Can Age Structure Forecast Inflation Trends?

Thomas Lindh and Bo Malmberg

The demographic age structure influences the aggregate of individual economic decisions. Standard macroeconomic models imply that inflation pressure will covary with the age distribution unless accommodated by monetary policy. We estimate the relation between inflation and age structure on annual OECD data 1960–1994 for 20 countries. The result is an age pattern of inflation effects consistent with the hypothesis that increases in the population of net savers dampen inflation, whereas especially the younger retirees fan inflation as they start consuming out of accumulated pension claims. This can be explained, for example, with life-cycle saving behavior combined with a cumulative process of inflation, but other mechanisms are also consistent with the results. In any case, the results suggest that demographic projections may be useful for long- and medium-term inflation forecasts. Forecasts from our panel model catch the general downward trend in OECD inflation in the 1990s. However, useful forecasts for individual countries need to incorporate more country-specific information. © 2000 Elsevier Science Inc.

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I. Introduction

In recent years there has been a growing interest in the effects of population change on the macroeconomy. It has been demonstrated that a number of macro variables—including gross domestic product (GDP) growth rates, national saving rates, interest rates, investment rates, the current account, and inflation—are influenced by changes in the age structure of the population.1

1 Age structure variables correlate with savings, Leff (1969), Fry and Mason (1982), Mason (1987), Horioka (1991), Kelley and Schmidt (1996), and Higgins (1998). For critical views see, e.g., Bosworth et al. (1991) and
The demonstration of age effects on macroeconomic variables has practical implications. It is possible to forecast age structure changes years and even decades into the future with reasonable precision. The age effects on macroeconomic variables, therefore, open up a new potential to improve upon medium- and long-term forecasts of macroeconomic trends. This article makes a starting contribution with respect to one key economic variable—the inflation rate. In a previous article, we have shown that in a panel of OECD countries there is an empirical correlation between inflation and the age group shares of the population. Here we extend this work in three directions.

First, we show that a log-linear regression model approximates a model of a Wicksellian cumulative inflation process model. The basic mechanism is that age structure affects saving and aggregate demand in a setting where contracts on factor inputs provide inertia that forces price adjustments when saving rates change.

Second, the model is tested and shown to be robust with respect to a number of conceivable econometric problems like dynamic bias, heterogeneity bias, endogeneity and omitted variables. The results are in agreement with the theoretical model and life-cycle saving behavior except for the oldest age groups.

Third, we evaluate the forecast performance of the model. Using both a pooled model and a model with fixed country effects and random time effects we forecast long-term inflation trends up to 2010. For most OECD countries projections of our estimates imply that the age structure induced inflation pressure will be very low for the next decade and may even generate deflationary pressures. Comparisons to out-of-sample inflation data 1995–1998 indicate that the pooled model is superior for forecasting. The forecasts are fairly accurate with respect to the direction of the general trend. However, age structure forecasts for individual countries need to be integrated into forecasting models that incorporate more country-specific information.

Age Structure and Inflation

Different theoretical frameworks can be used to derive hypotheses of age effects on the inflation rate. In Lindh and Malmberg (1998) age structure induced changes in the saving rate trigger an inflationary process unless the loan rate of interest is accommodating the changing inflation pressure. Similar effects would arise in the textbook IS-LM model. Here an age structure induced change in the saving rate will shift the IS curve and hence, given the monetary policy, influence both aggregate demand and the price level.

A third possibility is a real exchange rate argument. Age structure induced changes in the saving rate are—for obvious reasons—likely to have a stronger effect on the demand for domestic non-tradeables than on the demand for domestically produced tradeables. That will influence the domestic price level, at least during a fixed exchange rate regime [Gagnon et al. (1996)]. This line of argument is supported by the finding of current account effects of a changing age structure [Higgins and Williamson (1997); Higgins (1998); Herbertsson and Zoega (1999)].
A fourth argument is based on the finding of productivity effects of age structure change. As shown in Lindh and Malmberg (1999) the growth rate of GDP per worker is positively influenced when the population share of middle-aged adults is large. Large population shares of the age groups over 65 years, however, have a negative effect. Nominal wage increases that are non-inflationary in one period may, therefore, become inflationary when productivity growth declines due to an unfavorable development of the age structure.

A fifth possibility that has not been fully explored so far looks instead on the labor market effects of a changing age structure. Age composition may, for example, influence union policies and search behavior. In countries with a protective labor legislation young workers face both a higher risk of getting fired and a greater probability of being hired. They will, therefore, have less to lose than tenured workers from wage hikes that might endanger the survival of their firm. The consequence is a structural change in the NAIRU.

A sixth mechanism works through the public sector. A relative decrease in net savers increases the budget deficit by a decrease in income tax receipts and an increase in public expenses. Net savers have the highest incomes, receive less income transfers and utilize public services relatively less than other age groups. Dependent on how the deficit is financed it may reinforce the inflationary impulse.

Moreover, reinforcing combinations of these and other mechanisms are not at all unlikely. Net borrowers are mostly young people with a high relative demand for consumer durables and housing, while retirees consume either out of pension claims or own wealth. Excess demand for consumer durables and housing may well boost investment demand. Since housing is often subsidized in different ways this could simultaneously contribute to increasing deficits. Pay-as-you-go pension systems and health care for the elderly, of course, also increase budgetary pressures as the share of retirees increase. Blomquist and Wijkander (1994) show—in an overlapping generations model—that large fluctuations in the age distribution such as we observe in OECD data, may cause fundamental shifts in aggregate macro relations. For example the relation between real interest rates and aggregate saving may change sign. The same demographic shock that decreases the proportion of net savers implies that the labor force would be relatively less experienced, so the quality of average labor deteriorates. If nominal wages cannot be adjusted downwards and capital and labor are substitutes, the demand for investment funds increases as effective labor becomes more expensive. This creates a reinforcing inflationary impulse.

