note: at 60, bonuses (including indemnities and accessory earnings, etc.) represent 9% of total earnings for the teacher standard case.
Source: COR (2015a) pp. 146-148.

particular, effective changes in the index point value. Neither does it cover increases in bonus rates observed over the past 10 to 15 years (DGAFP, 2014b).

The effect of the rules of the different schemes

Table 1 presents the replacement rates at retirement, i.e. the ratio of the first pension payment to the final total salary (including bonuses, etc.) received by individuals covered by social security (both pension payments and salary are calculated net of social security contributions) for standard civil services cases of the 1955 generation, about to take their pension at the full rate in 2017. To give a general idea, final net salaries are around €2,600, €3,600 and €6,800 per month respectively for a category B civil servant, teacher and A+ manager standard cases. The replacement rates are calculated both according to the civil service plan rules (basic civil service plan and RAFP) and the private employee plan rules (Cnav and Arrco for the three standard cases considered, and also Agirc for the standard teacher and A+ manager cases).

For application of the private sector rules, a number of modelling assumptions were adopted according to the contribution rate applied in the Agirc and Arrco supplementary schemes and to whether the private sector rules to civil servants are applied to gross or net identical salaries at each age (Online supplement C1).

As mentioned above, the rules for calculating pensions under the civil service schemes are often seen as more generous, due to the 75% pension rate for a full career, as opposed to 50% for the general scheme, and calculation of the reference salary based on the final 6 months rather than the 25 best years across the career. But this apparent advantage is actually nuanced by the fact that the reference salary is only calculated on the basis of the salary *excluding bonuses* (the additional RAFP plan takes into account bonuses, but this has a very low impact on the replacement rate, because it is only partial and has only applied since 2005). For a given set of total earnings, the pension amount therefore systematically decreases as the bonus proportion of total earnings increases. For the generation that is preparing to retire (born in 1955), the replacement rate is therefore lower for A+ managers (54% replacement rate for an end of career bonus of 33%), than for category B staff (69% replacement for a bonus of 20%), which in turn is lower than for teachers (77% replacement rate for an end of career bonus of 10% of total earnings)[10].

The partial exclusion of bonuses in the calculation of civil service pensions may mean that private sector rules are less beneficial than public sector rules if the bonus proportion is low, and vice versa. The replacement rate for the teacher standard case (case no.6) is therefore higher

10. In the civil service, the bonus rate generally tends to increase with the salary level, and therefore with the qualifications of civil servants. This only applies, however, to employees other than teachers, who are highly qualified but have a low proportion of bonuses, and represent a high proportion of civil servants. Moreover, the correlation between salary level and bonus proportion does not seem to apply to teachers or active category civil servants (Flachère & Pouliquen, 2012).

Table 1
Net replacement rates at retirement as a percentage of final salary for civil servants standard cases according to various public and private sector pension calculation rules (generation born in 1955)

Standard case	Civil service rules	Cnav-Agirc-Arrco rules					
		for gross salary equivalence:			for net salary equivalence:		
		Arrco and Agirc contribution rate max.	Arrco and Agirc contribution rate min.	**Arrco and Agirc contribution rate mean**	Arrco and Agirc contribution rate max.	Arrco and Agirc contribution rate min.	**Arrco and Agirc contribution rate mean**
Category B (case no.5)	69	84	73	76	83	72	75
Teacher (case no.6)	77	76	65	69	75	65	69
A+ manager (case no.7)	54	56	49	52	55	49	51

Note: assuming full rate pension (taken at 62 for the three standard cases). Regulations as of June 2016.
Reading note: the net replacement rate at retirement for a category B civil servant (case no.5) born in 1955 is 69%. If we applied private sector pension rules to this standard case, assuming a net salary in the private sector equivalent to total net earnings (including bonuses), the net replacement rate at retirement would be 75%, assuming mean contribution rates to supplementary pension schemes (Arrco only for case no.5).
Source: CALIPER tool (Drees) and authors calculations.

than it would be if the Cnav, Arrco and Agirc rules were applied (between 65% and 76% depending on the conventions used), while the category B sedentary civil servant standard case (case no.5), whose bonus rate is double that of teachers, has a lower replacement rate than it would have under private sector plan rules (69% versus between 72% and 84%).

However, this relation does vary. Despite the higher bonus rate for category B civil servants and the resulting low public sector replacement rate, a category A+ manager (case no.7) born in 1955 would still be slightly better off under the civil service plan rules by comparison with private sector rules, unless these were applied assuming Agirc and Arrco contributions at the maximum rate (which would give a replacement rate of 55% of the final net salary, including bonuses, i.e. one percentage point more than under the civil service pension rules). This result is only surprising on the surface, because while this standard case presents career characteristics associated with a low civil service replacement rate, it also presents characteristics that lead to a lower replacement rate with private sector rules, i.e. a strongly ascending wage trajectory profile and a high proportion of earnings over the social security ceiling. Calculating the Cnav reference salary as a mean value over part of the career is actually unfavourable, in terms of replacement rate, to individuals for whom the difference is highest between the final salary and the mean reference salary, in particular those individuals with a strongly ascending wage trajectory. Moreover, the fact that the pension amount takes into account all career years under the Arrco and Agirc supplementary schemes, while it is based on just the 25 best years in the general scheme, results in a replacement rate that generally decreases as earnings increase over the social security ceiling, and therefore as the proportion of these supplementary schemes in the total pension increases (Duc & Lermerchin, 2011, p.25-27).

Replacement rates calculated using private sector plan rules also vary significantly depending on the assumptions used for Arrco and Agirc contribution rates (at the minimum, mean[11] or maximum rate). These variations are of the order of 4 to 8 replacement rate points, depending on the standard case considered. Until the mid-1990s, the differences between the minimum and maximum contribution rates were significant: 4 points for Arrco segment 1 (i.e. for the proportion of earnings below the social security ceiling), 8 points for Agirc segment B and 12 points for Arrco segment 2 (Figure IV). These differences dropped significantly between 1995 and 1999 as measures came in for raising minimum mandatory contribution rates, but they did not fully disappear after 1999, because some sectors covered by a collective labour agreement continue to set a contribution rate over the minimum mandatory rate. The longer the career before 1999, the greater the systematic effect of assumptions concerning the Arrco and Agirc contribution rates on the simulated replacement rates for the standard cases. For standard cases born in 1955, this portion represents a little over half the career.

It can be seen from the results above that public sector rules are not necessarily more generous than private sector plan rules (including when we take into account modifications to the Agirc-Arrco rules which will only come into effect as of 2019[12] – see Online supplement C4). This is particularly true when civil servant earnings include a high proportion of bonuses - although this alone is not an adequate condition either (as in the case of an A+ category civil servant). Either way, the preceding analyses for standard cases seek more to highlight the mechanisms at work than draw overall conclusions about whether public sector schemes are more generous than private sector schemes, a task which would be extremely complex given the diversity of civil service career profiles and the changes to these careers and the rules that are applied to them down the generations (see Online supplement C5).

A higher increase in pension with age in the civil service

In Table 1, replacement rates are calculated on the assumption of full rate pensions. Nevertheless, the pension amount and the replacement rate vary depending on the retirement age, in a way that varies depending on the scheme.

11. The mean rate is calculated by the Agirc-Arrco technical departments for all individual covered by these plans. Unfortunately, the mean Arrco rate is not calculated separately for managers and non-managers, so the same value has been used for both in the simulations.

12. The simulations are based on the 1955 generation and do not take into account changes under the October 2015 Agirc-Arrco agreement which will only come into effect from the 1957 generation, and include in particular, the implementation of temporary early retirement reduction factors (for 3 years) in the event of retirement at the full rate under the basic schemes. For this reason, the results were replicated for the 1960 generation in the Online supplement C4. This did not affect conclusions.

Under the Cnav and the civil service plan, working beyond the age at which pensions become payable and the period required for the full rate impacts the pension amount via the application of an extra pension for retiring late proportional to the period worked beyond the minimum retirement age, and to a lesser extent, by improvement to the reference salary (assuming end of career earnings are higher). Under the Agirc and Arrco supplementary schemes, there is no permanent extra pension for retiring late (paid until the death of the pensioner), but individuals continue to acquire pension points, which are paid as an additional pension. Finally, under the RAFP, a permanent extra pension is applied in the event of retirement after the age at which pensions become payable. Working beyond the age at which benefits become payable therefore increases the pension amount, both through a higher extra pension and more points.

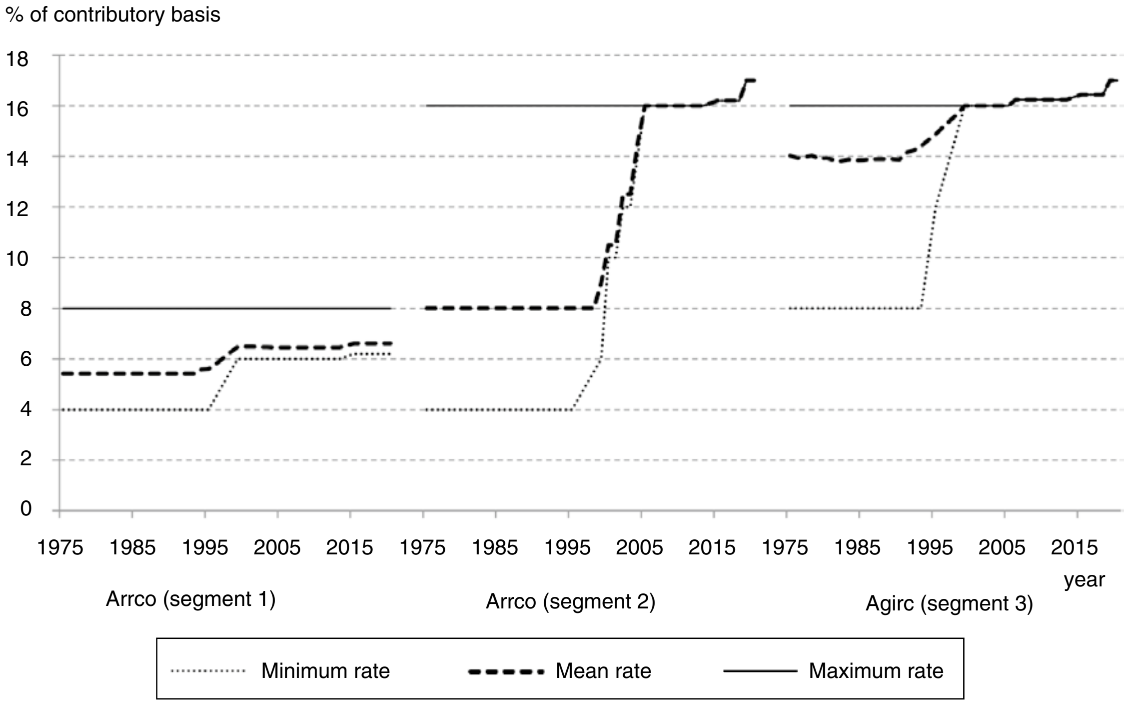
The increase in pension associated with working beyond the minimum age can therefore vary depending on the wage trajectory profile (Aubert, 2017). The scales can give an initial idea of the orders of magnitude involved. In the basic and integrated schemes, the increase in pension amount represents + 5% for an additional year of work (according to the extra pension for retiring late scale), plus any increase in the reference salary (the average increase is of + 1 percentage point for private sector employees). Under supplementary schemes, working for an extra year leads to additional points of around + 2.5% (e.g. ≈ 1/41 for a 41 year career), after which a term depending on the difference between the end of career salary and the mean salary across the career is added or deducted.

In practice, for the three standard cases, assuming an individual born in 1955, the increase in pension associated with working beyond the minimum age at which benefits become payable is higher with the civil service than the private sector rules. For example, for retirement at 67 rather than 62, depending on the standard case considered, the increase goes from + 26% to + 28% in the first case, versus + 17% to + 21% in the second case (case no. 5 and no. 6, respectively, in Figure V).

The impact of coverage under both public and private schemes over a career

In the same way that we applied public and private sector pension rules to the overall wage trajectories of the three standard COR civil service

Figure IV
Arrco and Agirc minimum, mean and maximum contribution rates



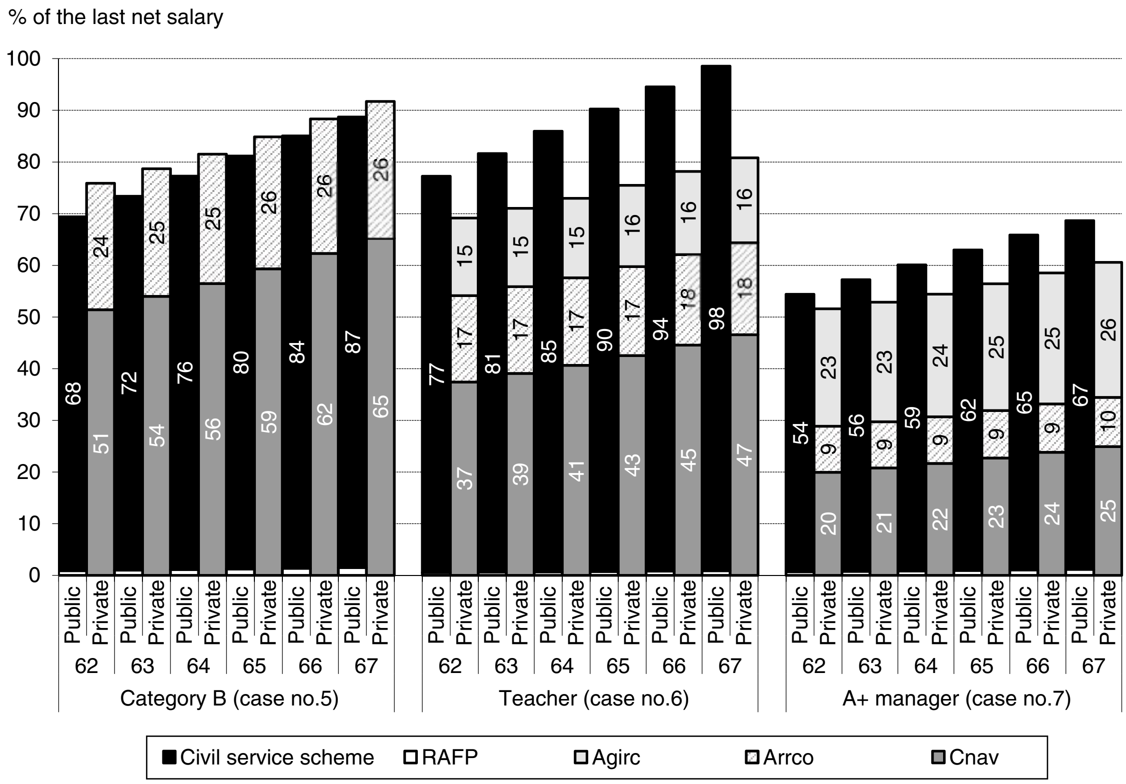
Note: regulations as of June 2016.
Source: Agirc-Arrco.

cases, it is also possible to apply them to just parts of these careers, in order to model, on a conventional basis, situations where employees have contributed to one and then another plan over the course of their career. These simulations are performed below, assuming a contribution rate at the mean Arrco and Agirc contribution rates and equivalent net salaries between the public and private sectors. The latter assumption supposes that an individual changing from one sector to another would maintain the same salary. This assumption, which is only justified here by the purpose of demonstrating the "pure" effect of the pension rules, is not always observed in reality. In practice, changing employment sector does not necessarily provide an immediate increase in salary (after one year), and can even lead to a slight decrease in salary in the short term, but there is often a medium-term salary increase (after 5 years) (Daussin-Benichou et al., 2014).

We simulate a number of profiles of individuals contributing to several successive schemes, with various lengths of employment in the private sector (5, 10, 15, … up to 35 years) and chronological sequence of contribution (public sector followed by private sector, or private sector followed by public sector).

In most situations where individuals have contributed to both public and private schemes, with identical periods worked and salary levels in each sector, replacement rates are higher when individuals finish their career in the private sector rather than in the public sector (Figure VI). The methods for calculating the reference salary for each sector have a strong impact on this result. As only the final salary (excluding bonuses) is taken into account for the civil service plan, while the 25 best years are taken into account under the general scheme (and the whole career for supplementary schemes), starting one's career in the public sector makes it possible to exclude salaries from the beginning of the career, which are the lowest, in calculating the pension amount,

Figure V
Net replacement rate at retirement depending on retirement age (generation born in 1955)



Note: regulations as of June 2016. Assuming net salaries would be identical under application of public or private rules. Assuming mean Agirc, Arrco and RAFP contribution rates across the period and constant projected yields (yearly pension increase of purchase and point values with inflation).
Reading note: retirement at 67 gives standard case no.5 a pension rate of 89% (87% for the civil service plan only). Retirement at the same age under application of private sector pension rules gives a pension rate of 92% (65% thanks to the Cnav pension and 26% from the Arrco pension).
Source: CALIPER tool (Drees) and author calculations.

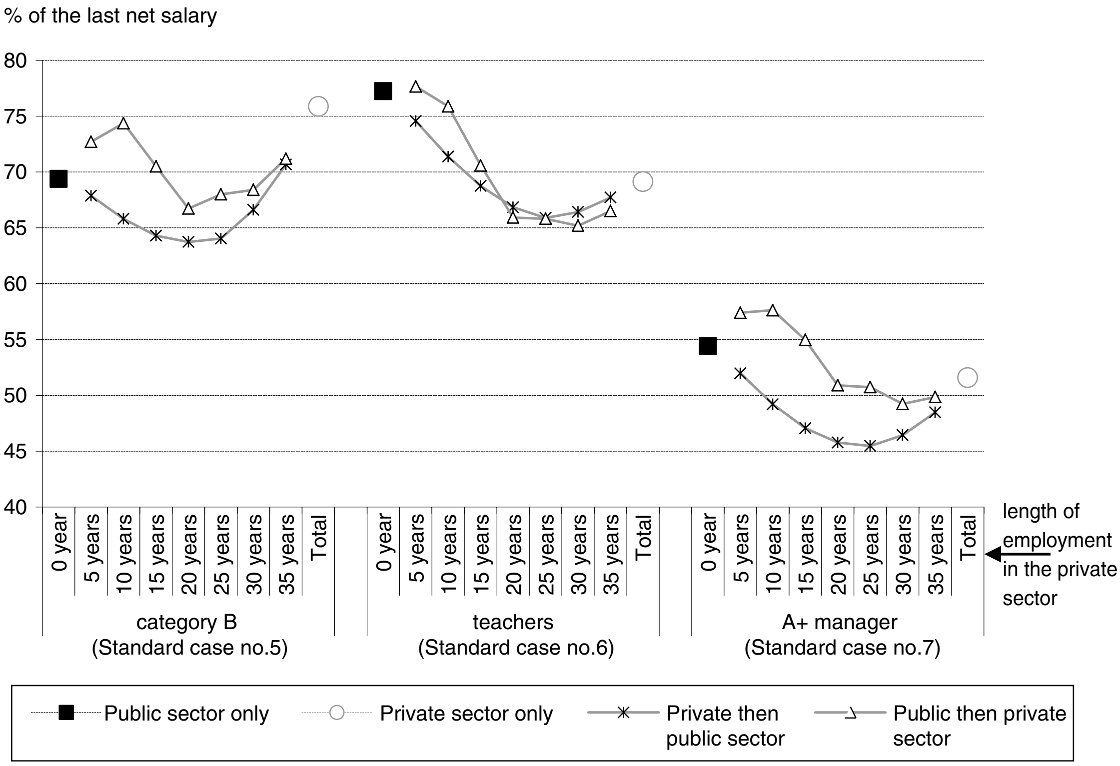
while starting one's career in the private sector means the low initial salaries are taken into account when calculating the reference salary[13], and therefore the pension amount[14].

However, this is not always the case. In the example of the standard case of a teacher born in 1955 having contributed to both public and private schemes, with a fairly long period in the private sector (20 years or more), the replacement rate seems a little higher for a career ending in the public sector, than for the other way round. The addition of a low bonus rate and earnings over the social security ceiling make this standard case the case with the greatest loss in pension under the rules of the Cnav, Arrco and Agirc rather than the civil service rules. A longer period in the private sector therefore has a more negative effect in the second part of the career than in the first part, because it is during this period that the highest salaries are paid, i.e. those that contribute the most to the total pension amount.

Finally, of the three standard cases studied, the highest and lowest replacement rates often correspond (given salary assumptions) to situations of multi-coverage, where the individuals have contributed to both public and private schemes. For example, for the category A+ manager standard case (case no.7), the highest replacement rate is obtained for a career that starts in the

13. If the period of employment in the private sector is under 25 years, all annual salaries are taken into account.
14. In addition to the selection of the years used for calculating the reference salary, the replacement rate also depends on the way in which pension benefits acquired under the initial plan increase yearly. If the change in employment sector occurs after 2004, these yearly pension increases are identical under the Cnav and civil service schemes. For a civil servant who leaves the public sector before retirement, the final salary increases yearly according to the same index used for pensions paid (in application of the final paragraph of Article L. 25 of the French Civil and Military Pension Code), i.e. since 2004, according to price changes apart from tobacco, as per the pensions and salaries under Cnav. However, for those leaving the public sector before 2004, the yearly pension increases applied up to this date correspond to changes in the civil service index point value, plus the effects of any possible yearly increases for categories. The simulations presented here do not, however, take into account such yearly increases for categories.

Figure VI
Net replacement rate at retirement for civil servant standard cases according to career period in private and public sectors (generation born in 1955)



Note: regulations as of June 2016. Assuming net salaries would be identical under application of public or private rules. Assuming mean Agirc and Arrco contribution rates across the period.
Reading note: if standard case no.5 had spent the first 20 years in the private sector rather than a complete career in the civil service (for identical net salaries), the replacement rate would be 64% rather than 69%.
Source: author calculations.

public sector and ends with 10 years in the private sector, while the lowest rate is achieved when the career starts with 25 years in the private sector, before going into the public sector. This final result underscores the fact that the impact of contributing to both public and private schemes on the pension amount may vary Depending on the career characteristics, it can make individuals either better or worse off.

* *
*

To summarise, the simulations performed on the examples of the three COR standard cases for sedentary categories of civil servants produce the following results.

First, the impact of applying private sector plan pension rules rather than civil service pension rules varies. For the generation preparing to retire (born in 1955), applying private sector rules would be more beneficial for a category B civil servant, but less beneficial for a teacher, and slightly less beneficial for an A+ category manager. However, these results vary depending on the private sector rules applied, and in particular, the Arrco and Agirc contribution rate used. These results correspond to the situation following the convergence of some public and private sector pension rules, which began under the 2003 pension reform. Second, this analysis is likely to change significantly in the future, even if legislation remains the same, in accordance with developments in the determinants under each set of rules, i.e. the bonus proportion of total end of career earnings in the civil service, and the mean rate of salary growth in the private sector. Monitoring these factors, among other things, is therefore essential for assessing the equity between schemes, which is constantly changing over time. Third, the pension amount also depends on the retirement age. For the three standard cases studied, the increase in pension due to working beyond the minimum retirement age seems higher under public sector rules, due to the extra pension for retiring late in the basic and integrated schemes (which would be more beneficial than accumulating points in private supplementary schemes). Finally, the impact on pension amount of individuals contributing to both public and private schemes over their career, assuming identical net salary at each age, varies. It can be either positive or negative by comparison with an individual under a single plan (using public or private rules). Of the three examples studied, for a given career period in each sector, moving from the public to the private sector usually leads to a higher replacement rate than moving from the private to the public sector, although this is not systematically true.

As we have stated on a number of occasions, the COR standard cases are not "representative" of the entire civil service, or even of all civil servants in their category. It is therefore impossible to extrapolate the results associated with them, and they need to be interpreted as three examples that illustrate the mechanisms in play, and test the sensitivity of the assumptions and conventions used. The analysis presented here therefore serves to supplement simulations on representative samples of individuals covered by social security, such as those performed by Beffy and Blanchet (2009) using the Insee DESTINIE microsimulation model or, more recently, Duc (2014) using the Drees contributor inter-scheme sample (EIC) data from 2009. These analyses of representative samples confirm the main lessons of these standard cases, in particular the fact that the impact of applying private sector pension plan rules, rather than public sector plan rules varies, and that the effects differ a great deal depending on civil servant characteristics. For example, according to Duc (2014), assuming constant net salaries, a little over half of civil servants born in 1958, would have a higher pension under private sector rules, while for other civil servants, the pension is highest under public sector rules (according to legislation in force when this study was conducted, and before the Agirc-Arrco agreement of 30 October 2015).

These kinds of representative sample analyses can also be used to put the results of these standard cases into perspective with regard to the diversity of actual civil service careers. In particular, the results of Duc (2014) suggest that the conclusions for the A+ manager standard case, for which the simulations suggest that public sector pension rules are slightly more beneficial, would only apply to a minority of civil servant managers in reality. Overall, around six out of ten civil servant managers born in 1958 would have a higher pension under private sector rules, for net equivalent salaries at all ages. More generally, the proportion of civil servants who would have a higher pension under Cnav, Agirc and Arrco rules is higher among women, sedentary categories, managers and individuals who started their career in the private sector and finished it in the public sector. It would, however, be below 50% for men, active categories, and individuals who started their career in the public sector but finish in the private sector (Table 2). These effects also take into account some differences in rules between private and public schemes which have no impact on the standard cases, due to the simplified nature of the careers they represent, and specifically the fact that they are assumed to have no children and that their length of employment (apart from the retirement year) always corresponds to full calendar years. In particular, for women, additional entitlements due to children provide longer periods of contributions under the Cnav rules (two years, rather than one year or six months, depending on whether the child is born before or after 2004, under civil service schemes[15]). However, for parents of four children or more, public sector plans apply higher pension increases than private schemes. Furthermore, for years that have only been partially worked, private sector schemes sometimes mean that four quarters can be counted (since the number of quarters used is defined based on the total annual salary). This is not the case with civil service schemes (quarters are counted depending on the calendar period worked).

Whether they are performed on standard careers or a representative sample, the simulations involving application of pension rules from one sector on another cannot be used to draw conclusions on the relative "generosity" of these sectors. Analysing the calculation of pension amounts, *assuming constant salaries and identical retirement age,* leaves open the question of earnings that would have been received in the public sector if other pension rules were in place, since higher pensions can, in some cases, serve as compensation for lower salaries, and the question of the retirement behaviour of individuals. The analysis also provides no information on differences in the rates of return on contributions. It compares pension levels without examining past contribution rates for the two groups of employees.

Whatever the case, considerations to align or standardize the rules applicable to the various French pension schemes must go beyond a simple comparison of the rules or their impact, all else being equal, just as the impression of equity or inequity held by some individuals covered by social security would not be justified by the results of such comparison. These reflections also raise the issue of comprehensibility and transparency that the legislator intends to give to the pension system, and more general reflections on its overall structure (see COR, 2015b, pages 11-12). ☐

15. Furthermore, additional entitlements due to children apply to calculation of both the reduced pension for retiring early/extra pension for retiring late and the prorata coefficient under Cnav, while it only applies to the reduced pension for retiring early/extra pension for retiring late for children born after 2004 under civil service schemes.

Table 2
Results of a simulation in which private sector pension rules are applied to a representative sample of civil servants born in 1958 (as per Duc, 2014)

	Mean variation (in %) of the pension amount under application of Cnav-Agirc-Arrco rules (for mean contribution rates) rather than civil service rules	Proportion of individuals covered by social security (%) for whom the most beneficial pension rules are…	
		… the civil service plan rules	… the Cnav, Agirc and Arrco plan rules (for mean contribution rates)
Total	+ 2.4	47	53
Men	+ 0.9	53	47
Women	+ 3.9	43	56
Sedentary category	+ 3.8	44	56
Active category	- 1.7	56	44
Non-managers	+ 1.0	50	50
Managers	+ 4.9	41	59
Public and private plans, primarily private	+ 0.7	54	44
Public and private plans, primarily public	+ 2.9	45	55
Public sector plans only	+ 3.0	48	52

Note: regulations as of April 2014. Assuming identical net salaries at all ages and retirement without a reduced pension for retiring early under the civil service plans. In the data used, careers were observed up to 51 years old (until 2009 for civil servants born in 1958). Changes after this age were simulated using the Drees TRAJECTOIRE model. The percentages do not add up to 100%. The difference corresponds to cases where the two types of rules give the same pension amounts.
Coverage: civilian public servants born in 1958, excluding military personnel and staff retired before 54.
Source: Duc (2014).

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Taking contributions into account in public-private comparisons of pensions

Antoine Bozio *

Comment on "Differences between public and private sector pensions: an analysis based on career profile simulations" by P. Aubert and C. Plouhinec

Comparisons between pension schemes are often distorted by the use of simple and misleading indicators, such as mean pensions, or mean replacement rates. Aubert and Plouhinec (this issue), comparing the calculation rules in the public and private sectors for given careers, highlight that replacement rate comparisons are not unequivocal. Such work makes it possible to gain a better understanding of the mechanisms involved, which are more complex than they first seem, and to highlight the heterogeneity of situations in the civil service. However, such comparisons do not make it possible to assess the relative generosity of the pension schemes because they are, to a large extent, contributory. This comment suggests that while comparing contribution efforts is, admittedly, complex, it is not beyond the realms of feasibility. Such a comparison would offer the advantage of being distinct from the related, but separate, issue of comparing total pay (immediate and deferred) between the public and private sectors. Finally, in the light of that work, recommendations for reforms are made, aiming to transform the "pensions" Special Account (*Compte d'Affectation Spéciale (CAS)* "*pensions*") into a pension fund for Central-Government civil servants, and gradually to incorporate bonuses into the contribution base on which civil servants' contributions are calculated.

JEL Codes: H55, J26.

Keywords: pension, replacement rate, public-private sector comparison.

The article by Patrick Aubert and Corentin Plouhinec is a welcome piece of work on the issue of the differences in pension entitlements between the public and private sectors. The issue is important not only for implementing good public policies on pensions, but also for policies on civil service pay. It is also a highly sensitive issue in the public debate, which is why rigorous studies giving a clearer vision are more than necessary.

In the minds of the general public, and indeed of many experts, it appears understood that public sector pensions are more generous in France than private sector pensions. This conviction results in strong oppositions between professional categories who mutually accuse each other of being "privileged". The fault lines are not limited to opposition between private and public. They also divide employees and self-employed, civilian and military public servants,

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Translated from « Comparer les efforts contributifs pour comparer les retraites entre public et privé ? ».

civil servants who receive large bonuses and teachers who have few bonuses, those who are entitled to special pension schemes and those who are not, etc. These feelings that the pensions system is unfair reduce trust in it, and lead to many calls for convergence of the public and private pension schemes[1].

This comment firstly revisits why commonly cited indicators are misleading, and then analyses the results proposed by the authors. Those findings are deemed to be robust and do not really deserve criticism, since the authors are well aware of the limitations of the exercise. Where this comment diverges from the conclusions of the article is on whether or not it is relevant to think that comparing pension schemes on the basis of contribution efforts would be more appropriate Finally, by way of a conclusion, we discuss the possible reforms that would make it possible to achieve greater fairness and transparency in the pension entitlements of the various mandatory schemes.

Simple but misleading indicators

The idea that pensions are more generous in the public sector than in the private sector is generally based on indicators that are simple, but very often misleading.

The mean public/private pension gap

The first indicator, often cited in the press, is the mean public/private pension gap. As Aubert and Plouhinec remind us, in 2013, retired Central Government civil servants received nearly 2,520 euros a month in pension while retired private-sector employees received, on average, a monthly pension of 1,770 euros (Drees, 2016). Naturally, that measurement does not reflect the difference in generosity of the pensions systems, but rather, quite simply, that, in a contributory system, pensions are proportional to the salaries and wages on which the contributions are based. The comparison could possibly have meant something in a world in which the public sector and the private sector were entirely comparable. In reality, public service production (health, education, etc.) requires staff who, on average, are more qualified than in the private sector, and civil servants therefore, on average, have higher pay than private-sector employees. This does not mean that, for the same levels of qualification and of responsibility, civil servants are better or less well paid than private-sector employees, and even less so that their pensions are more or less generous.

The calculation rules

A second item that is difficult to dispute is the fact that the rules for calculating the pensions are different. Without recapitulating all of the differences, the two main ones are the salary of reference and whether or not bonuses are taken into account. In the civil service, the salary of reference is the pay received for the last six months before retirement, whereas, in the private sector, the general basic pension scheme uses the salaries for the 25 best years, and the compulsory supplementary pension schemes use all of the contributions throughout the career. For a full and upward career, it is intuitive that a calculation based on the last salary will be more advantageous than a calculation based on the whole career. But here too, this reasoning does not take into account how the whole system fits together: for a given contribution rate, a calculation based on the last salary or on all of the salaries is not necessarily more or less advantageous. It is favourable to certain career paths (upward ones) and unfavourable to other, less dynamic career paths (with a complex mechanism of ceilings being applied). The fact that bonuses are taken into account only very partially in calculating civil service pensions is another argument, this time used in defense of civil servants: on average, bonuses account for 23% of civil servants' pay, civilian pensions mechanically represent a lower rate of replacement of the total pay than for private-sector workers (for whom bonuses are fully included in the contribution base and in the salary of reference). However, the argument remains limited in terms of unfairness: when contribution efforts are different (higher on basic pay, lower on bonuses, lower on self-employment income, etc.), is that synonymous with unfairness? Not really.

The mean replacement rates

The gap in replacement rates – that is, in the ratio of pension to last pay before retirement – has often been used to measure the relative

1. *For instance, we could cite the OECD (2016, chap. 6).*

generosity of pension schemes[2]. The work done by Drees (*Direction de la recherche, des études, de l'évaluation et des statistiques* – the Directorate for research, evaluation, and statistics of the French Ministry of Social Affairs and Health) based on the inter-scheme sample of retirees thus highlights that the replacement rates in the public sector are similar to (or indeed slightly lower than) the replacement rates in the private sector. Senghor (2015, Table 2, p. 5) thus finds a median replacement rate of 72.1% for former public-sector employees as against 73.8% for former private-sector employees. This indicator is more pertinent than comparing mean pensions, but it is not an indicator of pension generosity: firstly, when replacement rates are equal, it can easily be understood that an insured person who can retire earlier enjoys a pension that is more generous. Secondly the mean (or median) replacement rate, like mean pensions, is dependent on the structure of the studied population: a structure of contribution rates decreasing with income combined with non-contributory items intended for small pensions results in replacement rates decreasing with the pay level.

The study by Aubert and Plouhinec (2017)

The analysis conducted in the article by Patrick Aubert and Corentin Plouhinec consists in comparing the part played by pension calculation rules in the public sector and in the private sector by applying them to three standard career paths in the civil service (a "category B" grade employee with an end-of-career bonus rate of 20%, a teacher with a low bonus rate, and a "category A+" grade executive with an end-of-career bonus rate of 35%). The replacement rates are compared, but at given wage paths, making it possible to show the impact of differences in rule while controlling for composition effects between sectors.

Results illustrating the complexity of comparing pension calculation rules

The authors highlight that such comparisons are not unequivocal: the public sector rules clearly advantage the standard case of teachers (replacement rate of 77% as against 69%), marginally advantage the A+ executive with bonuses (54% as against 51-52%), and disadvantage the standard case of the category B employee with a bonus rate of 20% (69% as against 75-76%).

Those calculations are good illustrations of the interactions between different rules that make comparison particularly difficult: thus, although it can be intuitively understood that the bonus rate tends to depress the replacement rate with the public-sector rules, the role of the social security ceiling and the interaction of entitlements to the general basic scheme and in supplementary schemes explain why the A+ executive with a very high bonus rate obtains a similar replacement rate, whether the public-sector rules or the private-sector rules are applied.

Another interesting point is that the gains made by postponing retirement are greater with the civil-service calculation rules: the authors found that postponing retirement from 62 to 67 yielded a gain of from 26 to 28% with the public-sector rules, as against only 17 to 21% with the private-sector rules. This finding emphasises that, in spite of the rules being aligned in appearance (extra pension for retiring late, reduced pension for retiring early, and required length of the contribution period), differences in career path and in the definition of the salary of reference play major parts in explaining incentives for postponing retirement.

What conclusions can be drawn from the article?

Although the authors do indeed show that we cannot unequivocally conclude that the pensions calculation rules for the public sector are more generous, can we nevertheless conclude that they are more generous for certain sub-populations such as teachers? Should we conclude from this work that it would be legitimate to reduce the pensions of public-sector teachers and, to a certain extent, the pensions of A+ executives, and to increase those of category B employees? Or, conversely, should the bonus rate of teachers be increased (thereby lowering their replacement rate) or should the bonus rate of the other civil servant categories be lowered?

2. *See Andrieux and Chantel (2012),* Conseil d'orientation des retraites *(Pensions Advisory Council) (2014), and, more recently, Senghor (2015).*

These questions show that comparing replacement rates, even on standard careers, only gives little information about the compared generosity of pension calculation rules. Indeed, the authors do not claim anything different, reminding readers that their aims were, above all, to "shed light on the mechanisms involved".

Is it possible to compare the generosity of pension schemes?

The authors are aware of the above-mentioned limitations, but they deem that the option of comparing contribution efforts between schemes is not very pertinent. Well, I do not share that pessimism. Although I am aware of the difficulties involved in such an exercise, I think it is possible to achieve a pertinent comparison of pension schemes, provided that, conceptually, we clearly separate the issue of the relative "generosity" of the pension schemes from the comparison of total pay (net and deferred income) between sectors at equivalent levels of qualification, arduousness, and responsibility.

What is the "generosity" of a pension scheme?

The generosity of a pension scheme is often understood in ordinary language as being the amount of the pension or the total amount of spending on the pension. There is not much sense in using such a perspective for comparing two contributory pension schemes: for equivalent career paths, if some people choose a higher contribution rate and lower net salaries, that does not mean that their scheme is more "generous" but rather, simply, that their contribution efforts are higher. In a contributory pension system, the best indicator of the relative generosity of one scheme compared to another is the return on contributions (or, in more technical terms, the "internal rate of return"): to what extent one euro contributed to the civil service pension scheme yields more or less pension entitlements than in private-sector schemes. In such a context, if one scheme leads to higher replacement rates or to earlier retirement due to a higher contribution effort, and therefore due to lower net salaries, we should not find anything to complain about. The only thing that should count is the ratio between the benefits received and the contributions paid.

Does comparing contribution efforts really lack pertinence?

The authors rightly emphasise (cf. on-line supplement C2) that it is difficult to measure the real contribution effort of Central Government civil servants: the employer contribution rates given by the "pensions" Special Account (*"pensions" CAS*) are calculated as the subsidy for balancing the CAS relative to the contribution base (the salary). It is not an actual contribution rate, reflecting the contribution efforts of the civil servants, mainly but not only because the expenditure covered includes a large share of non-contributory expenditure which, in the private-sector schemes, is covered by taxation.

These very real difficulties are not insurmountable: there is indeed an actual contribution rate that could be calculated by subtracting the non-contributory entitlements from the "pensions" CAS expenditure, and that should be differentiated among civilian public servants between active and sedentary categories. Such a rate could be compared with the retirement pensions received, and compared with the return measured in the private-sector schemes[3]. Differences would definitely be possible, but personally I doubt whether they would be as significant as the gap in replacement rates mentioned earlier. However, what might happen is that we find contribution rates on gross civil service salaries to be higher than on gross salaries in the private sector, and contribution rates that are lower on the bonuses. If such an observation were found to be true, that would take us from the issue of the generosity of the pension scheme to comparison of overall pay.

What is the economic incidence of pension contributions?

