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Abstract:

We contrast the influence of demography and central bank independence on inflation. The recent demographic trends in developed countries are shown to weight more on inflation than central bank independence, while the contrary stands for the period from 1960 to 1979.

Keywords: Demography, Central Bank Independence, Inflation
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1. Introduction

Since the 1980s, in most of the OECD countries, the increase in the proportion of middle aged people and retirees has been accompanied by a reduction in inflation, and by monetary policies geared towards low inflation. This reveals differences in inflation aversion across generations[1], a fact that can easily be explained. For example, in France and the United States, retired people hold around 2.5 times more Treasury bonds than the working population, and they are about three times less indebted.[2] No wonder retirees' inflation aversion is higher than the young's one.

It is also remarkable that, at least since the last two decades, central bank independence has been considered as a quick fix against inflation, a relation that stands robustly for developed countries (see e.g. Brumm, 2002).

However, if central bank independence is a solution to a society's inflationary bias, and if older generations have higher degrees of inflation aversion, then the sheer fact of ageing should strongly weight on inflation. In which case, the negative relationship between inflation and central bank independence may no longer stand, once one considers demographic trends in the estimated equation.

Most of the empirical studies revealing the negative relation between central bank independence and inflation have been conducted in the 1990s, considering the 1960-1990 period. Yet, this period includes a first phase when baby-boomers have started to consume and contract debts and a second phase, when they have started to fret about the protection of their savings. Hence, the coefficients on central bank independence may have captured (and in fact hidden) changes in populations' preferences.

The question we here consider is thus: is the impact of central bank independence different in ageing societies? We test for the respective influences of central bank independence and demography on inflation for developed countries, considered under the 1960-1994 period. This period is the one for which central bank independence indexes have been shown to be the most influential (see notably Cukierman, 1992), permitting a direct comparison of our results with the literature. The conclusion however contains lessons to be drawn for the following periods.

The following section deals with the data sources and methods. The third section presents the results, while the last concludes.

[1] Shiller (1997) even shows that intergenerational differences in inflation aversion are more important than international ones.

[2] Data for France available from the INSEE website (http://www.insee.fr/fr/ppp/ir/accueil.asp?page=transm04/doc/documentation.htm); data for the US from the 2004 Survey of Consumer Finances. See Bucks et al. (2006).

2. Data and method

We estimate a reduced form model close to Lindh and Malmbergh (1998), where inflation is explained by five age group shares of the population: 15-29 years old, 30-49 years, 50-64 years, 65-74 years and 75 years and above[3].

We add to this equation a Central Bank Independence (CBI) index in order to distinguish between the influence of the latter and the part played by the elderly in the disinflation process. We also include a quadratic term for the elderly, to account for potential non-linearities in the relationship.

Eventually, we add two control variables traditionally used in the literature, namely an output variable and the openness rate of the economy. For an increase in the output gap is expected to create additional inflationary pressures. And, conversely, the extra competitive pressure related to openness has been shown to exert a negative influence on prices.

The estimated equation is:

$$infl_{it} = \beta_0 + \sum_{age=15\text{-}29}^{75+} \beta_{age}(\text{age share})_{it} + \beta_{quadr}(\text{share\_75plus})_{it}^{\ 2} + \beta_{cbi}CBI_i + \beta_g g_{it} + \beta_{or}or_{it} + \varepsilon_{it} \qquad (1)$$

where indexes *i* and *t* refer respectively to country and year, g_t is the output gap, computed as the difference between real output and potential output, divided by potential output. The latter has been generated using a Hodrick-Prescott filter, with a smoothing parameter equal to 100, as recommended by Hodrick and Prescott (1997) for annual data. or_{it} is the openness rate, defined as the ratio of the sum of exports and imports to GDP. The use of country-fixed effects is *de facto* impossible, since we use Cukierman's (1992) time invariant CBI index. He computed this index from the central bank charters of the 1980s and covers independence for the whole period. The use of the same index for the whole period is usual in the literature since the case for independence has been made (Alesina and Summers, 1993). For comparison purposes, we will also run regressions excluding the CBI index, in order to point out the exact influence of age structure on inflation:

$$infl_{it} = \beta_0 + \sum_{age=15\text{-}29}^{75+} \beta_{age}(\text{age share})_{it} + \beta_{quadr}(\text{share\_75plus})_{it}^{\ 2} + \beta_g g_{it} + \beta_{or}or_{it} + \varepsilon_{it} \qquad (2)$$