There is thus a wealth of potential mechanisms for age effects on inflation. The relative importance of these mechanisms will vary over countries due to differences in trade dependency, social welfare traditions, monetary regimes etc. We have chosen to focus our explanation of the patterns in data on a simple savings mechanism, and found the empirical results to be largely consistent with that model. However, we cannot exclude other possibilities.

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3 Shimer (1998) provides evidence of a close association between the share of young workers in the labor force and the aggregate unemployment rate in the U.S. He claims that the bulk of recent decreases in the U.S. unemployment rate is explained by an older work force.
Policy and Forecasts

Whether inflationary or deflationary impulses actually materialize in the price level, of course, depends on a multitude of other factors, many of which are institutional or political. Using a panel allows us to control for such country- and time-specific factors. Separate country regressions confirm the existence of a significant correlation between the age structure as a whole and inflation, but, as would be expected, with a varying and much less well-determined pattern. Inclusion of interest rate variables leave the age pattern essentially unchanged. That is in accordance with our maintained (and reasonable) hypothesis that monetary policy in our sample has not accommodated age structure changes.

The fertility fluctuations—known as baby booms or busts—that have left their imprints on the age distributions of the OECD countries are very large. In the period 1960-1994 the population share between 30 and 64 years old in these countries have varied between 34 and 48 percent. Now and for some years ahead this share of mainly net savers is near apex for most countries since the large postwar cohorts have not yet begun to retire. Thus, we see a low demographic inflation pressure for another 5 or 10 years to come. However, as the population in the OECD area continues to grow older this active population share will eventually decrease and around 2010 rekindle inflation pressures similar to those in the 1970s.

Our results suggest that changes in the age structure may account for 5-6 percentage points of the OECD variation in the rate of inflation. Thus monetary policy should be able to make good use of demographic projections in predicting medium-term inflation pressures and changes in the trend.

Our argument is organized in the following way. In Section II, we develop a simple model of age structure effects on inflation. Section III presents the OECD estimates and out-of-sample projections. Section IV finally sums up our conclusions.

II. An Inflation Model

Age structure changes would from a theoretical viewpoint be expected to be a major determinant of all aggregate macroeconomic relations. The wealth of probable and possible mechanisms pose methodological problems as we analyze the empirical evidence. Interaction between different mechanisms will sometimes blur and sometimes enhance the patterns in the data.

However, a consistent theoretical framework may still aid interpretation even if numerous complications have been ignored. In Lindh and Malmberg (1998) we model a savings mechanism version of Wicksell’s cumulative inflation process without attempting to relate it to the empirical model. Here we shortly recapitulate this theoretical model and proceed to show that a simple log-linear model can approximate it.

Saving is here assumed to be the only transmission mechanism for the age effects. Although it would be preferable to model saving and labor supply decisions of the households explicitly, this is too complicated for a heterogeneous population with changing age distribution. Therefore, we take labor supply as exogenously given and the aggregate saving rate as a function of an index of the age structure. We consider a case of pure inside money making it superfluous to model money explicitly. Price changes will in this case not affect the loan rate of interest since the money supply is infinitely elastic.
We also stick to very simple standard technology assumptions and allow for a variety of expectations assumptions.

**A Formal Model of a Cumulative Inflation Process**

Labor supply \( L \) is exogenous. Saving rates vary with the age structure through the specification \( s = B \prod n_{ik} \), a log-linear index of age group shares of the population, \( n_{ik} \), where \( a_k \) and \( B \) are constant parameters. The production sector is represented by a single optimizing firm. Gross production \( Q \) of an aggregate good includes the replacement value of the capital stock \( K \) brought forward to the next period. We assume full employment. Let gross production in period \( t \) be described by

\[
Q_t = AK_t^a L_t^b
\]

Households hold all assets, in this case loan claims on investment. The gross interest is \( R_t = 1 + r_t \), where \( r_t \) is the nominal loan rate of interest. Both interest and wages, \( w_t \), to be paid by the firm in the current period are determined by contracts in the preceding period. The gross nominal income of households is then

\[
Y_t = w_t L_t + R_t P_t K_t
\]

because the principal was invested in capital goods in the preceding period at then current prices \( P_{t-1} \). With \( P_t \) denoting the expected price in the next period, competitive conditions for the representative firm ensure that expected firm profit vanishes

\[
P_t Q_t - Y_t = 0
\]

The constant-returns assumption implies that first-order profit maximizing conditions only determine optimal capital intensity. It also makes it convenient to proceed using lower case letters for input and output per labor unit. The first-order condition is

\[
aP_t AK_t^{a-1} = R_t P_t
\]

so the optimal planned capital intensity is

\[
k^*_t = \left( \frac{R_t P_{t-1}}{aP_t^a} \right)^{\frac{1}{a-1}}
\]

Given the saving rate \( s_t \), market equilibrium for exchanged quantities of the aggregate good then requires that

\[
(1 - s_t) y_t + P_t k_{t+1}^{L_{t+1}} L_t = P_t A
\]

To simplify we henceforth assume that \( L_{t+1} = L_t \) but it would be straightforward to allow labor supply to differ between periods. Using the optimal capital intensity in equation (5) and the zero-profit condition equation (3), we can rewrite equation (6) as

\[
(1 - s) P_t A \left( \frac{R_t P_{t-1}}{aP_t^a} \right)^{\frac{a}{a-1}} = P_t \left[ A \left( \frac{R_t P_{t-1}}{aP_t^a} \right)^{\frac{a}{a-1}} - \left( \frac{R_{t+1} P_t}{aP_t^a} \right)^{\frac{1}{a-1}} \right]
\]
an expression that can be rearranged using $\pi_t = P_t/P_{t-1}$ and eliminating $A$

$$
(1 - s_t) = \frac{P_t}{P_{t-1}} \left[ 1 - a \left( \frac{R_t^e}{R_{t+1}} \right)^{\frac{1}{\alpha}} \left( \frac{\pi^e_{t+1}}{\pi_t^e} \right)^{\frac{1}{\beta}} \right]
$$

It can be shown that an increase in the saving rate generates an at least temporary decrease in the rate of inflation over a broad range of expectational assumptions, although not if the increase is perfectly anticipated. However, a decrease in the nominal loan rate of interest would counteract the deflation impulse from an increased saving rate.