Before we discuss the ways of comparing the overall pay between sectors, it is necessary to go back over the issue of contributory effort from an economic point of view, and not only because of the problems of measuring the employer contribution rate in the Central Government civil service. Both in the private sector, and in the

3. The other option is to calculate, for a representative sample, the present values of the pensions, for the public and private sectors, and then to compare net salaries and pension entitlements while checking for the characteristics of each sector (Colin et al. 1999, Disney et al. 2009).

public sector, we need to address the issue of the economic incidence of pension contributions: who ultimately pays the contributions? The employee or the employer? The issue of the economic incidence of social contributions is one of the major issues of public economics, for anyone wanting to study the impact and the effectiveness of social insurance systems. Economic analysis generally leads to the idea that the social contributions that fund contributory entitlements are fully borne by the employees in the form of lower net salaries. In the standard framework of labour market analysis (labour supply versus labour demand), although employees do incorporate the expected pension entitlements in their total pay, their labour supply is not affected by a rise in pension contributions, and such a rise results in a reduction in net salary (Summers, 1989; Kotlikoff & Summer, 2002). Empirical analyses remain scarce, but the most convincing evidence of social contributions having an incidence on employees comes from cases where the contributory relationship is visible and obvious (Gruber, 1997). Recent work on French data reinforces this observation (Bozio, Breda & Grenet, 2017). However, do pension contributions in the public sector have an economic incidence similar to what has been shown for the private sector? Nothing makes it possible to think so, since pay processes in the civil service follow quite distinct mechanisms. This is where comparing overall pay (or labour cost, that is the addition of net salary, employee and employer social contributions), between public-sector and private-sector employees, could be relevant: to what extent are gaps in total pay, at given qualification and job characteristics, detectable?

Comparing total pay, both immediate and deferred?

This subject is addressed in a considerable amount of literature on pay in the public and private sectors, in France and elsewhere[4]. This comment does not aim to review that literature but merely to point out that comparing total pay might correspond better to addressing the question asked by certain commentators about the compared generosity of pension schemes. Fundamentally, the exercise is difficult because beyond checking for qualifications, experience, and job location, it is difficult for economists to check for the degree of disutility of the respective jobs. Generally, the most convincing work uses occupations held by people in both the public and the private sectors (e.g. nurses, teachers), and uses panel data to estimate the total pay in each of the sectors, and the impact on mobility that pay gaps can have.

What reforms could reduce inequalities and the feeling of unfairness?

The preceding discussion should not conceal the fact that, in order to avoid unjustified inter-category conflict, it is possible to improve the transparency and legibility of the French pension system. Here, I am largely using ideas put forward in Newsletter No. 12 from the Pensions Advisory Council (COR 2015).

Transform the "pensions" CAS into a Central-Government civil servants' pension fund

The first essential thing to do is to leave behind the budgetary logic of the "pensions" CAS, which leads to the of employer contribution rates that do not have any real economic sense. This reinforces the feeling that Central-Government civil servants enjoy a privileged or preferential pension scheme, and fuels speculations as to the extreme generosity of the system, speculations that Aubert and Plouhinecshow to be ill-founded. Within such a pension fund, it would be necessary to distinguish between the common basis for all Central-Government civil servants (common contribution rate), and specific additional contribution rates for military personnel and active civil servants. Non-contributory entitlements should not be included in the financing of such a fund, but rather they should be funded through tax, as in the private sector – this is purely a bookkeeping change, since the employer contribution rate is naturally ultimately funded by tax. The advantage of such a reform would greatly clarify the Central-Government public service pension system, and, in addition to the resulting transparency gain, it would lead to facilitating mobility between the various components of the civil service.

4. See, for example, Postel-Vinay and Turon (2007) for an example, or Gregory and Borland (1999) for a review of the literature.

Include bonuses in the contributory basis

The current situation, in which the State (Central Government) organises in its own way the payment of bonuses not liable for contributions, is absurd and a source of multiple malfunctions, both for pensions but also for pay policy in the civil service. The inclusion of bonuses in the contributory basis can be done in various ways – probably progressively – that do not all have the same consequences for the civil servants in question and for public finances. Firstly, it is possible to increase the "employee" contributions on bonuses, thereby lowering the net bonuses paid, but proportionally increasing the pension entitlements on those bonuses. This operation would be neutral for public finances, and for the civil servants, but not necessarily desired by the latter. Alternatively, it is possible to increase the contribution rate on bonuses (employer and employee contributions) until parity is obtained with the contribution rate on basic salary. Such a choice would lead to an extra cost for public finances and to a gain in terms of pension entitlements for civil servants who receive significantly large proportions of their pay in the form of bonuses, thereby accentuating the total pay gap with civil servants who receive small percentages of their pay in the form of bonuses, that is, teachers.

Regardless of the option chosen, it would be necessary, on that occasion, to review the differences in pay structures between the different categories of civil servants. ☐

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Lifecycle deficit in France: an assessment for the period 1979-2011

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National Transfer Accounts (NTA) measure the way in which individuals produce, consume, save, and share resources at each age. They make it possible to identify the periods for which private and public consumption (education, healthcare, etc.) is not funded by labour income, before identifying the transfers between the ages that enable such consumption to be funded. This article presents individual age profiles of consumption and labour income in France, as established using that method, and how they changed from 1979 to 2011. The profiles are also calculated at aggregate age level, highlighting the importance of changes in the demographic structures over time. We also reconstruct partial cohort trajectories, thereby providing a generational reading of the changes.

In 2011, consumption by old people was higher than consumption by young people, which was not the case in 1979. The rise in consumption at each age, observed generation on generation, slowed down as from the cohort born in 1950. The range of ages at which labour incomes are received has narrowed, while the age at which labour income reaches its highest level has shifted from 36 to 46 over the years. The increase in labour incomes, observed at each age in the generations from 1930 to 1950, seems to have been interrupted momentarily between the 1950 and 1960 generations, at least at the beginning of working life. It resumed in the generations from 1970 onwards, but to a less pronounced extent. In 2011, the ages at which consumption exceeded labour income, corresponding to a deficit, ran from 0 to 24 and from 59 to 82. With the rise in life expectancy in France, the number of years in a deficit situation at high ages has increased considerably, going from 14 to 24 years between 1979 and 2011. Finally, the labour income and consumption profiles for France are very similar to those of the other European countries.

JEL Codes: E21, E24, J10, J11.

Keywords: consumption, labour income, lifecycle, age profile, National Transfer Accounts, intergenerational transfers.

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Evolution in the magnitude of transfers between generations and between ages is a recurrent issue in the public debate. It is even more crucial in times of economic slowdown or low growth, uncertainty about the sustainability of welfare systems, and profound demographic transformation, which tend to characterise France today. The ambition of the National Transfer Accounts (NTA) Project is to measure all public and private transfers between ages and between generations with a breakdown of these economic variables by age.

This article presents the results of the first phase of the project, consisting in calculating the age profiles for consumption and for labour income. Comparing these two profiles makes it possible, by subtraction, to obtain the ages for which the total individual consumption (private and public) is not funded by labour income and thus relies on transfers or asset-based reallocations between ages. The methodology of National Transfer Accounts (NTA), whose origins are to be found in the work of Lee (1980) and Mason (1988), is described in a reference manual published by the UN (United Nations, 2013). The principles and the results that have been brought to light so far have been the subjects of various recent publications (Lee & Mason, 2011; Lee et al., 2014; d'Albis et al., 2015; d'Albis & Moosa, 2015). This age-specific accounting offers multiple advantages.

Initially, it establishes mean values by age for economic variables, highlighting any inequalities between ages or generations. This approach also facilitates the economic assessment of the effects of demographic changes. NTA provide a new analytical framework for analysing a society on the basis of the economic relationships between generations, thereby revitalising and supplementing the conventional frameworks that are based, for example, on relationships between supply and demand on markets. As a statistical database, NTA appear useful for economists who use age-structured models such as lifecycle or overlapping-generations models. Finally, they offer the advantage of presenting a set of data that are consistent with the National Accounts and constructed similarly from one country to another[1].

In France, NTA supplement the work already done in the field of age-specific inequalities in resources. As early as the 1980s, Masson (1986) proposed measures of labour income by age for the period from 1949 to 1967, making it possible to compare not only age groups over time, but also cohorts at given ages. In 2002, a special edition of the journal *Économie et Prévision* was devoted to generational accounting (Accardo, 2002; Bonnet, 2002). More recently, Arrondel and Masson (2007) have quantified public and private transfers between two large age groups for a single year, around the pivot point of people aged 60.

The aim of generation-specific accounting was to compute, essentially for prospective analysis purposes, the balance of the State's net transfers, i.e. the difference between the benefits received and the taxes, duties, and contributions paid, over the lifecycle of each generation. This gave rise to a number of criticisms, from being based on the strong assumption that the social and tax legislation will be maintained for all current generations to the results, which are highly sensitive to the assumptions made (Bonnet, 2002). Although NTA tie in with similar literature on studying economic flows between ages and generations, the method and goal differ from the methods and goals in such literature. NTA look at all of the economic flows and they aim firstly to compare what each age (and possibly each generation whenever the NTA have been available for a sufficiently large number of years) consume and produce, before studying the way in which consumption is funded at each age when it is not funded by labour income.

The first phase of the French NTA project is dedicated exclusively to calculating the labour income and total consumption profiles. It sheds light on how the gap between consumption and labour income has been changing in France over the last three decades, from 1979 to 2011. This choice of period can be explained by the fact that, in order to construct NTA, it is necessary to have individual data relating to the consumption and to the labour income of households[2]. The data are mapped with the French System

1. *Today, about 70 national teams compile these accounts using the same methodology. See the National Transfer Accounts website for a presentation of the entire network of national teams: http://www.ntaccounts.org/web/nta/show/.*
2. *1979 and 2011 correspond to the earliest year and to the most recent year for which the French Household Expenditure Surveys (*Budget de famille*) are available. The choice was made to estimate the labour income and private consumption profiles on the basis of the same statistical survey, and thus on the basis of the same sample for any given survey year. It is quite possible that other surveys might lead to somewhat different estimated age profiles, for reasons of sampling, for example. The other surveys available in France do not include data relating both to private consumption and also to labour income. Only the* Budget de famille *Survey collects information about private consumption expenditure (conversely, other statistical sources do exist for income).*

of National Accounts data to determine, at each age, the mean levels of consumption and income for a given individual and for the population as a whole. Implementing NTA for France has produced some significant results.

Between 1979 and 2011, the level of labour income of people aged from 50 to 60 and the level of consumption of people aged 40 and over increased faster than the corresponding levels in younger age groups. Analysis of the profiles by cohort shows that the generations born up until 1940 have seen their level of consumption increase markedly compared with the generation born ten years earlier, and it also shows that the baby-boom generations have enjoyed a very significant increase in their level of labour income when compared with the generation born ten years earlier. Overall, the period of lifecycle surplus, i.e. the ages at which labour income exceeds consumption, has shortened over the period studied. It was 39 years in 1979 and only 34 years in 2011, even though the lengthening in life expectancy is mechanically increasing the funding needs during the retirement period. On an international level, comparing the French profiles with the profiles from other European countries reveals similarity in consumption, labour income, and lifecycle deficit profiles.

In the remainder of the article, we study the age profiles in 2011 for the most recent year of construction, and then the changes in consumption and labour income over time from 1979 to 2011[3]. The results also undergo comparative analyses, be it between cohorts or indeed at international level.

National Transfer Accounts

NTA quantify the acquisition and the use of economic resources at each age (Lee & Mason, 2011). They are based on a unified international methodology that consists in introducing age into National Accounts (United Nations, 2013). These accounts serve to understand the way in which economic flows move between the various age groups of a population for a country and for a given year. For any given year, determining the age profiles requires calculating the mean levels of consumption and of labour income in the population, for each age. Such profiles also specify the different sources of income (labour and capital) and the different uses of that income in terms of whether it is used for private and public consumptions or for savings.

During their lives, individuals consume at all ages. Conversely, they produce economic wealth in working adulthood only. During youth and old age, consumption therefore exceeds labour income. The difference between the total consumption and labour income age profiles corresponds to the lifecycle deficit using the NTA international methodology (United Nations, 2013). Initially, this difference or gap makes it possible to define surplus and deficit situations without this being for normative purposes. The aim is to distinguish between the periods for which labour suffices to fund consumption at a given age and the periods during which labour income is insufficient[4].

The way the lifecycle is organised results in reallocations of resources that can be voluntary or be organised by the public decision-makers. These reallocations go from the surplus period during which the gap between consumption and labour income is negative, i.e. working adulthood, to the deficit periods during which that gap is positive, i.e. during youth and old age. The different public policies clearly influence the ages at which private and public consumption is greater than or less than labour income, e.g. through education or retirement choices. Demography also plays a part in determining the lengths of these periods, through the increase in life expectancy.

The NTA are based on an accounting identity such that, at each age a, resources must be equal to the uses that are made of them (United Nations, 2013):

(1)
$$Y^L(a)+Y^K(a)+T^I(a)=C(a)+S(a)+T^O(a)$$

The sum of labour income $Y^L(a)$, capital income $Y^K(a)$ and transfer inflows $T^I(a)$ must be equal to the sum of private and public consumption $C(a)$, savings $S(a)$ and transfer outflows

3. The database used for this article and the detailed technical manual for constructing the profiles are available on the website dedicated to NTA in France: ctn.site.ined.fr.
4. From a terminology point of view, the concept of lifecycle deficit can be confusing. On the one hand it would suggest that the age groups in deficit necessarily have a negative impact. If we take the case of the young ages, for example, the deficit is due solely to the fact that children are not able to participate in the labour market. On the other hand, it explicitly refers to the lifecycle even though the deficit is instantaneous: it is computed for all ages for an observed population and for a given year (cross-cutting approach) and not for individuals that are monitored all through their lives (longitudinal approach). Despite its limitations, the choice has been made to use this concept of lifecycle deficit that has imposed itself in the international NTA network.

$T^O(a)$. This accounting identity shows the gap between consumption and labour income $C(a)-Y^L(a)$, which, at each age, corresponds to the life cycle deficit (Lee, 1994) :

(2)

$$\left(C(a)-Y^L(a)\right)=\left(Y^K(a)-S(a)\right)+\left(T^I(a)-T^O(a)\right)$$

The difference between consumption and labour income results in resource reallocations being made between the ages, in the form either of net public or private transfers $T^N(a)=T^I(a)-T^O(a)$, or of asset-based reallocations, which refer to asset income net of savings $Y^K(a)-S(a)$. For each of these components, the methodology chosen includes three stages.

- The first stage consists in calculating an age profile for a given flow and for a given year. This profile $f(a)$ is obtained from survey data.

- In a second stage, the profile undergoes smoothing of the statistical series $\tilde{f}(a)$ over the ages. Although this profile is computed at individual level, it is also possible to obtain the aggregate profile that takes into account the overall age structure of the population. With the number of people at each age a in the population being noted $N(a)$, the aggregate flow F is $F=\sum \tilde{f}(a)N(a)$.

- Finally, the last stage consists in adjustment on the basis of the National Accounts, so that the aggregate flow F coincides with the corresponding book aggregate C for the year in question. The corrective term $c=F/C$ is then calculated and applied to the individual and aggregate smoothed series. The corrected profiles are $\tilde{f}^c=\tilde{f}/c$ at individual level and $F^c=F/c$ at aggregate level.

The NTA for France have been computed by using the data from the French System of National Accounts for determining the aggregates, from data collected through surveys conducted on households, and from other sources of public statistics. The methodology and the various statistical sources uses are described in detail in the on-line supplement. In view of the availability of the various editions of the French Household Expenditure survey (*Budget de famille*), NTA have been constructed for the years 1979, 1984, 1989, 1995, 2000, 2005 and 2011. That period, which came after "*les trente glorieuses*" (France's thirty years of post-war boom), began with the second oil crisis and ended in the aftermath of the financial crisis of 2007-2010. Overall, it corresponds to a period of fairly low economic growth up until the mid-1990s, followed by even lower growth (Bergeaud et al., 2014)[5].

The lifecycle deficit in 2011 in France

Consumption higher in retirement than in working adulthood

Total consumption spending accounted for 1,425 billion euros in France in 2011. That spending breaks down as follows: 65.9% for private consumption, and 34.1% for public consumption. The spending structure differs considerably between the two types of consumption. Of the private consumption, spending on education and health accounted for very small percentages, namely 1.1% and 3.8% respectively. Of the public consumption, education accounted for 18.8%, health for 29.8%, spending related to the elderly[6] for 4.2%, housing benefits for 3.4%, and other non-assignable spending such as defence, justice, or public administration for 43.8%.

In 2011, the consumption per capita profile shows that the total private and public consumption increased strongly during the youth years[7], rising from 10,601 euros at age 0, between birth and the first birthday, to 22,810 euros at age 20 (Figure I)[8]. Then, the level of consumption remained relatively stable until the age of 50 (about 21,500 euros), whereupon the total spending increased almost linearly to the age of 66. At that age, the sum of private and public consumptions was at its maximum (27,202 euros). Beyond that age, consumption swung between 25,500 euros and 28,000 euros, without any real downward or upward trend emerging. This age profile shows two important things. Firstly, for any given year, the levels of

5. In France, GDP per capita growth was 1.8% per annum from 1979 to 1995, and 1.0% per annum from 1995 to 2011.

*6. The "old people" item includes spending that is specific for this age group, in particular personal independence allowance (*allocation personnalisée d'autonomie *or "APA") (see details in the on-line supplement).*

7. Readers are reminded that a cross-cutting approach is used in this part of the article, by describing the age profiles at a given date, namely 2011. This is not a lifecycle approach in which the individuals are monitored as they advance in age.

8. The rise in private consumption excluding healthcare and education, which accounts for nearly one half of total consumption during youth (45.6% from 0 to 9, and 46% from 10 to 19) is highly dependent on the rule used for breaking down intra-household private consumption (excluding healthcare and education). The relative weight of the children is assumed to be equal to 0.4 until the age of 4 inclusive, and then to grow proportionally to the age of 20 to reach 1, and to remain constant thereafter.

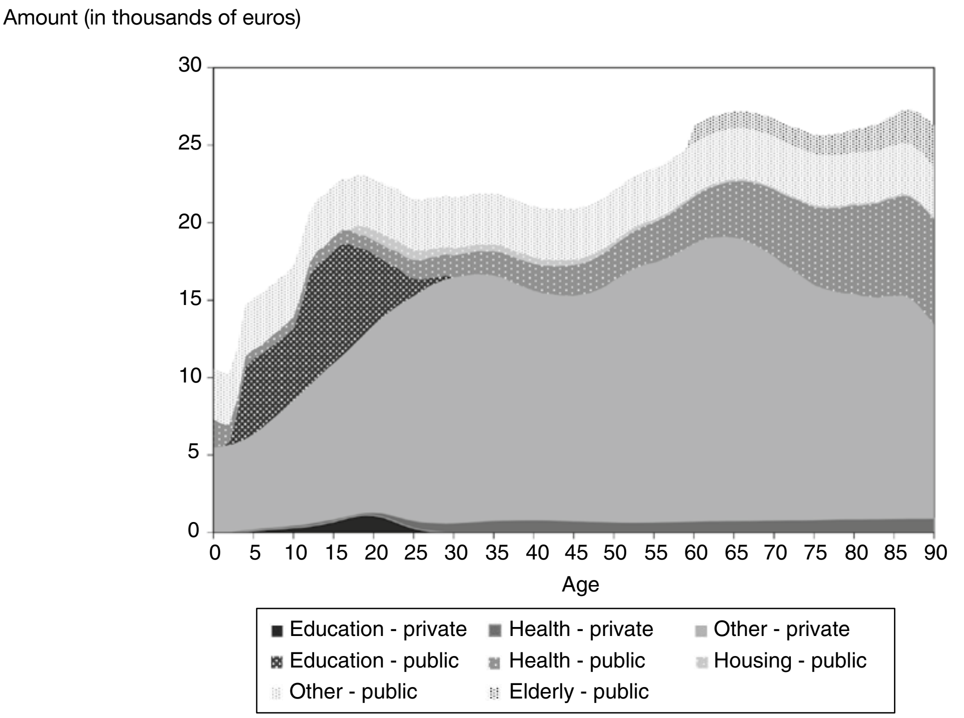
total consumption that are observed for retirees substantially exceed the levels observed for working-age adults. Secondly, mean consumption is relatively stable at high ages.

While the breakdown of private consumption by age depended mainly on spending excluding education and heath, due to the very low weights of those two items, public spending increased very strongly at the young ages through education spending and at the high ages under the influence of the "old people" spending item and of healthcare spending. Public expenditure per capita was at its maximum at the highest ages, with a mean amount of 12,837 euros at the age of 90. That sum was twice as high as the public spending in the 30-40 age bracket (5,285 euros on average). It was also higher than public consumption at the age of 15 (11,455 euros).

In 2011, the weight of public consumption in total consumption was 53% for the 0-9 age group and 50% for the 10-19 age group (Table 1). The reduction observed for the next age groups resulted from the large increase in private consumption. The contribution from public spending varied from 24.3% to 28.8% from age 30 to age 69. Looking in more detail, public spending on education represented 30.2% of total consumption for the 10-19 age group, but only 8.5% for the 20-29. The share represented by public spending on health was at its minimum for the 10-19 age group (4%). Compared with this age group, the weight of public spending on health was nearly five times larger for the 70-79 year-olds, nearly six times larger for the 80-89 age group, and indeed more than six times larger for the over 90s[9]. As a result of this growing healthcare spending at higher ages, and due to the spending related to the "old people" item, the relative significance of private spending in total consumption declined with increasing age: 74.1% for 50-59 year-olds, 63% for 70-79 year-olds, and 50,5% for people aged 90 and over.

9. By way of comparison, the weight of private spending on health was 3.2% for 70-79 year-olds, 3.3% for 80-89 year-olds, and 3.4% for people aged 90 and over.

Figure I
Consumption spending over age – per capita profiles – France 2011



Reading note: in France, mean public and private consumption was 26,197 euros at the age of 60 for the year 2011.
Coverage: Metropolitan France and French Overseas Départements*.*
Source: 2011 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 2008 permanent sample of people insured under state health insurance schemes and public statistics data, authors' calculations.*

The M-shaped age profile of private consumption observed in 2011, with a first mode at 33, a second mode at 64, and a low point at 45 between these two modes, can be observed for several countries taking part in the NTA Project (Tung, 2011). The V-shape from 33 to 64 corresponds to the ages at which the individuals have children in their households, their presence resulting in downward transfers within the households in order to fund children's consumption. The reduction in consumption after 64 can be explained by liquidity constraints, precautionary saving or motives for transmission (Deaton, 1992). An alternative explanation for this drop can be found in the fact that the consumption profile obtained in 2011 mixes generations born between the 1930s and the beginning of the baby boom. Those generations have experienced periods of war and of shortage that have marked their consumption behaviours throughout their life-cycles (Bodier, 1999).

The aggregate profile, which takes into account the population numbers, shows a sudden drop in consumption for the ages over 65[10]. This break is due to demographic changes, since the individuals aged 65 years or younger belong to the many baby-boom generations. The aggregate consumption is at its maximum for the ages ranging from 60 to 63, at about 22 billion euros per age, under the effect of two phenomena. Firstly, consumption per capita is high at those ages. Secondly, the population sizes associated with those ages that correspond to the first cohorts after the end of the Second World War, born between 1948 and 1951, are large. The aggregate level of consumption increases considerably from the age of 3 (9.6 billion euros) to the age of 19 (18.8 billion euros), due to the rise in public spending on education, and then increases at a lower rate during the working period. At higher ages, the aggregate consumption is 12.8 billion euros at 70, 11 billion at 80, and 4.2 billion at 90.

10. Figure C2-1 of the on-line supplement C2.

Table 1
Breakdown of total consumption by ten-year age group – France 2011

In %

Age Group	Private Consumption				Public Consumption					
	Education	Health	Other	Total	Education	Health	Elderly	Housing	Other	Total
0-9	0.8	0.6	45.6	47.0	22.3	6.8	0.0	0.0	23.9	53.0
10-19	3.3	0.7	46.0	50.0	30.2	4.0	0.0	0.5	15.3	50.0
20-29	2.1	1.9	64.4	68.4	8.5	5.2	0.0	2.9	14.9	31.6
30-39	0.0	3.4	72.3	75.7	0.0	7.2	0.0	2.0	15.1	24.3
40-49	0.0	3.6	69.9	73.5	0.0	9.3	0.0	1.7	15.6	26.5
50-59	0.0	2.9	71.2	74.1	0.0	11.0	0.0	0.9	14.0	25.9
60-69	0.0	2.8	67.3	70.2	0.0	13.0	4.0	0.5	12.2	29.8
70-79	0.0	3.2	59.9	63.0	0.0	19.0	4.9	0.5	12.6	37.0
80-89	0.0	3.3	53.6	57.0	0.0	23.2	7.0	0.5	12.3	43.0
90+	0.0	3.4	47.1	50.5	0.0	25.4	11.4	0.4	12.3	49.5
Overall	0.7	2.5	62.7	65.9	6.4	10.2	1.4	1.2	14.9	34.1

*Note: private consumption for education includes schooling fees and charges borne by the household (private school fees and higher education enrolment charges) and purchases of school equipment paid for by the household. Private consumption for health is what remains to be paid by the household after state health insurance cover. The other private consumption corresponds to the other items of private consumption (food and soft drinks, alcoholic drinks and tobacco, clothing and footwear, housing - including imputed rents, furniture, articles for everyday upkeep of the home, transport, communications, leisure and culture, hotels, cafés, bars and restaurants, and miscellaneous goods and services). Public consumption of education includes public spending for primary, secondary, and higher education. Public consumption of healthcare corresponds to state health insurance spending. Public spending for dependency is not included in the "health" or "healthcare" item, but rather in the "elderly" item (see on-line supplement). The "housing" item corresponds to personal housing benefit (*aide personnalisée au logement – APL*). Finally, the other public consumption spending corresponds to all of the public spending that cannot be allocated by age to individuals (defence, justice, public administration, etc.).*

Reading note: in France, public health consumption represented 11% of total consumption for the 50-59 age group for the year 2011.
Coverage: Metropolitan France and French Overseas Départements.
Source: 2011 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 2008 permanent sample of people insured under state health insurance schemes and public statistics data, authors' calculations.*

A concentration of labour income from 30 to 55

In France, the sum of labour income totalled 1,214.1 billion euros for the year 2011. These resources corresponded for the most part to employee earnings (68.4%) and, to a lesser extent, to employer social contributions (24.8%), the share accounted for by self-employment income being more limited (6.8%).

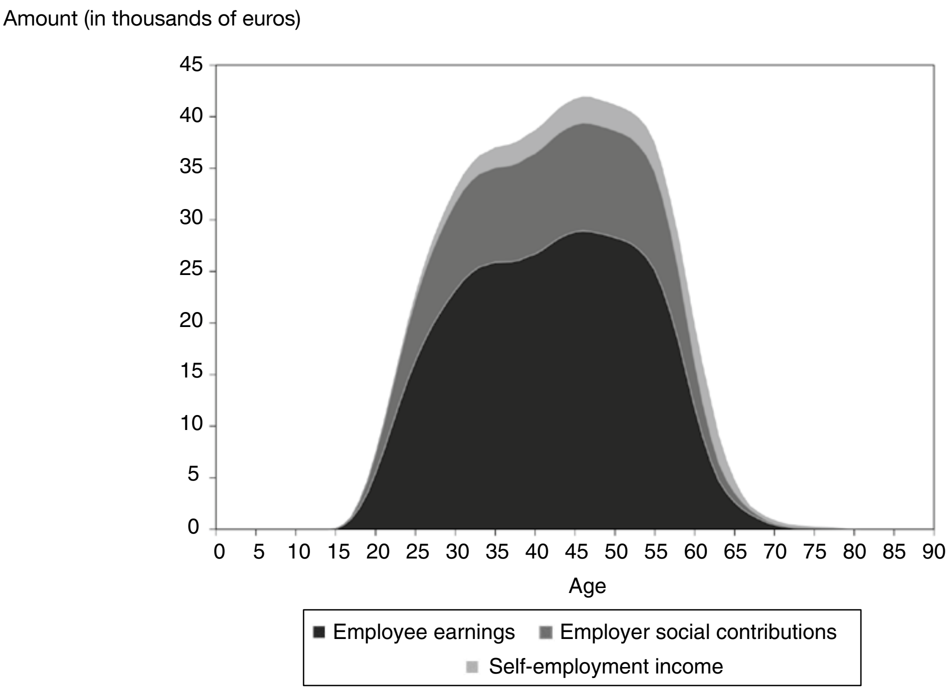
The age profile for labour income at individual level approximately forms an inverted U-shaped curve (Figure II). There are three distinct periods. Firstly, the income increases very steeply for the ages ranging from 20 to 35, by which age the mean income equals 37,023 euros. Then, the mean income continues to grow with increasing age, but at a much slower pace until the age of 45. At that age, labour income remains relatively stable for about 5 years, with a mean amount of 42,000 euros. Finally, after 54, labour income starts falling suddenly: 37,453 euros at 55, then 28,326 euros at 58, then 19,872 euros at 60, then 12,657 euros at 62, then 6,737 euros at 64, and 3,325 euros at 66.

Quite a high concentration of labour income results from this profile: the 18 highest-income years account for one half of the labour income, while the 30 highest-income years account for 80% of it. Probable explanations for this concentration of labour income lie firstly in the increased length of time spent studying, and in the difficulties encountered by young people for integrating the labour market, resulting in very low mean earnings at young ages, and secondly in the retirement age that was, on average, 59.3 years for men and 59.6 years for women in France for the year 2011 according to the OECD[11].

At aggregate level, the age profile of the labour income looks somewhat different from the individual profile[12]. The effects of seniority that result in regular increases in employees' earnings from 25 to 40 at individual level are attenuated. From 30 to 34, the share contributed to

11. The actual retirement ages calculated by the OECD correspond to weighted means taken over 5-year periods for workers aged 40 and over. For 2011, the period taken into consideration is 2006-2011.
12. Figure C2-2 of the on-line supplement C2.

Figure II
Labour income – per capita profiles – France 2011



Reading note: in France, mean labour income represented 41,948 euros at the age of 46 for the year 2011.
Coverage: Metropolitan France and French Overseas Départements.
Source: 2011 French Household Expenditure Survey (Insee, Budget de famille*) and public statistics data, authors' calculations.*

the total aggregate income by each of these ages is roughly stable, at about 28 billion euros for the year 2011. Labour income then increases steeply until the age of 40. The aggregate profile highlights the major contribution made to total income by individuals aged from 40 to 50 inclusive: the share contributed by this age group represents exactly one-third of total income. Finally, the first generations of the baby boom are now contributing very little to labour income because of them retiring.

More years of deficit than of surplus

At each age *a*, the difference between total consumption $C(a)$ and labour income $Y^L(a)$ (i.e. the share of consumption that is not funded by income from work) is equal to public transfer inflows minus public transfer outflows $TPU^I(a) - TPU^O(a)$ plus private transfer inflows minus private transfer outflows $TPR^I(a) - TPR^O(a)$ plus the private and public asset income net of private and public saving $Y^K(a) - S(a)$[13]. Public transfer inflows include public consumption and public cash transfers (retirement pensions, unemployment benefit, family allowance, etc.), while public transfer otuflows correspond to the total tax (i.e. the compulsory levies comprising employee and employer social contributions, and all taxes and duties). Private transfers include intra-household transfers (funding of consumption and transfers of imputed rents) and inter-household transfers (financial and in-kind assistance, excluding inheritances and excluding gifts or donations).

At aggregate level, the gap between consumption and labour income $D = C - Y^L$ totalled 211 billion euros in France in 2011, i.e. 10.2% of GDP. This overall deficit was funded by asset income net of saving (asset-based reallocations) $Y^K - S$ for an amount of 251.6 billion and by net public or private transfers for an amount of – 40.6 billion. The public components of the asset income and of the savings are negative, – 35.7 billion and – 76.4 billion euros respectively, which can be explained by public debt. The fact that the net public or private transfers T^N are negative corresponds to a situation in which the transfers given to the rest of the world exceed the transfers received from the rest of the world.

The per capita profile of the lifecycle deficit by age for the year 2011 follows the course of the major periods of life (Figure III). At the young ages, the maximum gap between consumption and labour income is observed at 16, and is equal to 22,344 euros. At the retirement ages, this gap remains roughly stable as from 68, at about 26,500 euros. The ages at which the gap between consumption and labour income is negative range from 25 to 58. Thus, the lengths of the periods during which consumption exceeds labour income are equal to 25 years at the young ages (from 0 to 24) and to 24 years at the retirement ages (from 59 to 82), on the basis of a life expectancy at birth of 82, as observed in 2011[14].

The length (49 years) of the cumulative period for which the difference between consumption and labour income is positive is less than the length (34 years) of the period for which the gap is negative during working adulthood (from 25 to 58). The latter period thus represents 40% of mean length of life in 2011. The largest surplus, equal to 20,952 euros, is observed at the age of 46 years. It exceeds 15,000 euros per annum over a relatively short period of 20 years, in the age range 35 to 54.

Comparing the per capita and aggregate profiles reveals gaps that can be observed above all for the high ages[15]. As the population sizes decline due to mortality, there is a mechanical decrease in the aggregate amount of the gap between consumption and labour income. For the old-age period, the maximum gap is reached at the age of 64 (for an amount of 15.7 billion euros), which corresponds to the cohort born in 1947. The annual amount of the deficit then declines slowly to the age of 80 (11 billion euros), whereupon it decreases much faster to 90 years (4.2 billion euros). Beyond that age, it is small in view of the small sizes of the very old populations in 2011. At aggregate level, the ages at which the labour income is greater than consumption remain equal to 25 and to 58.

13. More precisely, net public savings corresponds to gross savings by public administrations (or "PAs" for short) minus fixed-capital consumption by PAs. PA gross saving is composed by the difference between inflows (gross national income of the PAs, current taxes on net income and wealth of the PAs, and other current transfer inflows) and outflows (public transfers in cash and in kind, other current transfer outflows). Such public saving does not have any counterpart in the statistics that are usually presented in public finance. The public asset income (before savings are deducted, but net of fixed-capital consumption) is composed of capital income and of property income of public administrations. Such property income corresponds to income from assets owned by the public administrations. Public capital income is equal to the net operating surplus of the public administrations.

14. Life expectancy at birth was 78.4 years for men and 85 years for women in 2011 (Beaumel & Bellamy, 2013).

15. Figure C2-3 of the on-line supplement C2.

The dynamics of the lifecycle deficit

A deficit that is gradually increasing

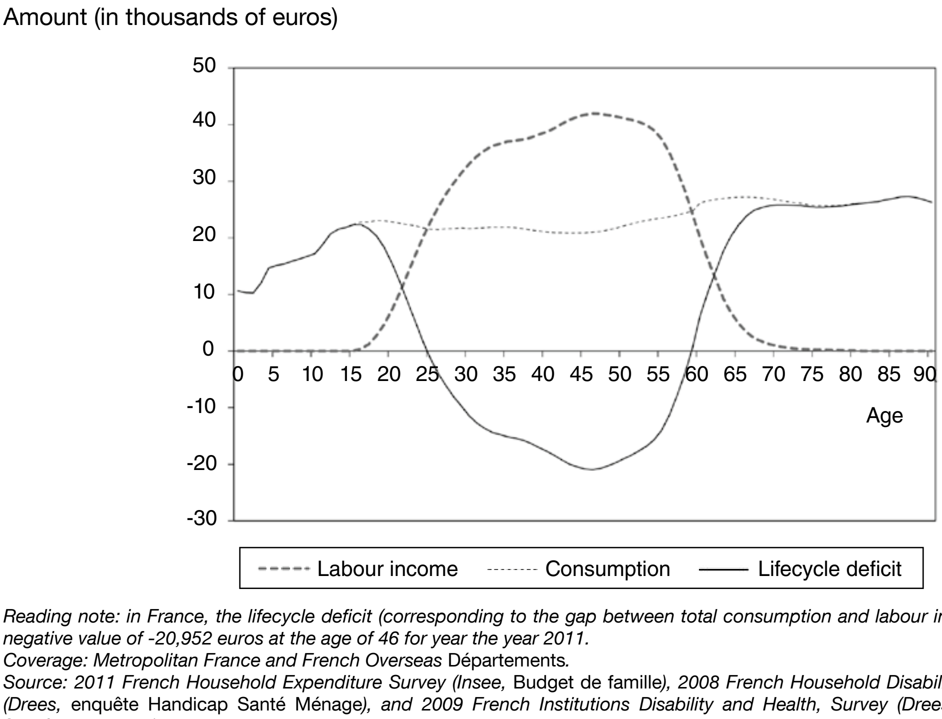
Over the last three decades, life expectancy in France has risen from 74 years in 1980 to 82 years in 2011, and the structure of the population has changed with the advancing ages of the baby-boom generations. The mean age was 40.3 in 2011, after being 36.9 in 1991. The French economy has gone through several economic crises, in particular in 1979-1981 (2nd oil crisis), in 1993 (EMS crisis), and more recently with the financial crisis that began in 2008 and then the euro zone crisis in 2010.

France has also undergone profound societal transformations. For example, the number of years of study has increased considerably because the school life expectancy between the ages of 2 and 29 rose from 16.9 years in 1985-1986 to 18.8 years in 1995-1996, before decreasing slightly until 2013-2014, when it reached 18.3 years (French Ministry of National Education, 2016). There have also been significant changes in the length of the contribution period and in the retirement age. In 1982, the pension entitlement age was lowered to 60 years with an insurance period of 37.5 years for full pension entitlement. The "Balladur" reform of 1993 then increased that insurance period to 160 quarters. The "Fillon" reform of 2003 aligned the insurance period for civil servants with the insurance period for private-sector employees, before the 2010 reform came and gradually increased the full pension entitlement age to 62. These changes alter the gap between consumption and labour income, now studied over the period going from 1979 to 2011 (Table 2)[16].

At aggregate level, there are two distinct periods. During a first stage, the total deficit grew steeply from 1979 to 1989. While labour income was 15.8 billion euros higher than total consumption in 1979, the gap between consumption and labour income then deteriorated abruptly. It became positive as of 1981, and then increased steeply to reach 100.7 billion in 1989. That amount represented 15.3% of the total consumption for that year. Labour income was then

16. The amounts are expressed in 2011 euros. So far, this dynamic aspect in the NTA Project has been addressed only in the United States (Donehower et al., 2011), in Sweden (Lindh et al., 2011) and in Taiwan (Lai & Tung, 2015).

Figure III
Life cycle deficit – per capita profiles – France 2011



Reading note: in France, the lifecycle deficit (corresponding to the gap between total consumption and labour income) represented a negative value of -20,952 euros at the age of 46 for year the year 2011.
Coverage: Metropolitan France and French Overseas Départements*.*
Source: 2011 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 2008 permanent sample of people insured under state health insurance schemes and public statistics data, authors' calculations.*

no longer sufficient to cover total consumption, which had to be funded otherwise, in particular by asset income net of savings (both public and private). We can thus observe an increase in the share of that income in funding consumption, from 0% in 1979 to 6.8% in 1984, and to 12.2% in 1989.

During a second stage, the ratio of consumption to labour income saw its growth slow down considerably, going from 1.12 in 1989 to 1.17 in 2011. Since 1989, the total gap between consumption and labour income has accounted for about 15% of the amount of private and public consumption. Over the last decade, growth in the lifecycle deficit has significantly slowed but it remains high (+ 23.1% from 2000 to 2005, and + 17.6% from 2005 to 2011).