We use annual data of consumer price index (CPI), of real GDP, exports and imports for 20 OECD countries[4], all from the International Monetary Fund (*International Financial Statistics*). Age distribution data are from the United Nations population report (2004). Since age data refer to end-of-the year estimates, the inflation variable is constructed as inflation in year *t* equal to $\log(CPI_{t+1}/CPI_t)$.

Estimates are performed over the 1960-1994 period, because it covers the ones used in most of the literature on the effects of central bank independence (see Hayo and Hefeker, 2002). To test for the changing relationship between inflation and age structure/CBI index, we divide our sample period in three sub-periods: 1960-1972 (*significant but controlled inflation*), 1973-

[3] Motivations for the suppression of the 0-14 year old share are provided by Lindh and Malmberg (1998).

[4] Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, New Zealand, Norway, Spain, Portugal, Sweden, Switzerland, United Kingdom, United States.

1979 (*stagflation*) and 1980-1994 (*sustained disinflation*). Besides their obvious economic foundations, Hausman/Wald tests[5] support their relevance[6].

3. Results

Preliminary investigations pointed out two problems. Firstly, Greece and Portugal appear as consistent outliers according to current mean/standard deviation criteria for the 1980-1994 period. Interestingly, no significant presence of outliers could be highlighted by the same criteria or the Hadi (1994) test for the two other sub-periods, or for the whole 1960-1994 period, apart from one or two observations. Secondly, a potential endogeneity problem has to be considered. Indeed, while both Hausman and Nakamura-Nakamura tests[7] accept the hypothesis of exogeneity for age shares and CBI index, they support strong evidence of endogeneity for the output gap and the openness rate. Instrumenting the output gap and the openness rate with their own lags and estimating by two stage least square (2SLS) provided an adequate solution[8]. Depending of the considered sub-periods, the number of lags used as instruments turned out to be two or three[9]. Finally, residuals have been corrected for heteroskedasticity using the White estimator of variance, and for autocorrelation using Bartlett kernel (Newey-West)-based estimations.

The results appear in Table 1. Interestingly, they differ depending on the period / sub-period considered, revealing a varying impact of central bank independence and of age cohorts on inflation. For each period and sub-period, we report estimates for equations (1) and (2). We first compare the results for the whole period to those for the sub-periods, and then focus on the switches revealed by sub-periods analyses.

----- Insert Table 1 here -----

First column in Table 1 shows the results for the whole 1960-1994 period. The set of age structure variables and the central bank independence index are significant with signs conform to economic intuition. The two control variables are also almost always significant and rightly signed, i.e. positive for the output gap and negative for the openness rate.

Furthermore, the influence of the different age shares is qualitatively identical over the global period and the sub-periods, except for the 15-29 category. For this class, the influence on inflation turns negative in the 1973-1979 sub-periods. Though young adults are likely to consume more than save, in these years their behavior has probably been credit-constrained and/or, for those already in debts, their attention has probably been brought to inflation[10]. It is also worth emphasizing that these outcomes are qualitatively very similar to the ones of Lindh