**Approximation to a Log-Linear Empirical Model**

The reason to keep saving behavior exogenous here is not that we believe that it will be constant and unaffected by interest and inflation. But we do believe that the saving rate response to these variables will be dominated by changes in the age distribution of the population. We cut through this by the assumption that $s$ is exogenous, like we have cut through other aggregation difficulties by assuming a representative firm on the production side.

To derive an estimable model further assumptions on the expectation formation are needed. We prefer to abstain from explicit modeling of the process of expectation formation. Instead we assume that $P_{t-1}/P_t$ is approximately equal to one. This is compatible with a rational expectations hypothesis but also allow for other hypotheses with reasonably small expectation errors one period ahead. The equilibrium condition, equation (8), can then be approximated by

$$
B\Pi_n = a \left( \frac{R_t^e}{R_{t+1}} \right)^{\frac{1}{\alpha}} \left( \frac{\pi^e_{t+1}}{\pi_t^e} \right)^{\frac{1}{\beta}} E_t
$$

where $E_t$ is a multiplicative approximation error. Taking the logarithm of this expression and using the continuous approximations of the discrete gross rates of interest $R \approx e^i$, and the gross rate of inflation $\pi \approx e^i$ where $i$ is the average net rate of inflation over the period, we get

$$
(1 - a)\log B + (1 - a) \sum_k a_k \log n_{kt} = (1 - a)\log a - r_{t+1} + ar_t - ai_t + i_{t+1}^r + e_t
$$

We assume expected inflation to be formed by a linear combination of current and historical inflation. To simplify we limit backward looking to the preceding period such that $i_{t+1}^r = b_1 j_t + b_2 j_{t-1} + c_r$, where the last term reflect unspecified information that could be both forward-looking and historical. We can then rearrange

$$
i_t = \frac{1}{a - b_1} \left[ b_2 i_{t-1} + (1 - a) \left[ \log a - \log B - \sum_k a_k \log n_{kt} \right] - r_{t+1} + ar_t + c_t + e_t \right]
$$

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4 If $r_{t+1}$ is a policy variable chosen to keep inflation constant at a given level $i^*$ the long-run rate of interest compatible with the inflation target is $r^* = i^* + \log a - \log B - \sum_k a_k \log n_{kt}$

With a changing age distribution this target rate of interest varies with the age index.
Thus, inflation in the current period depends on lagged inflation, the age savings index, the current interest rate and the one-period ahead interest rate, a constant and a time-specific (and possibly country-specific) error term. From theory, we would expect net saving age groups to have a negative effect relative to net consumers. If $a > b_1$, the lead of interest should have a negative coefficient and current period interest a positive coefficient, because the former decreases investment demand and the latter increases income and thereby consumption demand.

This model is, of course, neither precise nor very realistic. Real money is not pure inside money, so there would be some interaction between the loan rate and the rate of inflation. Household savings, of course, would also interact with interest rates. Interaction with the government sector is potentially very important. But the model catches essential characteristics of an inflationary or deflationary process and it can be estimated with robust standard methods.

We would expect to observe inflationary and deflationary impulses induced by age structure changes in empirical data, not necessarily in the form of actual inflation or deflation—because appropriate monetary policies would counteract the impulses—but as changes in the underlying inflation pressure that monetary authorities have not foreseen or decided not to accommodate. In the next section we proceed to demonstrate that the empirical evidence points in the direction that monetary policy in the OECD has not on average sterilized the age-structure impulses to inflation.

### III. Estimating the Relation Between Inflation and Age Structure

Most economic macro variables are endogenous with respect to inflation. For example, government debt, monetary aggregates, rates of interest, aggregate demand and supply as well as saving and growth are jointly determined with inflation by economic equilibrium processes. All of the mentioned variables are theoretically expected to correlate with the age structure as well. The difference is that, although there is a feed-back from these variables to demography (as numerous micro-studies show), this feed-back changes the overall age structure rather slowly and mainly affects fertility and migration in the short run. Fertility and migration in turn feed back into the age structure mainly in the younger cohorts. The age structure above around age 30 years is, therefore, an unusually good approximation of a truly predetermined variable in relation to contemporary annual inflation.

A reduced form regression model where inflation is explained by age group shares is thus much more likely to approximately satisfy the exogeneity assumptions than the average macro model is. Although equation (11) is derived from a consistent equilibrium model, there are by far too many simplifications on the way to validate any strict structural interpretation. However, that probably holds for every tractable macro model.

Furthermore, we clearly have differences between countries and over time that are hard to measure quantitatively, like political institutions, business cycles, etc., which also pose a problem in making inferences from regressions. Panel estimations allow us to control to some extent for these time- and country-specific factors.

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5 Note that in the model this rate is the rate of return on nominal investment funds made available in the current period. Thus, this variable is part of the current period information set.
Data

We use annual data on consumer price indexes, CPI, from 20 OECD countries taken from Main Economic Indicators, OECD (1996). Turkey and new OECD members were excluded due to lack of data. Iceland and Luxembourg were excluded due to small populations. More unfortunate is that Germany also had to be excluded, due to the reunification. Age distribution data are taken from United Nations (1994) in the form of annual estimates per 31 December of the population in 5-year age groups. To ensure that age variables are predetermined we lagged the raw data so our age variables for year $t$ refer to 31 December in year $t - 1$. Inflation in year $t$ equals $\log \frac{CPI_t}{CPI_{t-1}}$ leaving us with usable data 1961–1994. That is 34 time-series observations per country, except for Japan where CPI data were available only from 1970. In total, we then have 670 usable observations.