The age profiles of lifecycle deficit per capita are characterised by a lowercase v-shape, over the whole of the studied period, from 1979 to 2011 (Figure IV). Regardless of the year considered, the difference between consumption and labour income is positive for the young ages and for old people, while the intermediate age groups who are working have more income than they consume. Comparison of

Table 2
Variation in National Transfer Account aggregates – France 1979-2011

(in real terms, 2011 constant euros)

Aggregate	1979	1984	1989	1995	2000	2005	2011
1. Lifecycle deficit							
Consumption (in billions of euros)	761.0	848.6	975.6	1056.7	1182.9	1317.8	1425.0
Private consumption (in %)	68.1	67.0	68.2	65.9	66.5	66.1	65.9
Education (in %)	0.8	0.8	1.0	0.9	0.9	0.9	1.1
Health (in %)	2.1	2.4	2.9	3.4	3.3	3.5	3.8
Other (in %)	97.1	96.8	96.2	95.7	95.8	95.5	95.1
Public consumption (in %)	31.9	33.0	31.8	34.1	33.5	33.9	34.1
Education (in %)	22.9	22.0	20.5	22.0	22.0	20.6	18.8
Health (in %)	24.2	24.1	25.7	26.1	26.4	29.2	29.8
Housing (in %)	2.1	3.1	3.5	4.1	4.0	3.5	3.4
Old people (in %)	3.8	3.8	4.0	4.0	3.7	3.9	4.2
Other (in %)	46.9	47.0	46.3	43.8	43.9	42.8	43.8
Labour income (in billions of euros)	776.8	805.7	874.8	925.8	1037.2	1138.5	1214.1
Employee earnings (in %)	63.8	63.2	63.7	66.0	67.8	68.0	68.4
Employer social contributions (in%)	22.6	23.6	24.4	24.9	24.5	24.3	24.8
Self-employment income (in %)	13.6	13.2	11.9	9.1	7.7	7.7	6.8
Ratio of consumption to labour income	0.98	1.05	1.12	1.14	1.14	1.16	1.17
Lifecycle deficit (in billions of euros)	- 15.8	42.9	100.7	130.8	145.7	179.3	211.0
Lifecycle deficit (in % of consumption)	- 6.2	8.9	15.3	15.9	14.9	15.0	14.8
Lifecycle deficit (variation in % [t -(t-n)] / t-n)	-	- 371.4	134.8	29.9	11.4	23.0	17.6
2. Funding of the lifecycle deficit							
Net public or private transfers	- 15.6	- 14.7	- 18.1	- 20.2	- 28.4	- 33.1	- 40.6
Asset income (in billions of euros)	121.3	112.1	234.2	241.2	327.5	327.2	316.4
Private assets (in %)	101.9	110.2	107.0	113.0	111.1	111.4	111.3
Public assets (in %)	- 1.9	- 10.2	- 7.0	- 13.0	- 11.1	- 11.4	- 11.3
Savings (in billions of euros)	121.5	54.5	115.4	90.2	153.4	114.7	64.7
Private savings (in %)	82.0	117.9	94.7	150.1	99.5	131.1	217.9
Pubic savings (in %)	18.0	- 17.9	5.3	- 50.1	0.5	- 31.1	- 117.9
Ratio of asset income to savings	1.00	2.06	2.03	2.67	2.14	2.85	4.89
Ratio of assets net of savings to consumption	0.0	6.8	12.2	14.3	14.7	16.1	17.7

Reading note 1: in France, the share of public consumption in total consumption rose from 31.9% in 1979 to 34.1% in 2011.
Reading note 2: in France, the lifecycle deficit in real terms (in constant euros) increased by 17.6% from 2005 to 2011.
Coverage: Metropolitan France and French Overseas Départements.
Source: data from public statistics (French System of National Accounts).

the four profiles presented (1979, 1989, 2000, 2011) clearly shows that the gap between consumption and labour income has widened increasingly over the recent period. Expressed in constant euros, i.e. in real terms, the widest gap observed for the young ages has been multiplied by about 1.6 between 1979 (14,249 euros) and 2011 (22,344 euros). For the elderly, this gap almost doubled over the same period, from 13,979 euros in 1979 to 27,571 euros in 2011. This faster growth in the lifecycle deficit for the old ages compared with the young ages can be explained by the dynamics of the increase in consumption, which is more pronounced for the 60 years old and over from 1979 to 2011.

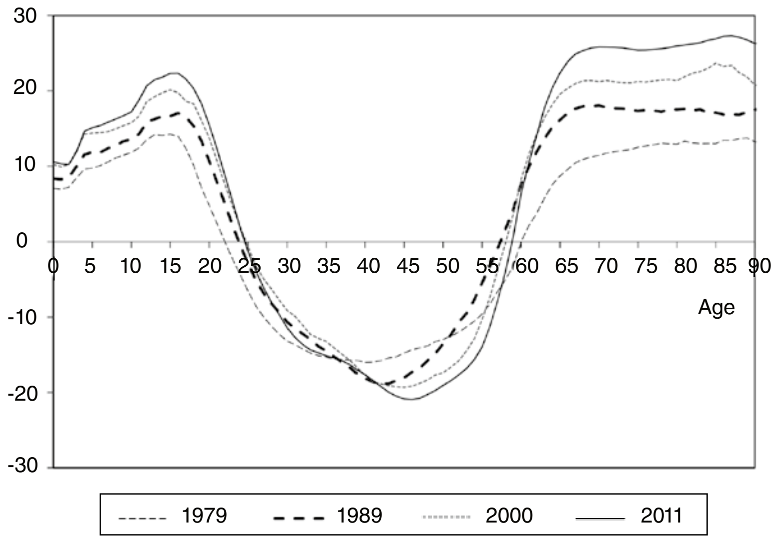
In parallel, the increase in the maximum surplus is of much smaller magnitude, going from 16,006 euros in 1979 to 20,951 euros in 2011 (i.e. a rise of 30%). The lengths of the periods for which consumption is greater or less than labour income have changed accordingly over time (Table 3). The number of years for which consumption exceeds labour income during youth increased significantly from 1979 to 1995 (going from 22 years to 26 years), and then remained stable at from 2000 to 2011 (at 25 years). The age at which consumption becomes greater than labour income again is 58 for the majority of the years considered, except for 1979, 2000, and 2011. A given individual consumed more than they produced at the age of 61 in 1979, and at the age of 59 in 2000 and in 2011.

With the continuous increase in life expectancy in France, the number of years in a deficit situation at high ages has increased considerably, going from 14 years in 1979 to 24 years in and 2011. Gradually, the number of years for which the gap between consumption and income is positive during old age is approaching the number observed during youth. Due to the concomitant lengthening of the deficit periods at young and old ages, the ratio of the ages for which consumption exceeds labour income to the ages for which labour income is higher than consumption has risen from 0.92 in 1979 to 1.44 in 2011. In 1979, 49% of the ages were characterised by a deficit, for a life expectancy equal to 74 years. That ratio then increased before becoming stabilised at about 60% from 1995 onwards.

At aggregate level, the lifecycle deficit profile continues to have a lowercase v-shape for the

Figure IV
Variation in life cycle deficit over age – per capita profiles – France 1979-2011

Amount in real terms (in thousands of 2011 constant euros)



Reading note: in France, at the age of 70, the mean lifecycle deficit grew from 11,445 euros in 1979 to 18,068 euros in 1989, to 21,221 euros in 2000, and then to 25,811 euros in 2011 (in real terms, 2011 constant euros).
Coverage: Metropolitan France and French Overseas Départements.
Source: 1979, 1989, 2000 and 2011 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 1992 and 1998 Irdes Health and Welfare Surveys (Irdes*, enquêtes Santé et Protection Sociale*), 2000, 2002, 2004, 2006, and 2008 permanent samples of people insured under state health insurance schemes, and data from public statistics, calculations by the authors.*

various years used[17]. For the year 1979, labour income exceeded consumption by 11 billion euros for the ages from 29 to 32, corresponding to the first cohorts of the baby boom, born from 1947 to 1950. Those cohorts are also those for which the difference between consumption and labour income was at its minimum in 1989 (they were then aged from 39 to 42), but they do not stand out from the other cohorts in 2000. Conversely, the gap during the old age period is at its maximum (more than 14 billion euros) for the 1947 and 1948 cohorts in 2011, when they were aged respectively 64 and 63. The increase in the mean gap, which particularly affects the high ages, has a major impact on the aggregate gap in a demographic context in which the share of the elderly population is increasing.

An improvement in the relative situation of people aged 60 and over

The composition of consumption has changed substantially over the period. The weight of private consumption has decreased in favour of public consumption, going from 68.1% in 1979 to 65.9% in 2011 (Table 2). However, this proportion has been remarkably stable since 1995, at about 66%. At a finer level, private education spending is very low whereas private health spending has been tending to rise steadily (2.1% of private consumption in 1979, 3.4% in 1995, and 3.8% in 2011). Public health spending has also increased considerably over the period, going from 24.2% of public consumption in 1979 to 29.8% of public consumption in 2011. Meanwhile, the share of public consumption devoted to education has tended to decline over the last decade (22% in 2000, 20.6% in 2005 and 18.8% in 2011).

At individual level, the age profile of total consumption is characterised by two main transformations. Firstly, the annual profiles have shifted upwards over time. The consumption levels have been systematically higher at each age since 1979 (Figure V). Secondly, the general shape of this profile has changed over the period. In 1979, consumption increased steeply from the ages of 0 to 16, and then the profile varied very little from the ages of 20 to 60. As from 2000, variations in consumption have been more marked during working-age adulthood. Since 1989, the level of consumption has been characterised by a first peak at about the age of 18. Beyond that age, a slight reduction in consumption is observed until about the age of 40, whereupon the level of consumption starts to rise again, and the magnitude of that growth has increased over the recent period. This upturn in total consumption in the second part of the working life coincides with ages when parents no longer have to provide for their children financially.

Comparison of the mean levels of consumption of the three main age groups (young adults, old people) highlights this relative improvement in the situation of the elderly. In 1979, people aged

17. Figure C2-4 of the on-line supplement C2.

Table 3
Characterisation of the gap between consumption and labour income at individual level – France 1979-2011

Consumption – labour income	1979	1984	1989	1995	2000	2005	2011
Youth – last age at which $C > Y^L$	21	22	23	25	24	24	24
Youth – number of years for which $C > Y^L$	22	23	24	26	25	25	25
Old age – first age at which $C > Y^L$	61	58	58	58	59	58	59
Old age – number of years for which $C > Y^L$	14	18	20	21	21	23	24
Total number of years for which $C > Y^L$	36	41	44	47	46	48	49
Total number of years for which $C > Y^L$	39	35	34	32	34	33	34
Ratio of years of $C > Y^L$ to years of $C < Y^L$	0.92	1.17	1.29	1.47	1.35	1.45	1.44
Ratio of years of $C > Y^L$ to life expectancy	0.49	0.55	0.57	0.60	0.58	0.60	0.60

Note: the number of years for which public and private consumption exceeds labour income during old age is given by the difference between life expectancy and the first age at which (inclusive). Life expectancy at birth was 74 years in 1979, 75 in 1984, 77 in 1989, 78 in 1995, 79 in 2000, 80 in 2005 and 82 in 2011.

Coverage: Metropolitan France and French Overseas Départements.
Source: 1979, 1989, 2000 and 2011 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 1992 and 1998 Irdes Health and Welfare Surveys (Irdes*, enquêtes Santé et Protection Sociale*), 2000, 2002, 2004, 2006, and 2008 permanent samples of people insured under state health insurance schemes, and data from public statistics, calculations by the authors.*

from 60 to 79 are characterized by a level of consumption that was greater by 1.7% on average than the consumption of the 20-59 age group. This difference has been accentuated over the period as a whole: + 7.5% in 1989, + 8.7% in 2000, and + 17% in 2011. Conversely, over the period as a whole, consumption of the 20-59 age group remained, on average, in the range 22% to 28% greater than consumption of young people aged 0 to 19. The dynamics of the consumption of 60-79 year-olds can be explained essentially by an increase in their level of private consumption relative to the younger age groups, because the relative level of public consumption between age groups remained stable from 1979 to 2011. The ratio of the private consumption of the 60-79 age group relative to the 20-59 age group went from 0.88 in 1979 to 1.11 in 2011.

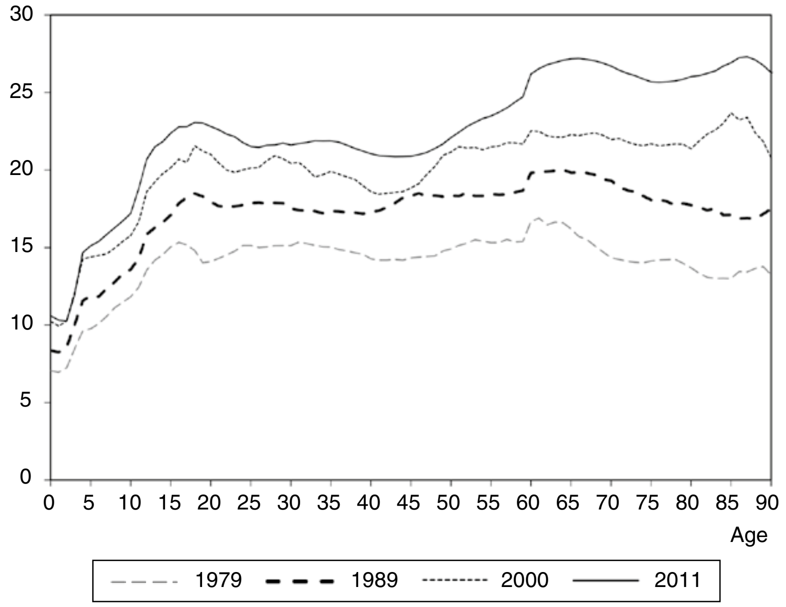
At the same time, the ratio between those two age groups for public consumption went from 1.49 to 1.46. This result might seem surprising, because public spending related to health is accounting for an increasing share of the total consumption of 60-79 year-olds (13.3% in 1979, 15.3% in 2000 and 15.4% in 2011). However, public spending on health is also occupying an increasing share for adults aged from 20 to 59 (7% in 1979, 7.1% in 2000, and 8.3% in 2011), resulting in a tendency for the relative ratio of public consumption between these two age groups to remain stable.

At aggregate level, the increase in the length of life that can be observed throughout the period is reinforcing the share contributed by for the high ages to total consumption. People aged 60 and over accounted for 18.1% of private and public consumption in 1979, 20.8% in 1989, 23% in 2000, and 27.9% in 2011. This significant increase at the very end of the period results from the fact that the cohorts born from 1946 to 1950 were at least 60 years old in 2011. As the baby-boom cohorts grow older, the mode of the aggregate profile is shifting rapidly rightwards[18]. Since the aggregate profiles are deformed due to time-related variations in the individual

18. *Figure C2-5 of the on-line supplement C2.*

Figure V
Variation in total consumption spending over age – per capita profiles – France 1979-2011

Amount in real terms (in thousands of 2011 constant euros)



Reading note: in France, mean public and private consumption at the age of 60 went from 16,680 euros in 1979 to 19,821 euros in 1989, to 22,527 euros in 2000, and to 26,197 euros in 2011 (in real terms, 2011 constant euros).
Coverage: Metropolitan France and French Overseas Départements.
Source: 1979, 1989, 2000 and 2011 French Household Expenditure Survey (Insee, Budget de famille), 2008 French Household Disability and Health Survey (Drees, enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 1992 and 1998 Irdes Health and Welfare Surveys (Irdes*, enquêtes Santé et Protection Sociale*), 2000, 2002, 2004, 2006, and 2008 permanent samples of people insured under state health insurance schemes, and data from public statistics, calculations by the authors.*

profiles and to demographic changes, the effects related to the increase in the length of life can be neutralised by thinking in terms of an unchanged population structure (Lee & Mason, 2011).

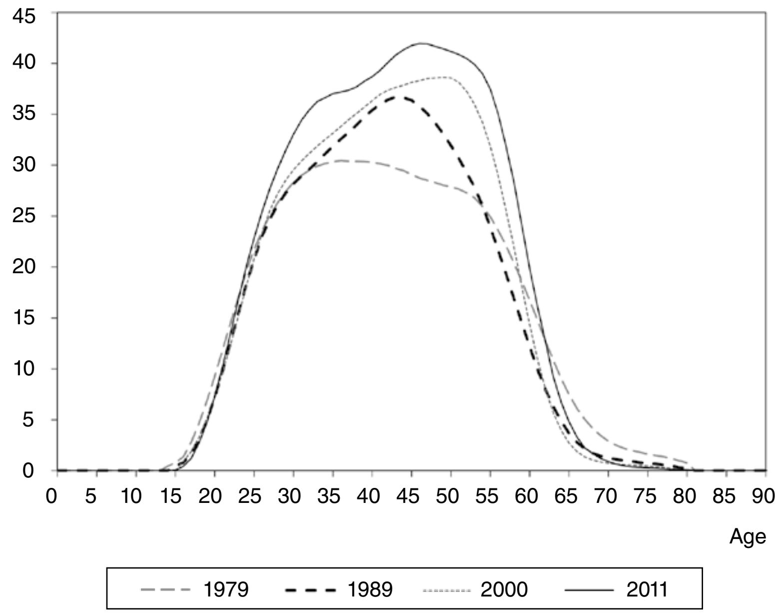
The mean age $\overline{a}_C$ at which one euro is consumed in France for the various years analysed is such that $\overline{a}_C = \sum aC(a) / \sum C(a)$ where $C(a)$ is the aggregate consumption at age a computed for the age structure of the population for that year. That age increased by 17.5% over the period as a whole (36.8 years in 1979, 38.5 years in 1989, 40.5 years in 2000, and 43.2 years in 2011). This rise appears much more moderate when the calculation of the mean age is based on the age structure of the population for the year 2011. Net of the effect of the increase in the length of life, the mean age at which one euro is consumed went from 41.6 in 1979 to 43.2 in 2011, i.e. a rise of only 3.9%. This would thus suggest that the demographic effect is the main factor in explaining the rise in the mean age at which one euro is consumed.

The main change observed for income relates to the marked reduction in the share contributed by the self-employed, in particular at the beginning of the period (Table 2). In 1979 and 1984, self-employment income accounted for more than 13% of labour income. This proportion was only 7.7% in 2000 and in 2005, and even 6.8% in 2011. Labour income was multiplied by more than 1.5 from 1979 to 2011. In addition to the higher levels of income at each age over time, at least for the ages from 25 to 55, the individual age profiles have been transformed (Figure VI). The modal age has varied significantly in 30 years by shifting rightwards[19]. In 1979, the mean labour income was higher at the age of 36. In that year, individuals aged from 30 to 39 earned, on average, 20% more than 50-59 year-olds. In 1989, the modal age had risen to 43 years, and the highest-earning age group was the 40-49 one. The modal age continued to rise in 2000 (49 years), but then fell back to 46 years during the year 2011. Overall, the profiles tended gradually to become more vertical, both at the beginning and at the end of working life.

19. The modal age corresponds to the cohort born in 1943 for the year 1979, to the cohort born in 1946 for the year 1989, to the cohort born in 1951 for the year 2000, and to the cohort born in 1965 for the year 2011.

Figure VI
Variation in labour income over age – per capita profiles – France 1979-2011

Amount in real terms (in thousands of 2011 constant euros)



Reading note: in France, mean labour income at the age of 40 went from 30,281 euros in 1979 to 38,690 euros in 2011 (in real terms, 2011 constant euros).
Coverage: Metropolitan France and French Overseas Départements.
Source: 1979, 1989, 2000 and 2011 French Household Expenditure Survey (Insee, Budget de famille*) and data from public statistics, calculations by the authors.*

At aggregate level, the rise in the modal age observed for the individual profile for income, and the ageing of baby-boom cohorts have led to an increase in the ages at which most of the labour income is received. The modal age associated with the highest aggregate income went from 31 in 1979 to 49 in 2000[20]. In 1979, the five-year age group that contributed the most to the total aggregate income corresponded to the 30-34 age group (with a proportion of 15%). Since 1996, the modal age group has been represented by the 45-49 age group (17% in 1996 and 15.7% in 2011). The mean age at which one euro is earned rose between 1979 (40) and 2011 (42.6), by 6.5%. This increase is due above all to the change in the age structure of the population over the period. For the age structure of the French population in 2011, the rise in the mean age at which one euro is earned is very small, going from 42.1 in 1979 to 42.2 in 2000.

A generational analysis

The variations observed from 1979 to 2011 would suggest that the resources have been shifting in favour of the older individuals in France. Although private and public consumption has increased at each age over time, it is the over 60s who have had the highest levels of consumption since 2000. In 1979, the cumulative amount of the deficit during the young ages was twice as high as during the high ages. This ratio then decreased considerably, going to 1.4 in 1989 and 1.1 in 2000. In 2011, the cumulative amount of deficit in old age exceeded by 7.8% the cumulative amount in youth.

These resource reallocations across the ages are, in part, attributable to changes in the age structure of the French population. At aggregate level, the magnitude of the total lifecycle deficit for any given year increases mechanically as the number of old people increases. The effect of this demographic factor is neutralised by applying the age structure of the population as observed in 2011. If the age structure of the population in 1979 had been as in 2011, then the cumulative amount of total deficit at the young ages would have been only 22.3% greater than the cumulative amount in old age. The total deficit that characterises youth would have been relatively lower than the total deficit for old people as of 1984 (– 3.7%) and would have been significantly lower in 2005 (– 9%) and in 2011 (– 7.8%).

The issue of interest is then to determine whether that resource reallocation between age groups at a given date, and taking place gradually towards the highest ages, changes the relative situations of the various generations. By definition, the lifecycle deficit approach consists in comparing, at a given date, different ages, and therefore different generations. Constructing the NTA for France over three decades makes it possible to shed new light at generational level through the formation of cohorts. According to the age profiles of lifecycle deficit for the cohorts born from 1900 to 2000 (with spacing of 10 years between each cohort), superposing the various curves by generation does indeed yield a gap that generally has the shape of a lower–case v (Figure VII). This observation is not surprising insofar as individuals do not have any labour income at the beginnings and at the ends of their lifecycles, regardless of the period in question, even if the length of those episodes varies depending on the generations.

Analysing the situations of the successive cohorts shows that the most noticeable intergenerational gaps are observed at the high ages. Comparing the cohorts born in 1900, 1910, 1920, and 1930 shows that the level of deficit has increased rapidly at the various end-of-life ages. At the age of 80, the level of deficit of a person born in 1900 was 13,850 euros (in real terms, 2011 constant euros). At the same age, this level was 29% higher for the generation born in 1910, 54.2% higher for the 1920 generation, and 90% higher for the 1930 generation. This rapid increase can also be observed at the age of 65. The level of deficit at that age for the 1940 cohort was 16% higher than for the 1930 cohort, the rise being by 34.1% between the 1930 and 1920 cohorts.

These variations are solely due to the dynamics of consumption, given that income is very low after 60 years and zero after 80 years. The rapid growth in consumption by a generation at high ages could be explained by three assumptions: a considerable reduction in the rate of saving from one generation to another, the rise of the pay-as-you-go pension system, or the increase in asset income. The first assumption can be discarded from the outset, because the generations born between the first and second World Wars are characterized by a high level of savings (Mathé et al., 2012).

The other two explanations would appear more plausible for explaining the large increase in

20. Figure C2-6 of the on-line supplement C2.

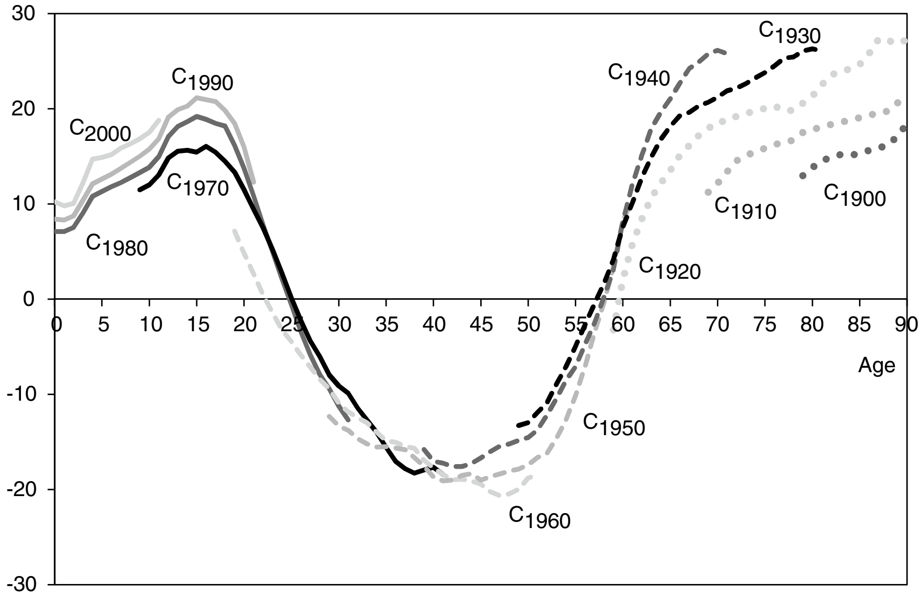
the level of consumption from one generation to another at the high ages. Firstly, the retirement pensions improved considerably from one generation to another. Their mean amount thus progressed faster than the labour income in the economy as a whole, almost continuously from the generations 1939 to 1947 (COR, 2014). Secondly, the level of wealth is increasingly concentrated at high ages (Arrondel et al., 2014), resulting in an increase in the level of asset income over the generations (Navaux, 2016). In addition, a clear break appears for the 1920 and 1930 cohorts when consideration is given to the age at which the deficit becomes strictly positive at the end of the working life. It is 60 for the 1920 cohort and 58 for the 1930 cohort, reflecting the reduction in the retirement age implemented in April 1983 and the rolling-out of the pre-retirement schemes (Burricand & Roth, 2000).

At the young ages, the amounts of the deficit at any given age have also tended to increase with the successive generations. At the age of 10, the amount of the deficit per capita that was, on average, 12,027 euros for the 1970 cohort has increased: + 15.3% for the 1980 cohort, + 31.1% for the 1990 cohort and + 45.9% for the 2000 cohort. The order of magnitude of these increases appears smaller in comparison with the earlier generations, in relation to the increasing weight of public health spending that primarily benefits the elderly, and in relation to per capita private spending that increases more rapidly at the high ages. There is also a break between the 1960 cohort and the 1970 cohort. For the former, the gap between consumption and labour income becomes negative at the age of 23, while for the 1970 cohort, it becomes negative at 26. This shift might be due to the increase in the number of years of study or to the economic context of the time that mades access to employment more or less easy.

The total consumption increases to a much larger extent with age at cohort level in comparison with the profile obtained for the various years of observation (Figure VIII-A). The cohorts born in 1940 and in 1950 have seen their mean amount of consumption multiplied by more

Figure VII
Variation in lifecycle deficit by birth cohorts – per capita profiles – France 1979-2011

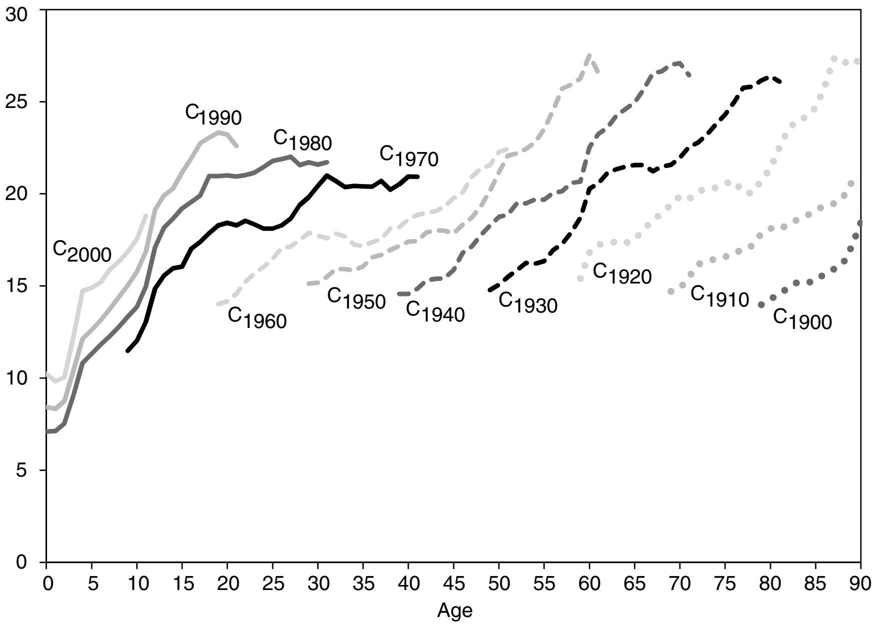
Amount in real terms (in thousands of 2011 constant euros)



Reading note: in France the mean annual lifecycle deficit at the age of 75 went from 15,866 euros for the generation born in 1910, to 20,064 euros for the generation born in 1920, and to 23,811 euros for the generation born in 1930 (in real terms, 2011 constant euros).
Coverage: Metropolitan France and French Overseas Départements.
Source: 1979, 1989, 2000 and 2011 French Household Expenditure Survey (Insee, Budget de famille), 2008 French Household Disability and Health Survey (Drees, enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 1992 and 1998 Irdes Health and Welfare Surveys (Irdes*, enquêtes Santé et Protection Sociale*), 2000, 2002, 2004, 2006, and 2008 permanent samples of people insured under state health insurance schemes, and data from public statistics, calculations by the authors.*

Figure VIII-A
Variation in consumption by birth cohorts – per capita profiles – France 1979-2011

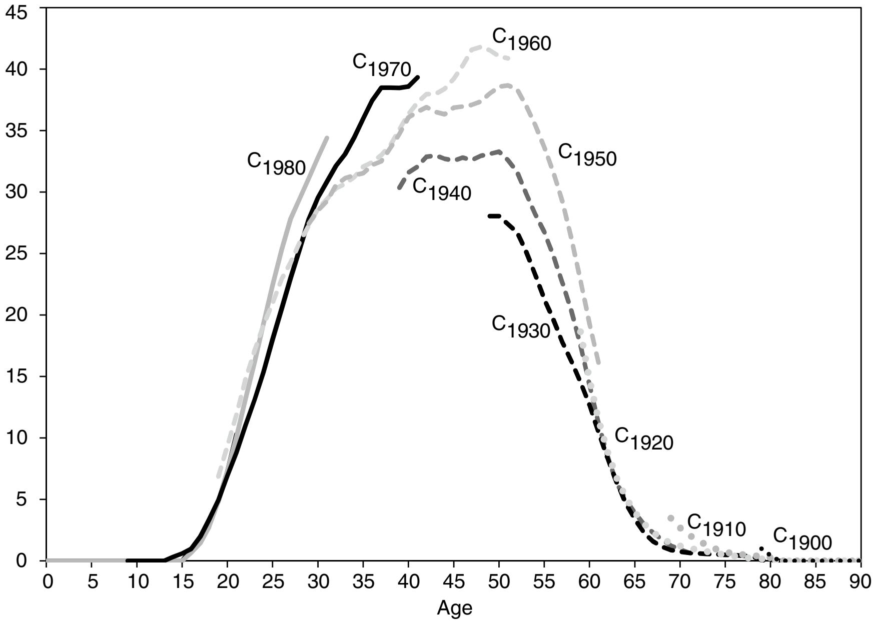
Amount in real terms (in thousands of 2011 constant euros)



Reading note: in France the mean annual public and private consumption at the age of 75 went from 16,576 euros for the generation born in 1910, to 20,550 euros for the generation born in 1920, and to 24,292 euros for the generation born in 1930 (in real terms, 2011 constant euros).
Coverage: Metropolitan France and French Overseas Départements*.*
Source: 1979, 1989, 2000 and 2011 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 1992 and 1998 Irdes Health and Welfare Surveys (Irdes*, enquêtes Santé et Protection Sociale*), 2000, 2002, 2004, 2006, and 2008 permanent samples of people insured under state health insurance schemes, and data from public statistics, calculations by the authors.*

Figure VIII-B
Variation in labour income by birth cohorts – per capita profiles – France 1979-2011

Amount in real terms (in thousands of 2011 constant euros)



Reading note: in France, mean annual labour income at the age of 75 went from 710 euros for the generation born in 1910 to 486 euros for the generation born in 1920 and to 481 euros for the generation born in 1930 (in real terms, 2011 constant euros).
Coverage: Metropolitan France and French Overseas Départements*.*
Source: 1979, 1989, 2000 and 2011 French Household Expenditure Survey (Insee, Budget de famille*) and data from public statistics, calculations by the authors.*

than 1.5 between the ages ranging from 40 to 60. The relative situation of the generations in questions has thus improved. If consideration is given to the way consumption changes from age 30 to age 40, the mean amount has increased by 14.7% for the 1950 cohort, by 5% for the 1960 cohort, and by 2.5% for the 1970 cohort. In other words, although the more recent generations are characterised by higher consumption earlier in the lifecycle, the improvement has been taking place at a decreasing pace.

The age profiles for labour income that are obtained from a representation of the cohorts are very close to those obtained from the cross-cutting analysis (Figure VIII-B). Here too, we can clearly see an improvement in the mean income at each age for the successive cohorts[21]. Although the profiles of the 1920, 1930, and 1940 cohorts tend to coincide for the ages from 60 to 65, the curve is somewhat shifted rightwards for the 1950 cohort who are going to have to work for longer. This explains the rise in the mean labour income at the ages close to retirement. The growth in labour income from one generation to another, that is very visible for the generations from 1930 to 1950 seems to have been momentarily interrupted between the 1950 and 1960 generations. Thus, up until the age of 40, the labour income for the 1950 and 1960 generations are identical in real terms (constant euros[22]), the growth resuming for the latter generation only after that age. The situation improves for the 1970 and later generations for whom the growth in labour income at each age resumes, even though that growth is slower than for the 1940 and 1950 generations[23].

Overall, these results shed light on the issue of intergenerational equality insofar as the generations preceding the baby boom appear to have benefited more in terms of consumption, and the baby-boom generations have enjoyed an increase in labour income between 50 and 60 years. However, in order to really understand how consumption and labour income vary from one generation to another, it is necessary to distinguish the effect that should be attributed to the birth cohort from the effects due to age or to the period of observation of the cohorts. D'Albis and Badji (in this issue) propose such an analysis and show that the relative situation of the cohorts born from 1901 to 1979 has improved and, in particular, that the baby-boom generation has not enjoyed a standard of living higher than the standard of living of the generations born in the 1970s.

France in a position similar to its European neighbours

The lowercase v-shaped age profile that has been highlighted in France for lifecycle deficit should be universal because survival requires consumption at each age while labour income is received only during working-age adulthood. However, this does not exclude the possibility of cross-country variations, e.g. in the number of years spent in a deficit situation at the various ages[24]. On the basis of the data available from the international NTA Project, the situation of France in 2005 is compared with the situations of the following countries: Germany (2003), Spain (2000), United States (2003), Finland (2004), Italy (2008), Japan (2004), United Kingdom (2007), and Sweden (2005). The comparison relates to the age profiles of private and public consumption and of labour income that are presented at individual level.

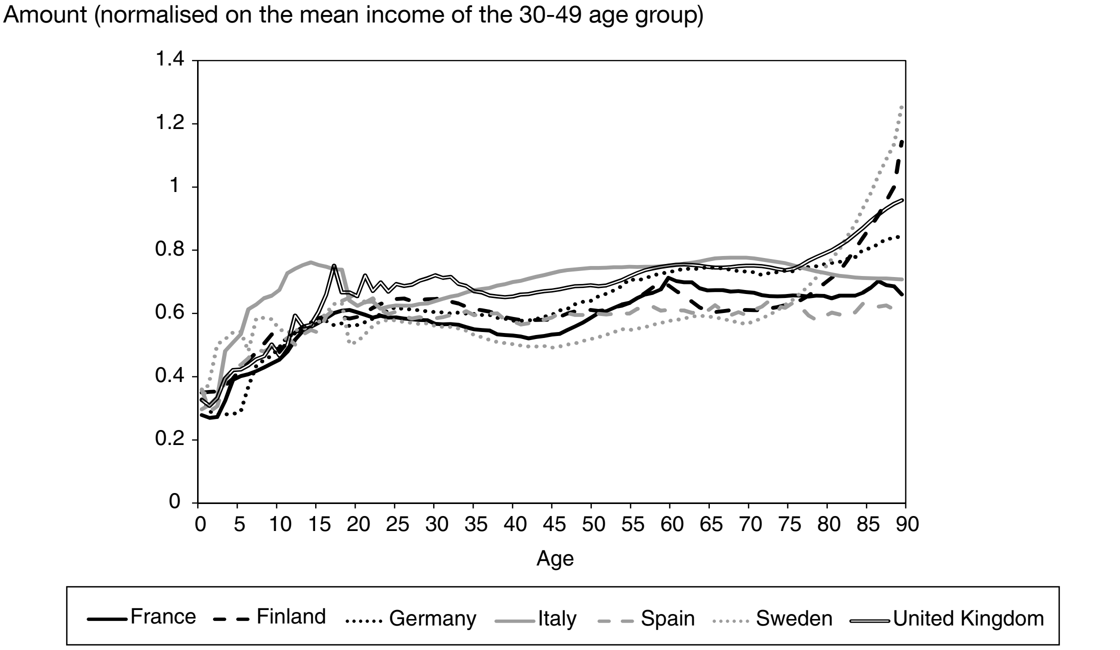
For the European countries, two distinct age profiles appear for total consumption (Figure IX-A)[25]. After a phase of quite fast growth in consumption at the young ages, and then a certain degree of stability during the working–age adult period, the countries of Northern Europe and of Southern Europe diverge at about the age of 75. For the North, the level of total consumption increases significantly at high ages, in particular in Sweden and in Finland. The most likely explanation for this is that it is due to the public spending devoted to dependency at such advanced ages (Fürnkranz-Prskawetz & Hammer, 2012). Conversely, the profile in France is similar to the profile observed in Germany, Spain, and Italy as from the age of 60. The total consumption profiles remain relatively stable during the period of old age, including after the age of 75. In this respect, the situation of France differs from the situation observed in Japan, and even more so from the situation in the United States, where it is the private health and dependency spending that can explain the very high growth

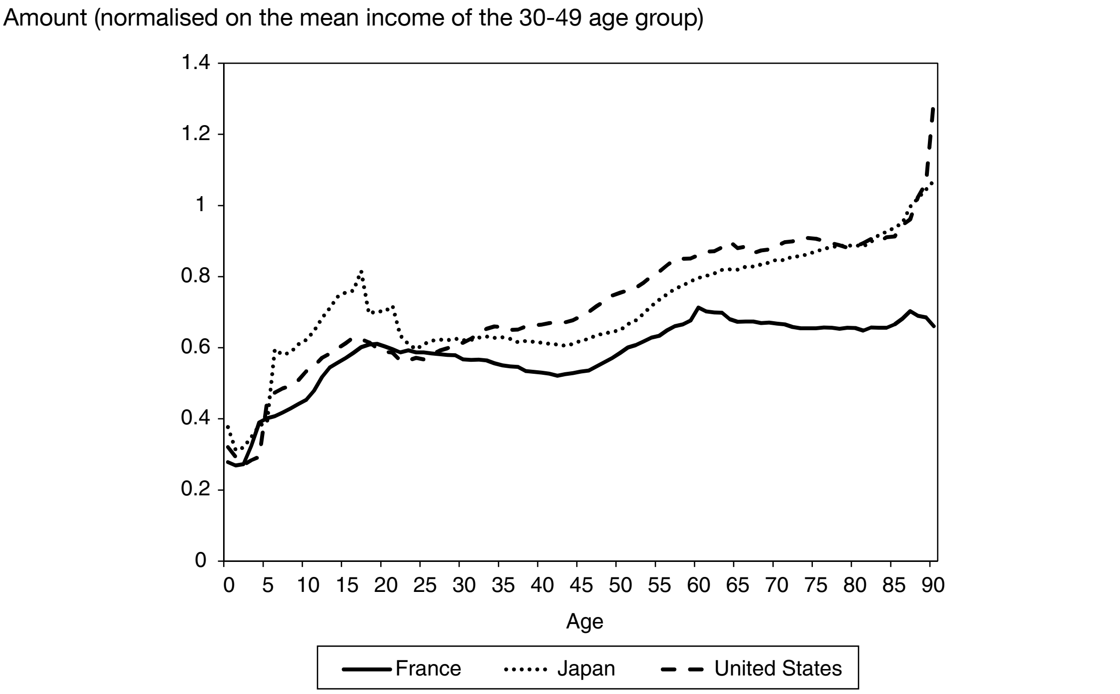
21. The particular situation of the 1910 cohort can be explained by a much later age of leaving the labour market, leading to mean labour income that is higher for the older ages.
22. Stability in "real terms" (constant euros) actually means a decline if account is taken of the general growth in income.
23. This observation echoes the one made by Clerc et al. (2011).
24. This is higher in a country in which the younger generations find it difficult to access the labour market or in which the older ones retire at an early age. A high life expectancy also increases the number of years in a situation of deficit at high ages.
25. Following the recommendations of Lee and Mason (2011), the profiles of this section have been normalised on the mean labour income for people aged from 30 to 49 in order to facilitate international comparisons. At each age, the per capita value is divided by the mean labour income of the 30-39 age group.

in consumption as from the age of 85 (Chawla et al., 2011).

