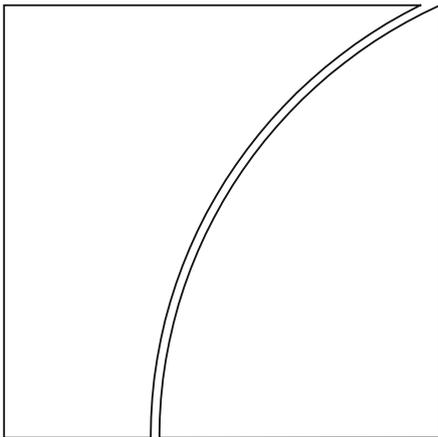




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## The enduring link between demography and inflation

by Mikael Juselius and Előd Takáts

Monetary and Economic Department

May 2018

JEL classification: demography, ageing, inflation,  
monetary policy

Keywords: E31, E52, J11

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# The enduring link between demography and inflation<sup>\*</sup>

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## Abstract

*Demographic shifts, such as population ageing, have been suggested as possible explanations for the past decade's low inflation. We exploit cross-country variation in a long panel to identify age structure effects in inflation, controlling for standard monetary factors. A robust relationship emerges that accords with the lifecycle hypothesis. That is, inflationary pressure rises when the share of dependants increases and, conversely, subsides when the share of working age population increases. This relationship accounts for the bulk of trend inflation, for instance, about 7 percentage points of US disinflation since the 1980s. It predicts rising inflation over the coming decades.*

Keywords: demography, ageing, inflation, monetary policy.

JEL classification: E31, E52, J11.

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## 1. Introduction

Our understanding of inflation is challenged by the recent experience of stubbornly low and unresponsive inflation rates in advanced countries (Draghi (2016) and Yellen (2017)). It is becoming increasingly difficult, for instance, to attribute this phenomenon solely to normal business cycle fluctuations. This opens up the possibility that other, slower-moving, forces also play a role (Faust and Leeper (2015)). A prominent line of argument in this respect is that current shifts in the population age structure can drive down inflation (eg Bullard et al (2012) and Summers (2014a,b)). If so, the implications could be significant, as demographic trends tend to be long-lasting. Indeed, nascent empirical work has provided some support for links between the age structure and inflation.<sup>1</sup> But this literature has not yet convincingly isolated possible age structure effects in inflation from those of more conventional monetary trends.

We use long panel data to overcome the challenge of identifying possible age structure effects in inflation. In addition to inflation and the age structure, our data include several monetary and real variables from 22 advanced economies from 1870 to 2016. The long time span covers several country-specific demographic cycles, allowing us to isolate potential age structure effects from other secular trends that may also account for long swings in inflation, such as changes in monetary policy frameworks. This contrasts with previous work relying on post-war samples, where demographic cycles are similar across countries, mostly due to the post-war baby boom, which makes it more difficult to isolate demographic effects.

We find a systematic relationship between the age structure and inflation: an increase in the share of the dependant population is generally associated with higher inflation, whereas an increase in the working age population has the opposite effect. However, the old population has an ambiguous effect. The effect is positive for most old cohorts except for the last open-ended age cohort (80+ year olds), for which it turns sharply negative. Interestingly, it is this cohort that is most strongly affected by longevity. The age structure effects are also economically meaningful, capturing the larger part of trend inflation at both the country-specific and global level. For example, they account for around a 7 percentage point increase in inflation from the 1950s to the mid-1970s in the United States, and a similarly sized decline thereafter.

The relationship between inflation and the age structure is robust. It is present in pre-war (1870–1913), interwar (1922–1938) and post-war (1950–1989) samples, and it extends into the most recent years (1990–2016) as well. It does not materially change when we control for economic variables, such as the output gap, real interest rates, money aggregates, government debt, and various other factors that may drive the saving-investment equilibrium. Instead, the age structure appears largely complementary to these factors. Similarly, the relationship does not depend on the particular estimation technique. For instance, it remains essentially the same whether or not we use static or dynamic models, or if we include or omit time effects, or use crude age-cohort shares or sophisticated population polynomials (eg Fair and Dominguez (1991)), or assume panel homogeneity or allow for full heterogeneity (eg Pesaran et al (1999)). The result also persists when we consider only five-year non-overlapping averages.

Our paper builds on recent empirical work. Focusing exclusively on ageing, Anderson et al (2014), Yoon et al (2014) and Bobeica et al (2017) find significant deflationary effects from an increasing share of older population cohorts. Juselius and Takáts (2015) and Aksoy et al (2015) take the age structure more fully into account and find that an increase in the number of dependants, young and old, is generally inflationary. Juselius and Takáts (2015) also show

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<sup>1</sup> Yoon et al (2014), Anderson et al (2014), Juselius and Takats (2015), Aksoy et al (2015), Goodhart et al (2015), and Bobeica et al (2017).

that the deflationary effects of ageing found in previous studies are driven primarily by the very old (80+ year old) cohort. A common feature of these studies, as noted above, is that they rely exclusively on post-war data, which makes it difficult to separate the age structure effect in inflation from other global secular factors that may be related to trend inflation.

The uncovered link is policy-relevant, because global ageing will substantially increase the share of the old-age population in almost all countries (eg Goodhart et al (2015)). Increased longevity and stagnant or declining birth rates will affect both advanced and emerging economies. While slow, such large-scale demographic shifts have the potential to materially affect trend inflation. For instance, we find that accounting for the age structure leads to substantially lower estimates of endogenous inflation persistence. Hence, past historical periods of high inflation persistence might have reflected, in part, persistent demographic changes. This implies that the role of conventional endogenous drivers, such as inflation expectations, may have been overstated. If so, this could account for the current conundrum with well-anchored long-term inflation expectations and persistently low inflation rates. The stability of the relationship furthermore suggests that this may help us forecast longer-term inflation trends, as previously noted by McMillan and Baesel (1990) and Lindh and Malmberg (2000). Our estimates indicate that inflationary pressures are likely to rise in future due to the increasing share of older population cohorts and a declining share of younger ones, which has not been emphasized in the literature.

While there is no standard theoretical explanation for how demography could affect trend inflation, the literature discusses at least two potential channels. Importantly, these explanations do not directly conflict with Friedman's (1963) dictum that "inflation is always and everywhere a monetary phenomenon", but rather suggest only how age structure could cause inflation within a given monetary framework. The first channel works through *the natural rate*, ie the real equilibrium interest rate. An increase in the share of the dependant population (ie the young and the old), lowers the savings rate and therefore drives up the natural rate, whereas increasing longevity has the opposite effect.<sup>2</sup> Such changes in the natural rate can lead to trends in inflation if monetary policy becomes constrained by the zero lower bound (eg Summers (2014a,b) and Eichengreen (2015)) or, more broadly, does not fully internalise them, for example, due to informational frictions (Gust et al (2015)). An alternative channel could work through *the political economy*, ie the old and the young might prefer different levels of inflation, which could drive central bank policies in turn (Bullard et al (2012)).<sup>3</sup> For instance, the young are often borrowers and therefore prefer inflation, whereas the opposite holds for the old.

Taken together, the signs of the estimated age cohort effects that we find accord with a life cycle explanation of the natural real interest rate. We find, for instance, that a rise in the dependency ratio, which should increase the natural rate, is inflationary. Moreover, increased longevity should have the opposite effect and therefore be deflationary. In line with this, we find that the very old (80+ year old) cohort, where such an increase would be most visible, has a negative effect on inflation. Yet, while much of the evidence points to a life cycle explanation, we also find that the age structure effect survives on its own without any reference to actual real interest rates. This is not fully consistent with the view that the age structure effect works mainly through movements in the natural rate. Hence, a more elaborate account of how life cycle behaviour can generate inflationary pressure may be needed to fully explain the puzzlingly strong link between inflation and the age structure that we uncover.

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<sup>2</sup> See, for instance, Carvalho et al (2016) and Eggertsson et al (2017), Lisacks et al (2017), and Rachel and Smith (2015)).

<sup>3</sup> In addition, monetary policy might affect inequality (for a recent treatment refer to Coibin et al (2017)), which can further affect low frequency inflation through political economic channels.

The rest of the paper is organised as follows. The next section presents the data and the main empirical results. The third section assesses the robustness of the estimates. The fourth section discusses the economic impact and its implications. The final section concludes.

## 2. Is there a link between demography and inflation?

A nascent empirical literature has started to investigate the potential link between demography and inflation. This line of inquiry has been motivated partly by some similarities between the global financial crisis and the Japanese crisis in the early 1990s: both crises occurred at a time when the dependency ratio bottomed and both were followed by low inflation as the share of old cohorts started to increase. Spurred by this similarity, several studies have focused on ageing as a possible agent of low inflation. For example, Anderson et al (2014), using the IMF's GIMF model, find that ageing may reduce inflation. Yoon et al (2014) find that an increasing share of old cohorts in the population (65+) is associated with lower inflation in data from 30 OECD economies between 1960 and 2013.

The disinflationary effect of ageing is not, however, a robust feature of the data: it changes signs when a finer division of the age cohorts is used or when the entire age structure is taken into account. This is demonstrated by Juselius and Takáts (2015), who model the entire age structure using a population polynomial in a panel of 22 countries from 1950 to 2013. They find that the young and the old are inflationary, while the working-age cohort is disinflationary. Aksoy et al (2015) and Goodhart (2015) also document similar effects in post-war panel data, using three age cohorts (young, working-age and old). Hence, omitting certain parts of the age structure or using too crude age cohorts can severely bias the results. Interestingly, these results would suggest that ageing alone is unlikely to fully explain the post-crisis low inflation.

Even though these early results are indicative of substantial age structure effects in inflation, the evidence is still far from conclusive. The challenge is to distinguish between the demographic effect and other factors that may have generated persistent low-frequency movements in inflation, such as oil price shocks and changes to monetary policy regimes. This identification is hard to make based on post-war data, because the time period is relatively short. Hence, the data contain at most one demographic cycle in each country. Furthermore, the demographic cycles across countries have been largely synchronous due to the baby boom in 1950s and 1960s and the subsequent baby bust. This implies that there is relatively little cross-country variation with respect to the age structure in post-war samples that can be used for identification. Put differently, there is a risk of misinterpreting a temporary correlation between global trends in inflation and the age structure as a meaningful relationship.

In this paper, we revisit the possible link between the age structure and inflation with the goal of overcoming these identification challenges. To do this, we extend past results in four directions. First, we use long panel data from 1870 to 2016 for 22 OECD countries. This gives us several country-specific demographic cycles from which to identify the effects of age structure on inflation. Second, we control directly for monetary and real factors that offer competing explanations for trend inflation. For instance, we control for money growth and real variables, such as life expectancy, that may drive real equilibrium interest rates. We also control for global factors, by adding time fixed effects to the model, so that the results are primarily driven by cross-country correlation rather than time correlation. Third, we use alternative measures of both inflation and inflation expectations to corroborate the age structure effect. Fourth, we investigate the entire age structure through the use of population polynomials as in Fair and Dominguez (1991) and Juselius and Takáts (2015).

## 2.1 Data

The data are annual and cover 22 advanced economies over the period 1870–2016.<sup>4</sup> Appendix A provides detailed variable definitions and data sources (Table A.1), as well as information on the country-time coverage (Table A.2). Given the long time span, data quality varies over the sample. For instance, data quality is likely to be lower in the early parts of the sample, suggesting that coefficient estimates may be reduced due to a potential attenuation bias. Hence, results for this part of the sample will generally be weaker and should be viewed with caution.

The main variable of interest is the yearly inflation rate, which we denote by  $\pi_{jt}$ , where  $j=1,\dots,N$  is the country index and  $t=1,\dots,T$  is the time index. We exclude observations during the two world wars (and the three years following them), as well as episodes of hyperinflation.<sup>5</sup> This is to ensure that extreme events, where inflation dynamics are likely to be substantially different, do not confound our estimates.