We have so far been unable to obtain data on interest rates that are comparable over both time and our sample of countries. In our basic estimations we, therefore, omit the interest variables. But we do have some interest data for most of the countries although with widely varying length of the time series that we use below, to check whether our omission of the interest rate in the basic estimations may cause serious problems. A priori it is, however, not clear that omitting an endogenous variable like interest is worse than including it with measurement errors and all.

Figure 1 documents the simple correlation pattern between the 5-year age groups and inflation in our estimation sample.

The correlation pattern is not quite consistent with the hypothesis that inflation is lower the more net savers there are in the population. People around 50 have a positive correlation to inflation and the very oldest have a negative correlation. However, none of the correlations between 45 and 64 are statistically significant on the 5 percent level. The age variables are mutually correlated so the simple correlations may be misleading as is well known. Using the share of people between 30 and 64 years as a rough measure of net...
savers and regressing inflation on this yields a slope coefficient of $-0.46$ with a $t$ value of $7.36$. This indicates a quite substantial impact because the group, 30–64 years, in, for example, Japan increases with about 10 percent over the period. Taken at face value, inflation in Japan then would have been 5 percent higher today if the Japanese population had remained stationary with respect to the age distribution in 1960.

The Regression Model

Because the evolution of the age shares is highly correlated, multi-collinearity adversely affects the precision of the estimates of individual parameters. For predictive purposes this is of less importance so long as the correlation pattern remains stable. Fair and Dominguez (1991) handle this problem by imposing a polynomial form on the age effects. Rather than impose functional restrictions on the effects we have preferred to let data speak for themselves as much as possible, by using age groups that are aggregated to reflect expected differences in savings behavior. Some experimentation with our sample show that the regression variables used to impose a polynomial restriction are, in fact, considerably more collinear than the age variables we use.

Inflation rates in percent are regressed on the logarithms of the following five age shares of the population: young adults 15–29 years old, mature adults 30–49 years, middle-aged 50–64, young retirees 65–74, and finally old people 75 years and above. This division was based on the hypothesis that the division of income between saving and consumption is the dominant mechanism behind the age share impact on inflation. We have attempted to identify age groups with similar saving behavior. Young adults are likely to have higher consumption than income, because they are engaged in higher education and if working have lower incomes. The consumption behavior of mature adults is dominated by the family situation, while the middle-aged are past their child-rearing years and are adapting behavior accordingly. Young retirees consume out of pension claims or personal wealth. Much empirical evidence point in the direction that the oldest group tends to save more than predicted by life-cycle consumption smoothing. This may be due to bequest motives or that they simply lack the health to consume actively.

The share 0–14 years old, i.e., children, were excluded for three reasons. First and foremost, some age share need to be dropped due to the high degree of multi-collinearity when they are all included. Second, the children have the highest partial correlations to other age groups, somewhat unexpectedly they are highly negatively correlated with the two oldest age groups. Dropping the children, therefore, substantially alleviates the collinearity problem. Third and more speculative: If there is some endogeneity in the age shares with respect to inflation, however unlikely, it seems most likely to show up in the age group where fertility changes immediately affect the size.

Thus the basic regression equation (omitting interest rates) that we derive from the theoretical approximation, equation (11) is

$$i_t = \beta_0 + \beta_1 i_{t-1} + \sum_{\text{age}=15-29} \beta_{\text{age}} \log n_{\text{age},t} + \eta_t + \nu_t + \varepsilon_{jt}$$

where we allow for country- and time-specific effects, $\eta_t$ and $\nu_t$, and $\varepsilon_{jt}$ is a stochastic disturbance with mean zero.
Estimates of Age Effects on Inflation

In Table 1, the first two columns report estimates from pooled data, first the straight OLS-estimates (650 observations) and then IV-estimates (590 observations) obtained by a GMM-estimator using the third and fourth lag of inflation to instrument for lagged inflation. The latter is motivated by the correlation that lagged inflation may have with the contemporaneous error, because we use it to proxy for expected inflation. This is likely to depend on other information that is not available to us. The tests for over-identifying restrictions do not reject the validity of the instruments. The estimates are essentially the same, although the sign of the coefficient on young adults changes. These coefficient estimates are, however, so imprecise that we cannot reject a restriction to zero in both cases. Endogeneity bias due to lagged inflation would not seem to be any major concern here.

The residuals show signs of country-specific effects and certainly time-specific effects. Heteroskedasticity is present and there are indications that higher-order serial correlation may be a problem, so the standard errors were estimated with the correction due to Newey and West (1987) to achieve heteroskedasticity and autocorrelation consistent errors.

Table 1. Inflation (CPI growth) explained by logarithms of age shares and lagged inflation. Absolute t-values in parentheses computed from Newey–West estimates of heteroskedasticity and autocorrelation consistent standard errors allowing for six lags; p-values for tests in brackets. Wald tests for the joint significance of the age shares. Time- and country-specific effects are tested on the residuals.