For labour income, the age profiles differ depending on the age span during which the income is received and the growth in the income due to the effects of seniority. In Europe, only the United Kingdom differs from the other countries at the beginning of working life, and then the income is, on average, significantly

Figure IX-A
International comparison of total consumption by age – per capital profiles



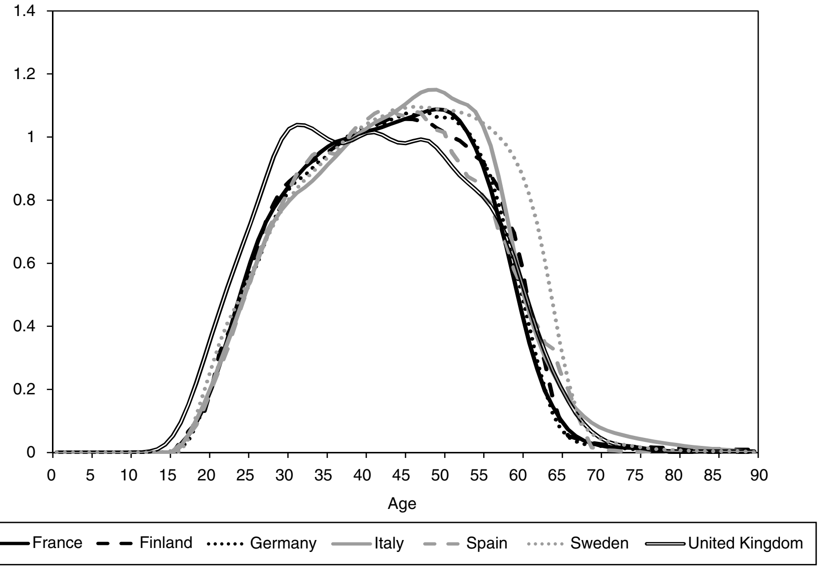


Reading note: In France, mean public and private consumption at the age of 60 accounted for 71.3% of the mean labour income received from the ages 30 to 49 for the year 2005.
Source: for France, 2005 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 2004 and 2006 permanent sample of people insured under state health insurance schemes and public statistics data, authors' calculations; for the other countries, international data from the National Transfer Accounts.*

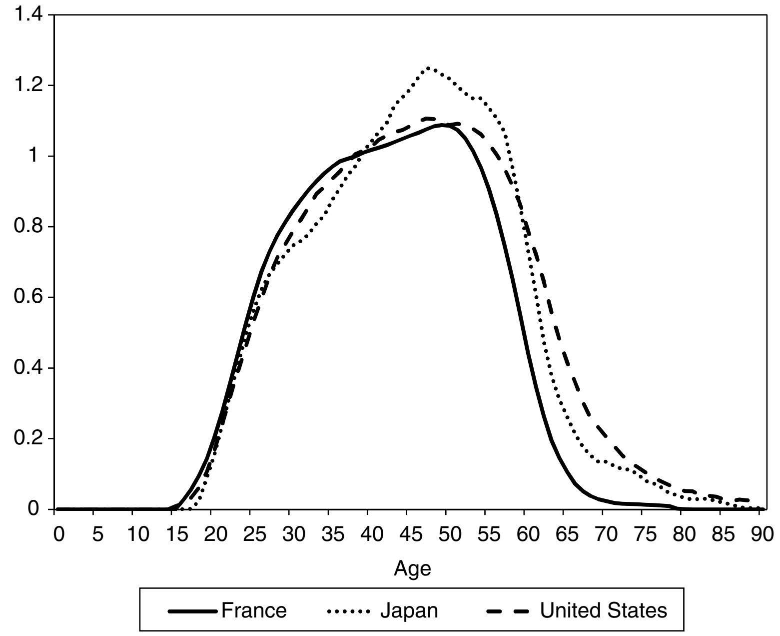
higher until the age of 30 (Figure IX-B) Beyond the age of 30, income tends to decrease slightly with age. This result contrasts with the situation observed in Italy, in France, or in Germany, where the income rises from the age of 30 to about the age of 50. The effects of seniority seem to be the largest in France and in Italy. Beyond the age of 60, France is one of the countries in which labour income is the lowest, in contrast to Sweden, where people leave the labour market later. Internationally, the levels of labour income at high ages are much higher in the United States than in France, in particular in the 60-65 age group.

Figure IX-B
International comparison of labour income by age – per capita profiles

Amount (normalised on the mean income of the 30-49 age group)



Amount (normalised on the mean income of the 30-49 age group)

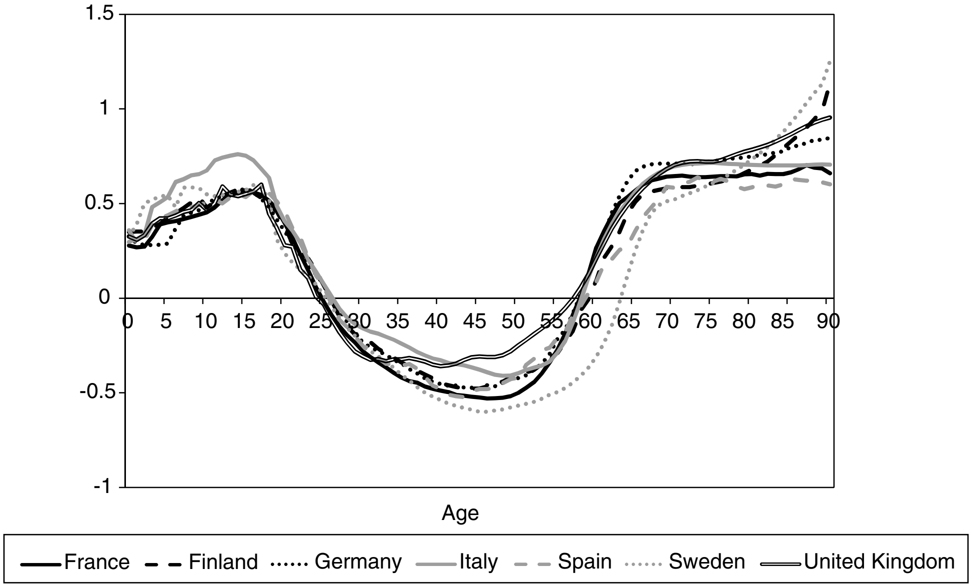


Reading note: In France, mean labour income at the age of 60 accounted for 44.4% of the mean labour income received from the ages 30 to 49 for the year 2005.
Source: for France, for France, 2005 French Household Expenditure Survey (Insee, Budget de famille*) and data from public statistics, calculations by the authors; for the other countries, international data from the National Transfer Accounts.*

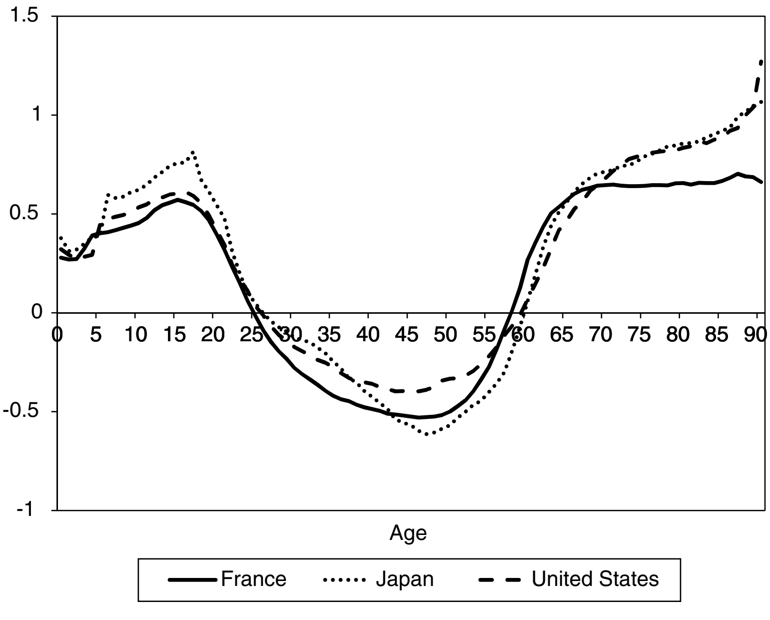
At international level, the lifecycle deficit age profiles are very similar (Figure IX-C). For all of the European countries in question, the differences are very minor until about the age of 30. The gap between consumption and labour income doubles for most countries between the ages of 0 and 18, and the first age at which that gap becomes negative varies between 25 (France and the United Kingdom) and 27 (Germany and Italy). The levels of maximum surplus are lower in Spain and in Italy. The gap becomes positive again at the age of 58 in

Figure IX-C
International comparison of the lifecycle deficit by age – per capital profiles

Amount (normalised on the mean income of the 30-49 age group)



Amount (normalised on the mean income of the 30-49 age group)



Reading note: In France, mean lifecycle deficit at the age of 60 accounted for 26.9% of the mean labour income received from the ages 30 to 49 for the year 2005.
Source: for France, 2005 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 2004 and 2006 permanent sample of people insured under state health insurance schemes and public statistics data, authors' calculations; for the other countries, international data from the National Transfer Accounts.*

France, Germany, and the United Kingdom, at the age of 59 in Italy and Spain, 60 in Finland, and 64 in Sweden (Table 4). Combined with cross-country variations in life expectancy, the total number of years in a deficit situation at high ages ranges from18 years for Sweden to 24 years for Italy. In the United States, the number of years in deficit for old age is also 18 years whereas it is 23 years in Japan due to the difference in life expectancy (82 for Japan, and 77 for the United States).

The comparison highlights similarities in the age profiles for total consumption and for labour income in the developed countries selected. During their life, each individual is in a surplus situation from the age of 24-26 to the age of 58-60 depending on the countries (except for Sweden). Although these variations in the threshold ages might appear limited, they are nonetheless quite substantial considering the average life span in those countries. In 2008, the number of years for which consumption exceeded labour income was 59% greater than the number of years spent in a surplus situation in Italy. Conversely, this difference was only 16% for Sweden in 2005 and 29% for the United States in 2003. France was in an intermediate situation, with a number of years in deficit 45% greater than the number of years in surplus.

* *
*

Implementing the first phase of the NTA for France has made it possible to show that the levels of consumption and of labour income have improved for all of the generations when they are compared at the same ages. The levels of consumption increased more rapidly from the age of 40 and above all after 60, and the rise in income was mainly enjoyed by the generations who were aged from 50 to 60 in the period from 1979 to 2011, without however calling into question the improvement in the standard of living from one cohort to another. Finally, the period for which labour income exceeds private

Table 4
Characterisation of the lifecycle deficit at individual level – international comparison

	France	Finland	Germany	Italy	Japan	Spain	Sweden	United Kingdom	United States
	2005	2004	2003	2008	2004	2000	2005	2007	2003
Life expectancy at birth	80	79	78	82	82	79	81	79	77
Youth – last age at which $C > Y^L$	24	25	26	26	25	25	25	24	25
Youth – number of years for which $C > Y^L$	25	26	27	27	26	26	26	25	26
Old age – first age at which $C > Y^L$	58	60	58	59	60	59	64	58	60
Old age – number of years for which $C > Y^L$	23	20	21	24	23	21	18	22	18
Total number of years for which $C > Y^L$	48	46	48	51	49	47	44	47	44
Total number of years for which $C < Y^L$	33	34	31	32	34	33	38	33	34
Ratio of years of $C > Y^L$ to years of $C < Y^L$	1.45	1.35	1.55	1.59	1.44	1.42	1.16	1.42	1.29
Ratio of years of $C > Y^L$ to life expectancy	0.60	0.58	0.62	0.62	0.60	0.59	0.54	0.59	0.57
Mean age at consumption of one euro	41.9	42.1	44.8	44.2	45.7	40.6	42.6	42.5	41.4
Mean age at production of one euro	41.9	43.0	42.1	43.3	45.0	40.8	44.1	40.8	43.6

Reading note: In 2005, the last age at which consumption is greater than labour income during youth is 24 in France and 25 in Sweden.

Source: for France, 2005 French Household Expenditure Survey (Insee, Budget de famille*), 2008 French Household Disability and Health Survey (Drees,* enquête Handicap Santé Ménage*), and 2009 French Institutions Disability and Health, Survey (Drees,* enquête Handicap Santé Institutions*), 2004 and 2006 permanent sample of people insured under state health insurance schemes and public statistics data, authors' calculations; for the other countries, international data from the National Transfer Accounts.*

and public consumption has tended to shrink, mainly because of the increase in the mean length of life. These findings raise questions about the way in which this lifecycle deficit is funded each year, and that will be the subject of the next phase of the NTA Project.

This is an important issue in a context when the population aged 60 and over should account for more than one-third of the French population as of 2060, according to Insee projections (Blanpain & Buisson, 2016). Understanding how the lifecycle deficit is funded requires age profiles to be calculated for asset income net of savings and for private and public transfers whose increasing weight has recently been emphasised for France (*Conseil des Prélèvements Obligatoires*, 2008; Piketty, 2011) After determining the funding of the NTA, it will be possible to compare the weight of each type of funding for non-working young people and for retirees.

These data are useful for proposing new information for diagnostics on the issue of intergenerational inequalities, the central utility of the NTA method being to incorporate all of the public and private flows between the generations. Although much research has been done in France into indicators of fairness between generations, each piece of research usually focuses on a single dimension, be it labour income (Chauvel & Schröder, 2014), public transfers or private transfers (Spilerman & Wolff, 2012; Arrondel et al., 2014). However, implementing intergenerational comparisons will still find itself constrained by the available data, which currently make it possible to reconstruct only portions of the lifecycle for each generation. ☐

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Intergenerational inequalities in standards of living in France

Hippolyte d'Albis * & Ikpidi Badji **

In this article, the effects of age (or lifecycle) and generation on the standard of living are estimated using a pseudo-panel developed from the various editions of the French Household Expenditure Survey *(Budget de famille - BdF)* between 1979 and 2011. The standard of living of households is calculated using the disposable income or the private consumption per consumption unit, including and excluding expenditure on housing and imputed rent. Using the identification strategy developed by Deaton and Paxson (1994) for Age-Period-Cohort (APC) models produces two main results. Firstly, the standard of living increases significantly with age from 25 to 64 years old. For example, consumption is 35% greater for 50-54 year olds than for 25-29 year olds. From 65 years old, changes depend on the living standard indicator considered. Furthermore, the standard of living of the baby boom generations is higher than generations born before the Second World War, but lower than or equal to the generations that follow. For example, the consumption of the cohort born in 1946 is 40% higher than the cohort born in 1926, but 20% lower than the cohort born in 1976. Considering all cohorts born between 1901 and 1979, no generation has been less fortunate than its ancestors. Discussion of these results demonstrates their robustness, particularly with regard to the results of other identification strategies, including the Age-Period-Cohort-Detrended (APCD) method which removes the linear trend from variables, and an original strategy, the Life Expectancy-Period-Cohort method (LEPC) which replaces the age variable with the life expectancy at each age. It shows the significance of economic growth in increasing the standard of living of generations and confirms that no generation has consumed less than the generations preceding it.

JEL codes: C23, D12, J14.

Keywords: income, consumption, generation, lifecycle, pseudo-panels.

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The subject of generations or age groups often features in public debate in the form of comparison. Some generations are considered fortunate and others are felt to have suffered. Essays on the topic attract a broad readership and generally insist on the special fortune of the post-war baby boom generation, sometimes even to the detriment of the following generations[1].

The aim of this article is to compare the standards of living of different age groups and generations in France using statistical data from the (*Budget de famille – BdF* hereafter). We will use two levels of comparison. The first assesses the standard of living as a function of age, in order to compare age groups with one another and assess whether "young people" are more or less fortunate than their elders. This first level of comparison primarily seeks to describe inequalities between age groups over a relatively long time period, eliminating the period effects, which could bias simple, cross-sectional analysis. However, it is difficult to draw normative conclusions. It is particularly unclear whether equal standards of living between age groups reflects household preference. Even in a context of complete markets, the lifecycle theory suggests that consumption increases with age if the return on savings is greater than the preference of an individual for the present (Yaari, 1965). Although the markets proposing life insurance in annuities are imperfect, it suggests that consumption follows an inverted U-shaped curve (Davis, 1981). The second level of comparison assesses standards of living as a function of the date of birth of individuals, by controlling the age and period effects. This compares generations and shows whether one generation has had a higher standard of living than others. However, it does not allow for analysis of the reasons behind any intergenerational inequalities and cannot be used to explain any cause-and-effect relationships between the good fortune of some and bad fortune of others. Despite this, comparison of inequalities by date of birth is better suited to normative discussion. It is obviously difficult to compare people born on different dates who have lived in very different contexts. Nevertheless, a first step can be taken by using the minimum sustainability rule which stipulates that the actions of present generations must not reduce the opportunities of future generations. A downwards trend in the standard of living of generations could therefore be considered unfair. It is difficult to take analysis further without drawing on ideological assumptions (Masson, 2009).

Comparison of standards of living between age groups and generations is complex for a number of reasons. The first concerns the choice of the variable of interest. Articles sometimes promote a specific variable such as youth unemployment or working income, which is important, but reflects just one aspect of the relative situation of the different generations (Gaini et al., 2013). In this article, we selected more general variables. We first use the total disposable income, which takes into account labour and capital income, and both public and private net transfer income. We also use a variable that describes private consumption. Using two variables is a pragmatic choice that avoids resolving the question of whether the standard of well-being is better measured using income or consumption. Using two variables also helps assess the robustness of our results. In addition, these two variables are broken down, specifying the share of expenditure on housing, particularly the imputed rent (rent that would be paid by occupant owners if they were renting their accommodation). This is used to analyse the robustness of results by removing rent from the variables studied. Finally, these variables are presented per consumption unit in the household. For the sake of simplicity, we use "standard of living" to refer to this set of variables.

The second difficulty concerns the data available. It would be ideal to have panels that follow individuals from various generations throughout their lives. In practice, we only have information on individuals who differ from one survey to another, which describes the behaviour of different generations at different moments in their lifecycle. We therefore use the seven editions of the *BdF* survey, carried out between 1979 and 2010, which we rework in order to develop a pseudo-panel to follow different cohorts throughout their lifecycle. This gives us 407 cohort observations, comprising an average of 164 individuals.

The third difficulty concerns the estimation method. Indeed, it is difficult to dissociate age effects from date of birth and period effects (assessed using the survey date). The sum of the estimated model's first two variables equals the third, making them collinear. We deal with this difficulty by setting restrictions on the period effects, which has been standard procedure since the article by Deaton and Paxson

1. The media success of these essays can be seen, for example, on the Guardian website, which presents data showing the loss of income for younger generations.

(1994). This identification strategy seems the most appropriate, but we nevertheless discuss our results using alternative strategies, including the Age-Period-Cohort-Detrended (APCD) method developed by Chauvel (2013) and an original strategy we propose, known as the Life Expectancy-Period-Cohort (LEPC) method. In this method, we estimate models that consider the “life expectancy at a given age” variable instead of the “age” variable. The clear advantage is that life expectancy is not collinear with the date of birth and date of observation. Introducing the life expectancy also takes into account the significant increase in length of human life (life expectancy for males at birth has increased by around 12% over the period studied). We therefore compare individuals of different ages but with the same life expectancy from one generation to another.

We obtained the following results with regard to changes in the standard of living as a function of age. Whatever the variable studied (income, consumption, including or excluding housing), a significant increase can be observed until the age of 60 if the effects of the date of birth and period are controlled. For example, the consumption of 50-54 year olds is 134.8% that of 25-29 year olds. The issue of the relative standard of living of the oldest individuals is more disputed in the literature. We show that there is no significant decline in the standard of living over 65 years old, except for consumption excluding expenditure on housing. Our estimations are generally consistent with previous work carried out for French households (Bossinot, 2007; Lelièvre et al., 2010), with profiles fairly similar to Belgian households (Lefèbvre, 2006) and quite different to American households, where the inverted U-shaped curve is more pronounced (Gourinchas and Parker, 2002; Fernández-Villaverde and Krueger, 2007; Aguiar and Hurst, 2013; Schulhofer-Wohl, 2015).

Our results clearly show an improved standard of living from one generation to another. Generations born later have a standard of living above or equal to that of the preceding generations and there are no “suffering” generations where one generation had a standard of living below that of its elders. The baby boomers therefore had a standard of living above that of generations born before the Second World War, but lower than or equal to generations born in the 1970s. For example, the consumption of the cohort born in 1946 is 40.6% higher than the cohort born in 1926, but 19.5% lower than the cohort born in 1976. However, the increase in the standard of living has not been continuous and a stagnation can be observed for cohorts born between the end of the Second World War and the end of the 1950s, who seem to have been more affected by the slowdown in economic growth from the 1970s.

Our results are consistent with those obtained by Lelièvre et al. (2010) based on tax revenue, and by Bernard and Berthet (2015) and Guillerm (2017) based on household wealth. However, our results differ from those of Chauvel (2013) and Chauvel and Schroeder (2014), who suggest that the baby boomers had a higher disposable income than other generations, once the trend of the variable of interest has been excluded. Although we are not convinced of the necessity of excluding the trend of the variable in order to compare generations, we wanted to reproduce the results of Chauvel and Schroeder (2014) using our data which have the advantage of consistency with the French System of National Accounts and covers a longer period. Using the same econometric specification, we do not find that baby boom cohorts were significantly more fortunate than the generations that followed. We get generally similar results with our LEPC identification strategy which substitutes life expectancy for age. This can be explained by the correlation between life expectancy and income. The results of Bernard and Berthet (2015) and Guillerm (2017) on wealth and our results on the standard of living suggest that the baby boomers were not more fortunate than the generations that followed.

The remainder of this article continues as follows. We begin by presenting our database, before detailing our identification strategy and then presenting and discussing our results.

Data and variables analysed

The *BdF* surveys

The data used are taken from the *BdF* surveys conducted in 1979, 1984, 1989, 1995, 2000, 2005 and 2010[2]. These surveys were carried

2. *Surveys are sometimes carried out over two years. In these instances, we retained just one of the two years without this choice affecting our results as we adjusted our variables in line with the French Sysem of National Accounts.*

out on over 10,000 households with the aim of reconstituting all household accounts by gathering information on their income and expenditure. It is worth noting that, in the survey, a household refers to a group of people, whether or not they are related, who ordinarily share a dwelling and have a shared budget. There may therefore be a number of "households-living unit" within the same dwelling. Information is collected over twelve months in order to eliminate the seasonal effects of some expenditure such as heating or certain food expenses.

In order compare data within a consistent time frame, it seems essential to adjust survey data in line with the French System of National Accounts (NA) aggregates. This adjustment is similar to the one carried out for the National Transfer Accounts (d'Albis et al., 2015, 2017) and aims to bring the consumption and aggregate disposable income of households into line with NA aggregates. In particular, we consider ordinary households residing in Metropolitan France. Before adjustment, we corrected differences in coverage and concept between the *BdF* survey and NA as much as possible[3].

Despite the quality of the surveys, it seems that the income and consumption from *BdF* surveys are different to the values in the National Accounts (NA). These differences can be explained first and foremost by the under-declaration or non-declaration of some consumption and income, and also by differences in coverage. The *BdF* survey only collects the income and consumption of individuals residing in France in ordinary households (i.e. excluding households residing in mobile or communal dwellings), whereas the NA considers all households. In addition, the *BdF* survey covers the consumption of French residents abroad, but does not include the consumption of foreign tourists in France, whereas the NA covers all consumption on French soil. The differences can also be explained by conceptual differences, particularly for some consumption items, which do not include the same types of expense. For example, for the housing item, the *BdF* survey only counts rent actually paid by tenants whereas the NA adds the imputed rent that homeowner households would have to pay if they were renting to their consumption.

Tables 1 and 2 show the *BdF* survey coverage rates compared to the NA for disposable income, which represents all income minus direct taxes, and consumption. Calculations take into account corrections associated with coverage and conceptual differences between the *BdF* surveys and NA data[4]. The disposable income of households was significantly underestimated in the *BdF* surveys before 1990, but coverage has improved since the 1995 survey. The trend is less clear for consumption.

The variables studied

Four variables are studied in this article.

- The first is the *disposable income of households.* The NA defines this as income after deduction of taxes and social security contributions. It therefore represents the income used by the household for consumption and savings. Income includes: *(i)* working income: salaries, self-employed income, etc.; *(ii)* income from household worth: dividends, interest, rent, etc. to which we add the imputed rents; *(iii)* social security benefits, including pensions and unemployment benefits; *(iv)* current transfers, particularly insurance indemnities minus premiums and transfers between households. We obtain the disposable income by adding all these sources of income and deducting any direct taxes paid (income tax, council tax, property tax). Note that the income declared in the *BdF* surveys is net of social security contributions (including CSG and CRDS payments).

- For the purposes of comparison, we also study the *disposable income excluding imputed rent.*

- The third variable is the private *consumption* of households. This is the sum of the 12 consumption items under the COICOP (*Classification of Individual Consumption by Purpose*). It excludes taxes, major maintenance work and loan repayments, but includes imputed rent.

- The final variable studied is *consumption excluding housing*, which represents the private consumption of households excluding expenditure on housing.

All the variables are deflated using the consumer price index.

Housing is an important aspect of the standard of living. In order to create consistent age and period comparisons, it is vital to take into account the value associated with the service provided by the housing of occupant

3. The corrections made and intermediate adjustment results are presented in the online supplement C1.
4. See the online supplement C1.

homeowners. Ignoring this variable would result in underestimating the standard of living of homeowner households. Imputed rent is the estimated rent that homeowners would have to pay if they were renting their accommodation. It can be considered both an income and additional consumption. Unfortunately, the *BdF* surveys from 1979 to 1995 do not provide figures for imputed rent. We had to estimate them using the characteristics of housing. The procedure is similar to the one used in Marquier (2003), Driant and Jacquot (2005) and d'Albis et al. (2015, 2017). Homeowners' imputed rent is calculated using the following equation:

$$loyer_i = exp(X_i' \hat{\beta} + residu_i)$$

where X_i is the vector of the variables (region, urban units, surface area, number of rooms, housing type, etc.) of the rent equation for observation i and where β is the vector of the estimated coefficients of the rent equation. In order to obtain correct rent distribution, the imputed residual must have the same distribution as the residuals taken from the rent equation. As the rent equation residuals are heteroscedastic and non-Gaussian, they cannot be expressed as a normal distribution. The appropriate residual imputation method is the Hot Deck method, which involves randomly selecting an estimated residual using the estimation from the rent equation. This residual is then imputed to housing "similar" to the one from which we selected the estimation residual and for which we have to calculate the imputed rent.

The surveys provide the level of income and consumption of households. During a lifecycle, changes to income and consumption particularly

Table 1
Comparison of the disposable income from the French Household Expenditure (*BdF*) surveys and National Accounts

	Disposable income in *BdF* (in billions of euros in nominal terms)	Disposable income in NA (in billions of euros in nominal terms)	Coverage rate (in %)
1979	168.1	250.0	67.2
1984	338.2	438.2	77.2
1989	437.0	588.6	74.2
1995	637.0	735.4	86.6
2000	784.4	867.4	90.4
2005	877.6	1045.9	83.9
2010	1104.67	1216.4	90.8

Note: data was adjusted for comparison between the BdF *surveys and NA.*

Reading note: the coverage rate is the ratio between the BdF *disposable income and the NA disposable income.*
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), French System of National Accounts, authors calculations.*

Table 2
Comparison of consumption from the French Household Expenditure (*BdF*) surveys and National Accounts

	BdF consumption (in billions of euros in nominal terms)	NA consumption (in billions of euros in nominal terms)	Coverage rate (in %)
1979	181.2	200.9	90.2
1984	352.4	369.5	95.4
1989	452.7	515.1	87.9
1995	605.0	620.0	97.6
2000	669.9	739.5	90.6
2005	785.7	894.7	87.8
2010	855.0	1024.3	83.5

Note: data was adjusted for comparison between the BdF *surveys and NA.*

Reading note: the coverage rate is the ratio between consumption in the BdF *and the NA surveys.*
Coverage: ordinary households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), French System of National Accounts (NA), authors calculations.*

reflect variations in the size of households, which changes according to the marital status and birth rate of the household. The size of the household throughout the lifecycle initially increases, reaching its maximum when the reference individual is approximately 40 years old, before decreasing. However, this trend varies from one survey to another (see the figure in Appendix 1). In order to better measure standards of living, we correct household income and consumption in line with these demographic variations, dividing the variables by the number of consumption units in the household. These consumption units give each member of the household a weighting depending on the age, in order to take into account economies of scale within households. This scale has changed over time in the *BdF* surveys. From 1979 to 1995, the Oxford scale was used (giving a weighting of 1 to the reference individual, 0.7 to individuals over 14 and 0.5 to individuals under 14), whereas from 2000 to 2010, the OECD-modified scale was used (1 for the reference individual, 0.5 for individuals over 14 and 0.3 for individuals under 14)[5]. It seemed more appropriate to use the same scale for all surveys in order to produce robust comparisons over time. We therefore weighted the variables from the surveys from 1979 to 1995 using the OECD scale. The decision to use the OECD scale is based in particular on the reasoning of Hourriez and Olier (1997) who show that the OECD scale is more appropriate than the Oxford scale in the 1990s for taking into account economies of scale[6]. However, the choice of scale is not insignificant and can influence estimations. Later in this article we test robustness by analysing the instances where the consumption unit is defined, as in the *BdF* surveys (Oxford scale from 1979 to 1995 and OECD-modified scale from 2000 to 2010), and as the square root of the number of individuals in the household. We also study instances where the variables are not weighted and where the number of consumption units is a control variable for the estimated model.

For the sake of simplicity, we refer to all four of our variables weighted by the number of consumption units using the term "standard of living", despite the fact that this terminology is usually used to refer to the disposable income per consumption unit. We are also well aware that our variables are an imperfect measure of "well-being" and that other variables such as health or environment are important. We also know that these are only mean values for each age, which do not take into account spreads that may affect the perception of the standard of living at each age.

Descriptive analysis

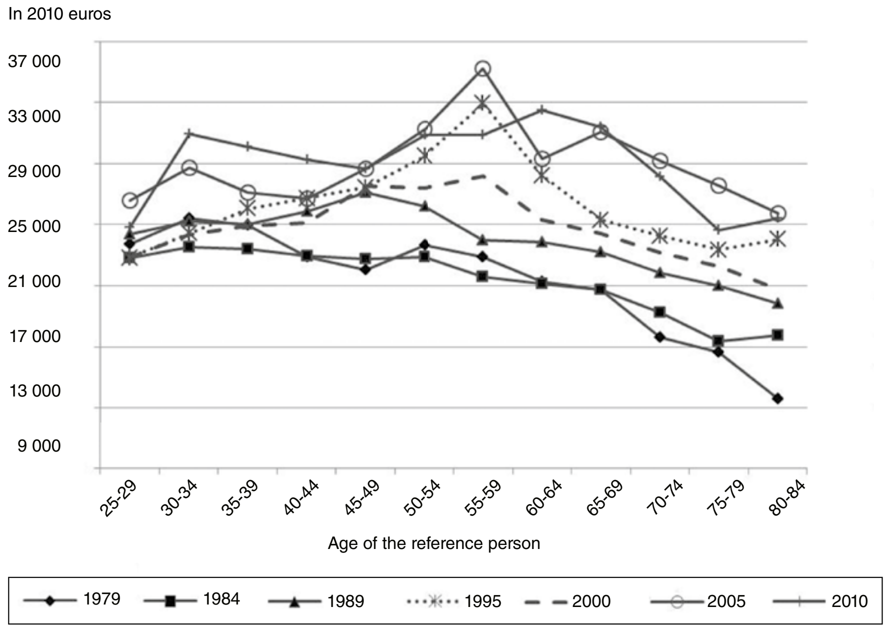
Reprocessed data may be presented synchronically or diachronically. Firstly, Figures I and III represent the standard of living (i.e. disposable income and consumption, both expressed by consumption unit) as a function of the age of the reference individual on the dates of the various surveys. This is used to compare the relative standards of living of the different age groups on a given date. Secondly, Figures II and IV represent the standard of living by age for 16 generations. These generations were constructed using seven cross-sectional databases (created from the seven *BdF* surveys). We first constructed 79 annual cohorts, defined according to the reference individual's date of birth, from the cohort born in 1901 to the cohort born in 1979. The generations were then defined using the mean of five consecutive cohorts (except for the first generation which consists of 4 cohorts).

Figures I and II regarding the disposable income per consumption unit firstly show a significant increase in the standard of living over the period considered. From one date to another, particularly between 2005 and 2010, a decrease in income can be observed for a given age, but across the entire period, the increase remains positive regardless of the age considered. However, the increase is very heterogeneous depending on the age groups. While the disposable income of 45-49 year olds increased by around 30%, it almost doubled for 70-74 year olds. The figures also seem to show relative stability in the standard of living as a function of age. Whatever date is considered, there are no major differences in income between the age groups. Between 25 and 74 years old, income is within a margin of 20% above or below the income of 45-49 year olds. For older age groups, the difference was initially greater, but has fallen throughout the period.

Analysis of consumption, with Figures III and IV confirms the analysis of income. A significant rise in consumption is observed over time, which increases as the individual grows older. In addition, the profile by age is fairly similar from one date to another and is characterised by

5. On equivalence scales, see the article by Martin in this issue.
6. A robustness test for our results regarding this choice is presented in the online supplement C3.

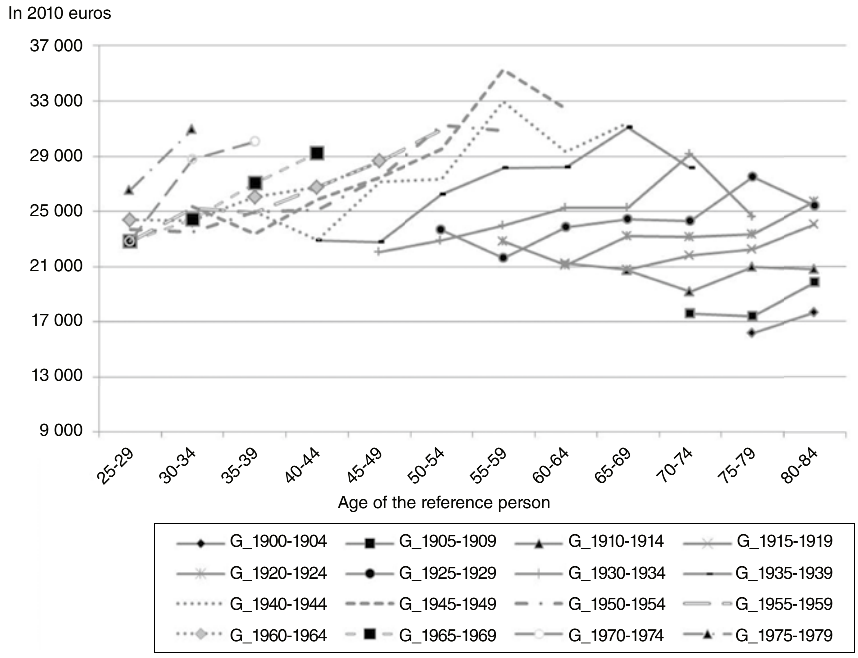
Figure I
Annual disposable income per consumption unit by age of the household reference person and the survey date



Note: disposable income is all household income (including imputed rent) after deducting taxes and social security contributions. Consumption units are calculated using the OECD-modified scale.

Reading note: in 2010, the mean disposable income per consumption unit for 25-29 year olds was €25,000.
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), authors calculations.*

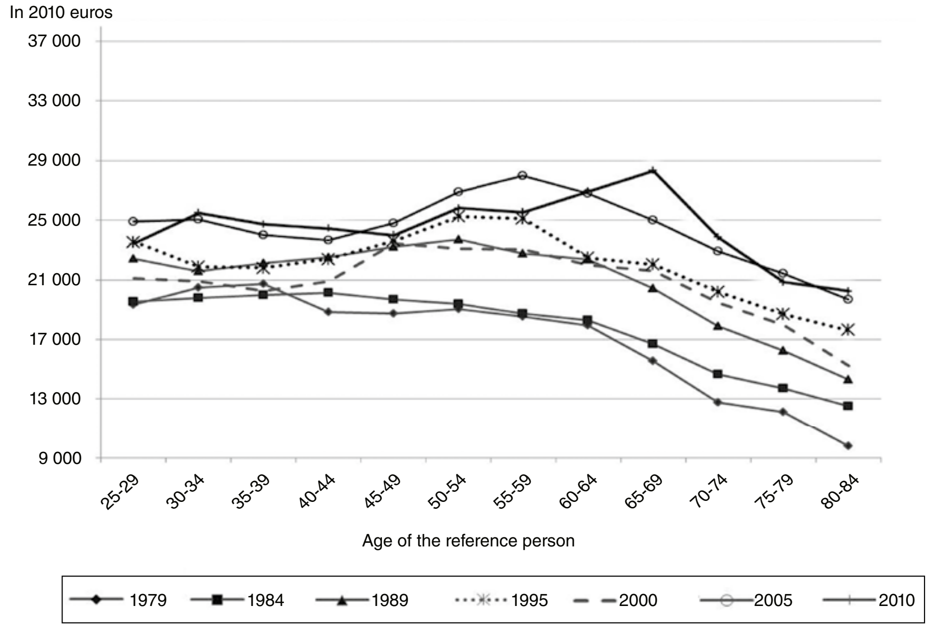
Figure II
Annual disposable income per consumption unit by the age and generation of the household reference person



Note: disposable income is all household income (including imputed rent) after deducting taxes and social security contributions. Consumption units are calculated using the OECD-modified scale.

Reading note: the mean disposable income per consumption unit for individuals born between 1975 and 1979 was €26,000 when they were 25-29 years old.
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), authors calculations.*

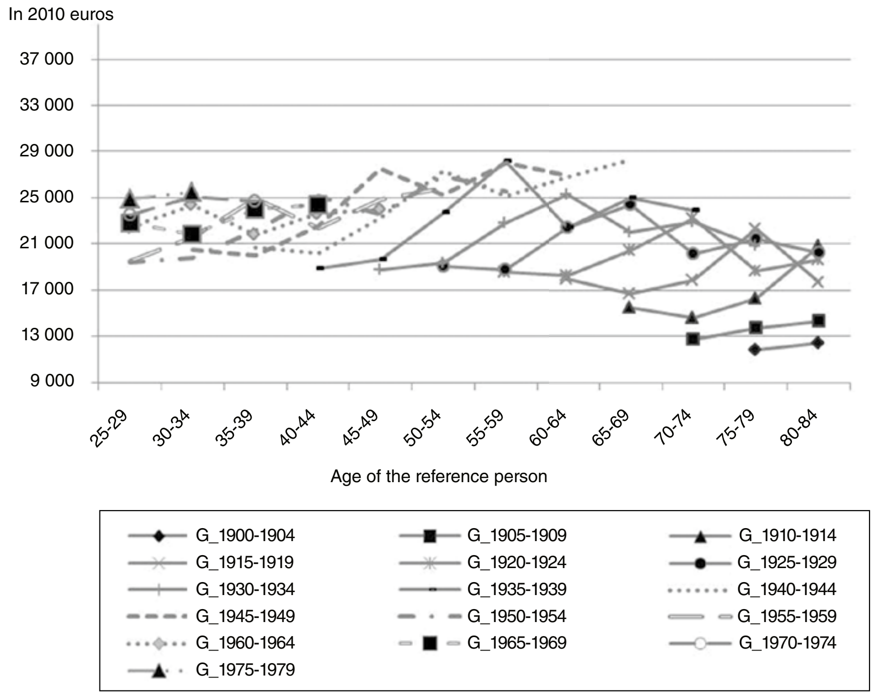
Figure III
Annual consumption per consumption unit by the age of the reference individual and the survey date



Note: Private consumption, including imputed rent. Consumption units are calculated using the OECD-modified scale.

Reading note: in 2010, the mean consumption per consumption unit for 25-29 year olds was €24,000.
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), authors calculations.*

Figure IV
Annual consumption per consumption unit by the age and generation of the reference individual



Note: private consumption, including imputed rent. Consumption units are calculated using the OECD-modified scale.