Looking at inflation rates over a long time span (Graph 1, left-hand panel, black solid line), produces several interesting facts. While inflation rates across countries display substantial dispersion, especially before the Second World War, there is also clear comovement globally throughout the sample. High-frequency variation in inflation seems more prevalent in the early parts of the sample, whereas comovement and persistence seem to increase after the war.

The second key factor of interest is the age structure of the population. To study its effect on inflation, we use data on the total number of persons in 17 different five-year age cohorts,  $N_{kjt}$ , where  $k = 1, \dots, 17$  corresponds to the cohorts 0–4, 5–9, 10–14, ..., 75–79 and 80+. We also denote the total population by  $N_{jt}$  and the share of cohort  $k$  in the total population,  $N_{kjt}/N_{jt}$ , by  $n_{kjt}$ .

Looking at broad demographic trends over the sample, we see that the share of young cohorts (0–19 years) declined throughout, reflecting declining birth rates with the exception of the relatively small reversal during the baby boom years (Graph 1, centre left-hand panel). In contrast, the share of the elderly (65+ years) has increased throughout, reflecting mostly increased longevity (right-hand panel). The share of working age cohorts (20–64 years), however, does not show such clear trends (centre right-hand panel): their share increased up until the end of the Second World War, mostly reflecting lower birth rates and fewer young people. This increase temporarily reversed during the baby boom, but picked up again when birth rates subsequently fell. Currently, we seem to be at the beginning of a new reversal as the baby boomers retire. While these trends are largely global, there is also substantial dispersion across countries, in particular with respect to the working age cohorts before the Second World War.

One potential issue with population data over very long samples is that the variation reflects at least three factors: (i) fluctuating, but in trend declining birth rates, (ii) declining infant mortality rates and (iii) increasing longevity. Since the economic effects may not be the same across these sources of variation, it is unclear at the outset which effect dominates in age cohorts that are more strongly affected by two or more of these factors simultaneously,

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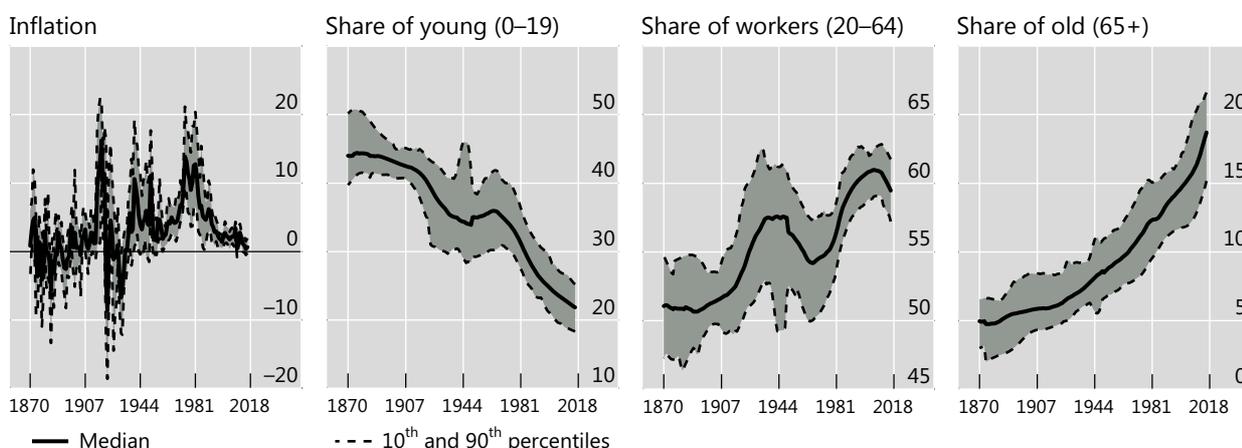
<sup>4</sup> The countries are: Austria, Australia, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Korea, the Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, the United Kingdom and the United States.

<sup>5</sup> We exclude observations where inflation is above +25% or below –25%. The results are essentially the same for higher cut-offs of +/-50% or +/-100%. The number of deleted observations are 84, 25, and 11, respectively.

such as the really young and really old.<sup>6</sup> This problem is exacerbated for the last age category, as it truncates the age distribution at 80+.

Age structure and inflation in the 1870–2016 data

Graph 1



To capture the effect of monetary policy, we need a measure of the gap between the real interest rate and the natural rate. As discussed above, the age structure may already in itself be a sufficient proxy for the latter. In this case, we only need to add a real short-term interest rate to capture the relevant effects. The same holds if the natural rate is, in fact, constant. For this reason our baseline specification includes the real interest rate, defined as  $r_{jt} = i_{jt} - E_{jt}\pi_{jt+1}$ , where  $i_{jt}$  is the nominal short-term money market rate and  $E_{jt}\pi_{jt+1}$  is the expected inflation rate at year  $t+1$ . Following Hamilton et al (2015) and Lunsford and West (2017), we proxy  $E_{jt}\pi_{jt+1}$  with one-year-ahead projections from rolling AR(1) estimates of inflation for each country separately. But if there are other relevant long-term drivers of the natural rate, we also need to include them. In an alternative specification, we try technological growth, population growth, life expectancy and income inequality for this purpose (see eg Eggertsson et al (2017)). We also consider broad money growth in excess of real GDP growth as an alternative direct measure of monetary policy.

Another key control variable is the output gap that may provide additional information, beyond the real interest rate, on for instance supply conditions that may affect inflation. We measure it as  $\hat{y}_{jt} = y_{jt} - y_{jt}^*$ , where  $y_{jt}$  is real GDP and  $y_{jt}^*$  is an estimate of its potential obtained from the Hodrick-Prescott filter (with  $\lambda = 100$ ). Together, the real interest rate and the output gap capture the central information from standard monetary policy frameworks and, therefore, serve as our baseline controls.

In addition, we use a number of other controls that may be important for low-frequency inflation. To control for slow-moving labour market changes, we use labour's share of income and annual hours worked per person. We also use the fiscal balance to capture potential effects that might arise under the fiscal theory of the price level. Our final control variable is a survey-based measure of one-year-ahead inflation expectations from Consensus Forecasts, which are available from 1990 onward. We use this measure both as an explanatory variable and to form an alternative measure of the ex-ante real interest rate.

<sup>6</sup> Under the real equilibrium interest rate channel discussed earlier, an increase in the dependant population drives the real interest rate up if it is generated by an increased birth rate, and down if it is generated by increased longevity (REF).

Since we will use time effects consistently in our regressions throughout this section and the next, we do not include any global variables such as oil and commodity prices.

## 2.2 Modelling the age structure effect

We capture the potential effects of the age structure on inflation in a panel regression setup. Throughout, we aim to avoid confounding a potential age structure effect with concurrent slow-moving country-specific or global factors. To this end, we consistently include both time (year) and country fixed effects. This reduces the risk, for instance, that the impact of oil price shocks (whose timing is close to the entry of the baby boomers into the workforce) is mistaken for an age structure impact.

Capturing the effect of the age structure on inflation involves some methodological issues. In principle, one way would be to include the cohort shares at each point in time in addition to various other control variables:

$$\pi_{jt} = \mu + \mu_j + \mu_t + \sum_{k=1}^{17} \beta_{1k} n_{kjt} + \beta_2' x_{jt} + \varepsilon_{jt} \quad (1)$$

where  $\mu_j$  are country fixed effects,  $\mu_t$  are time fixed effects and  $x_{jt}$  is a vector of controls. With the time fixed effect, we are in essence regressing the country-specific components in inflation, ie deviations from average inflation across countries, on country-specific age structure components. We also regularly cluster the residual along the country and time dimensions, to account for serial correlation and potential global trends that have uneven effects across panels.

Estimating equation (1) directly involves three econometric issues. First, the precision of the estimates becomes weaker if the number of population cohorts is large compared with the number of time periods. Second, the finer the division of the total population, the larger the correlation between consecutive cohorts' shares. Third, since the cohort shares sum to one, there is perfect collinearity with respect to the constant.

An elegant way of overcoming these estimation problems is suggested by Fair and Dominguez (1991) and applied later by Higgins (1998) and more recently by Arnott and Chaves (2012). The idea is to restrict the population coefficients,  $\beta_{1k}$ , to lie on a  $P$ :th degree polynomial ( $P < K$ ) of the form

$$\beta_{1k} = \sum_{p=0}^P \gamma_p k^p \quad (2)$$

where the gammas are the coefficients of the polynomial. Intuitively, this restriction also ensures that neighbouring age cohort estimates cannot differ "too much" from each other.

Combining equation (1) and (2), together with the restriction  $\sum_{k=1}^{17} \beta_{1k} = 0$ , which removes the perfect collinearity between the constant and the cohort shares, yields

$$\pi_{jt} = \mu + \mu_j + \mu_t + \sum_{p=1}^P \gamma_p \tilde{n}_{pjt} + \beta_2' x_{jt} + \varepsilon_{jt} \quad (3)$$

where  $\tilde{n}_{pjt} = \sum_{k=1}^{17} (k^p n_{kjt} - k^p / 17)$ . Once estimates of the  $\gamma_p$  coefficients have been obtained, the  $\beta_{1k}$  coefficients can be recovered from equation (2). In addition, since the  $\beta_{1k}$ :s are linear transforms of the  $\gamma_p$ :s, their standard errors are easy to calculate. Appendix B derives the relevant formulas.

Equation (3) with  $x_{jt} = (r_{jt}, \hat{y}_{jt})'$  and  $P = 4$  forms our baseline estimation equation in the subsequent analysis. We will also consider several modifications of (3) as robustness checks. These include adding dynamic terms, allowing for higher-order polynomials, including additional controls, replacing the polynomial terms with crude age cohort shares, and allowing for heterogeneity across panels.

## 2.3 The link between age structure and inflation

Before we estimate our baseline equation (3), we first estimate a specification that only includes the two main macroeconomic control variables, the real interest rate ( $r_{jt}$ ) and the output gap ( $\hat{y}_{jt}$ ), without any demographic terms (Table 1, Model 1). This helps us assess the value added of the age structure later on. As can be seen, the real interest rate and the output gap in Model 1 have significant coefficients with the correct expected signs and jointly explain around 17% of country-specific inflation.

The effect of age structure on inflation

Table 1

Model	1	2	3	4	5	6
Dependent var.:	$\pi_{jt}$	$\pi_{jt}$	$\pi_{jt}$	$\pi_{jt}$	$\pi_{jt}$	$\pi_{jt}$
$\tilde{n}_{1jt}(\times 1)$		<b>0.57</b> (2.23)	0.74 (1.41)	0.87 (1.84)	0.18 (0.60)	0.62 (3.55)
$\tilde{n}_{2jt}(\times 10)$		<b>-1.17</b> (-2.60)	-1.83 (-1.36)	-3.05 (-2.52)	-0.98 (-2.63)	-1.68 (-4.67)
$\tilde{n}_{3jt}(\times 10^2)$		<b>1.69</b> (2.60)	1.68 (1.30)	3.17 (2.62)	1.14 (2.96)	1.58 (4.38)
$\tilde{n}_{4jt}(\times 10^3)$		<b>-0.52</b> (-2.52)	-0.50 (-1.25)	-1.00 (-2.58)	-0.38 (-2.43)	-0.47 (-3.78)
$r_{jt}$	-0.52 (-3.35)	<b>-0.56</b> (-3.69)	-1.04 (-8.44)	-0.42 (-2.03)	-0.25 (-2.14)	-0.69 (-8.59)
$\hat{y}_{jt}$	0.08 (3.11)	<b>0.08</b> (3.13)	0.04 (1.35)	0.04 (0.99)	0.15 (2.80)	0.08 (2.23)
Countries	22	<b>22</b>	16	22	22	16
Time period <sup>1</sup>	1870–2016	<b>1870–2016</b>	1870–1913	1950–1989	1990–2016	1870–2016
Observations	2219	<b>2217</b>	540	846	584	1257
$R^2$	0.17	<b>0.22</b>	0.32	0.22	0.21	0.36
$R^2$ without age-str.	0.17	<b>0.17</b>	0.31	0.09	0.09	0.30
Age structure F-test <sup>2</sup>	N.A.	<b>0.00</b>	0.06	0.00	0.01	0.00
Contr.: natural rate <sup>3</sup>	No	<b>No</b>	No	No	No	Yes
Country effects	Yes	<b>Yes</b>	Yes	Yes	Yes	Yes
Time effects	Yes	<b>Yes</b>	Yes	Yes	Yes	Yes
Res. country cluster <sup>4</sup>	Yes	<b>Yes</b>	Yes	Yes	Yes	Yes
Res. time cluster <sup>5</sup>	Yes	<b>Yes</b>	Yes	Yes	Yes	Yes
Estimator	FE	<b>FE</b>	FE	FE	FE	FE