<table>
<thead>
<tr>
<th>dep var</th>
<th>CPI inflation</th>
<th>Pooled OLS</th>
<th>Pooled IV</th>
<th>Fixed effects</th>
<th>Fixed country and random time effects</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>-2.16</td>
<td>-6.65</td>
<td>—</td>
<td>—</td>
<td></td>
</tr>
<tr>
<td>(0.64)</td>
<td>(1.65)</td>
<td>(1.65)</td>
<td>(1.65)</td>
<td>(1.65)</td>
<td></td>
</tr>
<tr>
<td>Lagged inflation</td>
<td>0.80</td>
<td>0.84</td>
<td>0.65</td>
<td>0.68</td>
<td></td>
</tr>
<tr>
<td>(28.4)</td>
<td>(21.3)</td>
<td>(18.3)</td>
<td>(17.9)</td>
<td>(17.9)</td>
<td></td>
</tr>
<tr>
<td>Young adults, 15–29 years</td>
<td>0.78</td>
<td>-1.27</td>
<td>-3.64</td>
<td>-1.57</td>
<td></td>
</tr>
<tr>
<td>(0.59)</td>
<td>(0.85)</td>
<td>(1.83)</td>
<td>(1.03)</td>
<td>(1.03)</td>
<td></td>
</tr>
<tr>
<td>Mature adults, 30–49 years</td>
<td>-4.05</td>
<td>-5.1</td>
<td>-6.67</td>
<td>-7.34</td>
<td></td>
</tr>
<tr>
<td>(3.61)</td>
<td>(3.72)</td>
<td>(3.00)</td>
<td>(3.91)</td>
<td>(3.91)</td>
<td></td>
</tr>
<tr>
<td>Middle-aged, 50–64 years</td>
<td>-0.61</td>
<td>-0.58</td>
<td>-0.33</td>
<td>-1.38</td>
<td></td>
</tr>
<tr>
<td>(0.50)</td>
<td>(0.48)</td>
<td>(0.25)</td>
<td>(1.01)</td>
<td>(1.01)</td>
<td></td>
</tr>
<tr>
<td>Young retirees, 65–74 years</td>
<td>4.21</td>
<td>3.30</td>
<td>2.76</td>
<td>3.78</td>
<td></td>
</tr>
<tr>
<td>(3.20)</td>
<td>(2.74)</td>
<td>(2.09)</td>
<td>(2.98)</td>
<td>(2.98)</td>
<td></td>
</tr>
<tr>
<td>Old people, 75+ years</td>
<td>-2.90</td>
<td>-2.35</td>
<td>-3.50</td>
<td>-2.15</td>
<td></td>
</tr>
<tr>
<td>(3.91)</td>
<td>(3.15)</td>
<td>(2.28)</td>
<td>(2.32)</td>
<td>(2.32)</td>
<td></td>
</tr>
<tr>
<td>$\hat{R}^2$</td>
<td>0.74</td>
<td>0.74</td>
<td>0.51</td>
<td>0.55</td>
<td></td>
</tr>
<tr>
<td>$\chi^2$ (5) sign age shares</td>
<td>46.1 [0.00]</td>
<td>38.8 [0.00]</td>
<td>13.6 [0.02]</td>
<td>29.0 [0.00]</td>
<td></td>
</tr>
<tr>
<td>F(19) country effects</td>
<td>1.63 [0.04]</td>
<td>1.29 [0.18]</td>
<td>—</td>
<td>—</td>
<td></td>
</tr>
<tr>
<td>F(32) time effects</td>
<td>8.46 [0.00]</td>
<td>7.95 [0.00]</td>
<td>—</td>
<td>1.23 [0.18]</td>
<td></td>
</tr>
<tr>
<td>$\chi^2$(1) over-ident restr</td>
<td>—</td>
<td>0.68 [0.41]</td>
<td>Hausman</td>
<td>4.83 [0.09]</td>
<td></td>
</tr>
</tbody>
</table>

The two last columns in Table 1 report estimates where both country- and time-specific effects have been taken into account. It is often recommended in text books that a fixed effects model should be used when the sample population is identical to the population we wish to make inferences on, while the random effects model is more appropriate when we wish to make inferences outside the sample population. The country sample is indeed the one we wish to predict for. The sample over time is another matter, however, since we
want to make forecasts. Thus we report in the last column an estimate where time effects have been treated as random. The random effects estimator is more efficient and the Hausman test does not reject equality of the coefficients at the 5 percent level. On that evidence, the random effects estimate should be preferred. But the Hausman test depends on both estimators being consistent, and this cannot be asserted in this case.

It is well-known\(^6\) that the conventional least-squares estimator in a dynamic panel model suffers from a dynamic bias that can be quite considerable in panels with short time series. With 34 time series observations this particular source of bias is unlikely to be seriously misleading. The very precisely estimated lag coefficients differ by 20 percent between the pooled model and the fixed effects model. Since it is known that the dynamic bias is negative on this coefficient, the difference could be an indication of the size of the bias. But 20 percent is much more than theoretically expected. On the other hand, the failure to account for highly significant general time effects in the pooled model should cause omitted variable bias. If that is the cause of the difference the somewhat stronger age effects in the panel models should be more reliable.

As a further check we estimated a model where country effects were differenced away to correct for any dynamic bias in the fixed effects model [Anderson and Hsiao (1982)]. However, the consistent GMM estimator of this model is not very efficient [Kiviet (1995)]. The precision of the estimates deteriorated so much that we do not report them here. There is evidence that differencing introduces serious problems with serial correlation. First-order serial correlation is expected since differencing introduces a moving average of the error. That can be handled by using longer lags of either levels or differences as instruments. However, there are strong indications of serial correlation of much higher order, which invalidates the use of longer lags as instruments. Nevertheless, the pattern of the point estimates of the age coefficients is, in spite of these serious problems, not very different from those presented in Table 1. The coefficient of lagged inflation fares less well and changes sign.

\textit{Interpreting the Results}

The hypothesis from theory that net savers dampen inflation and net consumers fan it is consistent with the results in Table 1. The pattern is also largely consistent with the simple correlation pattern in Figure 1. It could be argued that the two groups with imprecisely estimated coefficients might indeed have ambiguous theoretical effects. If young people are to some extent credit constrained, as much empirical evidence indicate, we would expect less robust effects from this group. This age group is the one most affected by migration so endogeneity bias may also affect the results, even if we use the predetermined age variables, because the variables are serially correlated. The middle-aged have ambiguous effects already in the simple correlations in Figure 1. Although they are net savers on average there is microeconomic evidence that their saving change from real assets towards more liquid financial assets. Their propensity to consume out of wealth could therefore be increasing.\(^7\)

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\(^6\) See Nerlove (1971) and Nickell (1981). Nickell computes the bias on the lag coefficient to be of $O(1/T)$ where $T$ is the number of time series observations in the panel.