Reading note: the mean consumption per consumption unit for individuals born between 1975 and 1979 was €25,000 when they were 25-29 years old.
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), authors calculations.*

a greater drop towards the end of life than for income. The propensities to consume by age are fairly similar from one date to another, but tend to fall throughout the lifecycle.

Method and identification strategies

Estimation with pseudo-panel data

In order to dissociate the effects of age, cohort and period, it can be useful to use panel data as they follow households throughout their entire lifecycle. Our data are cross-sectional and we therefore established pseudo-panels. The idea is to identify households belonging to the same cohort and to monitor the mean behaviour of the cohorts established. As Bodier (1999) stresses, the results from pseudo-panels are not necessarily of lower quality than results obtained using panel data. The use of pseudo-panels has the advantage of avoiding selection biases associated with attrition effects (which increase with the number of periods) and biases associated with learning effects. Guillerm (2017) provides a recent and comprehensive presentation of the method.

We use the estimation technique proposed by Deaton (1985). Let us begin by stating that the estimation model used to control the individual effects that are constant over time for panel data is written as follows:

$$y_{it} = \beta_0 + \beta_1 x_{it} + \theta_i + \varepsilon_{it}$$

where y_{it} and x_{it} are explained and explanatory variables associated with individual i on date t and where θ_i is used to capture the effect of fixed individual characteristics over time. In some instances, these individual effects might correlate with the explanatory variables. It is therefore necessary to specify the type of effect (fixed or random) to include in the model. In the event of correlation between the individual effects and the explanatory variables, the fixed effects model is more appropriate. However, if the individual effects are orthogonal to the model's explanatory variables (i.e. no influence of non-observable individual characteristics on determining the level of the explanatory variables), using the random effects model is recommended. We used the Hausman test to choose between the fixed effects model and the random effects model.

Similarly, the estimation model to control individual effects for pseudo-panels is written as follows:

$$\bar{y}_{jt} = \beta_0 + \beta_1 \bar{x}_{jt} + \bar{\theta}_{jt} + \bar{\varepsilon}_{jt}$$

where $\bar{y}_{jt}$ and $\bar{x}_{jt}$ are the mean values of the explained and explanatory variables of individuals from cohort j on date t. Two types of problem tend to be generated by estimations made using pseudo-panels. The first concerns measurement errors for the different variables, which can lead to estimation biases. The model variables are not directly observed but are mean values calculated using survey data. Nevertheless, these are close to their true values when there is a large number of individuals in the cohort. Verbeek and Nijman (1993) show that measurement errors and estimation biases are negligible if the size of cohorts reaches 100. However, establishing large cohorts involves reducing the number of observations used (here the number of cohorts) across a given sample, which leads to less precise estimations. Reducing the number of cohorts can also increase the heterogeneity of individuals in a single unit and can therefore increase the variance of estimators, making them less effective. A compromise needs to be struck between sufficiently large cohorts to limit measurement errors, sufficiently homogeneous cohorts, and a sufficient number of observations to obtain adequately precise estimators.

We have seven cross-sectional databases (the 1979, 1984, 1989, 1995, 2000, 2005 and 2010 *BdF* surveys), each formed of 10,000 observations. We defined our cohorts using the "date of birth" variable, and thereby constituted 79 annual cohorts. The first cohort comprises households born in 1901 and the last cohort is formed of households born in 1979. Our pseudo-panel includes 407 observations of our cohorts, because not all cohorts are observed in each survey, and the mean size of an observed cohort is over 164 individuals (Table 3). Small numbers of observations mainly affect cohorts born up to 1917 (see detailed data in Appendix 2).

The second difficulty associated with the use of pseudo-panels concerns variation in the cohort effects which cannot be observed over time, unlike the individual effects of panel data, which, by definition, are constant. This is explained by the fact that the individuals observed from one survey to another are not the same. In order to apply the panel data estimation technique to pseudo-panels, the cohort effects must be assumed to be fixed over time. The acceptability of this assumption is based on the criteria used to define the cohorts, which must be stable over time. From this point of view, using the year of birth is optimal.

However, the simultaneous introduction of the “age”, “cohort” and “period” variables creates a collinearity problem because the survey year is equal to the sum of the “age” and “cohort” variables. Various solutions are proposed in the literature to resolve this problem. The first solution is to measure the three variables using different units by, for example, expressing the age in decades and the other two aspects in five-year periods. This is a fragile solution as it bypasses the collinearity problem without really resolving it. The results of this method have been proven unstable as they depend heavily on the units selected (Bodier, 1999). The second possibility involves replacing one of the three variables with a variable that is not collinear to the other two (Fienberg and Mason, 1985). For example, Bodier (1999) estimates consumption by replacing the date of survey variable with income, which captures economic changes over time (and is a key determiner of consumption). Nevertheless, this solution also has some limitations as income only partially reflects period effects. For the example of consumption, any changes to household consumption preferences would not be taken into account. In the discussion of our results, we propose an original identification strategy which involves replacing the age variable with a variable that measures life expectancy at each age, calculated using mortality tables of the time. This means that the three variables can be integrated simultaneously (life expectancy at each age, cohort, period) in the model without encountering collinearity problems.

The most common identification strategy involves placing restrictions on the estimated parameters. In this approach, Deaton and Paxson (1994) propose restricting period effects by assuming that the sum of the period effects is zero and that said effects are orthogonal to the long-term trend. Implicitly, the authors assume that macro-economic change can be broken down into a trend and a cycle. The cycle is fully imputed to the period effect whereas the trend is captured by the age and cohort effects. Nevertheless, their strategy has some limitations. In particular, the age and cohort effects incorporate the long-term trend due to the assumption made for the period effect. This therefore makes it difficult to isolate the age and cohort effect. Furthermore, the authors underline the fact that this procedure is risky if there are few surveys or if it is difficult to distinguish trend from transitory shocks. Despite its limitations, the Deaton and Paxson (1994) method seems the most appropriate for meeting our objectives.

Equations for the estimated models

We assume that the three effects (age, cohort and period) that we are seeking to estimate are additive. The model equation is written as follows:

$$log\overline{y}_{jt} = \mu + \sum_i \alpha_i 1_{a_{jt}} + \sum_c \beta_c 1_{j=c} + \sum_t \gamma_t 1_{t=p} + \overline{\varepsilon}_{jt}$$

where $\overline{y}_{jt}$ represents the explained variable associated with individuals from cohort j = 1901, 1902,…, 1979 on survey dates t = 1979, 1984,…, 2010 divided by the number of consumption units defined using the OECD-modified scale, $1_{a_{jt}}$ represent the indicators of the five-year age brackets from 25-29 years old to 80-84 years old[7] associated

7. We exclude people aged under 25 and over 84 as they are less representative of their generation in the BdF *survey than intermediary age categories. This is because the proportion of these people living in an institution or other household is greater and numbers in the various databases are lower.*

Table 3
Size of observed cohorts

Number of cohort observations	407
Mean size of cohorts observed	164.2
Minimum size of cohorts observed	30
Maximum size of cohorts observed	307
Proportion of cohorts observed larger than 100	85.7 %

Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille *- BdF), French System of National Accounts (NA), authors calculations.*

with cohort j on date t, $1_{j=c}$ represent the indicators of the cohorts (the fixed effects therefore correspond to the term $\sum_c \beta_c 1_{j=c}$), and $1_{t=p}$ represent the indicators associated with survey dates t.

Finally, in order to correct the heteroscedasticity potentially generated by the variation of numbers between the cohorts and, within the same cohort, from one date to another, the variables are multiplied by the square root of the size of cohorts.

In order to cancel out the collinearity relationship, we use the Deaton and Paxson (1994) method and require the sum of the period effects to be zero and orthogonal to the long-term trend. Formally, this gives:

$$\sum_t \gamma_t = 0 \text{ et } \sum_t (t \times \gamma_t) = 0$$

In concrete terms, this method involves introducing variables noted here as d^*_{ts}, rather than period indicators, into the estimated equations These variables are obtained using period indicators and the following relation:

$$d^*_{ts} = d_{ts} - \frac{ts - t1}{t2 - t1} \times d_{t2} + \frac{ts - t2}{t2 - t1} \times d_{t1} \text{ with } s \geq 3$$

and $d^*_{t1} = d^*_{t2} = 0$

where d_{ts} represent the survey years and ts represent the indicators relating to the different survey dates.

We estimated our equation for each of the four variables of interest. As shown in Table 4, in all instances, tests for fixed individual effects (cohort effects for pseudo-panels) are positive, which justifies our choice of a fixed effects model. More precisely, we estimate a Least Square Dummy Variable type fixed effects model.

Results

In the following section, we present our estimations of the effect of age on the standard of living and then our estimations of the effect of the cohort on the standard of living. Estimations of the period effect are not discussed here as they do not enter into the field of this study[8].

Comparison of standards of living between age groups

Our estimations of the standard of living as a function of the age of the reference individual are shown in Figure V[9]. The results are expressed in relation to a reference age group, 45-49 year olds.

Firstly, our estimations reveal an initial increase in the standard of living. There is significant growth in income at each age bracket until the 55-59 age bracket, and in total consumption until the 65-69 age bracket (consumption excluding housing only increases until 50-59 years old). There is a relatively large cumulative effect. For example, the consumption of 50-54 year olds is 134.8% that of 25-29 year olds. Housing slightly increases the differences between age groups. The difference in the previous example falls to 129.7% when expenditure on housing (imputed or otherwise) is removed. This increase in the standard of living does not appear in the descriptive statistics shown in Figures I and III, which, instead, suggest profile stability at the start of the lifecycle. This is an initial indication of the extent of the cohort effects that we study later in this article. After 55 years old, the standard of living does not decline, unless it is measured by

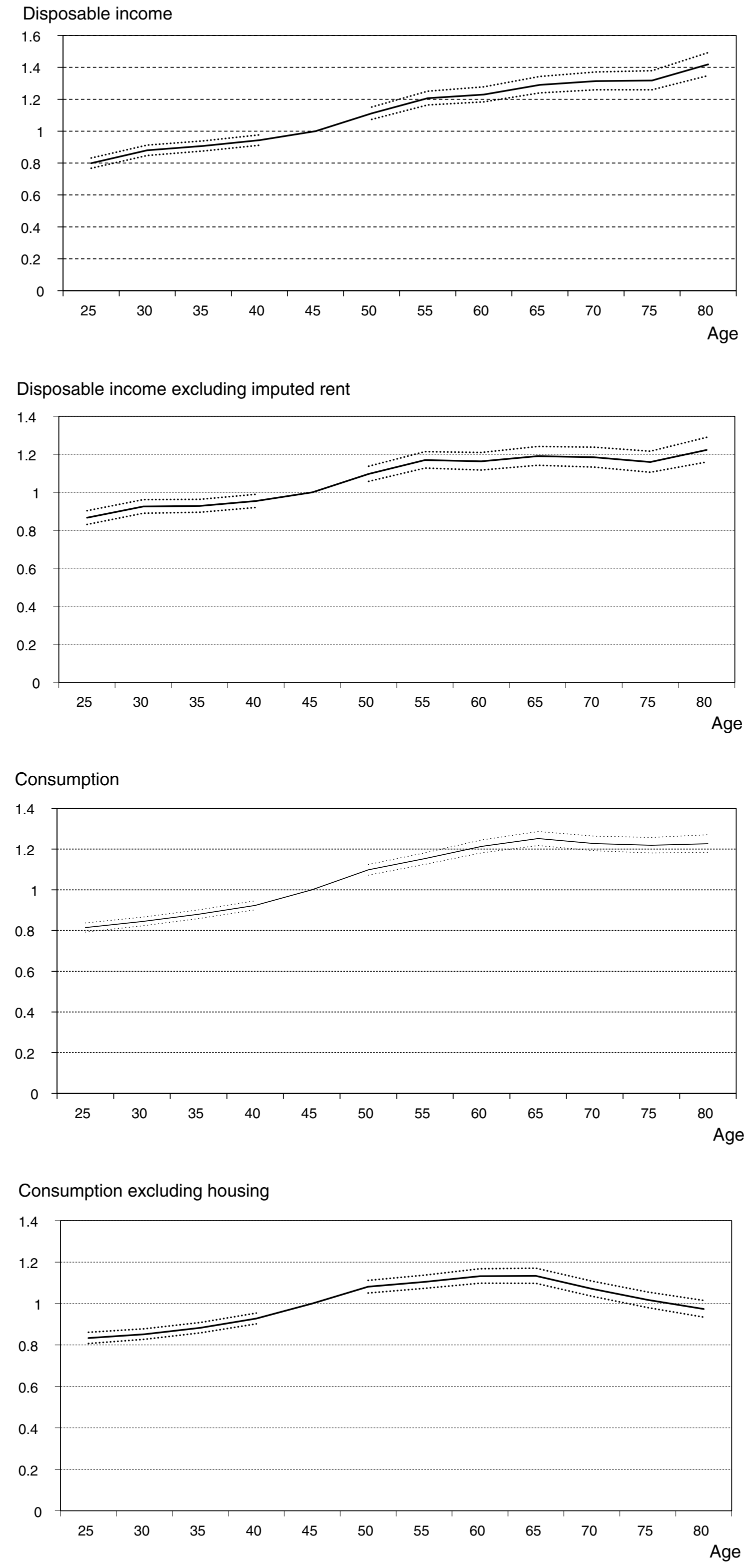
8. They are presented in the C2 online supplement (Table C2-3).
9. The coefficients are given in the C2-1 table in the C2 online supplement.

Table 4
Test for fixed individual effects and the Hausman test

	Individual effects test		Hausman test	
	F-statistic	P-value	F-statistic	P-value
Disposable income	15.21	0	297.79	0
Disposable income excluding imputed rent	8.77	0	250.4	0
Consumption	35.79	0	336.23	0
Consumption excluding housing	19.73	0	299.87	0

Reading note: the first two columns give the results of the test for individual effects. A P-value < 0.05 shows that the test for individual effects is positive at the 5% threshold. The next two columns give the results of the Hausman test. The fixed effects model is suitable for a P-value of < 0.05.

Figure V
Change to the standard of living as a function of the age group
(model controlled by the date of birth and the period)



Note: the standard of living is assessed using four variables (disposable income, disposable income excluding imputed rent, private consumption and private consumption excluding housing expenses), divided by the number of consumption units. The consumption unit is defined using the OECD-modified scale. Variables are standardised to 1 for the 45-49 age group. The dotted curves show the confidence intervals at 95%.

Reading note: the disposable income per consumption unit at 60-64 years old is 1.19 times higher than for 45-49 year olds.
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), authors calculations.*

consumption excluding housing expenditure. In this case, a significant decline is observed, which remains nonetheless moderate in size. Consumption excluding housing for 50-54 year olds is 11% greater than for 80-84 year olds.

Our estimations are similar to some results from the literature. For France, we can observe the decline in the consumption of nondurable goods at higher ages obtained by Boissinot (2007), but not the decline obtained by Lelièvre et al. (2010) for tax revenue. Our results are therefore consistent with Bodier (1999) and Herpin and Michel (2012) who demonstrated the decline in the propensity to consume after retirement. In comparison with other countries, our age profiles are fairly similar to those obtained for Belgium (Lefèbvre, 2006), but quite different to those obtained for the USA, which are characterised by a much sharper decline towards the end of the lifecycle (Gourinchas and Parker, 2002; Fernández-Villaverde and Krueger, 2007; Aguiar and Hurst, 2013; Schulhofer-Wohl, 2015).

Intergenerational comparison of standards of living

Our estimations of the standard of living as a function of the reference individual's date of birth are shown in Figure VI[10]. The results are expressed as a deviation from a reference cohort. We chose the cohort born in 1946, the date of the start of the baby boom. Although the birth rate remained high until the mid-1970s, baby boomers are generally considered to have been born between 1946 and, depending on the authors, 1955 or 1965. Furthermore, the 1946 cohort is one of the cohorts observed throughout all the surveys we have. All cohorts born between 1926 and 1954 are observed seven times (see Appendix 2). The further we move away from this group towards older or younger cohorts, the fewer observations we possess over their lifecycle. In particular, cohorts born up to 1905 and those born after 1975 are only observed twice. We will therefore naturally be more careful in interpreting the cohort effects the further we move away from the group of cohorts born between 1926 and 1954.

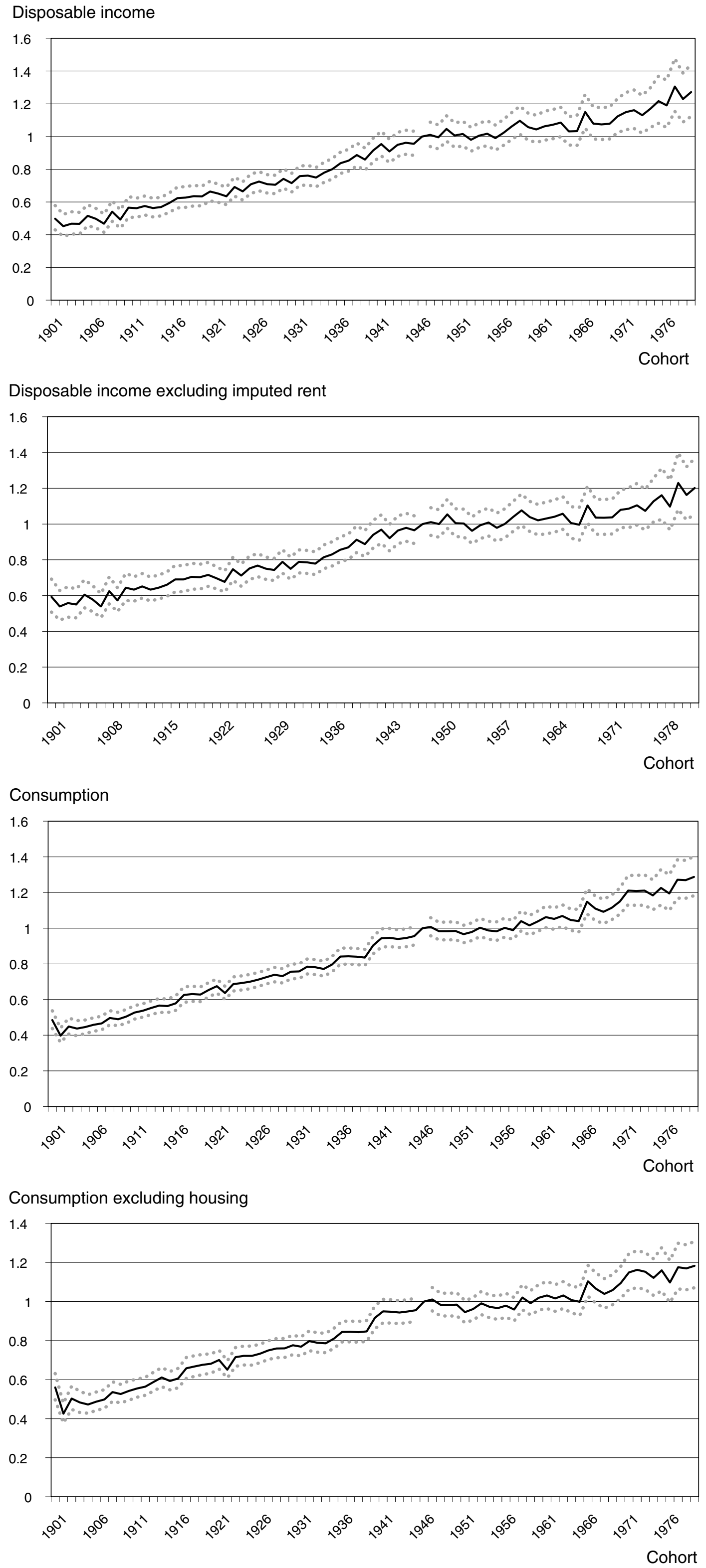
Figure VI clearly shows an improvement in the standard of living over time. Whatever variable is used, cohorts born later have a standard of living at least as high as the cohorts born before them. More detailed analysis reveals three phases in the development of the standard of living. In the first phase, the cohorts experienced a continuous increase in the standard of living. This is true of all cohorts born before the Second World War. The second phase sees a stagnation in the standard of living of cohorts, which, nevertheless, remains higher than that of the cohorts born before the war. This phase affects all cohorts born between 1945 and the end of the 1950s, if total consumption is used as the indicator, or until the end of the 1960s if income or consumption excluding housing are used as the indicator. It is therefore evident that the baby boomers and cohorts that immediately followed them had a higher standard of living than cohorts born before the war. The consumption of the cohorts born in 1926 and in 1936 at each age is estimated to represent 71.1% and 84.1% of the consumption of the cohort born in 1946, respectively. The third phase covers younger cohorts who once again saw an increase in the standard of living. The consumption of the cohorts born in 1966 and in 1976 is estimated to represent 114.9% and 119.5% of the consumption of the cohort born in 1946, respectively. Not taking into account housing only marginally modifies the differences between the generations. Furthermore, the differences are less pronounced when considering income rather than consumption. All things considered, this improvement in the standard of living is fairly consistent with the descriptive statistics presented above.

Careful interpretation of our results, only taking into account the cohorts observed seven times, concludes that the standard of living increased for all cohorts born up to the war, and then stagnated for those born later.

Our results are to be read against the backdrop of literature which had failed to reach a real consensus. The descriptive analyses of the late 1990s (Legris & Lollivier, 1996; Insee, 1998; Hourriez & Roux, 2001) found an increase in the standard of living of cohorts born before the war and observed a changing trend for those born in the 1950s. On the other hand, more recent studies (Bonnet, 2010; Clerc & Monso, 2011) show that the standard of living stopped falling for cohorts born after 1965. Furthermore, many recent articles have sought to estimate the cohort effect by differentiating it from the age and period effects. Lelièvre et al. (2010) use the French Tax Revenue Surveys (*enquêtes Revenus Fiscaux*) from 1996 to 2005. They found that cohorts born between 1942 and 1953

10. The estimated coefficients are given in Table C2-2 of the C2 online supplement.

Figure VI
Change to the standard of living as a function of the date of birth
(model controlled by the age group and the period)



Note: the standard of living is assessed using four variables (disposable income, disposable income excluding imputed rent, private consumption and private consumption excluding housing expenses), divided by the number of consumption units. The consumption unit is defined using the OECD-modified scale. The variables are standardised to 1 for the 1946 cohort. The dotted curves show the confidence intervals at 95%.

Reading note: the disposable income per consumption unit of the 1975 cohort is 1.2 times higher than the 1946 cohort.
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), authors calculations.*

were slightly more fortunate than the preceding and succeeding cohorts. However, this good fortune is reduced when transfers are taken into account. Chauvel and Schroeder (2014) use the *BdF* surveys provided by the Luxembourg Income Study (LIS) between 1985 and 2005. They state that the disposable income of baby boom cohorts is higher than for pre-war cohorts and cohorts born around 1970. We compare our results in greater detail with those of Chauvel (2013) and Chauvel & Schroeder (2014) later on. Our results are, however, consistent with those obtained by Bernard & Berthet (2015) and Guillerm (2017) for household wealth. Using the Deaton and Paxson (1994) method, they show that gross wealth increased for all cohorts born before the baby boom, before stagnating. In particular, they did not find that baby boomers were more fortunate than the generations that followed.

Robustness analysis

We assess the robustness of our results in two stages. Firstly, we check whether they are sensitive to our assumptions concerning the age group categories and the definitions of consumption units, while retaining the Deaton and Paxson (1994) method. We then discuss the implications of other identification strategies.

We checked if our results changed when individuals were not categorised by age group and if we used the age squared as a control, as per Guillerm (2017). We also checked their sensitivity if results were sensitive to the different ways of taking into account household size. Indeed, the literature is very disparate on the topic. Some authors use variables divided by consumption units, which can be defined in various ways (Clerc et al., 2010, use the *BdF* survey scales, whereas Chauvel, 2013, uses the square root of the number of individuals in the household). We also studied the case where variables are not weighted and the number of consumption units is a control variable of the estimated model, like Bodier (1999), Boissinot (2007) or Aguiar and Hurst (2013). Qualitatively, our results remain unchanged[11]. Improvement of the standard of living of generations appears to be very robust. In some instances, improvement of the relative situation of recent generations seems even more clear.

We then checked whether our results were dependant on our identification strategy. In particular, Chauvel and Schroeder (2014), who demonstrate that the baby boom generations had more disposable income than other generations and whose results differ from our own, use a different strategy based on Chauvel (2013). This is called the *Age-Period-Cohort-Detrended* (APCD) method and focuses on the fluctuations in the age, cohort and period effects around their respective linear trend. It cannot be used to compare cohorts with one another, only in relation to an unknown coefficient. We present this method in the C4 online supplement and we used the APCD module (available on Stata) with our data in an attempt to reproduce their results. Our disposable income excluding imputed rent variable is the closest to the variables they use. We find[12] that there are generally no significant differences between the cohorts born between 1920 and 1977. Only cohorts born between 1957 and 1960 have a disposable income that is (ever so slightly) significantly higher than the trend. Although the coefficient assigned to the baby boom cohorts is not significant, their income level is actually below the trend. One of the main reasons explaining the differences between the results of Chauvel and Schroeder (2014) and the results we reproduce in the C4 online supplement is the fact that the LIS *BdF* surveys do not seem to have been adjusted and that the 2010 *BdF* survey was not taken into account. When we apply the APCD method to our other variables (disposable income with imputed rent, private consumption and private consumption excluding housing expenses), we find that the only cohorts (slightly) more fortunate are those born in the late 1950s. We also find that the pre-war generations were less fortunate in terms of consumption.

One plausible explanation of the difference between the results we obtain using the Deaton and Paxson (1994) method and those we obtain with the APCD method is as follows. The first method allocates the cycle to the period effects and spreads the trend between the age and generation effects. On the other hand, the second method seeks to eliminate the trend to focus on non-linearities. The different estimations generated by implementing the Deaton and Paxson strategy therefore show that economic growth has benefited recent generations who have seen a rise in their standard of living. However, if the trend is removed, far fewer differences in the standard of living are detected between the

11. The figures concerning comparisons between cohorts for the different specifications are given in the C3 online supplement.
12. Our results are presented in Table C4-1 of this C4 online supplement.

generations, but no decline in the standard of living is observed. We explored this argument by proposing an original identification strategy.

Our idea is to replace the age variable by the life expectancy at a given age. This is a relatively simple way of eliminating the traditional problem of collinearity. We estimate the following LEPC model:

$$log\bar{y}_{jt} = \mu + \sum_i \alpha_i 1_{ev_{jt}} + \sum_c \beta_c 1_{j=c} + \sum_t \gamma_t 1_{t=p} + \bar{\varepsilon}_{jt}$$

where $1_{ev_{jt}}$ represent the indicators of life expectancy at each age associated with cohorts j and dates t. As previously, the individuals are broken down into age groups, but these are no longer defined by calendar age, but by life expectancy. Due to the increase in life expectancy, we place individuals of different (calendar) ages into the same age group when they belong to different cohorts. Individuals from a given cohort will therefore be older than individuals from cohorts born before them and younger than the cohorts born afterwards. This is not incongruous as an individual aged 70 is currently much "younger" than an individual of the same age thirty years ago (d'Albis & Collard, 2013) and life expectancy influences economic decisions throughout the lifecycle (Sánchez-Romero et al., 2016). Our estimations of the standard of living as a function of the reference individual's date of birth are given in Figure VII.

In terms of consumption, we find the same strong growth that characterises the pre-war cohorts, before a long stagnation. For income, the profile is quite different than the profile obtained using the Deaton and Paxson (1994) method as there are practically no longer any significant differences from one cohort to another. These results are relatively close to those obtained using the APCD method. This is due to the fact that life expectancy is strongly correlated with mean income. By controlling the life expectancy, the model allocates economic growth to the period effects. The benefits of economic growth for the generations are no longer taken into account. Removing growth clearly has a differing effect on consumption and income, which suggests a change to the propensity to consume over the generations.

The APCD and LEPC methods are ways of dealing with the collinearity problem without restricting the estimated parameters. However, they partially eliminate the effect of economic growth on the relative standard of living of cohorts. We therefore prefer the Deaton and Paxson (1994) approach, which appears the most relevant. Nevertheless, with these three identification strategies, we obtain the common result that the baby boom cohorts were not significantly more fortunate than the cohorts that followed.

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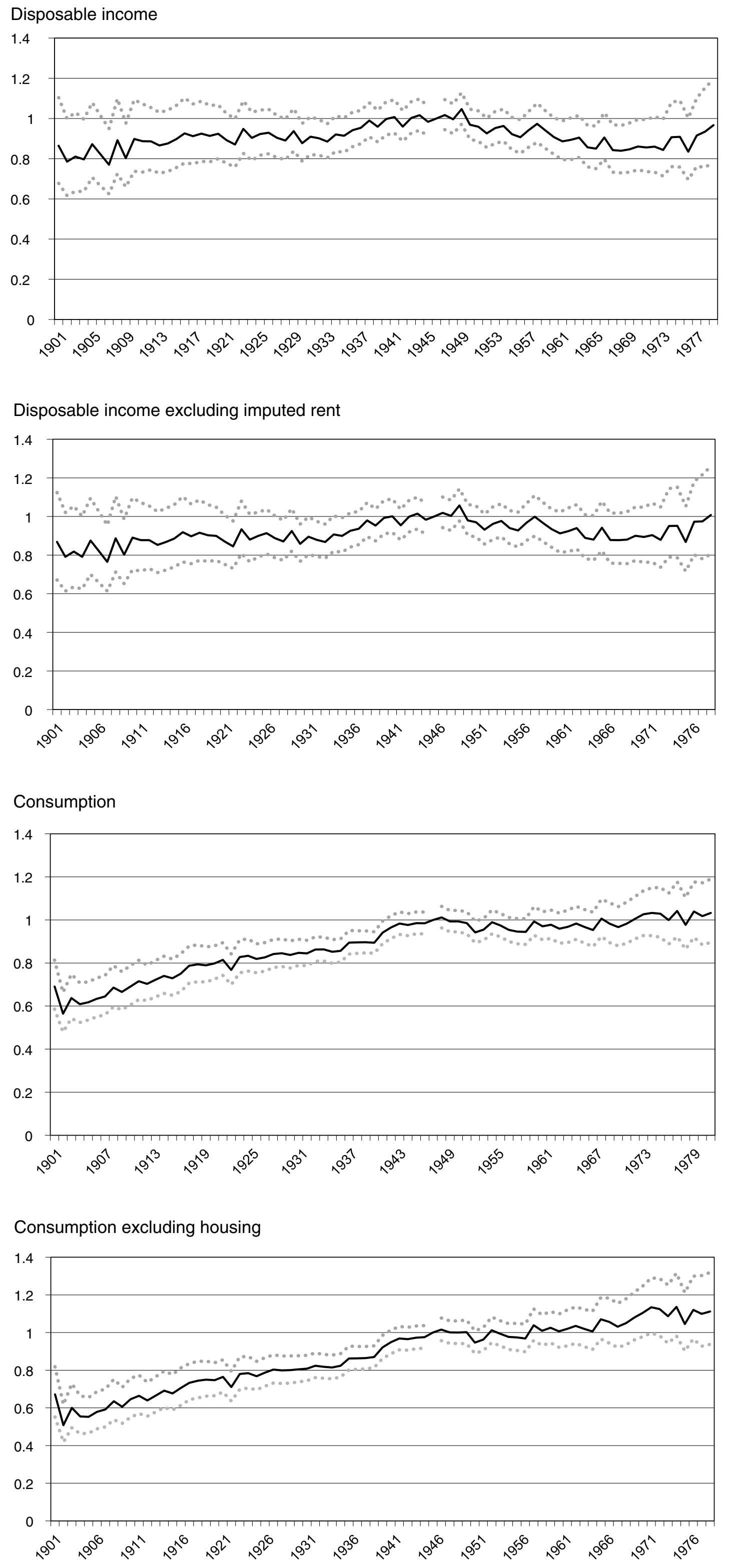
Using the *BdF* surveys conducted between 1979 and 2010, we estimated different models describing changes to the standard of living as a function of the age and date of birth of the reference individual. The aim was to measure inequalities between the age groups and generations in order to inform debate surrounding generational policies.

By analysing changes to the standard of living as a function of age, we move away from considerations concerning generations or the observation period. In line with previous studies on the topic, we demonstrated that the standard of living increases with age until around 60 years old. The dynamic then depends on the variable considered, as disposable income continues to rise whereas consumption stagnates. This change is strongly influenced by whether or not housing is included in the analysis. Imputed rent generates an increase in disposable income after retirement, which, otherwise, stagnates. Similarly, private consumption excluding housing expenditure (and imputed rent) falls after the age of 65.

By analysing changes to the standard of living as a function of the birth cohort, we move away from considerations of age or period. We showed that no generation had a level of consumption lower than the preceding generation. Regardless of the econometric specification selected, we found that no generation has "suffered" for the sake of its ancestors. In particular, we have not found that the baby boom generation had a higher level of consumption than the generations that followed. The result seems quite natural. Between 1979 and 2010, real consumption per head increased in France by over 85%. Individuals born later therefore live in an economy with higher average consumption. There would have needed to be considerable redistribution in favour of the baby

Figure VII
Change to the standard of living as a function of the date of birth
(model controlled by the age group defined using the life expectancy and the period)



Note: the standard of living is assessed using four variables (disposable income, disposable income excluding imputed rent, private consumption and private consumption excluding housing expenses), divided by the number of consumption units. The consumption unit is defined using the OECD-modified scale. The variables are standardised to 1 for the 1946 cohort. The dotted curves show the confidence intervals at 95%.

Reading note: the disposable income per consumption unit of the 1975 cohort is not significantly different to that of the 1940 cohort.
Coverage: private households living in Metropolitan France.
Source: Insee, 1979, 1984, 1989, 1995, 2000, 2005 and 2010 French Household Expenditure survey (enquêtes Budget de famille - BdF*), authors calculations.*

boomers to counterbalance this effect caused by economic growth.

Our findings could be explored further by work on two areas, the first of which is prospective. Debates around generational issues often feature the argument that the social welfare system is unsustainable, particularly its old-age and health insurance components primarily aimed at older people. It is clear that a decline in this transfer income could, in the future, call into question the estimated standard of living of generations born since the 1970s. Similarly, the increase in public debt or all the factors that have led to sustained slow growth may also compromise their standard of living. A second area for research would focus on inequalities within generations. It is possible that changes to intergenerational inequalities have been heterogeneous. Proof of an increase in inequalities among young people today could be one means of explaining the discontent often expressed by young people. ☐

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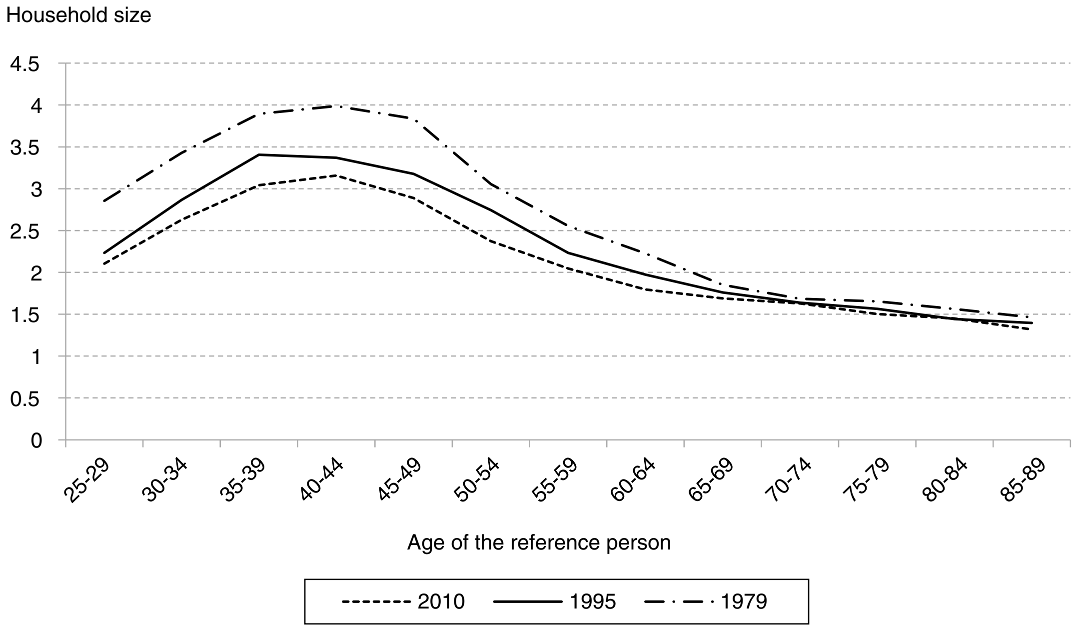
HOUSEHOLD SIZE BY THE AGE OF THE REFERENCE INDIVIDUAL

Figure
Household size by the age of the reference individual in the 1979, 1995 and 2010 *BdF* survey



Reading note: the household size increases and then decreases in line with the age of the reference individual.
Coverage: private households living in Metropolitan France.
Source: 1979, 1995 and 2010 French Household Expenditure survey (enquête Budget de famille - *BdF), authors calculations.*

APPENDIX 2

SIZE OF COHORTS BY DATE OF *BDF* SURVEY

Generation	1979	1984	1989	1995	2000	2005	2010
1901	82	40					
1902	64	62					
1903	63	71					
1904	88	71					
1905	80	81	45				
1906	103	89	40				
1907	87	104	54				
1908	99	100	80				
1909	114	142	89				
1910	124	128	79				
1911	130	110	96	48			
1912	157	160	89	55			
1913	139	150	109	55			
1914	150	159	99	73			
1915	115	147	72	52	38		
1916	82	95	56	46	30		
1917	90	93	66	52	39		
1918	113	106	74	52	49		
1919	130	111	84	61	108		
1920	232	133	139	94	121		
1921	196	203	139	146	112	54	
1922	240	217	164	148	118	56	
1923	232	221	167	128	114	81	
1924	231	223	138	135	140	90	
1925	217	212	138	127	138	79	
1926	251	204	133	138	135	98	68
1927	234	232	159	161	168	116	73
1928	232	207	146	152	147	107	72
1929	240	210	145	145	138	121	101
1930	240	213	150	143	144	118	112
1931	251	220	130	154	150	110	97
1932	243	195	174	146	142	123	103
1933	224	243	134	149	164	96	125
1934	221	216	138	149	160	117	118
1935	235	193	156	147	124	118	125
1936	212	191	152	156	145	124	132
1937	216	201	140	127	146	138	119
1938	202	179	145	151	140	135	105
1939	221	192	129	138	139	137	133
1940	179	218	138	133	130	118	114
1941	184	191	131	129	153	95	100
1942	218	160	126	122	169	130	124
1943	228	203	150	120	185	133	132
1944	215	217	164	155	163	141	131
1945	192	208	157	144	199	118	180
1946	265	226	201	156	215	171	193

➔

Generation	1979	1984	1989	1995	2000	2005	2010
1947	289	289	194	213	236	203	214
1948	276	307	214	251	213	219	206
1949	236	291	223	222	214	204	191
1950	241	277	218	223	254	187	189
1951	201	288	214	212	245	200	181
1952	204	269	213	217	196	212	230
1953	195	277	220	195	198	213	195
1954	177	273	208	225	209	251	220
1955		277	205	244	196	211	217
1956		259	203	202	245	187	219
1957		269	194	217	242	207	226
1958		221	159	208	219	228	192
1959		192	219	197	227	212	218
1960		187	159	206	192	195	214
1961			166	212	203	199	200
1962			138	199	209	216	216
1963			144	235	204	196	228
1964			130	179	198	226	210
1965				188	174	230	210
1966				176	189	223	220
1967				154	163	213	196
1968				130	167	214	212
1969				144	180	202	202
1970				116	172	207	183
1971					164	219	204
1972					141	179	196
1973					120	206	174
1974					91	194	169
1975						178	154
1976						145	154
1977						137	143
1978						143	149
1979						129	137

Reading note: in 1979, there are 82 observations in the cohort of individuals born in 1901.
Coverage: private households living in Metropolitan France.
Source: 1979 - 2010 French Household Expenditure surveys (enquêtes Budget de famille - BdF*), authors calculations.*

Calculating the standard of living of a household: one or several equivalence scales?