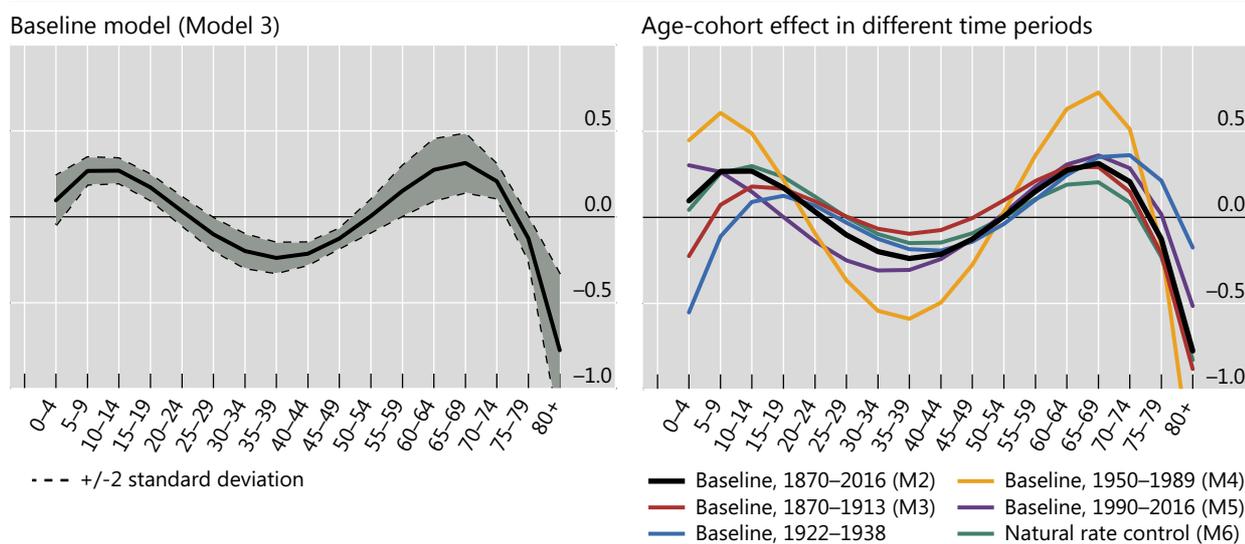
Notes: t-values in parenthesis.  $R^2$ -values refer to the within variation and do not include the fixed effects. <sup>1</sup> Maximum time span across panels reported. <sup>2</sup> F-test of the joint hypothesis that  $\tilde{n}_{pjt}$  for all  $p$ . <sup>3</sup> Natural rate controls: total factor productivity growth; population growth; life expectancy; income inequality. <sup>4</sup> Residuals clustered along the country dimension. <sup>5</sup> Residuals clustered along the time dimension.

We next estimate the baseline model, ie specification (3) with the real interest rate and output gap as controls (Table 1, Model 2). Including the population terms together with the more traditional variables leads to a strong age structure effect. The polynomial terms are highly significant both individually and jointly, and the explanatory power increases by 5 percentage points, from 17% of the variation to 22% (see lines  $R^2$  and  $R^2$  without age structure), compared with Model 1. It appears that, by removing some of the higher-frequency components in inflation, the real interest rate and the output gap help clarify the age structure effect. Notice also that both the real interest rate and the output gap remain highly significant and even the magnitudes remain almost unchanged. This suggests that the age structure effect is mostly orthogonal to the two macroeconomic variables.

An economic interpretation of the age structure effect can be obtained by converting the polynomial coefficients into age-cohort coefficients using equation (2). The age-cohort effects for the baseline estimates (Table 1, Model 2) are shown in the left-hand panel of Graph 2. They reveal a distinct pattern: the young age cohorts are inflationary, the working age cohorts are disinflationary, whereas the old are initially inflationary but turn highly disinflationary as they grow very old. The confidence interval (shown as the grey shaded area) suggests that these effects are generally significantly different from zero.

Estimated age-cohort effects ( $\beta_{1k}$  coefficients, equation (2))

Graph 2



The estimated age cohort effects are broadly in line with those that would arise in a conventional framework if (i) the age structure leads to slow-moving changes in the natural rate, and (ii) monetary policy does not fully internalise such changes. Under these conditions, an increase in the share of dependants, which drives up the natural rate, would be inflationary. Except for the negative impact of the very old, this is essentially the pattern in Graph 2. As noted earlier, the effects of the old cohorts can be ambiguous, as increased dependency and increased longevity have an opposite effect on savings. This might explain the negative effect of the very old, given that the open-ended 80+ cohort is likely to be most strongly affected by increased longevity. Other possible explanations are bequests from the very old to working-savers or fiscal transfers. Indeed, when we control for the fiscal balance, the negative effect of the very old becomes more muted (see Graph 3, lower left-hand panel, below). We also note that the negative effect of ageing (defined as 65+ year olds) found in Yoon et al (2014) among others, is mostly driven by the effect of the very old cohort. For finer age-cohort divisions and allowing for the entire age structure, the results resemble those that we report here.

Before putting too much stock in the baseline findings, one obvious reservation is that the age structure effect may be driven by some particular time period in the sample. For instance, even if we control for time fixed effects, it might still be spuriously related to the gold standard prevailing before 1913 or to different policy priorities in the 1970s. With this in mind, we re-estimate the baseline specification over three periods that are relatively free from extreme events: the pre-world war period, 1870–1913, the post-world war period, 1950–1989, and the post-1990 period to see if the age structure effect survives (Table 1, Models 3–5).

In all of these specifications the age structure remains statistically significant with age cohort effects that are roughly similar (Graph 2, right-hand panel). This is encouraging for the stability of the link between age structure and inflation, as these periods covers very different monetary policy regimes, ranging from a focus on convertibility during the gold standard to inflation targeting more recently. In fact, the age structure effect is also significantly present

in the interwar period, 1922–1938, despite the short sample that includes the Great Depression (Graph 2, right-hand panel, blue line).<sup>7</sup> This suggests that the results may not be spurious, at least in the statistical sense.<sup>8</sup> Nevertheless, there are also some differences across time periods. For example, the cohort effects are almost twice as large in the 1950–1989 sample. However, this does not necessarily imply that predicted inflation rates are larger in magnitude, but rather that inflation outcomes are more sensitive to changes in the age structure. Similarly, there are some indications that the cohort pattern can become slightly tilted in some of the sub-samples. The increase in explanatory power from adding the age structure is low in pre-1950 samples, where high-frequency components in inflation are dominant. However, the increase in explanatory power is much larger, more than 10 percentage points, in post-1950 samples (see  $R^2$  with and without the age structure in Table 1).

Finally, we control for variables other than the age structure that could determine long-term savings-investment equilibrium: total factor productivity growth, population growth, life expectancy and income inequality (Table 1, Model 7). We find that the age structure effect survives the inclusion of these variables and the age cohort impact remains stable (Graph 2, right-hand panel). This further indicates that potential changes in the natural rate that would result from these factors are unlikely to be the main explanation for low-frequency inflation. Furthermore, the findings also suggest that population growth on its own is not sufficient to capture the age structure effect as we control for it.

Taken together, the evidence suggests that the age structure effect is more than just a coincidence in some specific sample, say, as related to the baby boomers' entry to the workforce. To deepen this point, we undertake extensive robustness checks in the next section.

### 3. Robustness checks

We do a number of robustness checks to ensure that the age structure effect is not a coincidental feature of the data. We begin by considering a dynamic specification to ensure that the age structure effect is not spuriously correlated with sub-sample trends that have other origins. To do this, we include lags of inflation, the real interest rate and the output gap in equation (3) and rewrite the equation in error correction form:

$$\begin{aligned} \Delta\pi_{jt} = & \mu + \mu_t + \mu_j + \varphi_1\hat{y}_{jt} + \varphi_2\hat{y}_{jt-1} + \varphi_3\Delta r_{jt} \\ & -\alpha(\pi_{j,t-1} - \lambda_1 r_{j,t-1} - \sum_{p=1}^P \gamma_p \tilde{n}_{pj,t}) + \varepsilon_{jt} \end{aligned} \quad (4)$$

where the term in parenthesis captures deviations from an empirical long-run relationship between inflation, the real interest rate, and the population polynomial. We place the output gap outside the long-run relationship from the outset, as it is, by definition, of higher frequency.<sup>9</sup>

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<sup>7</sup> We also ran rolling regressions with a fixed window of 30 years for both the pre- and post-world war samples. The age structure pattern remains stable in all runs, suggesting that it is robust to minor changes in the sample (details available upon request).

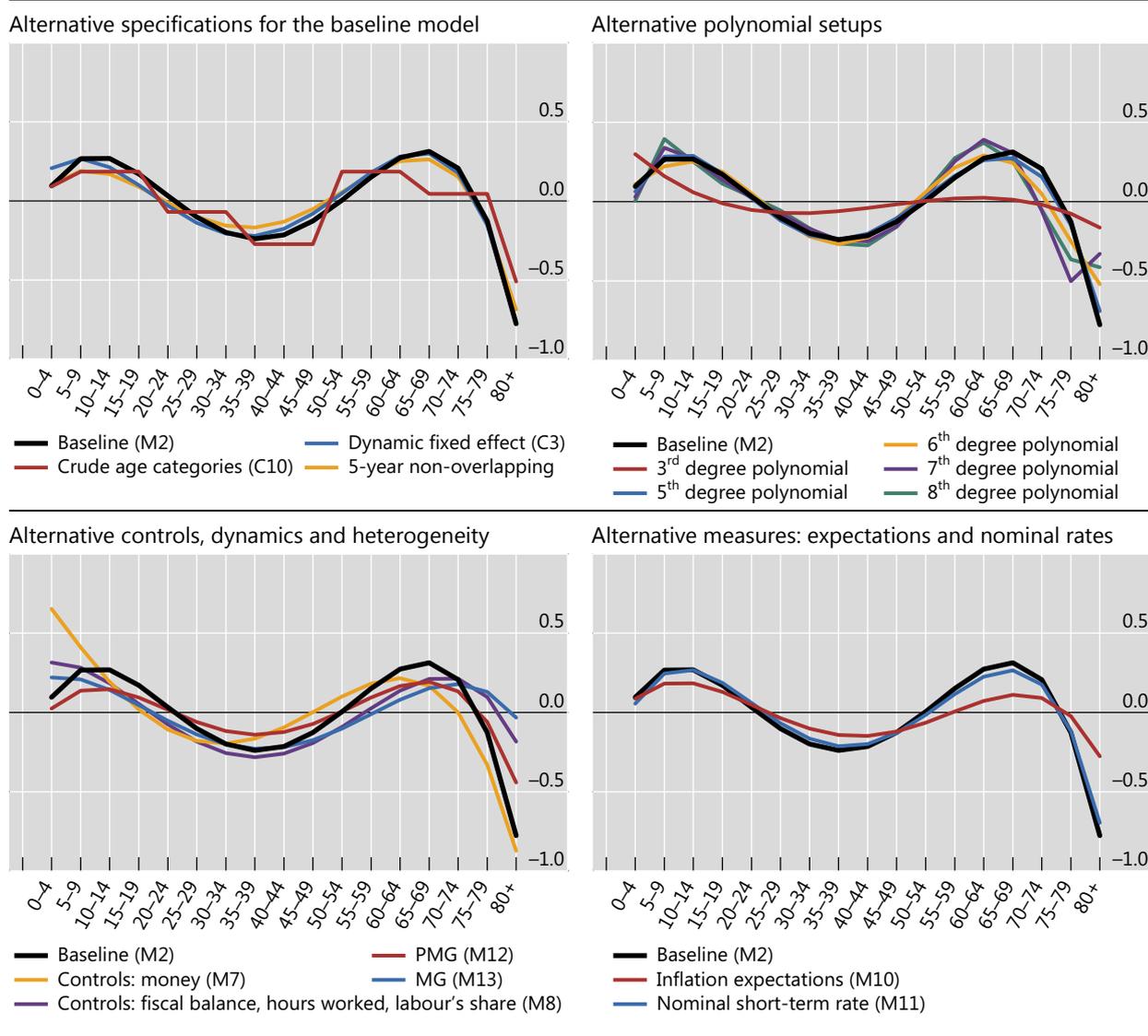
<sup>8</sup> Both the inflation rate and the population variables display dynamics which are sometimes hard to statistically distinguish from unit-root processes. This can in and of itself yield a lot of statistical power to identify an age structure effect, but it can also generate "spuriously" strong correlation between inflation and the age structure in specific sub-samples. But to the extent that such correlations do not reflect a true relationship, they are likely to break down in other sub-samples. We find the opposite: that the effect is reasonably stable over sub-samples.