\(^7\) Ekman (1996) provides micro evidence from Swedish household surveys showing that around their fifties people shift their wealth portfolio from real assets, mainly real estate, to financial assets. He also finds that the
However, the increase in financial asset demand may also stimulate investment, thus opening another channel for increases in aggregate demand as this group grows.

But the evidence presented here does not admit any strong conclusions on the issue of the ambiguous groups. The correlation between middle-aged and young retirees is 0.62 and the correlation between young adults and these two age groups is also high but negative. Collinearity may therefore confound the individual age group estimates. Time-specific effects may also interfere with age effects, as the time profiles of the age shares are similar in most OECD countries. We can, therefore, only draw the conclusion that these age groups for unknown reasons do not have any statistically significant effects in our sample.

In Figure 2 we plot actual inflation and the age structure predicted inflation using predicted inflation lags. We display predictions generated from the models in column 1 and 4 of Table 1. The latter include fixed country-specific effects and random time-specific effects over the period of estimation, which explain the more ragged appearance of this curve as compared to the smooth curve predicted from the pooled model. We have assumed that country-specific effects remain constant in the future and that time-specific effects are zero when we generate the long-term projection up to year 2010 from the panel model. On basis of the UN demographic projections and our estimates most OECD propensity to consume out of financial wealth is substantially higher than the propensity to consume out of real wealth. This is in accordance with results in, e.g., Skinner (1989).
countries should experience low inflation pressure and even deflation in the next decade although Japan, for example, exhibit a forecast of increasing inflation pressure.

From the figures, it is clear that the age variables only catch long-term variation in the data. While age structure explains much of the trend rise in inflation in the 1970s quite a lot of the short-term variation remains to be explained by for example oil price shocks and monetary policy. Although the model with country and time effects explains more and does not miss the level as much (see Greece, for example), this is not necessarily a good thing for the forecasts, as will be seen below.

Sensitivity of the Results

As noted above, we abstained from including interest rates in the basic estimation model, because we were unable to find comparable interest data for the whole sample. To check whether or not the omitted variable bias due to this is important we use an incomplete and not quite comparable set of interest data. Most interest rates refer to 3-month treasury bonds and 5- (or 10-) year government bonds. To get a reasonable coverage of the country sample we also included other interest rates on similar instruments when these were available. The only country that dropped out of the sample completely was Greece, but for two more, long rates were not available. For some other countries, only very short series were available. In sum, we got 395 usable observations for short interest rates and a non-overlapping sample with 451 observations of long interest rates. In Table 2, the results are presented.

Our stylized theoretical model predicts that appropriate steering of interest rates may counteract the age-induced inflation impulses. Comparing Table 1 and 2 it is evident, however, that at least the sample of interest rates we have had available do not affect the age variable estimates very much. Our maintained hypothesis that monetary policy have not accommodated the inflationary effects of changing age distributions seems valid.

In the pooled estimations, both long and short rates get the expected signs. Attempts to recover the parameters for the expectation process (which is theoretically possible in the complete model) gave absurd results, however. It is, therefore, not appropriate to make a strict structural interpretation of the interest rate coefficients in terms of weights in the expectation formation. Both when short and long interest rates are included, the hypothesis that intercepts are the same over countries cannot be rejected. Thus, we have no country-specific effects. When we introduce time dummies, the age variables lose joint significance but the point estimates are still reasonable. In an unbalanced panel the error estimates are biased. The inference is, therefore, uncertain. Moreover, the interest rates are likely to be simultaneously determined with inflation. Adding to this other known sources of bias strong conclusions from Table 2 cannot be drawn, nor is it meaningful to proceed to a more sensitive random time effects model. However, the omission of interest rates in the benchmark regressions in Table 1 should not be decisive for the estimated age effect patterns.

It is clear from Figure 2 that the age structure explains trend variations, but hardly any of the high frequency variation. This is to be expected because the changes in age structure

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8 Most of these data emanate from a commercial database, EcoWin, but for Ireland, Portugal, and New Zealand, IMF statistics were used. We are indebted to Jan Haggström and Mats Kinnwall for the compilation of these data.
are comparatively slow. Thus, the projections should not be expected to indicate yearly variation in the inflation rate. Rather, they indicate medium-term changes in the underlying inflation pressure. This provides another explanation for the poor performance of a differenced model because the difference operator removes most of the low-frequency variation from the data.

We have also estimated the relation on individual country time series. The coefficient patterns vary considerably between countries. In this case we cannot control for any unknown background variables. The variability, thus, may reflect omitted variable bias, heterogeneity in appropriate models, or both. Multi-collinearity between age shares becomes a more serious problem when the degrees of freedom are severely limited resulting in highly inflated coefficients. However, in all cases, except Switzerland, the age coefficients are jointly significant at the 1 percent level, with a wide margin. The \( p \) value for joint significance in the case of Switzerland is 0.022, so even in this case the age variables are significant at conventional 5 percent levels.