Henri Martin *

Equivalence scales, used to compare the standard of living of households of different size and composition, take into account the economies of scale resulting from pooling income and expenditure within households. Two approaches can be used to estimate these scales: an "objective" approach based on modelling household consumption expenditure, or a "subjective" approach based on how households perceive their standard of living. This article focuses on the latter.

Using data from the 1995 to 2011 editions of the French Household Expenditure survey *(Budget de famille)* by Insee, estimations of equivalence scales highlight the sensitivity of results to the model specification, estimation coverage, the choice of subjective living standard indicators and the conventions used to calculate the cost of dependent children.

The subjective approach does not give a robust identification of a single equivalence scale. It does, however, provide a set of possible equivalence scales; for instance, the adult equivalent for a child under 14 ranges from 0.15 to 0.8, while standard equivalence scales are based on a convention, such as 0.3 for the OECD-modified equivalence scale. Thus, for studies using these instruments, or for public policy, it may be preferable to consider a set of equivalence scales rather than just a single scale.

JEL Codes: D13, J18.

Keywords: equivalence scales, economies of scale, subjective method, child costs.

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This article is translated from « Calculer le niveau de vie d'un ménage : une ou plusieurs échelles d'équivalence ? »

In order to study topics such as poverty or inequality, economists calculate the standard of living of households, defined as groups of individuals who share their income and expenditure. Comparing the standard of living of households requires taking into account the demographic composition of these households, their disposable income, and also any economies of scale obtained by pooling their income and expenditure. For example, between a couple with no children and a monthly income of €1,500 and a couple with two dependent children and a monthly income of €2,100, who has the higher standard of living? The second couple may have more income, but they have greater expenditure due to dependent children. In practice, the most common equivalence scales assign a "weight" to each individual in a household, representing the additional income required by the household for each additional individual in order to obtain the same standard of living as a household composed of a single individual. Given economies of scale, i.e. the fact that the needs of two (or three, etc.) people living together are not twice (or three times, etc.) those of a person living alone, the ratio of additional individuals to the additional income required to maintain the same standard of living is less than one: hence the weight assigned to an additional individual is always below 1. A household's standard of living is calculated by comparing its disposable income[1] with the sum of these weights (sometimes called the number of consumption units or "adult equivalents"). Weight values differ depending on the methodology and approach.

This paper will briefly summarise the most common equivalence scales and associated criticisms, and then focus on the problem of estimating these equivalence scales. The data used is taken from the latest editions (1995, 2001, 2006 and 2011) of the French Household Expenditure survey (*Budget de famille*) by Insee. The most recent publication on this topic in France is an article from the late 1990s (Hourriez & Olier, 1997), where the authors examined the relevance of changing the equivalence scale used up until that time for household income statistics and studies. Following on from this publication, our study analyses the benefit of using a set of equivalence scales rather than a single equivalence scale. The aim is to show that currently available methods do not allow us to rule on a single equivalence scale, but only give a possible range of coefficients. This instrument must therefore be used carefully.

A brief review of equivalence scales

The issue of equivalence scales goes back to the aftermath of the Second World War and the implementation of public policies to fight poverty. Economists have long studied the topic and this period saw the first articles proposing estimation methods (Prais, 1953; Friedman, 1952; Prais & Houthakker, 1955). The "Oxford scale" dominated the literature from the 1950s (see Hourriez & Olier, 1997) and an OECD report recommended its use in 1982, explaining why the literature also refers to it as the "OECD scale". This scale assigns the first adult in a household a weight of 1, each additional adult a weight of 0.7, and each child (a person under 14) a weight of 0.5. The sum of these weights gives the number of adult equivalents in the household. However, in the early 1990s, after a review of empirical research on the topic, the OECD opted for a new scale of reference, assigning each household a number of adult equivalents equal to the square root of the number (N) of individuals living in the household. Knowing the age of individuals is not necessary for calculations based on the "square root of N" scale, making it easier to use. [1]

In parallel, Eurostat, who produces harmonised European statistics, gradually replaced the Oxford scale throughout the 1990s with what is known as the "OECD-modified" scale (although the OECD seldom uses it). The OECD-modified" scale also began to appear in the literature (see Hagenaars et al., 1994). Compared to the Oxford scale, the OECD-modified scale gives a lower weight to additional individuals (see Table 1). This scale was adopted by Insee in the late 1990s for computing the standard of living of households, which in turn was used to calculate the poverty threshold and poverty rates[2]. Although this scale is in widespread use across most European countries, some researchers still use the Oxford scale. Some countries also favour other methods of defining the poverty line. In the United States, the poverty line is defined by the US Census Bureau on the basis of a basket primarily composed of food items[3].

1. The sum of all household income (salaries and business income, property income, social security benefits, and net transfers from other households), net of compulsory contributions.
2. In France and for European statistics, the poverty threshold is currently set at 60% of the median standard of living. It was set at 50% until the late 1990s and many academic publications still use this threshold.
3. The value of this basket is modified each year in line with inflation. For an individual living alone, the poverty line was $1,026 dollars per month (€850 euros) in 2014 (excluding Alaska and Hawaii). In comparison, the poverty theshold was €1,007 per month in France the same year.

In theory, equivalence scales depend on the type of welfare system in question. They are influenced by the fact that some expenses are covered by the welfare system and others are borne by households. For example, in English-speaking countries where higher education is generally paid for by households, the cost of a child is probably higher (especially when aged over 18) than in a country like France where higher education is subsidised. Theoretically, an equivalence scale is valid for a given welfare and fiscal system, and may be distorted if the system changes. However, in practice, the OECD-modified scale and the Oxford scale have become international standards.

The equivalence scale concept and the assumptions on which it is based have long been the subject of various criticisms (e.g. see Lechêne, 1993). One criticism involves the implicit theoretical approach, referred to as the "unitary" approach to household behaviour, which assumes that the household maximises a utility function under a budget constraint. However, these two points can be challenged, as they contradict the principle of methodological individualism (Chiappori, 1992) and raise the question of how to combine the utilities of the different household members. On these issues, none of the solutions proposed in the literature is fully convincing (see Blackorby & Donaldson, 1993). Furthermore, the use of an equivalence scale implicitly assumes that the income of all members of the household is pooled and excludes the possibility of inequality within the household. All its members are presumed to have the same standard of living. This assumption can hide inequalities within households, e.g. between men and women or between parents and children. These diverse criticisms led to the emergence of "collective" approaches to the household (see, in particular, Chiappori, 1988; Browning et al., 1994; and for an informal presentation, Donni & Ponthieux, 2011). This approach explicitly recognises that households are composed of various individuals, each with their own preferences and utility function. A number of recent publications have introduced the concept of "indifference scale" which involves comparing the utility of a single individual in two different family contexts (Browning et al., 2013; Chiappori, 2016). European and French data on intra-household sharing of resources have shown that these resources might be pooled in significantly different ways and to various extent: in particular, couples are more likely to fully pool their income than other types of household. In France, for instance, less than two thirds of couples claim to fully pool their resources (Ponthieux, 2013).

Another common criticism of equivalence scales is that they tend to lead to the assumption that the "cost" of an additional individual is proportional to the household income. For example, using the OECD-modified scale, the cost of a child under 14 is estimated to be 0.3 times the income of someone living alone. In 2014, this represented approximately €300 per month for someone living alone with a standard of living close to the poverty line, €500 per month for a median standard of living and €900 per month for person at the ninth decile[4]. This assumption is highly debatable (Koulovatianos et al., 2004),

4. *Here the poverty line is considered to be 60% of the median income. In 2014, the median standard of living of the French population was €20,150 per year with the final decile at €37,300 per annum (Argouarc'h & Boiron, 2016).*

Table 1
Equivalence scales for various household composition

Household composition	Equivalence scale: Oxford		OECD-modified		Square root of N
Person living alone	1		1		1
Couple without children	1.7		1.5		1.41
Couple with children — Age of children	Under 14	14 and over	Under 14	14 and over	
Couple + 1 child	2.2	2.4	1.8	2.0	1.73
Couple + 2 children	2.7	3.1	2.1	2.5	2.00
Couple + 3 children	3.2	3.8	2.4	3.0	2.23
Single parent + 1 child	1.5	1.7	1.3	1.5	1.41
Single parent + 2 children	2.0	2.4	1.6	2.0	1.73
Single parent + 3 children	2.5	3.1	1.9	2.5	2.00

Reading: with the Oxford scale, a couple without children needs an income of 1.7 x R in order to achieve the standard of living of someone living alone with an income R.

but rejecting it would mean having to define as many scales as there are standards of living.

Although these criticisms are substantiated and well-known, as yet no alternative methodology has emerged to compute standards of living (Canberra Group, 2011). This paper will therefore remain within the standard framework of equivalence scales (i.e. the unitary approach to the household).

Two approaches to estimating an equivalence scale

In the literature, two methods have been developed for estimating these scales: an "objective" approach and a "subjective" approach.

The objective approach involves modelling household demand for various goods as a function of both income and household composition. However, in order to be identified, these models require an identifying assumption, but this assumption is not testable with survey data (Blundell & Lewbel, 1991). In practice, this means that statisticians must define their own measure of a household's standard of living. Two major assumptions have been proposed in the literature. The first is the Engel curve (1857), whereby the standard of living of a household depends on the share of its budget spent on food. The greater the budget share spent on food, the lower the standard of living. The budget coefficient for food therefore determines a household's standard of living. This assumption was credible in the 19th century when food represented up to 80% of the household budget, but is much less so today in a context where the structure of consumption has become much more varied. The second is Rothbarth's assumption, whereby expenditure on goods exclusively consumed by adults could be used to measure a household's standard of living. In other words, the more a household spends (in absolute value) on the purchase of goods for adults, the higher its standard of living. The problem for statisticians is then to distinguish, among the household expenses, those made exclusively for the adults. In the literature, the goods the most often used are adult clothing or tobacco and alcohol expenditure.

While these assumptions have the advantage of being based on objective data (household consumption expenditure), they are open to criticism on several levels. Firstly, the choice of a measure of standard of living is by-and-large conventional, and the estimated scale therefore reflects the definition of the standard of living set *ex ante* by the statistician. Next, these assumptions do not take into account the changing preferences of households as they grow in size. For example, the birth of a child may lead a couple to change their lifestyle and significantly reduce their expenditure on "adult goods," without leading to a reduction in its standard of living. But even if the assumption made in defining the standard of living seems generally credible, just a slight deviation from reality can result in biased estimations.

The "subjective" approach, proposed for the first time in the literature by Kapteyn and Van Praag (1976), has been adopted in this paper. Its main advantage is that (unlike the objective method) estimations do not need be based on a definition of the standard of living set arbitrarily by the statistician (Hourriez & Olier, 1997). The standard of living assigned to each household is based either on the household's opinion of its own standard of living or on the population's average opinion of its standard of living. The variables used are therefore not household expenditure, but result from questions on the household's perception of their standard of living. In general, this approach has been used less often in the literature by economists, who are rather inclined to give greater credit to what individuals really do rather than to what they say they do (Accardo, 2007). However, various authors have adopted a subjective approach, based on questioning households directly on their standard of living (Flik & Van Praag, 1991) or on the income level they consider to be minimum, average or comfortable for a household such as theirs (Van der Bosch, 1996). In France, various research projects have used this approach, with studies published by Bloch and Glaude (1983), Glaude and Moutardier (1991) and Hourriez and Olier (1997), all of which are based on different editions of the French Household Expenditure survey. This paper contributes to the existing literature in three ways. Firstly, estimations are based on the latest editions of the survey, allowing us to explore recent changes in the estimated coefficients. Secondly, unlike previous empirical work, confidence intervals are provided. Finally, a large number of robustness tests have been conducted.

Estimations based on a subjective approach

In the French Household Expenditure survey *Budget de famille*, three variables can be used to determine a household's perceived standard of living: AISE, NIVEAU and RMINI. The first two are based on asking households questions on how they perceive their financial situation and standard of living respectively. The third asks them to assess the minimum income they consider necessary, for a household like their own,to meet their needs (Box 1).

More precisely, the subjective method is based on modelling an indicator of the household's standard of living or unobserved utility function, U, which is an increasing function of its income R and a decreasing function of its size N. The parameters of this standard of living indicator U are estimated using an ordered logistic model for the variables AISE and NIVEAU. The indicator represents the model's latent variable. Socio-demographic variables are also introduced in order to control for the observed heterogeneity of households as well as possible. The standard of living indicator is written as:

$$U(R,N) = \alpha \cdot \log(R) + \beta \cdot N + \gamma \cdot \log N + controls + \varepsilon$$

The idea is to identify the additional income required to maintain the household's standard of living with an additional dependent person – in other words, find the multiplication factor $m(N)$ by which the income R of an individual living alone must be multiplied for him or her to maintain the same standard of living with N-1 additional dependent individuals (spouse or children). The following equation is solved:

$$U(R,1) = U(R \cdot m(N), N)$$

The following type of multiplication factors or equivalence scales are obtained:

$$m(N) = N^{\frac{-\gamma}{\alpha}} \cdot e^{(1-N)\frac{\beta}{\alpha}}$$

In order to take into account the age of the children, Hourriez and Olier define N as the "adjusted" household size, with $N_{under\ 14}$ representing the number of children under 14 in the household and $N_{14\ and\ over}$ representing the

Box

HOUSEHOLDS' PERCEPTION OF THEIR STANDARD OF LIVING IN THE FRENCH HOUSEHOLD EXPENDITURE SURVEY (*BUDGET DE FAMILLE*)

The French Household Budget survey (*Budget de famille*) is a survey conducted by Insee every five years since 1979, covering the population living in private households. The survey was conducted in 1979, 1985, 1989, 1995, 2001, 2006 and 2011. Its main aim is to study the income and consumption expenditure of households. There are also questions on the household's perception of its financial situation. Three variables, AISE, NIVEAU and RMINI, can be used to estimate an equivalence scale using the subjective method.

The *variable AISE* corresponds to the following question: "Please tell me which of the following options best describes your budget"

- You are comfortably off (10%)
- You manage (29%)
- You manage, but you have to be careful (43%)
- It's difficult to get by (16%)
- You cannot get by without contracting debts (3%)

The *variable NIVEAU* was introduced for the 1995 survey. It corresponds to the following question: "How would you qualify your standard of living?"

- Very high (0.6%)
- High (6%)
- Moderately high (46%)
- Moderately low (32%)
- Low (12%)
- Very low (4%)

For these two variables, the percentage of responses for each category in the 2011 survey are given in brackets.

The *variable RMINI* corresponds to the following question: "In your opinion, what is the minimum monthly income currently required for a household like your own to simply meet its needs?"

Unlike the two other variables, RMINI is a continuous variable and has been present in the same form in all editions of the survey since 1979. In the 2011 survey, the mean level of the variable was €2,230 per month (differing of course with the household size). This variable is more difficult to use since it does not directly refer to the household's perceived standard of living.

Hourriez (1996) shows that the answers given by households to these three questions are consistent with one another (strong correlations) and that they vary as expected according to some other economic or demographic variables (in particular, the income and the number of individuals living in the household).

number of people (children and adults) aged 14 or over in the same household. After estimating the weighting factor for children under 14, the authors adopted the following equation for the adjusted household size:

$$N = 0.55 \cdot N_{under14} + N_{14and\,over}$$

Initially, we will keep this definition of the adjusted household size with a weighting factor of 0.55 in order to use a methodology comparable to that of Hourriez and Olier. This choice of a weighting factor of 0.55 and an age threshold of 14 for children will be discussed in the second part of the study.

Unlike the AISE and NIVEAU variables, RMINI is a continuous variable. In this instance, following the method proposed by Kapetyn and Van Praag (1976), the indicator of the household's standard of living is defined as $U = \log\left(\frac{R}{RMINI}\right)$ where R remains the household income. The household's standard of living is therefore determined using the ratio between effective disposable income and the income deemed necessary to meet the household's needs. A household with an income below what it considers the minimum to meet its needs will be assigned a low standard of living. Likewise, if its income is much higher than this minimum income, its standard of living will be considered high. The estimated model is a linear regression:

$$\log(RMINI) = constant + \alpha \cdot \log(R) + \beta \cdot N \\ + \gamma \cdot log(N) + controls + \varepsilon$$

which is equivalent to:

$$U(R,N) = \log\left(\frac{R}{RMINI}\right) \\ = -constant + (1-\alpha) \cdot \log(R) \\ - \beta \cdot N - \gamma \cdot \log(N) - controls - \varepsilon$$

The associated equivalence scales take the form of:

$$m(N) = N^{\frac{-\gamma}{\alpha-1}} \cdot e^{(1-N)\frac{\beta}{\alpha-1}}$$

To begin, we use the exact same estimation method as the one used by Hourriez and Olier (hereafter H&O 1997). The models, variables used (Appendix 1) and the scope of the estimations are therefore identical. The sample includes all households consisting of an individual living alone or a couple with or without dependent children under 25. We also keep the same definition of income as the authors (that is, self-reported income before tax). The objective is two-fold. Firstly, to study changes in the equivalence scales over time by conducting the same estimations with more recent data. Secondly, to provide confidence intervals for the coefficients (which the other authors were not able to do). These confidence intervals are obtained using the delta method, which provides variance estimators for non-linear transformations of estimated parameters. These intervals are valuable for assessing changes in the equivalence scales between 1995 and 2011. H&O 1997 chose to focus primarily on the variable AISE in their calculations, which is why the results using this indicator are presented first (Table 2).

Given the confidence intervals, it is not possible to conclude that the equivalence scales follow a linear evolution between 1995 and 2011. Nevertheless, these confidence intervals give an idea of the accuracy of the estimations. For example, using data from the 2011 survey, the confidence interval for a couple with 2 dependent children over 14 is between 2.44 and 3.02.

To assess the robustness of estimations contributing to the choice of the standard of living indicator, similar estimates were made using the variables NIVEAU and RMINI for the most recent survey (Table 3). This shows that the results present a high level of sensitivity. The estimations are even sometimes contradictory when the confidence intervals do not match. This is especially true for estimations using the variable RMINI, where the confidence intervals are much narrower than when the estimations are based on other indicators[5]. The equivalence scale obtained using the variable AISE is relatively close to the OECD-modified scale, despite the wide confidence intervals. With the variable NIVEAU, however, a smaller number of consumption units is assigned to families, and NIVEAU seems to result in larger economies of scale than AISE. The variable RMINI shows a non-linear shape: each additional individual adds a significantly decreasing number of consumption units (0.48 for the first, then 0.26, 0.11 and finally 0.02 for the fourth).

5. This is explained by the continuous nature of the RMINI variable.

The sensitivity of the estimations to the chosen living standard indicator raises the question of the information captured by these variables. NIVEAU seems to question households more directly on their standard of living, but assumes that this complicated concept has been understood by respondents. The median categories ("relatively high" and "relatively low" standard of living) are chosen by the vast majority of households (almost 80%), which makes it difficult to distinguish between standards of living. The variable AISE poses other problems. It introduces almost objective considerations regarding management of the household finances, making direct reference to budget and the notion of debt. These considerations may be disconnected from a household's perception of its standard of living. For example, a well-off household may report a "high" or "very high" standard of living and reply "It's tight but we manage". Finally, for the variable RMINI, the respondents may have understood "income" to be a "perceived" income, including salary and major transfer income (unemployment benefits and pension), but ignoring other social benefits (family, housing and child benefits, etc.)[6]. The subjective approach choses to calculate equivalence scales using the "utility" (or standard of living) reported by the household. If a child is

6. This possibility was proposed by Jean-Michel Hourriez (1996).

Table 2
Equivalence scales estimated using the indicator AISE, H&O1997 method

Household composition	*OECD-modified scale*	H&O 1997 1995	1995	2001	2006	2011
Individual living alone	*1*	1	1	1	1	1
Couple without children	*1.5*	1.42	1.42 [1.33 ; 1.50]	1.44 [1.37 ; 1.52]	1.51 [1.43 ; 1.59]	1.51 [1.41 ; 1.61]
Couple + 1 child aged 14 or over	*2.0*	1.86	1.86 [1.72 ; 2.00]	1.87 [1.75 ; 1.98]	2.02 [1.90 ; 2.15]	2.08 [1.91 ; 2.24]
Couple + 2 children aged 14 or over	*2.5*	2.38	2.37 [2.16 ; 2.59]	2.31 [2.13 ; 2.49]	2.60 [2.38 ; 2.81]	2.73 [2.44 ; 3.02]
Couple + 3 children aged 14 or over	*3.0*	3.00	2.98 [2.59 ; 3.36]	2.79 [2.46 ; 3.11]	3.24 [2.85 ; 3.63]	3.51 [2.95 ; 4.06]

Note: 95% confidence intervals (delta method) are shown in square brackets.

Reading: using H&O 1997 method for the 2011 French Household Expenditure survey, a couple without children needs an income of 1.51 x R in order to achieve the standard of living of someone living alone with an income R.
Coverage: households composed of people living alone, couples without children or with children under 25, and single-parent families with children under 25, representing 8,820 households in 1995, 9,479 households in 2001, 9,539 households in 2006 and 14,053 households in 2011.
*Source: Hourriez & Olier (1997) and Insee, Household Expenditure survey (*Budget de famille*) 1995, 2001, 2006, 2011.*

Table 3
Equivalence scales estimated for the three indicators of living standard, H&O1997 method

Household composition	Equivalence scale			Estimation for the three indicators		
	Oxford	OECD-modified	Square root of N	RMINI	NIVEAU	AISE
Individual living alone	1	1	1	1	1	1
Couple without children	1.5	1.7	1.41	1.48 [1.47 ; 1.50]	1.32 [1.23 ; 1.41]	1.51 [1.41 ; 1.61]
Couple + 1 child aged 14 or over	2.0	2.4	1.73	1.74 [1.72 ; 1.76]	1.60 [1.48 ; 1.73]	2.08 [1.91 ; 2.24]
Couple + 2 children aged 14 or over	2.5	3.1	2.00	1.85 [1.83 ; 1.87]	1.89 [1.70 ; 2.07]	2.73 [2.44 ; 3.02]
Couple + 3 children aged 14 or over	3.0	3.8	2.24	1.87 [1.84 ; 1.89]	2.18 [1.87 ; 2.49]	3.51 [2.95 ; 4.06]

Note: 95% confidence intervals (delta method) are shown in square brackets.

Reading: using H&O 1997 estimation method with the indicator NIVEAU, a couple needs an income of 1.32 x R in order to achieve the standard of living of someone living alone with an income R.
Coverage: households composed of people living alone, couples without children or with children under 25, and single-parent families with children under 25, representing 14,053 households.
*Source: Insee, Household Expenditure survey (*Budget de famille*) 2011.*

wanted and creates surplus "utility", the "cost" of the child could be negative (if the increase in utility exceeds that of the expenditure associated with the child). The variable NIVEAU records a standard of living that could be considered equal to the total utility of a household, whereas the AISE variable focuses more on financial aspects. In the end, although it is less obvious than with the objective approach, statisticians implicitly define the standard of living within a subjective method by formulating survey questions in a certain way.

The second stage of this study aims to improve the methodology adopted by Hourriez and Olier. A previous analysis of the main determinants of perceived standard of living (Martin, 2015) lead to the following choices:

- The coverage of the estimations is restricted. Specifically, single-parent families and households with a reference individual aged over 64 have been excluded. These households present specific response behaviours to the questions AISE and NIVEAU (Martin & Périvier, 2015). The sample has therefore been limited to people living alone and couples whose reference individual was aged between 25 and 64 at the time of the survey.

- Secondly, the notion of disposable household income is preferred to pre-tax income. In the 2011 Household Expenditure survey, the household income is derived from tax data, ensuring greater reliability.

- Finally, two additional control variables have been introduced (Appendix 1) to take into account recent changes to the household's standard of living and its net worth. These variables are important in determining perceived standard of living (Martin 2015). These additions will be referred to hereafter as the "supplemented H&O" method.

Once again, the results show that the choice of one indicator over another produces different estimates. They also indicate that estimations are extremely sensitive to the model specification (see Tables 3 and 4). Even marginal changes in the sample, the definition of income or the control variables result in different estimations. For example, for the variable AISE, a household composed of a couple and two children aged 14 or over is attributed a coefficient of 3.27 with the supplemented H&O method, compared to 2.73 with the H&O 1997 method. To gain a better understanding of the reason behind the deviations between the two methods, we performed two other estimations, first varying only the sample, and then the sample and the definition of income. The results (Appendix 2) show that most variations can be explained by the control variables added.

The results obtained with the RMINI indicator are very different from those obtained with AISE and NIVEAU. With this variable, each additional person contributes a strongly decreasing number of consumption units. In the linear model estimation for the RMINI variable,

Table 4
Equivalence scales estimated for the three indicators of living standard, H&O 1997 supplemented method

Household composition	RMINI		NIVEAU		AISE	
Individual living alone	1		1		1	
Couple without children	1.43 [1.39 ; 1.48]		1.33 [1.20 ; 1.47]		1.56 [1.42 ; 1.70]	
Couple with children						
Age of children	Under 14	14 and over	Under 14	14 and over	Under 14	14 and over
Couple + 1 child	1.57 [1.52 ; 1.61]	1.64 [1.58 ; 1.69]	1.51 [1.42 ; 1.70]	1.68 [1.49 ; 1.87]	1.93 [1.74 ; 2.12]	2.29 [2.05 ; 2.63]
Couple + 2 children	1.65 [1.61 ; 1.70]	1.71 [1.64 ; 1.77]	1.70 [1.62 ; 1.79]	2.07 [1.80 ; 2.33]	2.38 [2.12 ; 2.63]	3.27 [2.84 ; 3.70]
Couple + 3 children	1.70 [1.65 ; 1.76]	1.69 [1.61 ; 1.78]	1.90 [1.81 ; 2.00]	2.51 [2.06 ; 2.95]	2.89 [2.55 ; 3.23]	4.60 [3.73 ; 5.48]

Note: 95% confidence intervals (delta method) are shown in square brackets.

Reading: using H&O supplemented method with the indicator NIVEAU, a couple needs an income of 1.33 x R in order to achieve the same standard of living the standard of living of someone living alone with an income R.
Coverage: all households composed of individuals living alone, couples without children or with at least one dependent child under 25. The person of reference must be older than 25 and younger than 64 at the time of the survey. The estimation includes 8,601 households for the variable RMINI, 9,020 households for the variable AISE and 8,932 households for the variable NIVEAU (differences are due to non-response).
*Source: Insee, Household Expenditure survey (*Budget de famille*) 2011.*

the β and γ parameters, associated with the *N* and *log(N)* variables respectively, are of opposite signs. The equivalence scale obtained does not therefore strictly increase as a function of N. This counter-intuitive result raises questions about the relevance of the indicator. In a study on RMINI, Gardes and Loisy (1997) demonstrate that the response behaviour of households varies significantly according to their income. For households at both ends of the income distribution (the least and most well-off households), RMINI is more an assessment of basic requirements. For intermediate households on the other hand, RMINI is more about demanding a higher standard of living. These differences suggest that the relationship between a household's income and RMINI is not a reliable indicator of the standard of living. For this reason, in the rest of this study we will focus on the indicators based on AISE and NIVEAU.

What age threshold for children?

The Oxford scale and OECD-modified scale assume a shift in the "cost of a child" at 14 years old. For example, for the OECD-modified scale, a child under 14 represents 60% of the cost of an adult[7]. From the age of 14, a child is assumed to "cost" as much as an adult. This threshold may have been relevant in the 1950s, when the household budget was primarily spent on food expenditure. The age of 14 represents the beginning of adolescence, when food requirements start to be comparable to that of adults. However, in recent years, expenditure on higher education has significantly increased, whereas only a limited number of households were affected in the 1950s. It therefore seems possible that children over 18 (in higher education) generate additional expenses for their family and that the cut-off point in the cost of a child is gradually shifting towards the age of 18.

All the results presented above are based on two major assumptions: the cut-off point for the cost of a child is set at the age of 14, and the relative cost of a child under 14 as compared to an adult is 0.55. These two assumptions must be examined to greater depth. H&O 1997 proposes the following ordered logistic model for the variable AISE in order to estimate the cost of children according to their age:

$$\begin{aligned} U(R,N) = {} & \alpha \cdot \log(R) + \beta_1 \cdot N_{0-4} + \beta_2 \cdot N_{5-9} \\ & + \beta_3 \cdot N_{10-14} + \beta_4 \cdot N_{15-19} + \beta_5 \cdot N_{19-24} \\ & + \beta_6 \cdot N_{adults} + controls + \varepsilon \end{aligned}$$

where the variables N_{x-y} represent the number of dependent children in the household aged between x and y, and *R* the household's pre-tax income. The cost of a child in the x-y age bracket corresponds to the factor *c* by which the income of a single parent with a dependent child in this age bracket should be multiplied for this household to have the same standard of living as an individual living alone. The equation to solve is:

$$U\left(Rc, N_{x-y} = 1\right) = U\left(R, N = 0\right)$$

to give: $c = e^{-\beta_{x-y}/\alpha}$, where β_{x-y} represents the parameter associated with the variable N_{x-y}. The parameter *c* gives the cost of a child as a percentage of the income of an individual living alone. In order to generate comparable results, the H&O 1997 method was used with data from the latest editions of the French Household Expenditure survey (Table 5).[7]

The confidence intervals are relatively wide, making it difficult to identify any significant change to the cost of a child per age group between 1979 and 2011. Nevertheless, the 2011 survey stands out with an especially high cost of a child aged 0 to 4 and a relatively low cost for 20-24 year olds. The same calculation was made using the indicator NIVEAU to test the robustness of these results (Table 6).

Once again, the two living standard indicators (AISE and NIVEAU) give different estimates. For example, for the 2011 survey, the perceived cost of a child aged 10 to 14 calculated using the variable NIVEAU is not significantly different from 0 (Table 6), unlike the estimate calculated using the variable AISE (Table 5). The perceived cost of a child is generally lower with the variable NIVEAU than with the variable AISE. There seem to be two age thresholds after which child costs increase. The first is around 14 years old and the second around 20. However, considering the values of the confidence intervals and the sensitivity of the estimation to the year of the survey, it is not possible to choose between one or the other of these thresholds. In the remainder of this study, the 14 year-old threshold limit has therefore been retained.

7. 0.3/0.5=0.6

The relative cost of a child under 14, expressed as μ, is used as follows for calculating the adjusted household size:

$$N = \mu \cdot N_{under 14} + N_{14\ and\ over}$$

μ represents the ratio between the cost of a child under 14 and the "cost" of an adult. $N_{under\ 14}$ represents the number of children aged under 14 in the household and $N_{14\ and\ over}$ represents the number of individuals (children and adults) aged 14 or over in the same household. Hourriez and Olier (1997) estimate μ based on the variable AISE and using the following ordered logistic model:

$$U(R,N) = \alpha \cdot \log(R) + \beta \cdot N_{14 and over} + \gamma \cdot N_{under 14} + controls + \varepsilon$$

Table 5
Estimation of the perceived cost of an additional dependent individual according to his or her age, H&O 1997 method with the indicator AISE (%)

	Hourriez and Olier (1997) results				Estimation using the method H&O 1997			
Additional individual aged:	1979	1985	1989	1995	1995	2001	2006	2011
Under 5	21	20	18	12	21 [14 ; 28]	18 [12 ; 25]	17 [10 ; 24]	32 [21 ; 42]
5 to 9	16	15	16	11	10 [4 ; 17]	17 [10 ; 23]	20 [14 ; 27]	22 [12 ; 31]
10 to 14	22	18	20	18	18 [14 ; 28]	13 [7 ; 18]	12 [6 ; 19]	17 [8 ; 26]
15 to 19	29	34	28	28	23 [15 ; 30]	19 [12 ; 25]	25 [18 ; 32]	29 [18, 40]
20 to 24	45	38	49	41	36 [26 ; 46]	37 [27 ; 48]	42 [30 ; 53]	32 [18 ; 46]
25 and over	43	47	45	44	42 [32 ; 52]	44 [36 ; 52]	50 [41 ; 58]	47 [37 ; 58]

Note: 95% confidence intervals (delta method) are shown in square brackets.

Reading: with the living standard indicator AISE, an individual with a child aged under 5 needs an income 32% higher than an individual living alone in order to achieve the same standard of living.
Coverage: households composed of people living alone, couples without children or with at least one dependent child under 25, and single-parent families with at least one dependent child under 25. The estimation includes 8,820 households for 1995, 9,479 households for 2001, 9,539 households for 2006 and 14,053 households for 2011.
*Source: Hourriez & Olier (1997) and Insee, Household Expenditure survey (*Budget de famille*) 1995, 2001, 2006, 2011.*

Table 6
Estimation of the perceived cost of an additional dependent individual according to his or her age, H&O 1997 method with the indicator NIVEAU (%)

Additional individual aged:	1995	2001	2006	2011
Under 5	20 [14 ; 27]	17 [11 ; 24]	14 [7 ; 21]	18 [9 ; 27]
5 to 9	8 [3 ; 13]	8 [3 ; 13]	12 [6 ; 18]	13 [5 ; 22]
10 to 14	18 [13 ; 24]	7 [2 ; 13]	10 [4 ; 17]	3 [-3 ; 11]
15 to 19	12 [6 ; 18]	16 [10 ; 21]	17 [10 ; 23]	18 [10 ; 26]
20 to 24	26 [18 ; 33]	19 [12 ; 26]	27 [18 ; 37]	21 [10 ; 33]
25 and over	39 [30 ; 47]	34 [27 ; 41]	35 [27 ; 43]	31 [22 ; 40]

Note: 95% confidence intervals (delta method) are shown in square brackets.

Reading: using the living standard indicator NIVEAU, an individual with a child under 5 needs an income 18% higher than an individual living alone in order to achieve the same standard of living.
Coverage: all households composed of people living alone, couples without children or with at least one dependent child under 25, and single-parent families with at least one dependent child under 25. The estimation includes 8,682 households for 1995, 9,422 households for 2001, 9,483 households for 2006 and 13,897 households for 2011.
*Source: Insee, Household Expenditure survey (*Budget de famille*) 1995, 2001, 2006, 2011.*

The relative cost of a child under 14 is given by the equation $\mu = \frac{\gamma}{\beta}$.

Working with the H&O 1997 methodology, estimations of μ obtained using the latest editions of the survey can be used to explore changes in this parameter over time (Table 7).

Hourriez and Olier finally adopted a value of 0.55 for the parameter μ. This study also initially adopted the same value in order to provide comparable results. However, the estimations underpinning this choice are fragile (Table 7). Depending on the survey edition, the living standard indicator adopted, and taking into account the confidence intervals, the possible values of the parameter μ vary between 0.35 and 0.96. The value selected for the parameter itself has a major impact on the estimation of equivalence scales, as it determines the adjusted household size. In order to assess the sensitivity of estimations to the value adopted for the parameter μ, a number of estimations were made for various values of the parameter. These estimations are based on the indicator AISE, using the latest edition of the survey and applying the "supplemented H&O method." This time, the results appear to be relatively robust (Table 8).

A final estimation, which aims to be as precise as possible (i.e. without considering the parameter μ as necessarily set at 0.55, and using the

Table 7
Estimation of the relative cost of a child under 14, H&O 1997 method with the indicators AISE and NIVEAU

	Results Hourriez and Olier (1997)				Estimations with the H&O 1997 method			
	1979	1985	1989	1995	1995	2001	2006	2011
AISE	0.54	0.55	0.56	0.44	0.54 [0.41 ; 0.67]	0.57 [0.44 ; 0.70]	0.56 [0.43 ; 0.59]	0.77 [0.57 ; 0.96]
NIVEAU	-	-	-	-	0.71 [0.54 ; 0.87]	0.54 [0.38 ; 0.71]	0.59 [0.42 ; 0.76]	0.60 [0.35 ; 0.86]

Note: 95% confidence intervals (delta method) are shown in square brackets.

Reading: in 2011, using the indicator NIVEAU and H&O 1997 method, the cost of a child under 14 relative to an adult is 0.60.
Coverage: all households composed of people living alone, couples without children or with at least one dependent child under 25, and single-parent families with at least one dependent child under 25. For AISE, the estimation includes 8,820 households for 1995, 9,479 households for 2001, 9,539 households for 2006 and 14,053 households for 2011. For NIVEAU, the estimation includes 8,682 households for 1995, 9,422 households for 2001, 9,483 households for 2006 and 13,897 households for 2011 (differences are due to non-response).
*Source: Hourriez & Olier (1997) and Insee, Household Expenditure survey (*Budget de famille*) 1995, 2001, 2006, 2011.*

Table 8
Estimation of equivalence scales using the indicator AISE and the H&O supplemented method with different values adopted for the parameter , 2011

Household composition	$\mu = 0.40$		$\mu = 0.55$		$\mu = 0.70$	
Individual living alone	1		1		1	
Couple without children	1.66 [1.54 ; 1.80]		1.56 [1.42 ; 1.70]		1.51 [1.36 ; 1.65]	
Couple with children						
Children age	Under14	14+	Under14	14+	Under14	14+
Couple + 1 child	1.95 [1.90 ; 2.01]	2.42 [2.17 ; 2.66]	1.93 [1.74 ; 2.12]	2.29 [2.05 ; 2.63]	1.95 [1.84 ; 2.05]	2.17 [1.93 ; 2.42]
Couple + 2 children	2.26 [2.15 ; 2.36]	3.31 [2.85 ; 3.76]	2.38 [2.12 ; 2.63]	3.27 [2.84 ; 3.70]	2.50 [2.32 ; 2.68]	3.08 [2.68 ; 3.47]
Couple + 3 children	2.58 [2.43 ; 2.73]	4.39 [3.49 ; 5.29]	2.89 [2.55 ; 3.23]	4.60 [3.73 ; 5.48]	3.19 [2.93 ; 3.44]	4.32 [3.60 ; 5.05]

Note: 95% confidence intervals (delta method) are shown in square brackets.

Reading: using H&O supplemented method with the indicator AISE, setting the relative cost of a child under 14 at 0.40, a couple needs an income of 1.66 x R in order to achieve the standard of living of someone living alone with an income R.
Coverage: all individuals living alone, couples without children or at least one dependent child under 25. The person of reference is over 25 and under 64 at the time of the survey. The estimation includes 9,020 households.
*Source: Insee, Household Expenditure survey (*Budget de famille*) 2011.*

supplemented H&O method) was made, using the latest edition of the survey. The parameter μ is estimated using the method presented above. Two confidence intervals are provided. The first is calculated using the delta method, considering the parameter μ fixed at the value resulting from its estimation. The second is obtained by bootstrap (999 replications). At each iteration, a sample is selected with replacement. A first estimation of the parameter μ is made using this sample, followed by a second to estimate the coefficients of the equivalence scale. For the second estimation, the adjusted household size (*N*) is recalculated using the estimate of μ. It is then possible to deduce a standard deviation. To estimate the confidence interval, this method takes into account the uncertainty inherent to estimation of the parameter μ. The deviation between the two intervals gives an assessment of uncertainty in the estimation of coefficients, linked to uncertainty in the estimation of the parameter μ (Table 9).

On the basis of this latest estimation, it is possible to propose a set of possible equivalence scales, between an upper scale and a lower scale (Table 10). This is achieved by taking into account both the confidence intervals and the sensitivity of the estimation to the chosen living standard indicator. For the sake of simplicity, a linear form is chosen, which dissociates the consumption units for children under 14 from those for adults. The upper scale is estimated by setting the upper confidence interval limit at 95% for each family situation. The number of consumption units considered is 1.64 for a

Table 9
Estimation of equivalence scales for the indicators NIVEAU and AISE, with prior estimation of the parameter (H&O 1997 supplemented method), 2011

Living standard indicator	NIVEAU		AISE	
Estimate of μ	0.74 [0.35 ; 1.12]		0.88 [0.65 ; 1.11]	
Individual living alone	1		1	
Couple without children CI, set value of μ CI, estimated value of μ	1.32 [1.18 ; 1.44] [1.17 ; 1.45]		1.50 [1.37 ; 1.64] [1.37 ; 1.64]	
Couple with children *Children Age*	*Under 14*	*14+*	*Under 14*	*14+*
Couple + 1 child CI, set value of μ CI, estimated value of μ	1.54 [1.43 ; 1.65] [1.36 ; 1.72]	1.62 [1.44 ; 1.81] [1.41 ; 1.84]	2.02 [1.90 ; 2.14] [1.81 ; 2.22]	2.09 [1.87 ; 2.31] [1.84 ; 2.35]
Couple + 2 children CI, set value of μ CI, estimated value of μ	1.78 [1.62 ; 1.94] [1.53 ; 2.03]	1.96 [1.72 ; 2.20] [1.65 ; 2.27]	2.63 [2.43 ; 2.83] [2.30 ; 2.97]	2.82 [2.49 ; 3.16] [2.39 ; 3.25]
Couple + 3 children CI, set value of μ CI, estimated value of μ	2.04 [1.84 ; 2.23] [1.68 ; 2.39]	2.33 [1.96 ; 2.70] [1.86 ; 2.81]	3.39 [3.10 ; 3.67] [2.82 ; 3.96]	3.74 [3.17 ; 4.31] [2.99 ; 4.50]

Note: two confidence intervals (CI) are presented. The first, for the set value of μ, is computed using the delta method. The second, for the estimated value of μ, incorporates the uncertainty inherent to the estimation of this parameter and is computed by bootstrap (999 replications).