<sup>9</sup> Including the output gap in the parenthesis does not change the estimated model since it just amounts to a different parametrisation, but it leads to a seemingly large output gap effect that will never materialise, as the steady-state output gap is zero by definition.

Since there cannot be any trends in the *change* of inflation, which is now the left-hand side variable, problems associated with spurious regression do not arise in equation (4). The key question is whether deviations from the empirical long-run relationship (in the parenthesis) matter for changes in inflation, ie whether the “error correction” coefficient  $\alpha$  is significant. This can only happen if the deviation from the long-run relationship is also stationary. The coefficient  $\alpha$  describes how fast deviations from the steady-state translate into changes in inflation. The remaining terms capture short-run dynamics. Note that we do not allow the population terms in equation (4) to have any short-term effects.

The age cohort effects in alternative specifications ( $\beta_{1k}$  coefficients, equation (2))

Graph 3



Adding dynamic terms does not materially change any of the age structure effects. In fact, when we re-estimate the models that appear in Table 1 using the dynamic specification in equation (4) (Table C1 in Appendix C), the age structure effect is always significant and similar to what we had before (Graph 3, upper left-hand panel). Moreover, the error correction term  $\alpha$  is large and highly significant, suggesting that deviations from the long-run steady state help explain changes in inflation. This suggests that the relationship is not coincidentally related to sub-sample trends.

Another way of teasing out the long-run correlation between inflation and the age structure is to redo the regressions on five-year non-overlapping averages of the data (Table

C2 in Appendix C). Furthermore, this also alleviates an implicit problem with overlapping samples: namely that consecutive five-year cohorts overlap between consecutive years. For instance, those in the 5–9 year cohort in 1980, who were five, six, seven and eight years old, will still belong to the same cohort in 1981. Again, with five-year averages, the findings remain essentially the same as in the static regression of Table 1.

One might also ask to which extent our findings depend on the use of the population polynomial. In order to answer this question, we again re-estimate the models in Table 1 with crude age categories in place of the polynomial terms (Table C3 in Appendix C). Specifically, we consider seven crude cohort shares corresponding to ages 0–4, 5–19, 20–34, 35–49, 50–64, 65–79, and 80+. The crude age impact also delivers statistically significant estimates that match the estimated patterns for the polynomials (Graph 3, upper left-hand panel). The only exceptions are cases where the crude age cohorts span sub-cohorts, which take both positive and negative signs – in such cases, the effects of the crude age cohort tend to be insignificant. This suggests that the finer division of cohorts, made possible by the population polynomial, helps to clarify the cohort effects.

A related concern may be that we have misspecified the order of the population polynomial. But all of the estimated cohort effects remain almost identical for higher, fifth-, sixth-, seventh- and eighth-order polynomials (Graph 3, upper right-hand panel). However, for lower-order polynomials, the pattern becomes very different and would not, for example, be compatible with the pattern from the crude age cohorts. This suggests that a fourth-order polynomial is the most parsimonious way of capturing the full effect.

Next, we try additional controls to preclude a potential omitted variable bias and take inflation expectations more fully into account (Table 2). We first control for excessive growth in the money supply (Model 7), ie growth in M2 in excess of GDP growth. The age structure coefficients remain robust and statistically significant. Furthermore, the inclusion of the population polynomial improves fit considerably: it doubles the  $R^2$  from 12% to 24%. The age cohort impact does not change materially (Graph 2, lower left-hand panel). Next, we try three additional variables simultaneously: the primary fiscal balance-to-GDP ratio, average hours worked per week, and labour's share of income (Model 8). With the first variable we intend to capture the impact of age structure through fiscal transfers. Average hours captures any potential effects from age structure affecting labour supply on the intensive margin. Finally, labour's share of income is used, as in the New Keynesian Philips curve literature, to measure marginal costs. Again, the age structure coefficients remain robust and statistically significant. Including the population polynomial substantially increases the  $R^2$  from 9% to 28%. And, as perhaps expected, the age cohort impact also remains virtually unchanged (Graph 2, lower left-hand panel).

So far we have not attempted to include inflation expectations, which are arguably one of the most important candidate explanations for low-frequency inflation. To do so, first we add survey-based inflation expectations to our model (Model 9). Unfortunately, such expectations have had widespread country coverage only from 1990 onwards. With inflation expectations in the model, the age structure effect disappears. Indeed, inflation expectations capture low-frequency inflation better than the age structure.

Yet, as it turns out, the expectations themselves seem to be driven by the age structure (Model 10). When we move inflation expectations to the left-hand side, the age structure becomes significant again. It has approximately the same cohort effects as when we use it to explain actual inflation (Graph 2, lower right-hand panel) and accounts for around 16% of the variation in inflation expectations. This is remarkable given the relative stability of both inflation and inflation expectation in this period. This leaves two interpretations: either the age structure must be a fundamental driver of inflation, since agents condition their forecast on it, or expectations are naïve and backward-looking. In the first case, the age structure effect

indirectly determines inflation through its effect on expectations. In the second case, age structure directly determines inflation and expectations only pick up on this impact.

Robustness: controls, inflation expectations, and country heterogeneity Table 2

Model	7	8	9	10	11	12	13
Dependent var.:	$\pi_{jt}$	$\pi_{jt}$	$\pi_{jt}$	$\pi_{jt}^e$	$i_{jt}$	$\Delta\pi_{jt}$	$\Delta\pi_{jt}$
$\tilde{n}_{1jt}(\times 1)$	-0.25 (-0.65)	0.14 (0.55)	-0.02 (-0.23)	0.31 (1.50)	0.59 (2.46)	0.36 (2.51)	0.13 (0.41)
$\tilde{n}_{2jt}(\times 10)$	-0.07 (-0.10)	-0.78 (-1.49)	0.06 (0.32)	-0.92 (-2.30)	-1.69 (-2.63)	-1.04 (-3.09)	-0.62 (-0.75)
$\tilde{n}_{3jt}(\times 10^2)$	0.43 (0.84)	0.86 (2.22)	-0.04 (-0.28)	0.86 (2.30)	1.61 (2.55)	1.01 (3.40)	0.66 (0.78)
$\tilde{n}_{4jt}(\times 10^3)$	-0.20 (-1.40)	-0.27 (-2.78)	0.00 (0.20)	-0.25 (-2.08)	-0.49 (-2.43)	-0.31 (-3.56)	-0.20 (-0.73)
$r_{jt}$	-0.20 (-1.34)	-0.25 (-3.89)	-0.04 (-1.18)			-0.68 (-15.04)	-0.63 (-6.76)
$\hat{y}_{jt}$	0.13 (2.33)	0.15 (2.74)	0.06 (2.05)			0.05 (2.58)	0.03 (2.16)
$\pi_{jt}^e$			1.22 (22.87)			0.08 (4.11)	0.10 (4.93)
$\hat{y}_{jt-1}$						0.08 (4.11)	0.10 (4.93)
$\Delta r_{jt}$						-0.52 (-8.20)	-0.43 (-9.25)
Error correction: $\alpha$						-0.44 (-10.32)	-0.61 (-15.77)
Countries	22	22	22	22	22	22	22
Time period <sup>1</sup>	1951–2016	1980–2016	1990–2016	1990–2016	1870–2016	1870–2016	1870–2016
Observations	988	535	563	571	2738	2093	2093
$R^2$	0.24	0.28	0.82	0.16	0.10	N.A.	N.A.
$R^2$ without age-str.	0.12	0.09	0.82	0.00	0.00	N.A.	N.A.
Age structure F-test <sup>2</sup>	0.00	0.00	0.07	0.00	0.00	0.00	0.37
Contr: money growth <sup>3</sup>	Yes	No	No	No	No	No	No
Contr: additional <sup>4</sup>	No	Yes	No	No	No	N.A.	N.A.
Time effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Res. country cluster <sup>5</sup>	Yes	Yes	Yes	Yes	Yes	N.A.	N.A.
Res. time cluster <sup>6</sup>	Yes	Yes	Yes	Yes	Yes	N.A.	N.A.
Estimator	FE	FE	FE	FE	FE	PMG	MG

Notes: t-values in parenthesis.  $R^2$ -values refer to the within variation and do not include the fixed effects. <sup>1</sup> Maximum time span across panels reported. <sup>2</sup> F-test of the joint hypothesis that  $\tilde{n}_{pjt}$  for all  $p$ . <sup>3</sup> M2 growth in excess of real GDP growth. <sup>4</sup> Additional controls: primary fiscal balance to GDP; average hours worked per week; labour's share of income. <sup>5</sup> Residuals clustered along the country dimension. <sup>6</sup> Residuals clustered along the time dimension.

Next, we turn to another, more implicit, measure of inflation expectation that is available for the full panel. Specifically, we use the short-term nominal interest rate as a proxy for inflation expectations. While the bulk of the variation in the nominal rate could reflect inflation expectations, low-frequency fluctuation in the real interest rate might also affect nominal rates. Hence, we should view the results with caution. Having said that, the age structure effect is also present in the model explaining short-term nominal interest rates (Model 11). Again, the age structure coefficients are significant and the age cohort effect is almost exactly the same shape as before (Graph 2, lower right-hand panel). Hence, to the degree that nominal rates proxy inflation expectations, this confirms the age structure effect.

As a final robustness check, we allow for both panel heterogeneity and dynamic effects. We first allow all short-run coefficients and the adjustment coefficient of equation (4) to vary with the country index. We estimate this model using the pooled mean group (PMG) estimator derived in Pesaran and Smith (1995) (Model 12). We then allow for full heterogeneity with respect to all the coefficients and estimate the model using the mean group (MG) estimator (Pesaran et al (1999)) (Model 13). Again, the estimates deliver similar and significant coefficient estimates and a similar cohort pattern as before (Graph 2, lower right-hand panel). However, in the MG case, the age structure coefficients are no longer significant, possibly due to the high number of estimated parameters.

## 4. Economic significance and implications

### 4.1 Age structure impact across countries and time

Given that we have established an age structure effect in the data, it is natural to ask how far it can account for actual observed inflation. To answer this question, it is necessary to drop the time fixed effects from the baseline specification (Equation (3) with  $x_{jt} = (r_{jt}, \hat{y}_{jt})'$ ). In the previous section we used time fixed effects as a conservative strategy to ensure that the estimated age structure effect does not reflect any concurrent global trends. However, from an econometric standpoint, only global trends that impinge both on inflation and the entire age structure simultaneously would bias the estimates.<sup>10</sup> It is hard to come up with economic factors that have this property. Worse still, adding time fixed effects can actually bias the estimates if sub-groups of countries experience different secular trends. Hence, dropping them can even lead to more accurate estimates.