If model parameters vary across countries, heterogeneity bias may arise in the panel estimates, see Pesaran et al. (1996). Under the assumption that the variation in coefficients is random, one may obtain consistent estimates of the average effect by simply averaging the coefficients estimated from individual time series. This mean-group estimator does show a similar pattern to Table 1 with a significant negative effect from the mature adults and a borderline significant positive effect from the young retirees. The coefficients are

### Table 2. Inflation (CPI growth) explained by logarithms of age shares and lagged inflation.

<table>
<thead>
<tr>
<th>Dep var.</th>
<th>CPI inflation</th>
<th>3-month rates time dum</th>
<th>5- or 10-year rates time dum</th>
</tr>
</thead>
<tbody>
<tr>
<td>Constant</td>
<td>4.32</td>
<td>−4.06</td>
<td>−1.12</td>
</tr>
<tr>
<td></td>
<td>(0.76)</td>
<td>(0.67)</td>
<td>(0.24)</td>
</tr>
<tr>
<td>Lagged inflation</td>
<td>0.76</td>
<td>0.69</td>
<td>0.78</td>
</tr>
<tr>
<td></td>
<td>(16.9)</td>
<td>(13.6)</td>
<td>(17.7)</td>
</tr>
<tr>
<td>Young adults, 15–29 years</td>
<td>3.09</td>
<td>−0.24</td>
<td>1.88</td>
</tr>
<tr>
<td></td>
<td>(1.72)</td>
<td>(0.13)</td>
<td>(1.13)</td>
</tr>
<tr>
<td>Mature adults, 30–49 years</td>
<td>−4.31</td>
<td>−4.01</td>
<td>−4.90</td>
</tr>
<tr>
<td></td>
<td>(2.18)</td>
<td>(1.84)</td>
<td>(2.98)</td>
</tr>
<tr>
<td>Middle-aged, 50–64 years</td>
<td>0.10</td>
<td>0.70</td>
<td>−1.94</td>
</tr>
<tr>
<td></td>
<td>(0.09)</td>
<td>(0.73)</td>
<td>(1.78)</td>
</tr>
<tr>
<td>Young retirees, 65–74 years</td>
<td>4.29</td>
<td>1.84</td>
<td>4.87</td>
</tr>
<tr>
<td></td>
<td>(3.19)</td>
<td>(1.68)</td>
<td>(3.92)</td>
</tr>
<tr>
<td>Old people, 75+ years</td>
<td>−2.56</td>
<td>−1.67</td>
<td>−2.60</td>
</tr>
<tr>
<td></td>
<td>(2.62)</td>
<td>(1.97)</td>
<td>(3.26)</td>
</tr>
<tr>
<td>Interest rate (( t + 1 ))</td>
<td>−0.18</td>
<td>−0.03</td>
<td>−0.32</td>
</tr>
<tr>
<td></td>
<td>(2.73)</td>
<td>(0.62)</td>
<td>(2.96)</td>
</tr>
<tr>
<td>Interest rate (( t ))</td>
<td>0.09</td>
<td>0.06</td>
<td>0.28</td>
</tr>
<tr>
<td></td>
<td>(1.46)</td>
<td>(1.01)</td>
<td>(2.98)</td>
</tr>
<tr>
<td>( R^2 )</td>
<td>0.72</td>
<td>0.80</td>
<td>0.71</td>
</tr>
<tr>
<td>( \chi^2 ) (5) sign age shares</td>
<td>31.0 [0.00]</td>
<td>8.95 [0.11]</td>
<td>40.0 [0.00]</td>
</tr>
<tr>
<td>( F(18/16) ) country effects</td>
<td>0.93 [0.54]</td>
<td>0.51 [0.95]</td>
<td>1.16 [0.30]</td>
</tr>
<tr>
<td>( F(31/32) ) time effects</td>
<td>5.88 [0.00]</td>
<td>—</td>
<td>7.38 [0.00]</td>
</tr>
</tbody>
</table>
blown up though with a factor of more than ten, due to the collinearity problems in the individual time series estimates.

**Forecasting Performance**

In Figure 3, we compare our forecasts beyond 1994 to the most recent OECD CPI data we could find, from May 1999, downloaded from the OECD homepage. Note that these forecasts differ from the age-predicted series in Figure 2 by using actual inflation in 1994 as the lag in the prediction for 1995, and only after that is then predicted lag used in the equation. These dynamic multi-step forecasts use only information available up to 1994. The annual average inflation is the solid line. The years 1995–1998 are demarcated by a break in the solid line. The performance of our projections is rather mixed. In some cases, such as Switzerland, we seem to catch the trend rather well, whereas Spain, for example, deviates considerably.

Evaluating these forecasts raises some non-trivial questions. First, because there is evidence of parameter heterogeneity, the estimated model is some kind of OECD average model rather than the best model for each country. We must, therefore, expect the model to fit better for the more representative economies, and the same will hold for the forecasts.

Second, projecting fixed country-specific effects into the future implies that the forecast is contingent on the constancy of unobserved factors in each country. Changes of monetary regime or exchange rate regimes are likely to render that assumption invalid. We know that several countries in fact have experienced such policy changes, e.g., in
connection to the introduction of the European Monetary Union and the convergence criteria that has imposed on many of the European countries.

Third, the time effects have been assumed zero, because the common intercept is accounted for and the expected value of random time effects will then be zero. But part of the time effects may be foreseeable, viz., the part that derives from the age group variation that is common to the whole OECD area. We made an attempt to use OECD averages of age shares in order to catch that, but ran into serious problems with multi-collinearity preventing us from getting reliable estimates of common age effects. Bad estimates of time effects may be worse than no estimates. We, therefore, abandoned that approach. However, in the pooled model, common age effects are affecting the estimates (even if the estimates are inconsistent).

Fourth, the standard loss functions, implied by measures of forecast performance like root mean square error or mean absolute error, penalize short-term errors around a trend. That is, it is presumed that high-frequency volatility should be well forecasted. However, if you want to forecast changes in the trend the appropriate loss function should measure the deviations from the trend movements. To do so require a measure of the trend. One could use concepts like the linear trend or the trend implied by a Hodrick–Prescott filter with some given parameters, or some other benchmark reference.9 However, there is no consensus on how to model the inflation trend and measures may be substantially different.