Reading: using H&O supplemented method with the living standard indicator AISE, the relative cost of a child under 14 is estimated at 0.88. Furthermore, when this cost is fixed at 0.88, in order to achieve the same standard of living as someone living alone with an income R, a couple needs an income of 1.50 x R.
Coverage: all individuals living alone, couples without children or at least one dependent child under the age of 25. The reference person must be older than 25 and younger than 64 at the time of the survey. The estimation includes 9,020 households for the variable AISE and 8,932 households for the variable NIVEAU (the difference is due to non-response).
*Source: Insee, Household Expenditure survey (*Budget de famille*) 2011.*

Table 10
Set of equivalence scales: coefficients for an additional individual age according to his or her age

	Child under 14	Adults and children aged 14 or over
Upper scale	0.8	0.9
Central scale (OECD-modified scale)	0.3	0.5
Lower scale	0.15	0.2

Reading: the upper scale assigns 0.8 consumption units to each dependent child under 14 in the household.

couple, 2.35 for a couple with one child aged 14 or over, 3.25 for a couple with two children 14 or over and 4.50 for a couple with three children 14 or over. The number of consumption units attributed to each additional adult or child 14 or over is therefore 0.64 for the first, 0.71 for the second, 0.9 for the third and 1.25 for the fourth, representing an average of 0.875, rounded up to 0.9. The same reasoning is used to deduce the number of consumption units for children under 14. A symmetric method is used to establish the lower scale. The centre scale is obtained using the mean estimates calculated on the basis of the indicators NIVEAU and AISE[8].

* *
*

The aim of this paper is to underline the limitations inherent to the estimation of an equivalence scale. The objective approach is problematic, as statisticians must choose their own definition of a household's standard of living. The subjective approach raises other problems. Firstly, the confidence intervals for the estimated coefficients are particularly wide. Next, the results of the estimations depend on which living standard indicator is chosen. It is difficult to choose one rather than the other. Finally, these estimates are also sensitive to model specifications. The subjective approach does not clearly identify a single equivalence scale. It would be preferable to speak of a set of possible estimates, defined by the confidence intervals obtained from several estimations and to use a set of equivalence scales, built according to the principles set out in this paper. This paper's major contribution is the development of such a set of scales. Researchers who use equivalence scales in their work should apply caution when using these instruments. When selecting one scale across the entire set of possible scales, it is preferable to systematically test the robustness of the results obtained. Legislators should also be aware that the choice of the OECD-modified scale – used to develop a number of social indicators (such as poverty rates or the definition of the poverty line) and certain public policies – is largely a conventional choice.[8] ☐

8. *For example, the number of units adopted for a couple without children is the mean of 1.32 and 1.50, i.e. 1.41. This type of calculation would have created a slightly higher coefficient for children under 14 (0.35), but we chose to retain 0.3 in line with the OECD-modified scale.*

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APPENDIX 1

CONTROL VARIABLES FOR THE ESTIMATED MODELS

The following control variables are used in estimations using the method Hourriez and Olier (H&O 1997):

- the employment status of the person of reference of the household: employment (reference) unemployment, retirement and non-employment – excluding retirement;
- the "socio-occupational" category of the person of reference of the household: managers, self-employed workers, intermediate professions, manual workers, blue collar workers (reference);
- the housing status of the household: homeowners with and without a mortgage, tenants (reference);
- age brackets: 18-29 years old, 30-39, 40-49 (reference), 50 and over;
- a dummy variable for single-parent families;
- a dummy variable for Paris residents;
- a dummy variable for one-earner families.

For the method "Hourriez and Olier supplemented", the following variables are added:

- the household's view of recent changes to its standard of living: significant drop, slight decrease, stability (reference), increase.
- the household's net worth: below €100,000, between €100,000 and €500,000 (reference), over €500,000.

ESTIMATIONS OF EQUIVALENCE SCALES WITH REFERENCE TO THE METHOD H&O 1997, CHANGING FIRST THE COVERAGE AND THEN THE COVERAGE AND THE DEFINITION OF INCOME

Table A.1
Change of the coverage only

Household composition	RMINI	NIVEAU	AISE
Individual living alone	1	1	1
Couple without children	1.48 [1.46 ; 1.50]	1.28 [1.17 ; 1.39]	1.43 [1.32 ; 1.54]
Couple + 1 child aged 14 or over	1.73 [1.71 ; 1.75]	1.54 [1.39 ; 1.69]	1.93 [1.75 ; 2.10]
Couple + 2 children aged 14 or over	1.84 [1.82 ; 1.86]	1.80 [1.60 ; 2.00]	2.54 [2.26 ; 2.80]
Couple + 3 children aged 14 or over	1.85 [1.82 ; 1.87]	2.08 [1.76 ; 2.40]	3.29 [2.79 ; 3.79]

Note: 95% confidence intervals (obtained using the delta method) are shown in square brackets.

Reading: When only modifying the range of the estimations but using all other aspects of the method H&O 1997, with the indicator NIVEAU, a couple needs an income 1.28 times that of an individual living alone (1.28 x R) in order to achieve the same standard of living.
Coverage: all individuals living alone, couples without children or with at least one dependent child under the age of 25. The person of reference must be between 25 and 64 years old at the time of the survey. The estimate includes 8,601 households for the RMINI variable, 9,020 households for the variable AISE and 8,932 households for the variable NIVEAU. The differences are due to non-response.
*Source: French Household Expenditure survey (*Budget de famille*) 2011, Insee.*

Table A.2
Change of the coverage and the definition of income

Household composition	RMINI	NIVEAU	AISE
Individual living alone	1	1	1
Couple without children	1.48 [1.45 ; 1.49]	1.28 [1.16 ; 1.41]	1.46 [1.33 ; 1.58]
Couple + 1 child aged 14 or over	1.72 [1.70 ; 1.73]	1.48 [1.41 ; 1.75]	2.02 [1.83 ; 2.21]
Couple + 2 children aged 14 or over	1.81 [1.79 ; 1.83]	1.91 [1.67 ; 2.13]	2.75 [2.44 ; 3.06]
Couple + 3 children aged 14 or over	1.81 [1.78 ; 1.84]	2.28 [1.90 ; 2.66]	3.70 [3.10 ; 4.30]

Note: see Table A.1.

Reading: When modifying both the range of the estimations and the definition of income, but following all other aspects of the method H&O 1997, with the indicator NIVEAU, a couple needs an income 1.28 times that of an individual living alone (1.28 x R) in order to achieve the same standard of living.
Coverage: see Table A.1.
*Source: French Household Expenditure survey (*Budget de famille*) 2011, Insee.*

Pseudo-panel methods and an example of application to Household Wealth data

Marine Guillerm *

Pseudo-panel methods are an alternative to using panel data for estimating fixed effects models when only independent repeated cross-sectional data are available. They are widely used to estimate price or income elasticities and carry out life-cycle analyses, for which long-term data are required, but panel data have limits in terms of availability over time and attrition.

Pseudo-panels observe cohorts, i.e. stable groups of individuals, rather than individuals over time. Individual variables are replaced by their intra-cohort means. Due to the linearity of this transformation, the linear model with individual fixed effect corresponds to its pseudo-panel data counterpart. The individual fixed effect is replaced by a cohort effect and the model is particularly simple to estimate if the cohort effect can be itself considered as a fixed effect. The criteria for forming the cohorts must therefore take into account a number of requirements. It must obviously be observable for all the individuals and form a partition of the population (each individual is classified into exactly one cohort); beyond this, it must correspond to a characteristic of the individuals that will not change over time (e.g. year of birth). Finally, the size of the cohorts results from a trade-off between bias and variance. It must be large enough to limit the extent of measurement error on intra-cohort variable means, that generates bias and imprecise estimators of the model parameters. However, increasing the size of the cohorts decreases the number of cohorts observed, which makes estimators less precise.

The extension to non-linear models is not direct and only introduced here. Finally, the article provides an application to the French Household Wealth Survey (*enquête Patrimoine*).

JEL codes: C21, C23, C25, D91.

Keywords: pseudo-panel, grouped data, fixed effects models, repeated cross-sectional data.

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Behavioural economics is generally confronted by the fact that many dimensions of the information needed to analyse behaviours cannot be observed in the available data. For example, consumer behaviours depend on individual preferences that are only imperfectly captured in statistical data. Income elasticity estimates are therefore biased. Sometimes it is difficult to dissociate the effects of several variables even though they are observed at the same time. Although age and generation are usually available, it will be impossible to distinguish what derives from one or the other on the basis of cross-sectional data (at a given date). This is particularly detrimental for life-cycle analysis. Take the example of examining variation in wage trajectories over the lifecycle. Cross-sectional data would provide observations on individuals of different ages and for this reason, at various stage of their careers. However, it is not possible, on the basis of this information, to establish that differences observed in wage trajectories result from an effect of age (or professional experience), rather than an effect of generation. The generation effect partially determines the time individuals spend on their education, the job market conditions when they begin their career, which are factors that also influence the wage.

It is standard to use panel data to answer these questions, using observations repeated over time for identical units with the aim of neutralising potentially specific individual characteristics. This usually involves introducing individual "fixed effects" to capture these specific characteristics. Repeated observations of the same variables at different dates helps also to address, at least partly, the aforementioned identification problems. Age varies with time, unlike the generation, which means that the same generation can be observed at different ages. However, this type of data is rare and often limited to small samples and covers short time periods, (this reduces their relevance for life-cycle analysis for example). This type of data is also subject to attrition or non-response problems, making it difficult to follow the same individuals over a long period of time. Over time, the representativeness of panel data can become problematic.

Pseudo-panel methods are one way of making up for the lack of panel data. Their use dates back to Deaton (1985), who was the first to suggest using panel methods on repeated cross-sectional data. The advantage of these data is their availability and the fact that they can cover long periods of time, many surveys being carried out at regular intervals over time. They generally include independent repeated cross-sections, i.e. different samples. Panel methods cannot be directly applied as the observed individuals change at each date. And even with exhaustive sources such as census surveys or certain administrative data, it is not always possible to follow individuals over time for reasons such as confidentiality. However, when the same individuals cannot be followed, types of individuals, generally referred to as "cohorts" or "cells" can be followed. These cohorts are identified by a set of observed characteristics that are stable over time (such as the generation or gender). In the estimations, this make it possible to capture, by a fixed "cohort" effect, some unobserved characteristics that could result in biased estimations. Pseudo-panels have been used to model a wide range of topics, including investment (Duhautois, 2001), consumption (Gardes, 1999; Gardes et al., 2005; Marical & Calvet, 2011), or long-term behavioural changes, such as wage trajectories (Koubi, 2003), women's participation in the labour market (Afsa & Buffeteau, 2005), subjective well-being (Afsa & Marcus, 2008) or living standards (Lelièvre et al., 2010), to mention just the most recent research. In practice, the use of these methods depends on the way in which cohorts are defined. In the case of linear models, standard estimation methods using panel data can be adapted quite easily.

This article provides an introduction to these techniques with an emphasis on practical aspects. After a brief recap of fixed effects models on panel data, it focuses on the principles that should guide the criteria applied for the definition of cohorts. The second part presents estimation methods. These first two sections only cover the case of linear models. The third part provides additional technical information and evokes the extension to dichotomous models. Finally, the last section provides a case study with an application to the French Household Wealth surveys (*enquêtes Patrimoine*).

Issues of implementation of statistical software are not addressed in the articles. Examples of SAS, R and Stata programmes are provided in Guillerm (2015), on which the article is based.

General principle: from individual fixed effects to cohort effects

Why use panel data and what to do when they are not available

The starting point for pseudo-panel models are fixed effects linear models, typically used with panel data. It is therefore useful to present them (for a more detailed presentation, see Magnac, 2005). In general, we want to model the influence of one or more explanatory variables on a variable of interest. We consider here the case of continuous variables of interest. For binary variables, specific methods need to be used (see section on "Estimation of dichotomous models"). The difficulty of estimating these types of models usually stems from the fact that the determinants of the variable of interest are not all observed. If these unobserved determinants are partially correlated with the explanatory variables of the model, there is a risk of incorrectly attributing part of their effect to these explanatory variables.

A classic illustration of this problem is the estimation of the income elasticity of a consumer good. For example, the actual price of food consumption is imperfectly observed: the time spent on the preparation and consumption of meals, which is not valued in the same way by each household, needs to be added to the price of the goods themselves. The value of time increases with income (Gardes et al., 2005). Not taking this value into account results in underestimating the income-elasticity of food consumption.

A typical solution is to use panel data (i.e. repeated observations of the same individuals over time), in order to control factors whose effect is supposed to be constant over time. An individual fixed effect is therefore added to the standard linear model, in order to capture the effect of individual characteristics that are constant over time on the variable of interest[1]:

$$y_{it} = x_{it}\beta + \alpha_i + \varepsilon_{it} \qquad (1)$$
$$i = 1,\ldots,N \quad t = 1,\ldots,T$$

where y_{it} is the variable of interest (in the example, the level of consumption of the good), x_{it} is a vector (line) of K explanatory variables observed for the individual i on the date t (in the example, individual or household income, age, etc.), β is the effect of these variables (i.e. a vector of parameters of dimension K).

α_i is the individual fixed effect. It captures all the determinants of the variable of interest that are fixed over time. Only the parameters associated with variables that are not constant over time can be identified if a fixed effect is introduced into the model. For example, an estimate of the intrinsic effect of gender cannot be obtained if the model includes a fixed effect. Finally, ε_{it} is a residual term, i.e. anything that the model does not take into account. Ignoring the fixed effect in the estimation leads to biased estimators of the effect of the explanatory variables considered when these variables are correlated with the fixed effect.

With repeated observations, the impact of explanatory variables can be estimated using the linear model by neutralising the impact of individual fixed effects. In practice, this can be done by using a transformation of the variables instead of their level, in order to eliminate the individual fixed effect. The most commonly used estimator (as it is the most efficient under certain assumptions) is obtained by carrying out a "within" transformation: at each date we use observations centred on the individual mean over the period, i.e. the transformed variables $z_{it} - \bar{z}_i$, where $\bar{z}_i = \frac{1}{T}\sum_{t=1}^{T} z_{it}$ is the mean of individual values of z over the entire observation period. Another solution would be to directly estimate the fixed effects as model parameters. However, this implies estimating a very large number of parameters (a fixed effect for each of the individuals observed in addition to explanatory variable parameters), which has no real interest for the interpretation[2].

This "within" estimator converges towards the true values of the parameters of interest insofar as the explanatory variables are not correlated

1. Random effects models are another type of modelling traditionally used on panel data. These models also include an individual effect and are another way of taking into account the fact that unobserved characteristics of the individual that are fixed over time have an effect on the variable of interest in modelling. However, unlike fixed effects models, they are based on the assumption that the individual effect is not correlated with the explanatory variables (the individual effect takes into account the correlation of different observations associated with a single individual without overestimating the precision of estimators). If we are able to make such an assumption, there is no point in using pseudo-panels. With independent cross-sections, there is no correlation between the observations, as each individual is only observed once. Models can therefore be estimated directly based on stacked individual data.

2. Especially since if few temporal observations per individual are available, fixed effects estimation lacks precision.

with the remaining residual terms. In other words, the individual impacts at each date, for a given individual, must not be linked to the realisation of any of the explanatory variables included in the model[3].

However, panel methods are based on the observation of the same individuals at different dates, which is rare. In many cases, we have repeated independent cross-sectional data. The principle of pseudo-panels is to follow cohorts (i.e. groups of individuals sharing a set of characteristics that are fixed over time), rather than individuals over time. The model will be considered in terms of these cohorts of individuals rather than the individuals in them. In practice, this means that the observed variables are replaced by the means of these variables within each cohort. These data are treated as panel data and, when possible, panel data estimation techniques are applied.

Life-cycle analysis is another example, as already mentioned with the estimation of income-elasticity and price-elasticity, where pseudo-panel methods are frequently used. If we want to study the accumulation of household wealth over the life cycle, a naïve analysis would study differences in wealth according to age using observations at a given date. However, many other individual characteristics explain the differences in wealth between individuals, such as variations in wage and career, education level, family resources, propensity to save, etc. Some characteristics are correlated with age. For example, this would be the case if some generations had experienced more favourable conditions than others at the beginning of their careers. Failing to take these determinants into account can lead to biased estimations of the effect of age on household wealth. A typical solution is to include these additional aspects (the effect of these variables is "controlled") in a linear model. However, although some of these determinants are usually available in most surveys, this is not always the case. It is therefore easy to obtain measures of age, education level or current salary, but it is more difficult to obtain precise information over the entire career, or on inherited assets, let alone determine if they are "ants" or "grasshoppers" in terms of their propensity to save. As described above, one solution is to estimate a fixed effects model similar to (1).

Life-cycle analysis and the estimation of income or price-elasticity are two examples of issues where pseudo-panels are often used for lack of panel data. Life-cycle analysis requires data over particularly long periods of time, and series of cross-sections provide this time dimension more often than panel data. This justifies the use of pseudo-panel estimations even when panel data are available. For example, Antman and McKenzie (2005) use a rotating panel to assess earnings mobility. Keeping only the new observations entering the panel each quarter (one fifth of the sample) provides them with a long-term data, while they would have been limited to a period of five quarters if they had used the panel. Furthermore, unlike panels, pseudo-panels do not raise issues of sample attrition associated with following households. In the example of earnings mobility, attrition raises problems because it may be related to a move, which itself may result from a change in earnings. Using panel data, Gardes et al. (2005) carry out an estimation of income-elasticity on panel data and pseudo-panels. In the example they use, they show that the estimations are quite close.

Formally, we are interested in $y^*_{ct} = E\left(y_{it} | i \in c,t\right)$, the expectation of the variable of interest in cohort c at date t. The following is obtained from the previous model (by its integration conditional to the date and cohort):

$$y^*_{ct} = x^*_{ct}\beta + \alpha^*_{ct} + \varepsilon^*_{ct} \qquad (2)$$
$$c = 1,\ldots,C \quad t = 1,\ldots,T$$

where for each variable z, $z^*_{ct} = E(z_{it} | i \in c,t)$.

Like the initial model at the individual level, the pseudo-panel model (2) is linear in its parameters, which means that, in principle, standard estimation techniques can be used for panel data. However, in practice, things are a little more complicated.

First, the "true" values y^*_{ct} and x^*_{ct} are not known. We only have an estimation, their empirical counterpart within the observed cohort: $\bar{y}_{ct} = \frac{1}{n_{ct}} \sum_{i \in c,t} y_{it}$ and $\bar{x}_{ct} = \frac{1}{n_{ct}} \sum_{i \in c,t} x_{it}$ (i.e., at each date, the means of observed values for the individuals of the sample belonging to the cohort). The estimation on this sub-sample of individuals may not correspond exactly with "true" values. Fluctuations in the sampling of individuals from a same cohort from one date to another are another problem. Since the observed individuals are not the same at each date, the

3. *In the fixed effects model, this residual term represents all the individual factors that are variable over time and not observed.*

mean of fixed effects $\bar{\alpha}_{ct}$ may vary over time, although in theory, it is constant.

Measurement errors raise different difficulties for estimating model (2), depending on whether they affect the covariates or the variable of interest. Measurement errors on the covariates result in biased estimators (for further details, see "Measurement error model" and Appendix B). The good thing is that the higher the number of individuals of the cohort in the sample, the closer the estimation will be to the true value and the higher the precision of the estimators of the mean values will be, making it possible to neglect measurement errors in the model. On the other hand, measurement errors on the variable of interest and the temporal variability of the cohort effect reduce the precision of estimators and lead to a problem of efficiency if the measurement error is heteroscedastic. Finally, the problem of the variability of cohort effects over time can also stem from how cohorts are defined: beforehand, the effects α^*_{ct} must be able to be considered as constant, otherwise there is a risk of producing biased estimators. These remarks guide the criteria that will be used when defining the cohorts of individuals.

Constructing cohorts

Firstly, the selection criterion must be observable for all the individuals and form a partition of the population (each individual is classified into exactly one cohort). Beyond this obvious point, the criteria for defining the cohorts must not be chosen at random. It must aim to make plausible the assumption that the cohort terms $\bar{\alpha}_{ct}$ are fixed over time. Two distinct factors can call this assumption into question. With survey data, only one sample of the true cohorts is observed. The first source of variation of $\bar{\alpha}_{ct}$ comes from sampling fluctuations: $\bar{\alpha}_{ct}$ corresponds to the mean of fixed effects on the observations of cohort *c* from the sample available at date *t*. It is an estimator of the true value α^*_{ct}, which is not observed. Even if the true cohort is stable, the individuals that represent it change over time. α^*_{ct} can also vary if the true cohort is made up of a population itself unstable over time, especially if the criterion adopted does not correspond to a characteristic of the individuals that is stable over time. This is the second potential source of variation of $\bar{\alpha}_{ct}$.

A stable criterion on a stable population

Choosing a selection criterion that makes α^*_{ct} constant over time eliminates one of the sources of variation of $\bar{\alpha}_{ct}$, to a certain extent. α^*_{ct} is fixed when the true cohorts contain the same individuals at each date. Two conditions are required: that cohorts are constructed on a stable population and on the basis of a stable criterion (otherwise it would mean that the profile of the individuals might change over time).

Year of birth is obviously an example of a selection criterion that corresponds to a stable characteristic of the individuals. In this case, generations of individuals are followed. This criterion is frequently used in pseudo-panel estimations. The term cohort does not imply that only this criterion is valid (some authors use the term "cell"). Other groupings are possible and several criteria can be combined. For example, Bodier (1999) constructs cohorts based on the generation and higher education level to study the effects of age on the level and structure of household consumption. Conversely, a selection criterion based on earnings or the labour market status would not be relevant *a priori* because, for a given individual, it is likely to change over time [4].

However this condition of criterion stability at the individual level is not sufficient. The cohort itself must not change over time either. This issue is particularly crucial for repeated survey data on different samples. In a survey, individuals with a particular profile form a sample of the entire cohort of interest. However in some cases, their representation in the survey may vary depending on the criteria applied to construct the cohort. For example, let us assume that cohorts are defined on the basis of the year of birth. Depending on the date of the survey, the different generations will be represented to varying degrees. They will progressively enter the cohort as they reach the minimum age required to be surveyed (or when young people form new households), whereas the oldest individuals will gradually leave (death, entry into retirement homes or care institutions if out of the scope of the survey). It is important to be aware of these composition effects for analysis if they are linked to the variable of interest. For example, let us assume that we are interested in

4. In practice, there are cases where pseudo-panels have been constructed using criteria that are unstable over time. The relevance of such pseudo-panels must be discussed on a case by case basis. For instance, Marical and Calvet (2011) construct a pseudo-panel based on household age to estimate fuel price elasticities. As age is not a stable characteristic of individuals, even with panel data, the cohorts would not contain the same individuals. However, a pseudo-panel by age can be used to follow households that do not age, and where the family composition (which is linked to fuel consumption) changes little over time.

the profile of the income of successive generations. Life expectancy and income are partially correlated (for example, see Blanpain, 2011). At an advanced age, individuals with the highest income are therefore overrepresented among the "surviving" individuals of a single generation. A cohort analysis that following a generation could suggest that the income of individuals from this generation increases with age, which might not be the case. In practice, a case by case analysis is necessary to assess whether the cohorts represent a stable population over time, even if it means limiting the scope of the analysis. For example, in a study on the effects of age and generation on the level and structure of consumption, Bodier (1999) limited the population to individuals aged 25 to 84, considering that households composed of people beyond these limits may no longer be representative of the population of their generation.

It has to be underlined that this problem is not specific to pseudo-panels, but it is particularly obvious when cohorts are followed over long periods where these entry and exit phenomena (entries onto the labour market, leaving the parents' home, business creation, death, migration, etc.) are likely to occur. However, unlike traditional panel data, attrition problems associated with the difficulty of following identical individuals over time (due e.g. to moving, refusal to answer the next wave of a survey,…) are not an issue.

Large enough cohorts…

The principle of pseudo-panels is to construct cohorts, i.e. profiles, that group together individuals with behaviours considered to be similar. This assumption is even more plausible if precise profiles are defined. However, this can come at a cost, especially with survey data. The smaller the cohort, the greater the extent of errors when measuring empirical means $\overline{y}_{ct}$ and $\overline{x}_{ct}$ and the greater the temporal variability of the means of individual effects $\overline{\alpha}_{ct}$. There will also be even more bias and imprecision issues with the standard estimator (within estimator) covered earlier (for further details, see "Measurement error model" and Appendix B).

Bias and imprecision of estimators can be limited by increasing the size of cohorts. In practice in empirical studies, it is generally considered that 100 individuals per cohort is enough to ignore sampling errors (and therefore simplify the estimation). This choice is based in particular on the studies of Verbeek and Nijman (1992, 1993). Using simulated data, they conclude that the assumption is reasonable (in the sense that the resulting bias is not too high) for categories with at least 100 individuals. However, they recommend cohorts twice as large to significantly reduce the risk of bias.

…while conserving variability

The larger the cohorts, the lower the extent of measurement errors and the bias and imprecision of the estimators that they generate. But the cohorts' size is not the only parameter to be taken into account. It is quite easy to see that for a given sample size, forming large cohorts means that the number of observations used for the pseudo-panel model will be reduced. For example, let us assume that the cohort is built on the criterion of the year of birth but that the repeated cross-sectional data contain few people from one generation at each date. To reduce potential sample fluctuations, one typical solution is to increase the size of the cohorts by broadening the generations (e.g. by five-year age brackets). However in this case, the variability of observations at a given date is reduced, as the final number of useful observations decreases. Grouping close but different generations also means that the variability of these means is reduced over time. These two elements (number of observations used for the estimation, low variability) are both factors that traditionally reduce the precision of the final estimator. Intuitively, the smaller the number of observations, the less precise the estimation is. However, it is also necessary to observe different values of the variables of interest (that is, to be able to observe their variation over time), in order to assess how strongly they are correlated. This reflects a classic bias-variance tradeoff. Forming large cohorts limits the bias of the estimator but causes variability to be lost, which reduces the precision of the estimators. Verbeek and Nijman (1992) show that the bias of the within estimator traditionally used (see below) can be large if the inter-temporal variability is low in relation to the measurement errors, even when the cohorts are large.

In short, a good selection criterion must: (1) be a characteristic that does not change over time on an individual basis, define a stable (sub-) population, and result from a tradeoff so that (2) large enough cohorts can be formed (3) without losing too much variability. These various constraints highly limit the choice of cohort selection criteria. In practice, many studies use the year of birth as this criterion meets many of these requirements, is often available in survey data and is stable. Furthermore, depending on the size of cross-sectional samples, close generations can be grouped to create larger or smaller cohorts. Finally, it is important to remember that this dimension is of interest itself in many studies. The cohort effect can then be directly interpreted as a generation effect, which can be interesting to study. In life-cycle analysis in particular, grouping individuals by generation preserves variability on the "age" variable.

Estimation of pseudo-panel models

When the cohort selection criterion has the qualities required to consider model (2) as a fixed effects model, the parameters are generally estimated based on standard panel data estimation techniques. In practice, the estimated model is therefore:

$$\bar{y}_{ct} = \bar{x}_{ct}\beta + \bar{\alpha}_c + \bar{\varepsilon}_{ct} \qquad (3)$$
$$c = 1 \quad t = 1,\ldots,T$$

We apply a within transformation evoked above, in which, for each cohort, the various variables are centred on the mean of the observed values for the cohort, for all the observation dates. We therefore regress $\bar{y}_{ct} - \bar{y}_c$ on $\bar{x}_{ct} - \bar{x}_c$, where for each variable z, $\bar{z}_c = \frac{1}{T}\sum_{t=1}^{T}\bar{z}_{ct}$. The within estimator is obtained:

$$\hat{\beta}_W = \left[\sum_{c=1}^{C}\sum_{t=1}^{T}(\bar{x}_{ct} - \bar{x}_c)'(\bar{x}_{ct} - \bar{x}_c)\right]^{-1}\sum_{c=1}^{C}\sum_{t=1}^{T}(\bar{x}_{ct} - \bar{x}_c)'(\bar{y}_{ct} - \bar{y}_c) \qquad (4)$$

This allows us to deduce the following cohort effect estimator:

$$\hat{\alpha}_c = \bar{y}_c - \bar{x}_c\hat{\beta}_W \qquad (5)$$

In practice, the within estimator is obtained by carrying out first a within transformation then calculating the least squares estimator on these centred variables. However, this has to be done carefully because, as transformed variables are being used, the standard estimator of the variance obtained with the ordinary least squares procedure does not correspond directly to the unbiased estimator of the variance of the within model. It underestimates it. A multiplying factor $(CT - K)$ / $(CT - C - K)$ needs to be taken into account, where C is the number of cohorts, T the number of observation dates and K the number of explanatory variables. In SAS, the Bwithin macro written by Duguet (1999) takes this problem into account (for other Stata and R procedures, see Guillerm, 2015).

The within estimator is obtained in the same way – either by including cohort dummies or via instrumentation. Including cohort dummies in model (3) can be used to directly obtain the fixed effects estimators[5], which sometimes are of interest in themselves. In a life-cycle analysis where cohorts consist of generations, the generation effect could be estimated directly. Again, it is important to be careful, since the estimation of these fixed effects will lack precision if the number of periods is not large enough.

Moffitt (1993) proposes an alternative estimation method using instrumentation. He shows that the within estimator (4) of the pseudo-panel model technically corresponds to the two-stage least squares estimator on individual data (explanatory variables and cohort dummies), where all cohort-time interaction dummies would be used as the instrument. The formal proof is provided in Appendix A. In order to understand the intuition, remember that in the first step of the two-stage least squares procedure the explanatory variables are projected onto the instruments. The projection of x_{it} onto cohort - date interaction dummies corresponds exactly to the empirical mean $\bar{x}_{ct}$, where c is the cohort to which individual i belongs. The second step involves replacing the instrumented variables in the initial model with their projection, in this case regressing y_{it} on $\bar{x}_{ct}$ and the cohort dummies. The estimator obtained is the same as the within estimator (4).

This can simplify the estimation, because we are working directly on the individual data. This analogy also serves as a basis for extending pseudo-panels to dichotomous models (see "Estimation of dichotomous models"). Another advantage of this approach is that other types of more parsimonious instruments can be used. For example, if the year of birth is adopted, a function of the year of birth (e.g. a polynomial function) can be used to build the instrument

5. Direct estimation of the fixed effects is not recommended with individual data as it requires estimating an extensive number of parameters. For pseudo-panels, the number of cohorts is generally limited. If each cohort has approximately 100 individuals, the number of fixed effects to estimate in the pseudo-panel model is divided by as much in relation to the panel model.

rather than dummy variables associated with a partition of the years of birth.

This approach can also be used to find the criteria for grouping individuals in cohorts[6]. Two conditions are required to construct a good instrument. It must first be correlated with the explanatory variables. This is due to the fact that cohorts must have enough variability to allow the estimation of the model at the aggregated level of cohorts. To understand the underlying intuition, we can use the extreme case where these cohort-date interaction dummies would be completely independent from the model's explanatory variables (i.e. that the distribution of these explanatory variables is identical at each date and from one cohort to another). In this case, the empirical means of these variables at a date and cohort level are very similar, which means that the model cannot be estimated. The other feature of a valid instrument is that it must not be correlated with the unobserved determinants of the variable of interest. Moffitt shows that this property is proven if the cohorts are constructed on the basis of a stable criterion and when the size of the cohorts tends to infinity.

Beyond the estimation itself, several remarks can be made, The first of which concerns the choice of explanatory variables. In the standard fixed effects linear model, only the parameters associated with variables that are not constant over time can be identified: the fixed effect "absorbs" the effect of constant variables. In a pseudo-panel model, the aggregation into cohorts artificially creates variability and gives the impression that the parameters associated with the fixed characteristics are identifiable. For example, a variable that is constant on an individual level such as the dummy variable "being a woman" becomes "the proportion of women in the cohort c on the date t" in the pseudo-panel data. The observed temporal variations (normally low) are only due to sampling errors. Introducing these types of variables in the analysis is therefore not recommended.

Some additional technical points

This section provides two extensions to the standard way of handling technical issues with pseudo-panel estimations: taking into account (1) the heteroscedasticity of residual terms and (2) the heteroscedasticity of measurement errors in the estimation. The models presented so far are only suitable if the variable of interest is continuous. With discrete variables, specific methods need to be used. An introduction to this aspect is presented in a third section.

Heteroscedasticity in pseudo-panels

In practice, cohorts vary in size from one to the other and for a given cohort, between one date and another. These size variations may result in heteroscedasticity in model (2). As the precision of the estimator directly depends on this number, varying degrees of error terms are introduced depending on the cohorts. In the presence of heteroscedasticity, the within estimator (4) is unbiased but the estimator of its precision is biased and the statistical tests are therefore invalid.

The efficient within estimator is obtained by weighting the observations by the cohort's size, which means a least squares estimation of the following model:

$$\sqrt{n_{ct}}\,\overline{y}_{ct} = \sqrt{n_{ct}}\,\overline{x}_{ct}\beta + \sqrt{n_{ct}}\,\alpha_c + \sqrt{n_{ct}}\,\overline{\varepsilon}_{ct} \qquad (6)$$

Just as with the homoscedastic model, $K + C$ parameters need to be estimated. This estimation is easy to implement unless the number of cohorts is too large, in which case a within transformation is generally used with the aim of eliminating the fixed effects before estimation. However in this model, a standard within transformation will not eliminate the cohort dummies because the weight assigned to each cohort (n_{ct}) varies over time. Gurgand et al. (1997) show that in this case the efficient within estimator is:

$$\hat{\beta}_{WP} = \left(X'(WDW)^{-} X\right)^{-1} \left(X'(WDW)^{-} y\right) \qquad (7)$$

where X a matrix of dimension CT × K stacks the line vectors $\overline{x}_{ct}$, y a vector of dimension CT stacks the values $\overline{y}_{ct}$, $(WDW)^{-}$ is the generalised inverse of the matrix WDW, with W the standard within matrix of dimension CT and D the diagonal matrix where the diagonal elements are $1/n_{ct}$.

Measurement error model

The estimation methods presented in the section above do not take into account the fact that the true intra-cohort means noted y^*_{ct} and x^*_{ct}

6. *For more information, see Moffitt (1993) and Verbeek (2008).*

are measured with errors using the means calculated on the sample (noted $\bar{y}_{ct}$ and $\bar{x}_{ct}$). As stated above, these measurement errors pose two problems: Error in the explanatory variables, results in biased estimators and error in the variable of interest as well as the variability of the cohort effect over time reduce the precision of the estimators. The estimation techniques presented above are implicitly based on the assumption that measurement errors can be overlooked. Otherwise, appropriate techniques are required. The estimators of model (2) proposed by Deaton (1985) therefore rely on measurement error models that take this problem into account. He adapts Fuller' theory (1986) to pseudo-panel estimation.

We write u_{ct} and v_{ct}, the measurement errors:

$$\bar{y}_{ct} = y^*_{ct} + u_{ct}$$

$$\bar{x}_{ct} = x^*_{ct} + v_{ct}$$

When they are integrated into model (2), we obtain:

$$\bar{y}_{ct} = \bar{x}_{ct}\beta + \alpha_c + \tilde{\varepsilon}_{ct} \quad (8)$$

$c = 1,\ldots,C \quad t = 1,\ldots,T$

where $\tilde{\varepsilon}_{ct} = \varepsilon^*_{ct} + u_{ct} - v_{ct}\beta$. We show that this residual value is correlated to $\bar{x}_{ct}$.

The estimator of the parameter β proposed by Verbeek and Nijman (1993) relies on a parametric specification of the measurement error and its correlation with the variable of interest (for more information, see Appendix B). It gives:

$$\tilde{\beta} = \left(\frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}(\bar{x}_{ct} - \bar{x}_c)'(\bar{x}_{ct} - \bar{x}_c) - \frac{T-1}{T} \times \frac{1}{n}\hat{\Sigma} \right)^{-1}$$
$$\left(\frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}(\bar{x}_{ct} - \bar{x}_c)'(\bar{y}_{ct} - \bar{y}_c) - \frac{T-1}{T} \times \frac{1}{n}\hat{\sigma} \right) \quad (9)$$

$\sum$ and σ correspond, respectively, to the variance-covariance matrix of measurement errors in x^*_{ct} and to the covariance between measurement errors in x^*_{ct} and y^*_{ct}. They are generally not known. Deaton suggests estimating them on the individual data:

$$\hat{\Sigma} = \frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}\hat{\Sigma}_{ct} \quad (10)$$

where $\hat{\Sigma}_{ct} = \frac{1}{n-1}\sum_{i\in c,t}(x_{it} - \bar{x}_{ct})'(x_{it} - \bar{x}_{ct})$

$$\hat{\sigma} = \frac{1}{CT}\sum_{c=1}^{C}\sum_{t=1}^{T}\hat{\sigma}_{ct} \quad (11)$$

where $\hat{\sigma}_{ct} = \frac{1}{n-1}\sum_{i\in c,t}(x_{it} - \bar{x}_{ct})'(y_{it} - \bar{y}_{ct})$

Several types of convergence can be considered in the case of pseudo-panel estimations as several parameters come into play: N the number of individuals observed at each date, C the number of cohorts, n_{ct} the size of cohorts and T the number of observation dates.

Intuitively when the cohorts' size increases, the larger the cohorts, the more the intra-cohort means – that is, the estimators of the true intra-cohort means – are precise. Measurement errors become negligible and we find the standard within estimator.

The within estimator has an asymptotic bias when the size of cohorts is fixed but a lower variance than the Verbeek and Nijman estimator (for more information, see Verbeek & Nijman, 1993). This reflects again a classic bias-variance tradeoff.

Estimating dichotomous models

The previous estimators are only suitable for linear models and not when the variable of interest is binary. For this, specific estimation techniques need to be used. With panel data, switching from linear to non-linear estimation of a fixed effects model is in itself difficult. The use of pseudo-panels makes the estimation even more complex. To date, few studies have implemented the estimation methods developed for such models. Only the broad principles are given here.

The model to be estimated appears in the following form:

$$\tilde{y}_{it} = x_{it}\beta + \alpha_i + \varepsilon_{it} \quad (12)$$

$i = 1,\ldots,N \quad t = 1,\ldots,T$

where $\tilde{y}_{it}$ is a latent variable (unobserved). The value of the observed binary variable y_{it} is 1 if $\tilde{y}_{it}$ is positive and 0 otherwise. x_{it} is a vector of explanatory variables, α_i is an individual fixed effect and ε_{it} an error term which is generally assumed to follow a logistic or a normal distribution.

As in the linear case, the goal is to estimate a fixed effects model. With panel data, there are two standard estimation techniques: the conditional *logit* which consists in transforming data to eliminate the fixed effect (see, for example, Davezies, 2011) or the Chamberlain approach, (Chamberlain, 1984).

The Chamberlain approach is the starting point of the estimation method using pseudo-panel data proposed by Collado (1998). It consists in writing the relationship between the individual fixed effect and the covariates:

$$\alpha_i = x_{i1}\lambda_1 + \ldots + x_{iT}\lambda_T + \theta_i \qquad (13)$$

where $E(\theta_i \mid x_{i1},\ldots,x_{iT}) = 0$.