Re-estimating the baseline specification without time fixed effects yields:

$$\pi_{jt} = \mu + \mu_j + \frac{2.13}{(7.37)} \tilde{n}_{1jt} + \frac{-0.49}{(-6.37)} \tilde{n}_{2jt} + \frac{0.04}{(6.07)} \tilde{n}_{3jt} + \frac{-0.00}{(-5.57)} \tilde{n}_{4jt} + \frac{-0.85}{(-4.41)} r_{jt} + \frac{0.07}{(7.32)} \hat{y}_{jt} + \varepsilon_{jt}$$

where t-values are reported in parenthesis.

The most noticeable change is that the estimated age structure coefficients are larger in magnitude and more significant statistically. Moreover, the age structure increases explanatory power by 15 percentage points (from 33% to 48%), compared with a model with just the real interest rate and the output gap. The reason is that now we obtain our estimates not exclusively from the cross-sectional variation, but also fully use the time variation.

The age cohort impact also becomes more pronounced: the young and old disinflationary impact along with the inflationary impact of the working age cohort all become stronger (Appendix C, Graph C1, left-hand panel). Furthermore, the previously large negative value for the very old age cohort becomes much more muted. The pattern also shifts slightly to the left.

<sup>10</sup> Consider the following special case of the setup in Pesaran (2006), where a global factor would lead to a bias:

$$\pi_{jt} = \mu + \mu_j + \sum_{p=1}^P \gamma_p \tilde{n}_{pjt} + \beta_2' x_{jt} + e_{jt}$$

$$e_{jt} = \psi f_t + \varepsilon_{jt}$$

$$\tilde{n}_{pjt} = \Psi_p f_t + v_{pjt}, \text{ for } p = 1, \dots, P$$

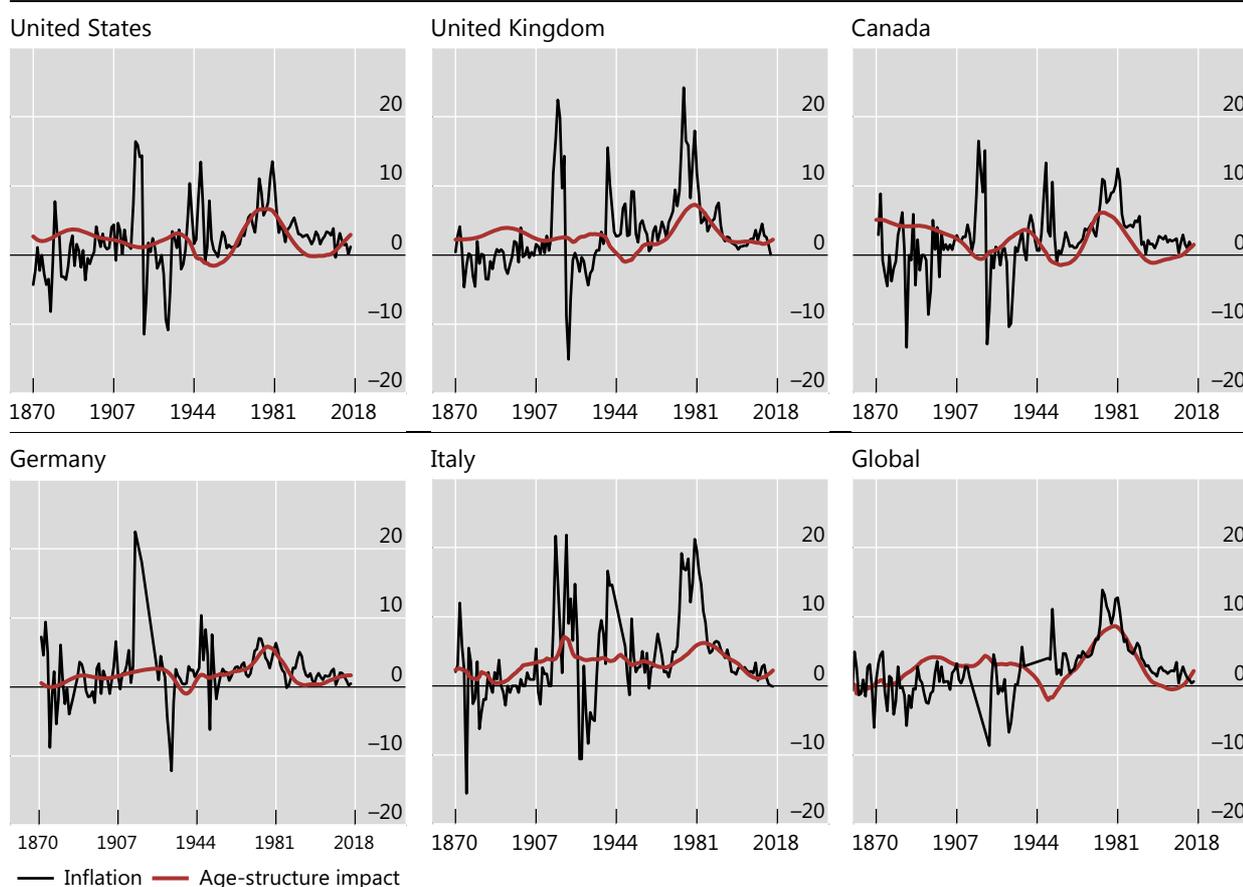
where  $f_t$  is a global factor and  $v_{pjt}$  is some error process. An endogeneity bias with respect to the age structure results if both  $\psi \neq 0$  and  $\Psi_p \neq 0$  (for some  $p$ ), ie if the global factor affects both the age structure and inflation at the same time since the  $\tilde{n}_{pjt}$  population regressors in (5) become correlated with the residual,  $e_{jt}$ , in this case.

This shift is more consistent with a lifecycle-related hypothesis: the new estimates suggest that major net saving cohorts, such as the 50–54 or 55–59 age cohorts, are disinflationary.

The age structure effects account for a large share of low-frequency inflation across countries. For example, comparing actual inflation with the estimated age structure effect in three English-speaking countries (Graph 4, upper panels) and two continental European countries (lower panels), shows a strikingly good fit except for extraordinary events, such as wars or oil price shocks. This is despite the fact that we have used the panel coefficients to calculate the impact for each individual countries. In other words, differences in the fitted effects (red lines) reflect only different age structures across countries, which themselves display substantial heterogeneity, particularly in the early part of the sample. The age structure impact also has a good fit with low-frequency inflation in most of the remaining countries (Appendix C, Graph C2).

Age structure effect accurately describes low-frequency inflation

Graph 4



The fitted demographic effects from the benchmark model are normalised to have the same mean as actual inflation. Figures in percent.

The age structure effect is particularly strong in the United States, which has seen large demographic shifts during the baby boom and bust. The age structure accounts for around a 7 percentage point increase in inflation from 1950s to the 1970s and a similar reduction from the 1970s to the 2000s (Graph 4, upper left-hand panel). Furthermore, demographic developments seem to explain much of the cross-country variation in low-frequency inflation. For instance, the larger swings in US low-frequency inflation compared with German inflation movements (lower panel, first column) seem to reflect mainly the larger demographic changes in the United States.

Interestingly, the estimates show quite stable inflationary pressures for Japan over the past three decades (Appendix C, Graph C2, first row, first column). The lack of deflationary

pressures arises because the deflationary impact of the declining share of the young cohorts roughly offsets the inflationary impact of the increasing share of the old. This also suggests that the low inflation seen in Japan did not necessarily arise from the growing share of old age cohorts but rather from the declining share of the young, as opposed to the argument in Anderson et al (2014).

However, while the panel-based age structure impact has a good fit with inflation trends for most countries, some outliers remain. The fit is weaker than average in Greece and Portugal, for instance. Yet, this should not be surprising, because the transition from military rule to democracy was associated with some economic disturbance in these countries – which could, in turn, have driven inflation trends over the medium term.

The fact that the age structure effect becomes stronger when we drop the time fixed effects suggests that it might also account for global inflation, at least to some degree. This is relevant because a number of recent papers have suggested that a global factor might account for a large share of inflation movements across countries (see, for instance, Ciccarelli and Mojon (2010) or Medel et al (2016)).

To provide some cursory evidence for the potential demographic impact on global inflation, we estimate our baseline model using cross-country averages of the variables as crude measures of their corresponding global components. This gives us 126 annual observations at the global level. We add lagged inflation on the right-hand side, since persistence may be a greater concern in a single time-series context. This increases the precision, but the results also go through without the lagged term. The estimated equation is:

$$\bar{\pi}_t = \mu + \frac{0.43}{(7.29)} \bar{\pi}_{t-1} + \frac{1.85}{(5.31)} \bar{n}_{1t} + \frac{-0.39}{(-4.68)} \bar{n}_{2t} + \frac{0.03}{(3.98)} \bar{n}_{3t} + \frac{-0.00}{(-3.32)} \bar{n}_{4t} + \frac{-0.89}{(-6.84)} \bar{r}_t + \frac{0.09}{(1.03)} \bar{y}_t + \varepsilon_t$$

where the bars denote cross-country averages.

From this analysis, the global age structure appears to be an important determinant of global inflation. The polynomial terms are significant both individually and jointly (F-test p-value 0.005). The estimated global age-cohort pattern is also very similar to that obtained from the panel without fixed effects, albeit of larger magnitude (Appendix C, Graph C1, right-hand panel). Moreover, the global age structure effect accounts for a large share of persistent movements in global inflation (Graph 4, lower right-hand panel). Compared with a model with only the baseline controls (dropping lagged inflation from both models), the explanatory power increases by 29 percentage points from 42% to 71%. These results serve as a first indication that the age structure effects may also be important when seeking to explain global inflation. However, more careful measurement of the global factors is clearly warranted before robust conclusions can be made.

## 4.2 The age structure and endogenous persistence

If the age structure can account for much of low-frequency inflation, how much endogenous persistence is left? To assess this, we add lagged inflation to the right-hand side of the specification in (3). Formally, we run the regression below:

$$\pi_{jt} = \rho \pi_{jt-1} + \mu + \mu_j + \mu_t + \sum_{p=1}^P \gamma_p \bar{n}_{pjt} + \beta_2' x_{jt} + \varepsilon_{jt} \quad (6)$$

where  $\rho$  captures the degree of endogenous inflation persistence. We consider three different specifications of (6): (i) a simple AR(1) for inflation, ie  $\gamma_p = 0$  and  $\beta_2 = 0$  (ii) and AR(1) with the baseline controls, ie with  $x_{jt} = (r_{jt}, \hat{y}_{jt})'$  but no age structure, and (iii) the full specification in (6) with the baseline controls.

Controlling for the age structure lowers estimated inflation persistence consistently across all time periods (Table 2). Consider first the full sample (left-hand column). Adding the population polynomial reduces inflation persistence from 0.49 to 0.35 compared with a model

with only the real interest rate and the output gap. The latter variables, in contrast, do not reduce inflation persistence by much compared with the simple AR(1) specification. The persistence is lower in the pre-1950 sample (centre panel), and most of it can be attributed to the two macroeconomic controls. In contrast, we see much a higher persistence in the post-1950 sample (right-hand column), and it remains high regardless of whether we include the controls or not. Including the age structure in this period reduces endogenous persistence substantially from 0.71 to 0.52. This likely reflects the post-war baby boom that led to large and fairly synchronous changes in the age structure across countries. These results are not confined to the time periods highlighted on Table 3: when we estimate equation (6) with rolling regressions of 40-year windows we typically see a similar drop in estimated endogenous persistence throughout the entire sample.

Model /sample	Full sample	1870–1949	1950–2016
Inflation AR(1)	0.56 (12.33)	0.44 (8.08)	0.71 (14.82)
Inflation AR(1) with baseline controls	0.49 (10.91)	0.08 (4.51)	0.71 (15.33)
Inflation AR(1) with controls and age structure	0.35 (7.66)	0.06 (3.58)	0.52 (10.22)
Estimator	Arellano-Bond	Arellano-Bond	Arellano-Bond

Notes: estimated auto-regressive parameters. The numbers in parenthesis are t-values based on robust residuals.