We do not have a generally applicable solution to these problems. The statistical error measures do, however, provide some guidance for evaluation. Therefore, in Table 3, we present the cross-section averages of standard measures of prediction errors reported for each year.

As can be seen, the pooled model consistently performs better than the panel model with fixed country effects and random time effects, and also outperforms the naive model in terms of mean error but not in terms of mean absolute error or root mean square error.

Table 3: Cross section averages of out-of-sample prediction errors. ME is the mean error, MAE the mean absolute error and RMSE the root mean squared error.

The mean error is a more appropriate performance criterion for a trend forecast than the mean absolute error or the mean squared error. After all, our intuitive understanding of a

<table>
<thead>
<tr>
<th></th>
<th>Naive prediction</th>
<th>Pooled prediction</th>
<th>Fixed country effects</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>ME</td>
<td>MAE</td>
<td>RMSE</td>
</tr>
<tr>
<td>1999</td>
<td>0.36</td>
<td>0.91</td>
<td>1.16</td>
</tr>
<tr>
<td>1996</td>
<td>−0.34</td>
<td>0.71</td>
<td>0.93</td>
</tr>
<tr>
<td>1997</td>
<td>−0.79</td>
<td>1.26</td>
<td>1.68</td>
</tr>
<tr>
<td>1998</td>
<td>−0.00</td>
<td>1.30</td>
<td>1.79</td>
</tr>
</tbody>
</table>

ME: mean error, MAE: mean absolute error, and RMSE: the root mean squared error.

---

Shimer (1998) actually uses a similar approach to evaluate a model of the relation unemployment and age structure.
trend is something around which a stationary error fluctuates, and the volatility around that
trend should cancel out. In that respect, the pooled-model forecasts perform best. Sys-
tematic bias can be suspected when the mean absolute error is close in magnitude to the
mean error. Looking at the ratio ME/MAE in Table 3, we see that the pooled model is less
likely to suffer from systematic forecast bias than either one of the other forecasts.

Repeating the mean error calculations within each country yields a more mixed picture.
For some countries, the naive prediction is best; for some, the pooled-model predictor is
best; and for still others, the best predictor is the model with fixed country effects and
random time effects.

It is, therefore, natural to assume that better forecasts could be obtained from individual
time series. When we estimate an age model for individual countries, however, the
forecasts have a tendency to “explode” in some direction. The much larger coefficients in
the time series estimates are unreasonably large and the result is unrealistically wild
swings in the forecasts.

Thus, the panel dimension does improve on the estimated coefficients by controlling
for some of the omitted variables. However, the forecasting performance is clearly better
for the pooled model, which we take to imply that the interference between age variables
and time effects is more of a problem when we try to estimate time-varying intercepts.
Estimation of country-specific and time-specific effects seems to add more to forecasting
bias than is won by explicit controls for differences in intercepts. Structural shifts over the
period may well be a factor behind this counter-intuitive fact. The cross-sectional variation
in the sample seems sufficient to prevent omitted variables from seriously biasing the age
coefficients.

To make use of the age structure correlations to improve forecasting for individual
countries obviously requires that more country- and time-specific information is added.
Probably this is best done by modeling each country separately, taking care to account for
structural shifts and interdependence between countries. Such an approach takes us,
however, far beyond the scope of this paper, and is likely to require more observations
than the 34 time-series observations that we had available.

**IV. Conclusions**

Our empirical study shows that age structure variations in the OECD countries are
strongly correlated with inflation. Net savers have a negative correlation with inflation,
whereas net consumers have a positive effect. These results are consistent with a model
relying on age effects on saving and factor price inertia to generate cumulative inflation
processes.

Although our results do not constitute a discriminating test of this particular model,
they do point to a promising method to improve forecasts of the inflation trend by taking
account of the correlation between inflation and age variables. Age distribution variables
are promising forecasting variables. They can reasonably be regarded as exogenous to the
inflation process and still forecasted with comparatively good precision by independent
demographic techniques.

The forecasts generated by the model do catch the general downward inflation trend in
the OECD area. Moreover, forecasts using age structure projections up to 2010 imply that
this downward trend will continue in most OECD countries. Our evaluation of the
forecasts against out-of-sample data yields the conclusion that forecasts from a pooled
model catches the average trend of inflation better than a model with country- and
time-specific effects. However, forecasting models need to incorporate more country-specific information if the results are to be reliable for individual countries.

It stands to reason that slightly different models should be used to forecast for individual countries. Time-series forecasting models, on the other hand, cannot account well for factors that are common across countries nor those that are common within countries. Thus, they are liable to omitted variables bias that is magnified by multicollinearity. The result is worse forecasts that tend to explode in one direction or the other. To our knowledge there has been no research on how to strike a balance between the advantages and disadvantages of panels versus time series in forecasting inflation.

Further research would be needed to take account of different interest rate regimes and monetary policies. There are also purely econometric questions to resolve. For example, it is not entirely clear how to handle the low-frequency co-variation between the age variables and the rate of inflation in a panel context. It is also not clear how standard forecast criteria should be interpreted when it is the trend rather than the variation around the trend that we wish to forecast. The age effects on inflation explain up to 5–6 percentage points of the variation in inflation rates. From a policy point of view, such effects should be taken very seriously. Inflation-fighting policies at a time when demographics predict deflation pressures, might be misguided and potentially dangerous. Equally misguided would Keynesian policies be when demographics predict inflation pressures. Keeping an eye on the age distribution and forecasting these pressures would help to stabilize stabilization policies.

We are grateful to Anders Vredin for an important suggestion. Helpful comments have also been received from seminar participants at the Central Bank of Sweden, at the International Symposium on Forecasting 1998, Edinburgh, at the Trade Union Institute for Economic Research (FIEF) in Stockholm, and at the Workshop on Age Effects 1999, Stockholm. Research funding from the Bank of Sweden Tercentenary Foundation is gratefully acknowledged.

References