Substituting (13) into (12) gives the reduced form:

$$\tilde{y}_{it} = x_{i1}\pi_{t1} + \ldots + x_{iT}\pi_{tT} + \theta_i + \varepsilon_{it} \qquad (14)$$

$i = 1,\ldots,N \qquad t = 1,\ldots,T$

where $\pi_{ts} = \beta + \lambda_s$ if $s = t$ or $\pi_{ts} = \lambda_s$ otherwise. The error term $\theta_i + \varepsilon_{it}$ is not correlated with the covariates.

In the absence of panel data, the complete series of covariates is not available for a single individual. Model (14) can therefore not be directly estimated. Collado (1998) suggests estimating this model by replacing in (14) each individual value of the covariates x_{it} with the cohort mean of the individual's cohort, i.e. $\bar{x}_{ct}$. Here the cohorts are constructed following the same rules as those presented in the linear framework (see above). It should be noted that the variable of interest y_{it} is not aggregated.

Substituting individual observations with the intra-cohort means of explanatory variables introduces measurement errors into the model (the sum of the individual deviation, the intra-cohort mean and the sampling error mean) and a correlation between the error term and the covariates. Collado proposes two estimators for the β parameter. These estimators are calculated in two steps. The first, applied to both estimators, involves a quasi-maximum likelihood estimate of the π_{ts} parameters. The two proposed estimators of the β parameter are then deduced from the estimator of the π_{ts} parameters. One is calculated by minimum distance and the other by doing a within transformation on the data. The within estimator has the advantage of being easier to calculate but is not efficient, unlike the minimum distance estimator.

Moffitt (1993) proposes an alternative estimation technique, based on the parallel drawn between pseudo-panel estimation and instrumentation (see above). In the linear framework, estimating the model using the pseudo-panel method is equivalent to instrumenting using cohort-date interaction dummies. Moffitt proposes this same instrumentation to estimate model (12).

An example of pseudo-panel application: effect of age and generation on household wealth

There are many examples of pseudo-panels being used in econometric work on consumption (e.g. Gardes et al., 2005; Marical & Calvet, 2011) and in life-cycle analysis (see box). Here we propose a basic application of pseudo-panel methods to estimate age effects on household wealth. This application is highly simplified with respect to the issue of wealth accumulation and is only meant to provide a practical example of these methods. A more comprehensive analysis of this issue can be found in Lamarche and Salembier (2012).

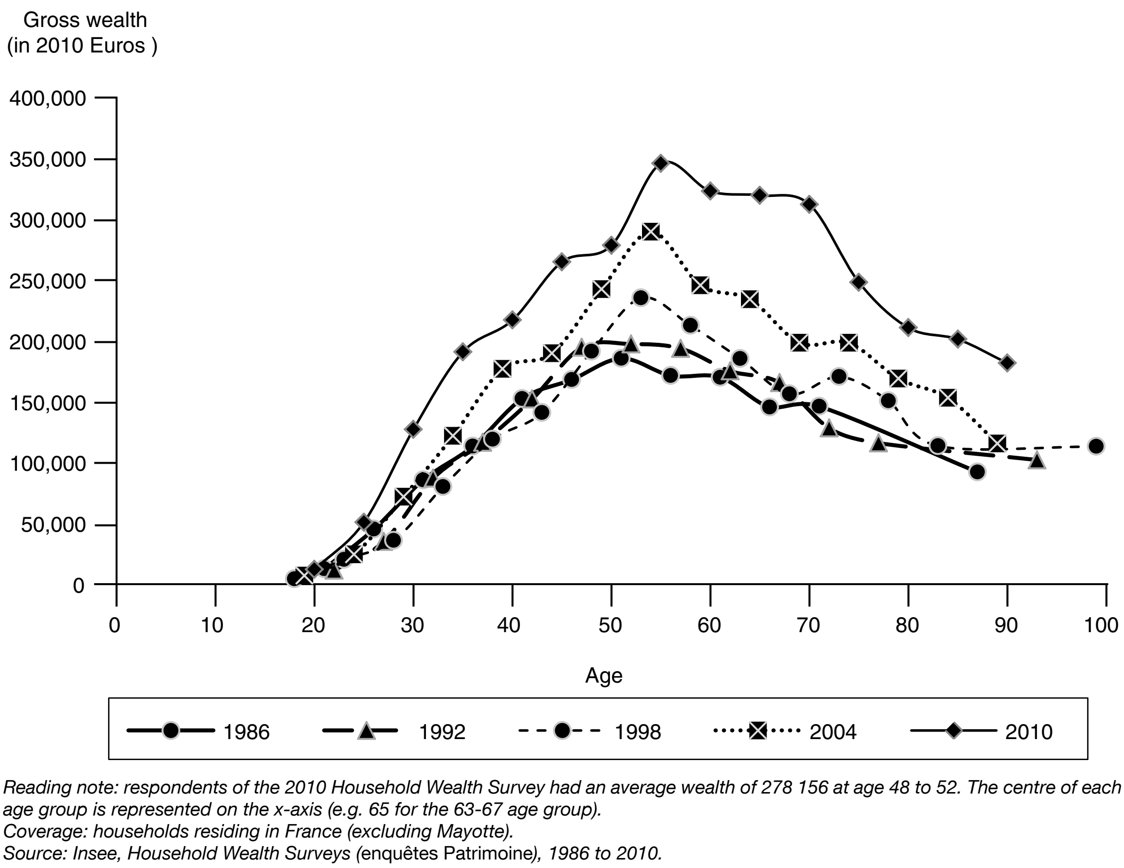
We will use the French Household Wealth surveys (*enquête Patrimoine*), conducted every six years since 1986[7], which provides five observation dates (1986, 1992, 1998, 2004 and 2010). In the survey, households are asked about their real estate, financial and professional assets. The sum of these assets provide the gross wealth (calculated in constant 2010 Euros). In 2010, the survey underwent major changes to better assess households' wealth. In particular, the categories of households with the highest wealth were oversampled and assets such as cars, household equipment, jewellery, and artwork were taken into account. To avoid biasing the changes between 2004 and 2010, these methodological changes were for the most part neutralised in the wealth calculations.

7. In 1986 and 1992, the name of the French Household Wealth survey was enquête Actifs financiers.

To briefly describe the issue, the aim is to study saving patterns at different ages. In the initial version put forward by Modigliani and Brumberg (1954), the life-cycle theory assumes that people adopt an intertemporal approach to allocating their income. Over their lifetimes, they experience three periods during which their earnings, and their savings and consumption behaviours differ. At the beginning of their career their income tends to be low and they spend more than they earn (dissaving). Then, throughout their career, their income increases, they save and accumulate wealth as they prepare for their income to drop when they retire. Wealth accumulation therefore follows a bell curve pattern with age. It is difficult to test the life-cycle theory, by estimating for instance changes in wealth with age. This type of estimation would require the same individuals to be followed over a very long period, which is quite impossible. As stated earlier, a cross-sectional estimation would not be relevant since it does not allow the distinction to be made between the effects of age and generation. With this very simple case of estimating the effect of age, the next two graphs can be used as a starting point for a typical exploratory approach. Each Household Wealth survey is used to represent the change in mean gross wealth according to age (Figure I). The profiles obtained seem to confirm the life-cycle theory beyond a doubt. A bell curve is obvious with an increase in gross wealth until about 60 years of age, followed by a drop. However, part of this profile can be explained by the fact that different generations are observed at each date. Economic context, the age at which people begin working, and taxes are all characteristics shared by the individuals of a same generation that have an effect on accumulated household wealth. They also explain differences in wealth at the same age between different generations. Long term data are required to separate these two effects.

To attempt to capture this "generation" dimension, all the surveys are stacked so as to obtain observations for individuals from identical generations at different dates (and therefore different ages). We obtain five observations,

Figure I
Household wealth according to age in 1986, 1992, 1998, 2004 and 2010



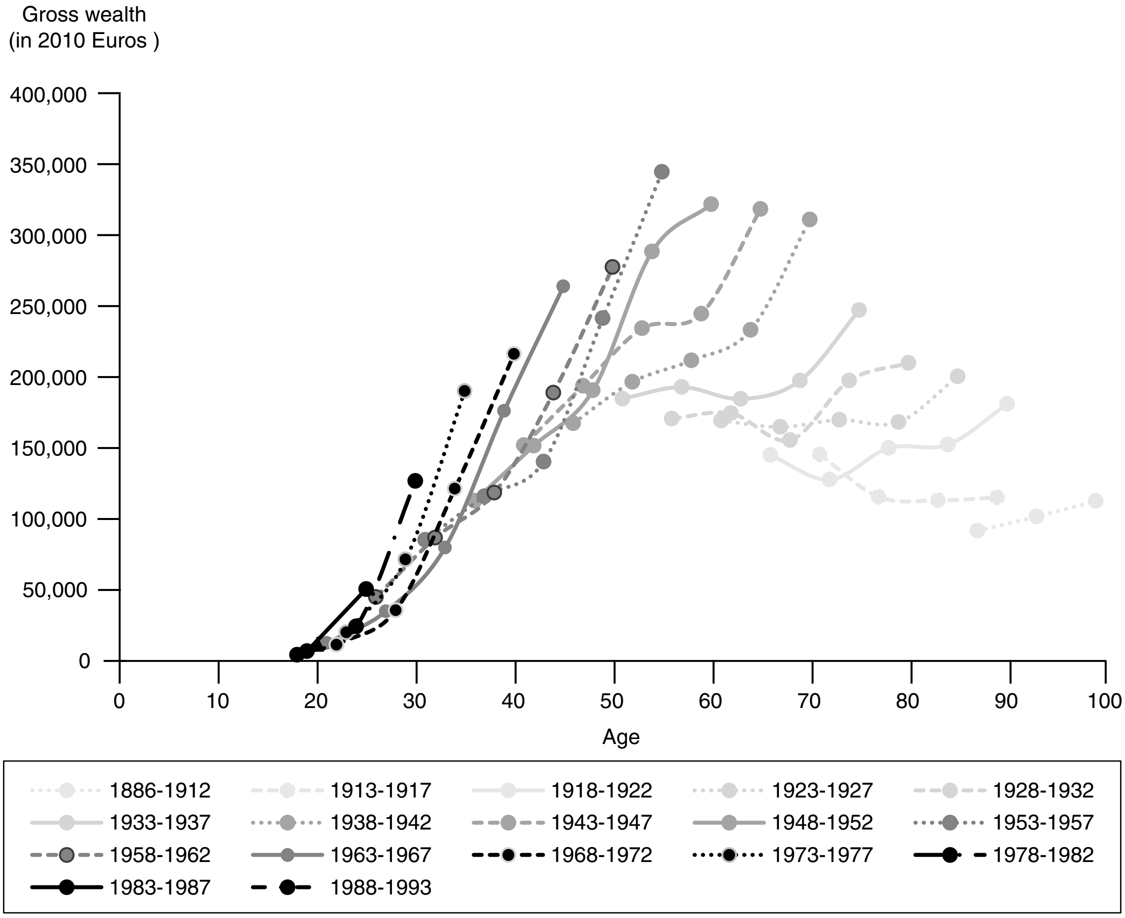
Reading note: respondents of the 2010 Household Wealth Survey had an average wealth of 278 156 at age 48 to 52. The centre of each age group is represented on the x-axis (e.g. 65 for the 63-67 age group).
Coverage: households residing in France (excluding Mayotte).
*Source: Insee, Household Wealth Surveys (*enquêtes Patrimoine*), 1986 to 2010.*

corresponding to the average wealth at five different ages for almost all the generations (except for the youngest or oldest). In theory, one profile for all the generations, defined by year of birth, could be represented. However in practice, we are confronted with the problem that in a survey sample, the number of individuals from a given generation is not very high. These estimations are therefore very imprecise. To offset this problem, we define cohorts as the grouping of adjacent generations (five in Figure II).

Figure II shows, for each cohort, the profile of wealth accumulation by age. It is very different from the profile presented using only the cross-sectional dimension. Contrary to what Figure I suggests, wealth continues to grow well over the age of 60. As underlined by Lamarche and Salembier (2012), several factors explain this stylised fact. Even beyond retirement, households may want to save in order to leave an inheritance or simply build up contingency savings (should they become dependent). Furthermore, the most elderly may decide not to sell their real estate assets to avoid moving and the particularly high cost that this entails (see Angelini & Laferrère, 2012). It should also be highlighted that the accumulation of wealth with age partially results from changes in generation composition observed at extreme ages. The scope of the survey only examines private households and therefore does not include elderly people in retirement homes. Wealthier households also have a longer life expectancy than others (and likely more assets).

Figure II compares the average wealth of different cohorts at the same age. There are

Figure II
Household wealth according to age from one generation to another



Reading note: respondents of the 2010 Household Wealth Survey had an average wealth of 278 156 at age 48 to 52. The centre of each age group is represented on the x-axis (e.g. 65 for the 63-67 age group).
Coverage: households residing in France (excluding Mayotte).
*Source: Insee, Household Wealth Surveys (*enquêtes Patrimoine*), 1986 to 2010.*

sometimes significant differences. The vertical deviation between the curves corresponds to the generation effect and a period effect. For example, let us assume that these period effects, which correspond to the increase in household wealth over time (again, we are working in constant 2010 Euros to avoid including inflation) are negligible. This resolves the problem of identifying age, cohort and period effects (see Box). Under this assumption, the graph suggests that, at the same age, each generation has accumulated more wealth than the previous. The difference is considerable between generations born in the 1950s who experienced the post-World War II economic boom (1945-1975) and previous wartime generations. The decrease in wealth after the age of 60 observed in Figure I likely stems more from significant differences in wealth between these two generations than dissaving at retirement.

Pseudo-panel econometric modelling provides a more accurate quantification of the age effects seen in Figure II. It is based on a model written out on an individual basis, as follows:

$$\log Pat_{it} = \beta_1 age_{it} + \beta_2 age_{it}^2 + \alpha_i + \varepsilon_{it} \qquad (15)$$
$$i = 1,\ldots,N \quad t = 1,\ldots,T$$

log Pat_{it} is the logarithm for the wealth of the individual i on date t, age_{it} is their age on date t. Let us assume here that the effect of age on wealth is identical for all generations and that it has a quadratic profile[8]. α_i is an individual fixed effect. It estimates the impact of unobserved fixed characteristics of individual i on his/her wealth.

The pseudo-panel model that is estimated in practice is as follows:

$$\left(\log Pat\right)_{gt} = \beta_1 age_{gt} + \beta_2 age_{gt}^2 + \alpha_g + \varepsilon_{gt} \qquad (16)$$
$$g = 1,\ldots,G \quad t = 1,\ldots,T$$

where for each variable z, $z_{gt} = E(z_{it} \mid i \in g,t)$. These values are not observed. They are estimated by the intra-cohort means $\overline{z}_{gt} = \frac{1}{n_{gt}} \sum_{i \in g,t} z_{it}$ calculated from available data, where n_{gt} is the number of individuals of cohort g observed on date t.

Two practical remarks need to be made. The first concerns the composition of the sample. The estimation relies on the fact that $\overline{\alpha}_{gt}$ is fixed over time. This can be called into question. As mentioned above, for the oldest generations, two composition effects come into play. First, the wealthiest households have a longer average life expectancy, and secondly,

8. The accumulation of wealth with age between different generations only differs in level. The model could be made more complex by integrating interaction terms between age and generation.

Box

AGE, COHORT AND PERIOD EFFECTS

Simultaneously estimating an age, cohort and period effect is a recurring problem that already existed before pseudo-panels, but which is raised in the same way for individual data and pseudo-panel data. The difficulty stems from the collinearity between the three variables (age + cohort = period), i.e. from the fact that individuals of the same age and the same generation cannot be observed at different dates.

It is generally resolved by treating age, cohort and period effects as additive. The model therefore simply includes a set of age, cohort and period dummies without interaction terms. This additivity assumption is significant. It leads us to assume that the age effect, for instance, is common to all generations. In the case of this model, the literature proposes two primary solutions for resolving the identification problem. The first involves imposing identifying constraints on the model (in addition to the nullity of a coefficient for each dimension and an identifying constraint in the presence of a constant in the model). Mason et al. (1973) show that we can simply assume that two coefficients from a single dimension (age, cohort, or period) are equal. Different identifying constraints lead to different estimations and must be discussed on a case by case basis. Rodgers (1982) disagrees with this practice and proposes replacing one of the effects with variables that correlate with it, for example macro-economic variables in the place of the period effect. Readers interested in this issue may refer to Hall et al. (2007) for a literature review on the subject, or Yang and Land (2013).

the Household Wealth survey does not survey people in retirement homes. On the other end, the Household Wealth survey only includes a small number of very young households, which are probably very specific. To work on a stable population, we limit ourselves to households over the age of 26 and under 80[9]. The second remark concerns the size of cohorts. Cohorts group together several successive generations. Limiting the number of these successive generations reduces the risk of aggregating heterogeneous behaviours. However this means that estimations are based on very few observations per cohort and therefore risk being very imprecise. To illustrate this issue, the model was estimated using relatively broad cohorts (three, five and ten years) (Table C1 in the Appendix).

The table below shows the results of pseudo-panel estimations. For comparative purposes, the results obtained from cross-sectional regression (the data from the five successive surveys are stacked) and the estimations taking measurement errors into account are also presented.

Figure III shows the effect of age on wealth as estimated using both the cross-sectional and pseudo-panel approaches[10]. The two estimations show a bell curve relationship between wealth and age. From cross-sectional data, we estimate that wealth begins to decrease at age 58. The pseudo-panel estimation gives a much higher turning point, around age 70. So when the generation effect is taken into account, the decrease in wealth is observed much later than a cross-section approach suggests.

As the model is log-linear, 100 x [exp(α_g) - 1], where α_g is the coefficient associated with the generation g in the model (Table C2 in the Appendix and Figure IV below), corresponds to the effect on wealth (measured in %) of belonging to generation g rather than to the

9. Furthermore, as means are sensitive to extreme values, some very high net worth households were removed from the analysis. The few observations corresponding to zero net worth were also removed since logarithm modelling is used.
10. The polynomial of degree 2 is therefore represented: $\beta_0 + \beta_1 age + \beta_2 age^2$ where the coefficients are estimated from cross-sectional data and pseudo-panel data.

Table
Estimation of age effects

		Pseudo-panel Estimations		
	Cross-sectional data	3-year generations	5-year generations	10-year generations
		Within estimator		
Intercept	4.59***	4.80***	4.65***	4.89***
	(0.127)	(0.383)	(0.437)	(0.542)
Age	0.223***	0.197***	0.199***	0.193***
	(0.0052)	(0.0142)	(0.016)	(0.0212)
Age^2	- 0.0019***	- 0.00140***	- 0.00136***	- 0.00136***
	(0.0000493)	(0.000135)	(0.000145)	(0.0002)
		Measurement error model		
		Verbeek and Nijman estimator (9)		
Intercept		4.63***	5.05***	5.63***
		(0.279)	(0.307)	(0.398)
Age		0.203***	0.187***	0.162***
		(0.0104)	(0.0127)	(0.0172)
Age^2		- 0.00143***	- 0.00128***	- 0.00102***
		(0.000092)	(0.00012)	(0.00016)
Number of observations	43 117	94	57	31

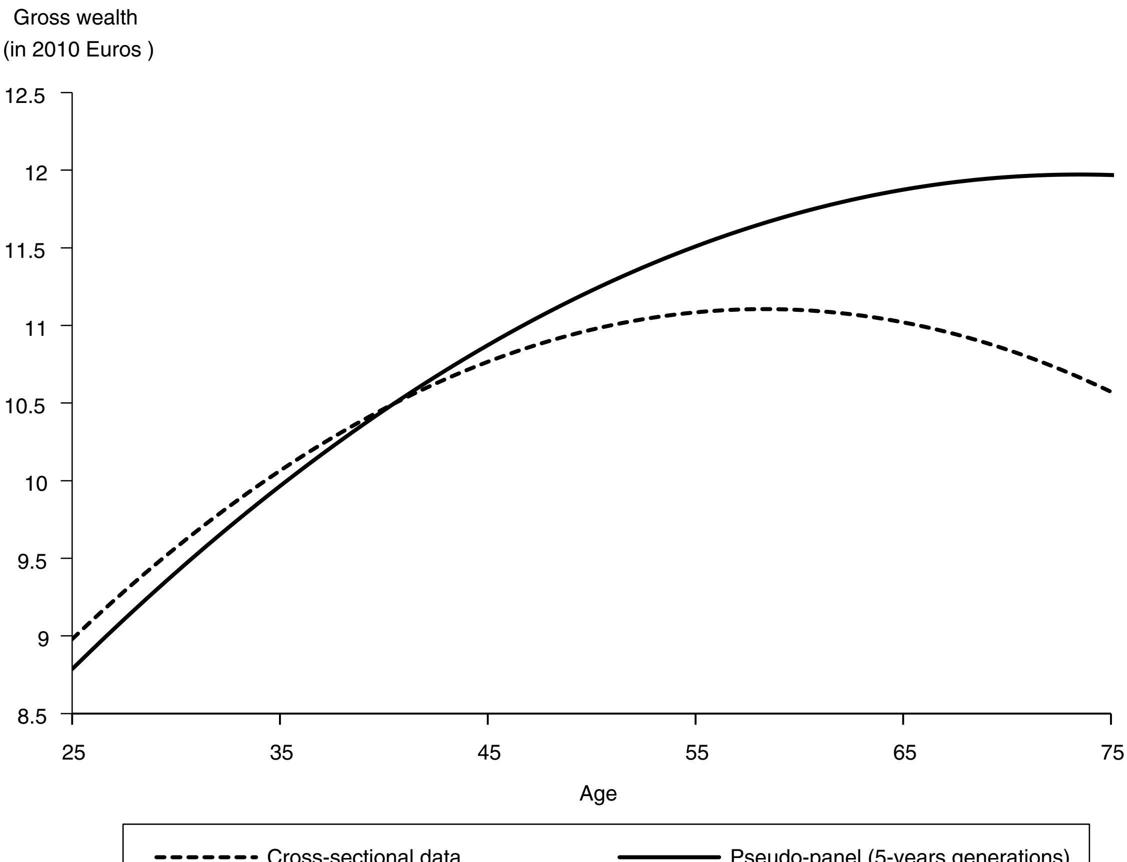
Note: the constant is calculated using the birth years 1951-1953 as the baseline generation for 3-year generations, 1953-1957 for 5-year generations and 1953-1962 for 10-year generations. Standard deviations were calculated by bootstrapping for the measurement error model.
****. **. * indicate the significance level of the coefficients at 1%, 5% and 10% respectively. The number of individuals observed in the different generations is presented in Table C2 in the Appendix.*
Coverage: households residing in France (excluding Mayotte).
*Source: estimation based on the French Household Wealth Surveys (*enquêtes Patrimoine*).*

1951-1953 generation (generation of reference). For example, being born between 1939 and 1941 rather than between 1951 and 1953 has a negative effect on household wealth, estimated at 100 x [exp(– 0.44) – 1] = – 35.6%. We estimate that between the 1939-1941 and 1951-1953 generations, household wealth increased on average by 3.7% annually. Its growth then slowed down.

The sensitivity of estimations to the cohort grouping criteria does not seem too high in this case. Figure IV shows the generation effects estimated using the pseudo-panel methods based on three ranges chosen to construct the generations. Unsurprisingly, the greater the range, the smoother the profile. In all cases, we observe a significant increase in the wealth of successive generations until the baby-boom generations, followed by stagnation. For the youngest generations, the diagnosis seems to diverge depending on the selection criteria, but these changes are never significant (see Table C2 in the Appendix). This uncertainty stems from the fact that estimations are based on smaller samples (these generations are not observed in the older surveys), as shown in Table C1 (Appendix C). It can also be seen that, as expected, the precision of the estimators of coefficients β_1 and β_2 is greater for three-year generations than for five or ten-year generations.

The Verbeek and Nijman estimator, which takes into account measurement errors, was also calculated directly with the estimator formula. As the estimator formula suggests, in theory, the direction of bias is not known and changes depending on the range adopted to define the generations. The estimations differ little from those obtained with the within estimator, except for the 10-year generations. ☐

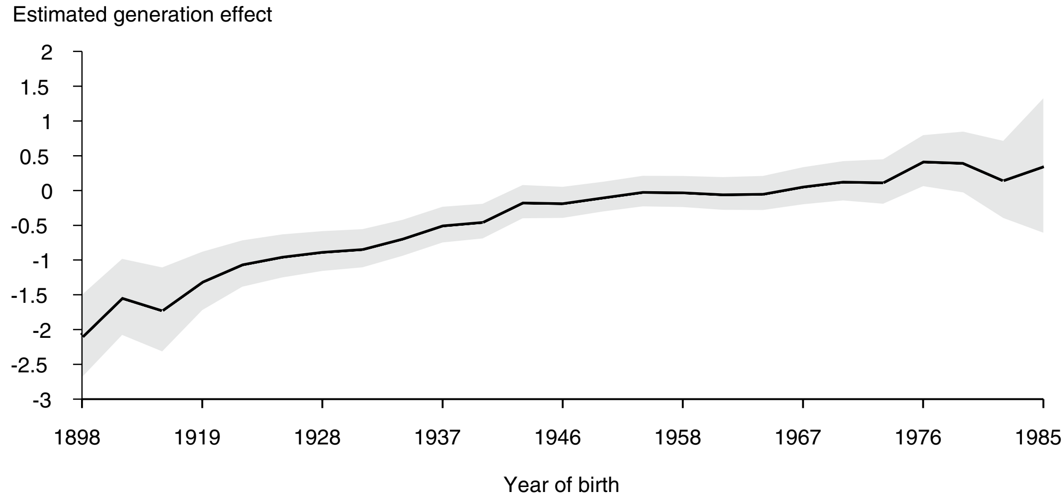
Figure III
Household wealth according to age as estimated by models



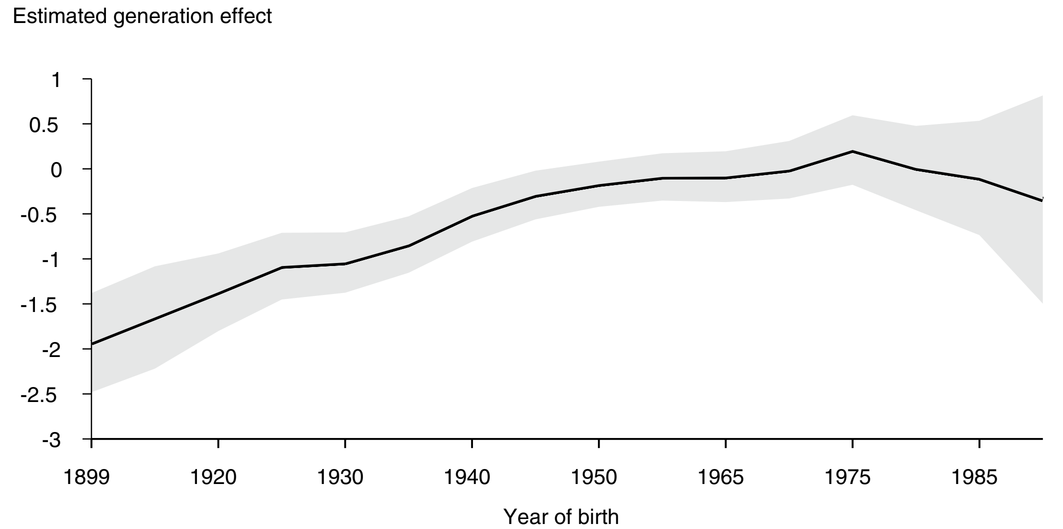
Reading note: at 65, the gross wealth logarithm as estimated by the pseudo-panel model is 11.87.
Coverage: households residing in France (excluding Mayotte).
*Source: estimation based on the French Household Wealth Surveys (*enquêtes Patrimoine*).*

Figure IV
Generation effects estimated using the pseudo-panel method

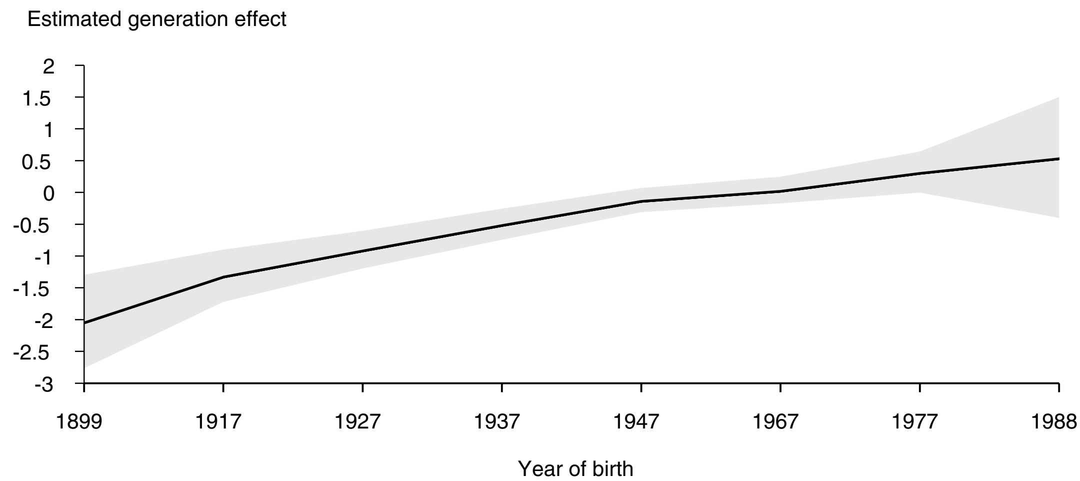
(a) 3-years generations



(b) 5-years generations



(c) 10-years generations



Reading note: the generation effect estimated by the pseudo-panel model (3-year generation) for the 1939-1941 generation is – 0.44, which corresponds to 35.6% lower gross wealth than the baseline generation (1951-1953). The grey area corresponds to a 5% confidence interval.
Coverage: households residing in France (excluding Mayotte).
*Source: estimation based on the French Household Wealth Surveys (*enquêtes Patrimoine*).*

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APPENDIX

A. PSEUDO-PANEL AND INSTRUMENTATION

Moffitt (1993) shows that estimation using the pseudo-panel approach and estimation by instrumenting using cohorts-date interaction dummies provide the same estimator.

Estimation via two-stage least squares follows the following two steps:

Step 1: Projection of explanatory variables onto the instrument.

If the individual fixed effect α_i is written as the sum of a fixed effect cohort α_c and an individual deviation $v_i = \alpha_i - \alpha_c$, model (1) would be as follows:

$$y_{it} = x_{it}\beta + \alpha_c + v_i + \varepsilon_{it} \tag{17}$$

x_{it} is potentially correlated to v_i. Therefore, x_{it} is instrumented using cohort indicators in interaction with the time indicators. The first step is to project x_{it} onto the instrument. The predicted value of x_{it} in this regression corresponds to the intra-cohort mean $\bar{x}_{ct}$.

Step 2:

x_{it} is replaced by its predicted value in (17). y_{it} is therefore regressed on $\bar{x}_{ct}$ and the cohort indicators, which gives the same estimator as the within estimator (4).

B. DETAILS ON THE ESTIMATION OF THE PARAMETERS OF A MEASUREMENT ERROR MODEL

$\bar{x}_{ct}$ and $\bar{y}_{ct}$ are observations with errors of the true intra-cohort means x^*_{ct} and y^*_{ct}. u_{ct} and v_{ct} are the measurement errors:

$$\bar{y}_{ct} = y^*_{ct} + u_{ct} \tag{18}$$

$$\bar{x}_{ct} = x^*_{ct} + v_{ct} \tag{19}$$

They are assumed to be normally distributed:

$$\begin{pmatrix} u_{ct} \\ v'_{ct} \end{pmatrix} \sim N\left(\begin{pmatrix} 0 \\ 0 \end{pmatrix}; \frac{1}{n}\begin{pmatrix} \sigma_{00} & \sigma' \\ \sigma & \Sigma \end{pmatrix}\right) \tag{20}$$

where n is the size of the cohorts.

Integrating (18) and (19) into model (2) gives:

$$\bar{y}_{ct} = \bar{x}_{ct}\beta + \alpha_c + \tilde{\varepsilon}_{ct} \qquad c = 1,\ldots,C \qquad t = 1,\ldots,T \tag{21}$$

where $\tilde{\varepsilon}_{ct} = \varepsilon^*_{ct} + u_{ct} - v_{ct}\beta$.

The correlation between this residual value and the covariates gives:

$$E\left(\bar{x}'_{ct}\tilde{\varepsilon}_{ct}\right) = \frac{1}{n}(\sigma - \Sigma\beta)$$

In general it is not zero. The estimator of the least squares of $\bar{y}_{ct}$ on $\bar{x}_{ct}$ is therefore biased.

Model (21) is a fixed effects model. After a within transformation, model (21) becomes:

$$\bar{y}_{ct} - \bar{y}_c = \left(\bar{x}_{ct} - \bar{x}_c\right)\beta + \tilde{\varepsilon}_{ct} - \bar{\tilde{\varepsilon}}_c \quad \text{where } \bar{\tilde{\varepsilon}}_c = \frac{1}{T}\sum_{t=1}^{T}\tilde{\varepsilon}_{ct} \tag{22}$$

We show that:

$$E\left(\bar{x}_{ct} - \bar{x}_c\right)'\left(\bar{y}_{ct} - \bar{y}_c\right) = E\left(\bar{x}_{ct} - \bar{x}_c\right)'\left(\bar{x}_{ct} - \bar{x}_c\right)\beta + \frac{T-1}{T} \times \frac{1}{n}(\sigma - \Sigma\beta)$$

From this equation, an expression of β is deduced:

$$\beta = \left[E\left(\bar{x}_{ct} - \bar{x}_c\right)'\left(\bar{x}_{ct} - \bar{x}_c\right) - \frac{T-1}{T} \times \frac{1}{n}\Sigma\right]^{-1} \left[E\left(\bar{x}_{ct} - \bar{x}_c\right)'\left(\bar{y}_{ct} - \bar{y}_c\right) - \frac{T-1}{T} \times \frac{1}{n}\sigma\right]$$

Estimator (9) is the empirical counterpart of this expression.

When only the explained variable is observed with error, the within estimator is without bias but it is less precise than a model without measurement errors. When the measurement error only relates to the explanatory variables, it leads to an attenuation bias (the absolute value of the within estimator converges towards a lower value than the absolute value of parameter β).

C. APPLICATION OF PSEUDO-PANELS TO THE FRENCH HOUSEHOLD WEALTH SURVEY (*ENQUÊTE PATRIMOINE*)

Table C1
Cohorts' size

3-year generations

Generation (year of birth)	Year				
	1986	1992	1998	2004	2010
1886-1911	267				
1912-1914	191	124			
1915-1917	109	132			
1918-1920	179	268	153		
1921-1923	321	431	375		
1924-1926	278	502	397	228	
1927-1929	305	544	440	421	
1930-1932	301	498	469	444	336
1933-1935	282	522	512	468	555
1936-1938	287	426	456	413	593
1939-1941	284	430	488	445	569
1942-1944	317	481	502	392	704
1945-1947	372	614	654	467	804
1948-1950	408	727	728	562	894
1951-1953	391	683	680	570	838
1954-1956	373	731	626	554	756
1957-1959	292	704	652	560	774
1960-1962	77	569	582	544	723
1963-1965		407	582	552	743
1966-1968		124	465	506	654
1969-1971			463	511	599
1972-1974			132	426	541
1975-1977				367	414
1978-1980				112	396
1981-1983					290
1984-1986					85

Reading note: in the 1986 French Household Wealth Survey, 373 individuals born between 1954 and 1956 were surveyed.
Coverage: households residing in France (excluding Mayotte).
*Source: Insee, French Household Wealth surveys (*enquêtes Patrimoine*).*

5-year generations

Generation (year of birth)	Year				
	1986	1992	1998	2004	2010
1886-1912	344				
1913-1917	223	256			
1918-1922	391	551	395		
1923-1927	477	831	672	359	
1928-1932	516	861	767	734	336
1933-1937	476	787	804	744	964
1938-1942	478	742	815	707	954
1943-1947	588	944	993	734	1307
1948-1952	678	1181	1192	938	1457
1953-1957	615	1213	1068	964	1295
1958-1962	248	1020	1008	888	1233
1963-1967		531	915	877	1209
1968-1972			727	842	1000
1973-1977				643	742
1978-1982				112	598
1983-1987					173

Reading note: in the 1986 French Household Wealth Survey, 615 individuals born between 1953 and 1957 were surveyed.
Coverage: households residing in France (excluding Mayotte).
*Source: Insee, French Household Wealth surveys (*enquêtes Patrimoine*).*

10-year generations

Generation (year of birth)	Year				
	1986	1992	1998	2004	2010
1886-1912	344				
1913-1922	614	807	395		
1923-1932	993	1692	1439	1093	336
1933-1942	954	1529	1619	1451	1918
1943-1952	1266	2125	2185	1672	2764
1953-1962	863	2233	2076	1852	2528
1963-1972		531	1642	1719	2209
1973-1982				755	1340
1983-1993					173

Reading note: in the 1986 French Household Wealth Survey, 863 individuals born between 1953 and 1962 were surveyed.
Coverage: households residing in France (excluding Mayotte).
*Source: Insee, French Household Wealth surveys (*enquêtes Patrimoine*).*

Table C2
Estimated generation effects

3-year generations	
1886-1911	- 2.09***
	(0.302)
1912- 1914	- 1.53***
	(0.279)
1915-1917	- 1.71***
	(0.308)
1918-1920	- 1.30***
	(0.214)
1921-1923	- 1.05***
	(0.170)
1924-1926	- 0.94***
	(0.158)
1927-1929	- 0.87***
	(0.146)
1930-1932	- 0.83***
	(0.140)
1933-1935	- 0.68***
	(0.132)
1936-1938	- 0.49***
	(0.131)
1939-1941	- 0.44***
	(0.127)
1942-1944	- 0.16
	(0.122)
1945-1947	- 0.17
	(0.114)
1948-1950	- 0.088
	(0.109)
1951-1953	*ref.*
1954-1956	- 0.0078
	(0.112)
1957-1959	- 0.014
	(0.114)
1960-1962	- 0.042
	(0.120)
1963-1965	- 0.035
	(0.125)
1966-1968	0.069
	(0.136)
1969-1971	0.14
	(0.144)
1972-1974	0.13
	(0.163)
1975-1977	0.43
	(0.187)
1978-1980	0.41
	(0.223)
1981-1983	0.16
	(0.283)
1984-1986	0.36
	(0.493)

5-year generations	
1886-1912	- 1.93***
	(0.281)
1913-1917	- 1.65***
	(0.290)
1918-1922	- 1.37***
	(0.220)
1923-1927	- 1.08***
	(0.189)
1928-1932	- 1.04***
	(0.171)
1933-1937	- 0.84***
	(0.160)
1938-1942	- 0.51***
	(0.152)
1943-1947	- 0.29**
	(0.138)
1948-1952	- 0.17
	(0.128)
1953-1957	*ref.*
1958-1962	- 0.089
	(0.134)
1963-1967	- 0.0086
	(0.144)
1968-1972	- 0.0089
	(0.163)
1973-1977	0.21
	(0.197)
1978- 1982	0.0098
	(0.239)
1983-1987	- 0.10
	(0.324)
1988-1993	- 0.34
	(0.590)

10-year generations	
1886-1912	- 2.03***
	(0.375)
1913-1922	- 1.31***
	(0.210)
1923-1932	- 0.90***
	(0.151)
1933-1942	- 0.50***
	(0.125)
1943-1952	- 0.12
	(0.097)
1953-1962	*ref.*
1963-1972	0.038
	(0.107)
1973-1982	0.32*
	(0.165)
1983-1993	0.55
	(0.485)

*Reading note: The estimated coefficient of the 1939-1941 generation in the model is – 0.44, which means that being born between 1939 and 1941 rather than between 1951 and 1953 (reference generation) has a negative effect on wealth, estimated at 100 x [exp(– 0.44) – 1] = – 35.6 %. ***, **, * indicate a significance level of the coefficients at 1%, 5% and 10% respectively.*
Coverage: households residing in France (excluding Mayotte).
*Source: Insee, French Household Wealth surveys (*enquêtes Patrimoine*).*