We also see similar drops in persistence at the global level. For instance, starting from the simple AR(1) specification on the full sample, and successively adding first the two macroeconomic controls and thereafter the demographic terms yields estimates of the auto-regressive parameter of 0.78, 0.64, and 0.43, respectively.

The finding that the age structure reduces inflation persistence has implications for monetary policy, because it implies that endogenous factors, such as inflation expectations, might be less influential than hitherto thought. Hence, the importance of factors such as commitment and credibility, which have traditionally been viewed as critical for inflation control (see Williams (2006)), may have been overstated. If so, the current conundrum with well-anchored long-term inflation expectations and persistently low inflation rates may not be so surprising after all.

### 4.3 Forecasting inflationary pressure

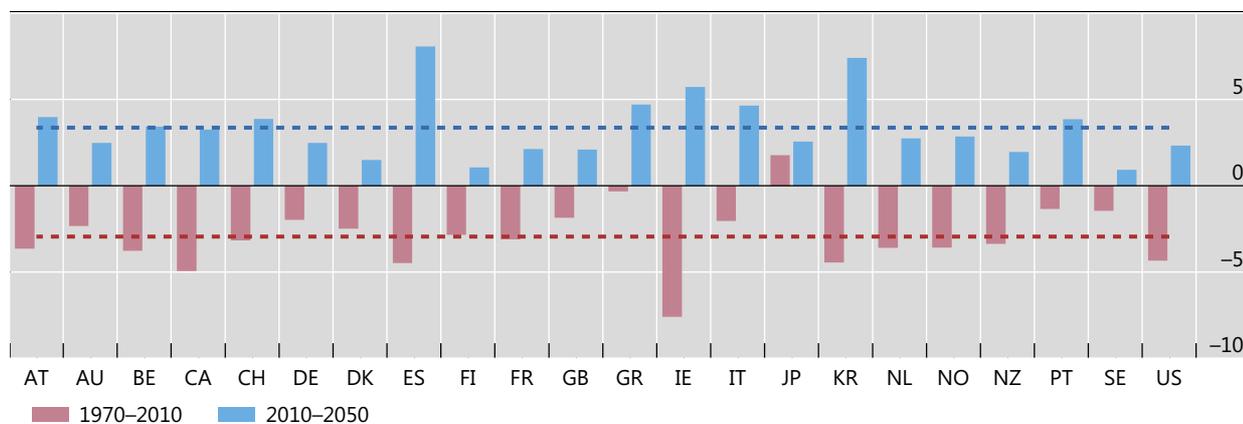
Given the relative stability of the age structure effect, it might help us forecast long-term inflation pressure. The reason is that the age structure can be forecasted several years into the future with reasonable accuracy. For instance, projecting the evolution of the 10–14 years age cohort a decade into the future is not difficult, as the members of that future cohort have already been born.

Our findings suggest that the deflationary effect that the age structure has had on inflation for the past four decades will reverse over the coming decades and become inflationary. This can be seen from Graph 5, which combines our estimates with medium fertility-based population forecasts up until 2050 from the UN population database. Over the past 50 years, the increasing share of working age population has lowered average inflationary pressures by around 3 percentage points (red dotted line below the line). Currently, the shrinking number of young cohorts largely offsets the increasing number of old ones – which holds inflation pressures steady at historically low levels. Over the next half-century, the growing share of the old will dominate and increase inflationary pressure by approximately 3

percentage points on average (blue dotted line). Although country-specific estimates vary, the reversal of past disinflationary pressure into inflationary pressure is present in all individual country estimates (see red and blue bars).

Age structure effect: a turn from disinflationary to inflationary pressure

Graph 5



AT = Austria; AU = Australia; BE = Belgium; CA = Canada; CH = Switzerland; DE = Germany; DK = Denmark; ES = Spain; FI = Finland; FR = France; GB = United Kingdom; GR = Greece; IE = Ireland; IT = Italy; JP = Japan; KR = Korea; NL = Netherlands; NO = Norway; NZ = New Zealand; PT = Portugal; SE = Sweden; US = United States.

The dashed lines show averages of the above economies.

These figures should not, however, be treated as a projection of future inflation, but rather as an indication that we need to better understand the age structure impact.

## 5. Conclusion

Our analysis reveals a stable pattern in data from 1870 to 2016. This takes the shape of a statistically significant relationship between the age structure of the population and low-frequency inflation. Specifically, the young and old are generally associated with higher inflation while working age cohorts are associated with lower inflation. This relationship is robust to changes in estimation methodology and sample.

The results are also economically significant. The age structure accounts for a substantial part of low-frequency inflation variation, both at the country-specific and global level. It may also provide a partial answer to why inflation has been so low in the past decade. Our findings suggest that the inflationary pressure from the increasing share of the old is not yet strong enough to offset the disinflationary pressures from the declining share of the young. Yet, the Great Recession may also have had large and lasting additional effects on inflation, which are difficult to disentangle from those deriving from the age structure.

The robust link between age structure and inflation has at least two direct policy implications. First, since the age structure can account for a substantial share of low-frequency inflation, controlling for it substantially lowers estimates of endogenous inflation persistence. This indicates that some key elements may still be lacking from our current understanding of the inflation process.

Second, the stability of the age structure effect implies that future slow-moving inflationary pressures are at least partly predictable. Using public population projections together with our estimates suggests that inflationary pressures will increase substantially in the coming decades due to population ageing. And such pressures will be difficult to distinguish from, for instance, the delayed effects of the unprecedented monetary easing in

the wake of the Great Recession, if seen through the lens of existing monetary policy frameworks.

The robust empirical link between inflation and the age structure cannot be fully rationalised within existing theoretical models. While the pattern is consistent with a life cycle explanation of slow changes in the natural interest rate that are not fully internalised by monetary policy, two aspects argue against this interpretation. First, the age structure link survives on its own and not only in relation to the real interest rate. Second, from a more conceptual perspective, it is hard to believe that central banks would not eventually have detected slow-moving changes in the natural rate and, if they had done so, would not have fully reacted to them. A political economy explanation, where the median voter's age-based preferences drive the inflation target, as detailed in Bullard et al (2012), does not seem to fit with the evidence either, because its effects are the opposite to what we find: the elderly, except for the very old, are inflationary not disinflationary.

Gaining a deeper understanding of the puzzling link between the age structure and inflation would not only be of theoretical interest but it might also provide guidance for monetary policy. If, for instance, past episodes of high inflation were driven largely by exogenous changes in the age structure, as our evidence suggests, the current conundrum of low inflation, as highlighted by Yellen (2017), may also be related to such changes. Clarifying the channels through which this link arises is crucial for analysing how monetary policy should respond to changes in the age structure. This is all the more pressing as major trends with respect to the age structure over the past decades are about to reverse: ageing is set to increase the share of dependants, through the share of old cohorts, globally. We hope that our findings will stimulate further research in this direction.

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## Appendix

### Appendix A: variable definitions and data sources

Variable definitions and data sources			Table A.1
Variable	Series	Data sources	
$\pi_{jt}$	CPI annual growth	The Global Financial Data; Mitchell's International Statistics; national authorities	
$N_{kjt}$	Number of people in cohort $k = 1, \dots, 17$ , where the age-brackets are 0–4, 5–9, 10–14, ..., 75–79 and 80+	United Nations; Human Mortality Database; Mitchell's International Historical Statistics	
$N_{jt}$	Total population; sum of $N_{kjt}$ over $k$	See $N_{kjt}$ above	
$n_{kjt}$	$N_{kjt}/N_{jt}$	See $N_{kjt}$ above	
$\tilde{n}_{pjt}$	$\sum_{k=1}^{17} (k^p n_{kjt} - k^p / 17)$	See $N_{kjt}$ above	
$i_{jt}$	Short-term interest rates (three-month government bill yields, or closest proxies)	Global Financial Data; Jordà, Schularick and Taylor (2017); Bordo et al (2001); national authorities	
$E_t \pi_{jt+1}$	Projected one-year-ahead rolling estimates (20-year window) of a AR(1) process capped at 0.9 for $\pi_{jt}$	See $\pi_{jt}$ above	
$r_{jt}$	$i_{jt} - E_t \pi_{jt+1}$	See $i_{jt}$ and $\pi_{jt}$ above	
$y_{jt}$	Real GDP	The Global Financial Data; the Maddison Project; national authorities; OECD <i>Economic Outlook</i> ; IMF WEO; Datastream	
$y_{jt}^*$	Hodrick-Prescott-filtered $y_{jt}$ with $\lambda = 100$	See $y_{jt}$ above	
$\hat{y}_{jt}$	$y_{jt} - y_{jt}^*$	See $y_{jt}$ and $y_{jt}^*$ above	
$\pi_{jt}^e$	Survey-based expectations of one-year-ahead inflation	Consensus Forecasts	
Productivity	Total factor productivity	Bergeud et al (2016)	
Population growth	$N_{jt}$ annual growth	See $N_{jt}$ above	
Life expectancy	Life expectancy at birth	Human Mortality Database; Our World in Data; The Human Life-Table Database	
Inequality	Top 1% income share, or closest proxies	Roine & Waldenström (2015); World Wealth & Income Database; Lindert (2000); Chartbook of Economic Inequality	
Broad money	M2 or closest equivalent	Jordà, Schularick and Taylor (2017); European Central Bank; OECD Economic Outlook; IMF IFS; Global Financial Data; national authorities	
Money growth	Broad money annual growth minus $y_{jt}$ growth	See Broad money and $y_{jt}$ above	
Fiscal balance	Fiscal balance as a share of GDP	IMF WEO	
Hours worked	Hours worked per person	Conference Board Total Economy Database	
Labour's share	Share of wages in national income	OECD Economic Outlook; Datastream; national authorities	

Data coverage: series start dates

Table A.2

Countries	AU	AT	BE	CA	CH	DE	DK	ES	FI	FR	GB
$\pi_{jt}$	1864	1862	1871	1871	1851	1851	1851	1851	1901	1851	1851
$N_{kjt}$	1869	1861	1850	1851	1860	1871	1850	1877	1850	1850	1851
$i_{jt}$	1850	1851	1850	1934	1850	1850	1875	1880	1870	1860	1850
$y_{jt}$	1870	1850	1850	1870	1851	1850	1850	1850	1860	1850	1850
$\pi_{jt}^e$	1990	1990	1990	1990	1990	1990	1990	1990	1990	1990	2004
Productivity		1891	1891	1891	1891	1891	1891	1891	1891	1891	1891
Life expectancy <sup>1</sup>	1870	1885	1850	1850	1876	1875	1850	1882	1850	1850	1850
Inequality <sup>2</sup>		1921		1920	1933	1891	1870	1981	1865	1900	1850
Broad money <sup>3</sup>	1959	1959	1969	1968	1975	1950	1962	1969	1980	1961	1982
Fiscal balance <sup>3</sup>	1988	1988	1980	1980	1983	1991	1980	1980	1980	1980	1980
Hours worked <sup>4</sup>	1950	1950	1950	1950	1950	1950	1950	1950	1950	1950	1950
Labour's share <sup>4</sup>	1970	1960	1970	1981	1990	1991	1981	1964	1975	1960	1975
Countries	GR	IE	IT	JP	KR	NL	NO	NZ	PT	SE	US
$\pi_{jt}$	1950	1950	1862	1870	1956	1851	1851	1908	1931	1851	1851
$N_{kjt}$	1950	1950	1861	1884	1950	1850	1850	1874	1864	1850	1870
$i_{jt}$	1950	1950	1885	1879	1951	1860	1870	1950	1880	1870	1850
$y_{jt}$	1950	1950	1850	1870	1953	1850	1850	1870	1865	1850	1850
$\pi_{jt}^e$	1993	1990	1990	1990	1990	1990	1990	1990	1990	1990	1990
Productivity			1891	1891		1891	1891		1891	1891	1891
Life expectancy <sup>1</sup>			1872	1865		1850	1850	1901	1940	1850	1880
Inequality <sup>2</sup>			1901	1886		1914	1875	1921	1976	1903	1913
Broad money <sup>3</sup>	1980	1960	1950	1955	1960	1956	1950	1988	1979	1961	1950
Fiscal balance <sup>3</sup>	1980	1980	1988	1980	1995	1995	1980	1985	1986	1980	1980
Hours worked <sup>4</sup>	1950	1950	1950	1950	1950	1950	1950	1950	1950	1950	1950
Labour's share <sup>4</sup>	2000	2002	1961	1960	1975	1968	1978	1986	1995	1960	1960

Notes: <sup>1</sup> Sample ends in 2015. <sup>2</sup> Sample ends in 2014. <sup>3</sup> Sample ends in 2010. <sup>4</sup> Sample ends in 2013.

## Appendix B: population polynomial

Consider the population regression in Equation 1 with  $K$  cohort shares and without control variables and fixed effects for ease of exposition:

$$\pi_{jt} = \mu + \sum_{k=1}^K \beta_{1k} n_{kjt} + \varepsilon_{jt} \quad (\text{B1})$$

As mentioned earlier, there are at least three difficulties associated with this regression. First, the precision of the estimates becomes weaker if the number of population cohorts is large compared with the number of time periods. Second, the finer division of the total population, the larger the correlation between consecutive cohorts' shares. Third, since the cohort shares sum to one, there is perfect collinearity with respect to the constant.

The first two difficulties can be resolved by restricting the population coefficients,  $\beta_{1k}$ , to lie on a  $P$ :th degree polynomial ( $P < K$ ) of the form

$$\beta_{1k} = \sum_{p=0}^P \gamma_p k^p \quad (\text{B2})$$

where the  $\gamma_p$ 's are the coefficients of the polynomial. Substituting A2 into A1 yields

$$\begin{aligned} \pi_{jt} &= \mu + \sum_{k=1}^K \sum_{p=0}^P \gamma_p k^p n_{kjt} + \varepsilon_{jt} \\ &= \mu + \sum_{p=0}^P \gamma_p \sum_{k=1}^K k^p n_{kjt} + \varepsilon_{jt} \\ &= \mu + \gamma_0 + \sum_{p=1}^P \gamma_p \sum_{k=1}^K k^p n_{kjt} + \varepsilon_{jt} \end{aligned} \quad (\text{B3})$$

where the last step uses  $\sum_{k=1}^K k^0 n_{kjt} = 1$ .

The third difficulty can be resolved by imposing the restriction  $\sum_{k=1}^K \beta_{1k} = 0$ . Substituting (B2) in the sum  $\sum_{k=1}^K \beta_{1k}$  yields

$$\begin{aligned} \sum_{k=1}^K \beta_{1k} &= \sum_{k=1}^K \sum_{p=0}^P \gamma_p k^p \\ &= \gamma_0 K + \sum_{p=1}^P \gamma_p \sum_{k=1}^K k^p \end{aligned} \quad (\text{B4})$$

where the last line uses the fact that  $\sum_{k=1}^K k^0 = K$ . Setting this expression to zero yields

$$\gamma_0 = - \sum_{p=1}^P \gamma_p \sum_{k=1}^K (k^p / K) \quad (\text{B5})$$

and substituting into (B3) yields

$$\pi_{jt} = \mu_{j0} + \sum_{p=1}^P \gamma_p \sum_{k=1}^K (k^p n_{kjt} - k^p / K) + \varepsilon_{jt} \quad (\text{B6})$$

which is as in the main text if we define  $\tilde{n}_{pjt} = \sum_{k=1}^K (k^p n_{kjt} - k^p / K)$  and set  $K = 17$ .

Given estimates of the  $\gamma_p$ :s one can easily calculate the  $\beta_{1k}$ :s directly from (B2). It is also possible to calculate the variance of the  $\beta_{1k}$  estimates. To do this we substitute B5 into B2 to get

$$\beta_{1k} = \sum_{p=1}^P \gamma_p (k^p - \sum_{h=1}^K h^p / K) \quad (\text{B7})$$

where we have changed the index from  $k$  to  $h$  on the sum in the parenthesis to avoid ambiguity. Equation A7 shows that the  $\beta_{1k}$ :s are linear transforms of the estimated  $\gamma_p$ :s. Collecting all the  $\beta_{1k}$ :s and  $\gamma_p$ :s in vector format we can write A7 as

$$\beta_1 = \Psi \gamma \quad (\text{B8})$$

where  $\Psi$  is a  $K \times P$  matrix with typical element  $\Psi_{kp} = (k^p - \sum_{h=1}^K h^p / K)$ . From B8 we have

$$\text{var}(\beta_1) = \text{var}(\Psi \gamma) = \Psi \text{var}(\gamma) \Psi' \quad (\text{B9})$$

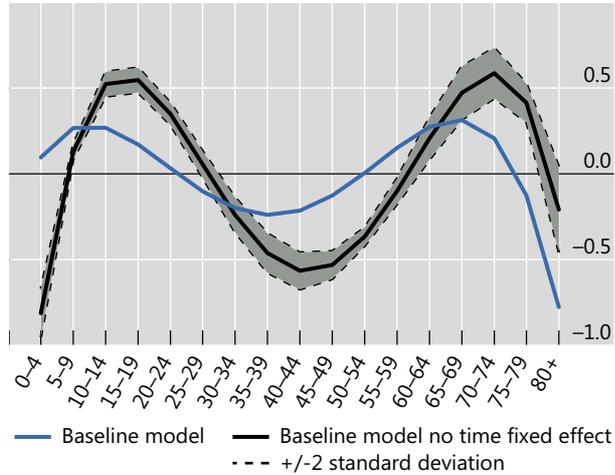
applying the standard formula.

## Appendix C: Auxiliary graphs and tables

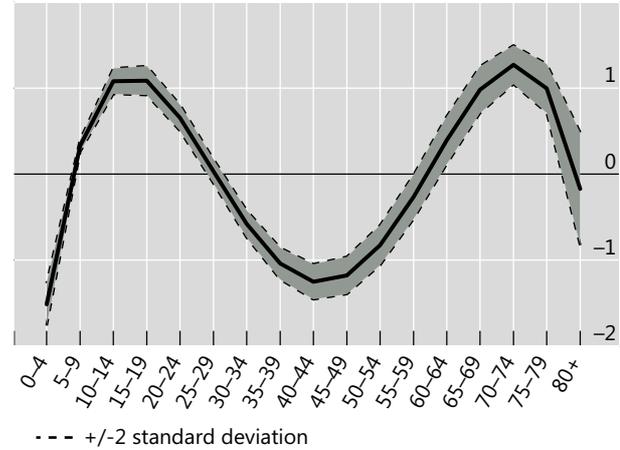
### Age structure impacts

Graph C1

Age structure impact without time fixed effects

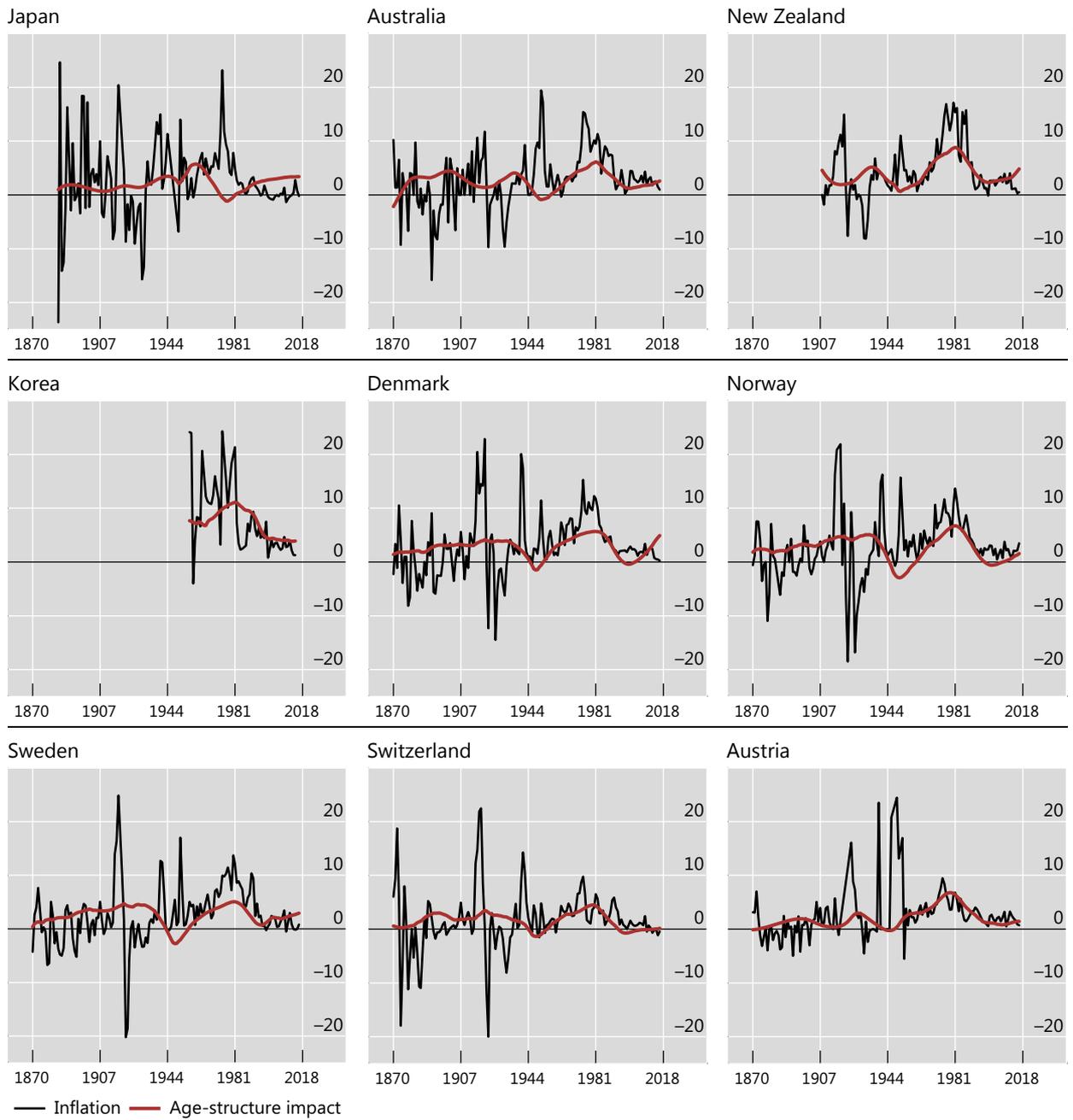


Global setup age structure impact



Age structure impact in individual countries

Graph C2

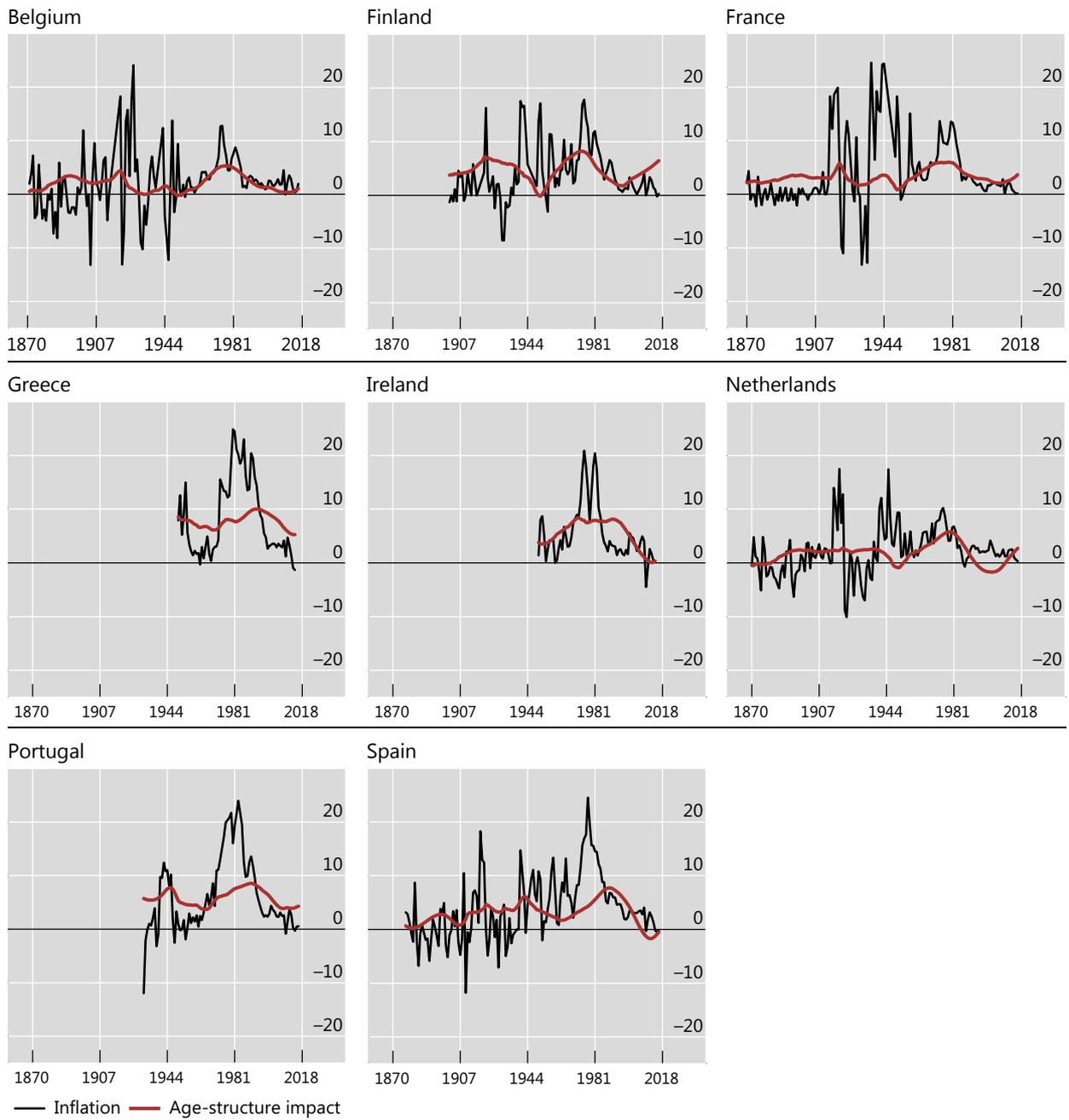


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## Age structure impact in individual countries

continued

Graph C2



## Age structure effect in dynamic models

Table C1

Model	C1	C2	C3	C4	C5	C6	C7
Dependent var.:	$\Delta\pi_{jt}$	$\Delta\pi_{jt}$	$\Delta\pi_{jt}$	$\Delta\pi_{jt}$	$\Delta\pi_{jt}$	$\Delta\pi_{jt}$	$\Delta\pi_{jt}$
$\tilde{n}_{1jt}(\times 1)$		0.15 (0.98)	0.35 (1.41)	0.90 (1.51)	0.79 (1.32)	0.13 (0.34)	0.53 (1.92)
$\tilde{n}_{2jt}(\times 10)$		-0.65 (-1.66)	-1.13 (-2.17)	-2.10 (-1.39)	-2.91 (-2.22)	-0.94 (-2.05)	-1.55 (-3.34)
$\tilde{n}_{3jt}(\times 10^2)$		0.76 (1.84)	1.33 (2.45)	1.83 (1.30)	3.07 (2.64)	1.15 (2.42)	1.52 (4.73)
$\tilde{n}_{4jt}(\times 10^3)$		-0.26 (-1.88)	-0.43 (-2.50)	-0.53 (-1.23)	-0.97 (-2.74)	-0.39 (-2.01)	-0.47 (-5.06)
$r_{jt}$	-0.44 (-2.40)		-0.47 (-2.68)	-0.94 (-10.16)	-0.39 (-1.53)	-0.43 (-2.05)	-0.59 (-6.75)
$\hat{y}_{jt}$	0.03 (1.05)		0.03 (1.09)	0.03 (1.20)	0.00 (0.08)	0.06 (1.04)	0.09 (2.47)
$\hat{y}_{jt-1}$	0.04 (2.20)		0.05 (2.34)	0.06 (2.48)	0.07 (2.12)	0.10 (1.38)	-0.02 (-0.62)
$\Delta r_{jt}$	-0.56 (-5.56)		-0.54 (-5.50)	-0.26 (-2.25)	-0.42 (-2.93)	-0.34 (-7.16)	-0.62 (-9.69)
Error correction: $\alpha$	-0.38 (-2.40)	-0.63 (-11.37)	-0.40 (-7.22)	-0.94 (-16.58)	-0.42 (-4.70)	-0.26 (-2.37)	-0.34 (-8.37)
Countries	22	22	22	16	22	22	16
Time period <sup>1</sup>	1870–2016	1870–2016	1870–2016	1870–2016	1950–1989	1990–2016	1870–2016
Observations	2093	2417	2093	490	823	584	1188
$R^2$	0.42	0.33	0.43	0.63	0.38	0.40	0.56
Age structure F-test <sup>2</sup>	N.A.	0.04	0.00	0.39	0.00	0.06	0.00
Contr.: natural rate <sup>3</sup>	No	No	No	Yes	No	No	Yes
Country effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Time effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Res. country cluster <sup>4</sup>	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Res. time cluster <sup>5</sup>	No	No	No	No	No	No	No
Estimator	DFE	DFE	DFE	DFE	DFE	DFE	DFE

Notes: t-values in parenthesis.  $R^2$ -values refer to the within variation and do not include the fixed effects. <sup>1</sup> Maximum time span across panels reported. <sup>2</sup> F-test of the joint hypothesis that  $\tilde{n}_{pjt}$  for all  $p$ . <sup>3</sup> Natural rate controls: total factor productivity growth; population growth; life expectancy; income inequality. <sup>4</sup> Residuals clustered along the country dimension. <sup>5</sup> Residuals clustered along the time dimension.

## Age structure impact in five-year non-overlapping windows

Table C2

Model	C15	C16	C17	C18	C19	C20	C21
Dependent var.:	$\pi_{jt}$						
$\tilde{n}_{1jt}(\times 1)$		0.32 (2.26)	0.36 (1.58)	1.24 (4.05)	0.65 (1.62)	-0.22 (-0.62)	0.41 (2.89)
$\tilde{n}_{2jt}(\times 10)$		-1.10 (-2.76)	-1.17 (-1.98)	-2.74 (-3.41)	-2.38 (-2.31)	-0.33 (-0.75)	-1.08 (-3.25)
$\tilde{n}_{3jt}(\times 10^2)$		1.19 (2.76)	1.20 (2.07)	2.29 (2.91)	2.60 (2.48)	0.77 (2.71)	1.01 (2.77)
$\tilde{n}_{4jt}(\times 10^3)$		-0.39 (-2.70)	-0.38 (-2.07)	-0.64 (-2.54)	-0.86 (-2.49)	-0.31 (-2.85)	-0.31 (-2.39)
$r_{jt}$	-0.39 (-2.92)		-0.40 (-3.18)	-0.76 (-5.59)	-0.32 (-1.70)	-0.05 (-0.30)	-0.45 (-4.52)
$\hat{y}_{jt}$	0.14 (4.45)		0.14 (4.16)	0.15 (3.25)	0.06 (1.36)	0.35 (3.02)	0.11 (2.58)
Countries	22	22	22	17	22	22	16
Time period <sup>1</sup>	1870–2016	1870–2016	1870–2016	1870–1949	1950–1989	1990–2016	1870–2016
Observations	508	551	508	189	168	151	295
$R^2$	0.13	0.04	0.18	0.34	0.28	0.23	0.27
Age structure F-test <sup>2</sup>	N.A.	0.01	0.03	0.00	0.19	0.01	0.04
Contr.: natural rate <sup>3</sup>	No	No	No	Yes	No	No	Yes
Country effects	Yes						
Time effects	Yes						
Res. country cluster <sup>4</sup>	Yes						
Res. time cluster <sup>5</sup>	No						
Estimator	FE						

Notes: t-values in parenthesis.  $R^2$ -values refer to the within variation and do not include the fixed effects. <sup>1</sup> Maximum time span across panels reported. <sup>2</sup> F-test of the joint hypothesis that  $\tilde{n}_{pjt}$  for all  $p$ . <sup>3</sup> Natural rate controls: total factor productivity growth; population growth; life expectancy; income inequality. <sup>4</sup> Residuals clustered along the country dimension. <sup>5</sup> Residuals clustered along the time dimension.

## Age structure impact in crude age categories

Table C3

Model	C8	C9	C10	C11	C12	C13	C14
Dependent var.:	$\pi_{jt}$						
ages 0 – 4		–0.06 (–0.33)	0.09 (0.56)	–0.10 (–0.35)	0.89 (2.17)	0.23 (0.64)	0.22 (1.52)
ages 5 – 19		0.13 (2.22)	0.19 (2.92)	0.15 (1.53)	0.27 (2.55)	0.25 (2.39)	0.35 (7.05)
ages 20 – 34		–0.11 (–0.93)	–0.69 (–0.58)	–0.06 (–0.29)	–0.26 (–0.99)	–0.13 (–0.67)	0.18 (0.99)
ages 35 – 49		–0.11 (–1.52)	–0.27 (–3.91)	0.10 (0.59)	–0.80 (–3.82)	–0.14 (–1.57)	–0.11 (–1.22)
ages 50 – 64		0.26 (1.51)	0.18 (1.00)	0.15 (0.59)	0.34 (1.29)	0.34 (1.17)	0.21 (1.70)
ages 65 – 79		–0.13 (–0.80)	0.05 (0.33)	–0.29 (–0.40)	0.83 (2.35)	0.06 (0.32)	0.30 (1.58)
ages 80 +		–0.47 (–0.86)	–0.51 (–1.04)	0.08 (0.06)	–2.90 (–1.74)	0.01 (0.02)	–1.13 (–1.67)
$r_{jt}$	–0.52 (–3.57)		–0.65 (–5.17)	–1.04 (–8.44)	–0.40 (–1.85)	–0.25 (–3.84)	–0.78 (–13.13)
$\hat{y}_{jt}$	0.08 (3.35)		0.04 (2.68)	0.03 (1.06)	0.03 (0.66)	0.13 (2.42)	0.04 (1.75)
Countries	22	22	22	16	22	22	22
Time period <sup>1</sup>	1870–2016	1870–2016	1870–2016	1870–1913	1950–1989	1990–2016	1870–2016
Observations	2219	2474	2217	540	846	584	1257
$R^2$	0.17	0.02	0.22	0.27	0.24	0.15	0.32
Age structure F-test <sup>2</sup>	N.A.	0.05	0.03	0.02	0.01	0.00	0.00
Contr.: natural rate <sup>3</sup>	No	No	No	No	No	No	Yes
Country effects	Yes						
Time effects	Yes						
Res. country cluster <sup>4</sup>	Yes						
Res. time cluster <sup>5</sup>	No						
Estimator	FE						

Notes: t-values in parenthesis.  $R^2$ -values refer to the within variation and do not include the fixed effects. <sup>1</sup> Maximum time span across panels reported. <sup>2</sup> F-test of the joint hypothesis that  $\tilde{\pi}_{pjt}$  for all  $p$ . <sup>3</sup> Natural rate controls: total factor productivity growth; population growth; life expectancy; income inequality. <sup>4</sup> Residuals clustered along the country dimension. <sup>5</sup> Residuals clustered along the time dimension.

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