[5] More details upon these tests available from the authors on request.
[6] An observation on non-stationarity concerns. The current panel unit root tests (Levin and Lin, 2002; Im *et al*., 2003; Maddala and Wu, 1999) often disagree on the persistence of the data, so that it would not be clear whether the variables in (1)/(2) are indeed integrated or co-integrated, whatever Pedroni's (2001) panel cointegration test would tell. Besides, the sub-periods we are considering are too short (5 to 14 years) for imposing any meaningful restrictions. We can therefore perform estimations in levels, since we are *de facto* interested in the short-term relationships between inflation and right-hand side variables.
[7] Results are available from the authors upon request.
[8] Moreover, a relatively small number of lags for the instruments minimizes the potential small sample bias that may arise when too many overidentifying restrictions are imposed.
[9] Hansen overidentification test did not reject our sets of instruments. Complete results available upon request.
[10] This phenomenon probably contributes to explain their subsequent fear of inflation.

and Malmbergh (1998), except for the share "75 and more". Though correctly signed, the share of the elderly does not significantly impact inflation on the whole period, whereas the CBI index displays the expected significant negative impact.

Turning to sub-periods analysis (columns (2) to (4)), it is striking to see that in the first two sub-periods, some age structure variables and central bank independence are both strongly significant. Everything thus happens, in the first two sub-periods, as if societies had to protect themselves from inflationary pressures. This result stands with Farvaque (2002), that inflation-adverse-societies build institutions fitting their feelings. Hence, during the period under review, central bank independence clearly appears as a rational institutional choice, adopted by inflation adverse societies confronted with rising inflationary pressures.

However, demography appears to dominate the influence of institutional frameworks over the 1980-1994 sub-period. Central bank independence is now insignificant, while the weight of the retirees on inflation is strongly significant and negatively signed. These results show that independence is no longer so strongly needed when people have matured enough to resist the siren's song of inflation.

Concerning the non-linear impact of the share of elderly, as shown in Table 1, the case for non-linearity is not overwhelming. The coefficient of the squared share of the elderly is in most cases negative but non-significant, and when significant (for the last sub-period, 1980-1994), it is only at the 10 % level.

To check the robustness of these results, we undertake several sensitivity checks. First, including over the last sub-period dummy variables for European Monetary System membership and the adoption of inflation targeting frameworks for some countries delivered similar results. Second, using GDP growth instead of output gap brings quantitative identical results, while sometimes affecting very marginally the significance. More interestingly, controlling for intra-group correlation[11] leads to only one major change in the results: for the whole period 1960-1994, the coefficient on CBI index becomes clearly non-significant (t-stat equal to -0.87), confirming the role of age as a better and stronger predictor of inflation.

4. Conclusion

As shown by Forder (2005), the belief in the benefits of central bank independence has spread quickly, and policymakers and politicians have endorsed the idea. In the years following the sample period under study, central banks have even become more independent.

As the countries in our sample have seen their people becoming older, they may have transformed their institutions beyond what is now required. However, institutions should exist to serve not just the current, but also the future generations. And, should the latter become more numerous, younger, and more turbulent, they would appreciate the bequest of central bank independence.

[11] Intra-group correlation means that the error terms are correlated within groups, but not across groups.

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Table 1: ***Inflation explained by age shares, CBI, output gap and openness rate***

Dep. var. infl.	1960-1994 2SLS, HAC residuals with CBI	1960-1994 2SLS, HAC residuals no CBI	1960-1972 2SLS, HAC residuals with CBI	1960-1972 2SLS, HAC residuals no CBI	1973-1979 2SLS, HAC residuals with CBI	1973-1979 2SLS, HAC residuals no CBI	1980-1994 2SLS, HAC residuals with CBI	1980-1994 2SLS, HAC residuals no CBI
Constant	0,056	0,056	-0,020	0,118	0,543	0,661	0,423	0,417
	1,120	*1,120*	*-0,220*	*1,320*	*3,310*	*3,930*	*3,770*	4,000
shr_15_29	0,472	0,461	0,449	0,277	-0,960	-1,116	-0,223	-0,206
	4,260	*4,170*	*3,190*	*1,990*	*-3,490*	*-4,260*	*-1,040*	*-1,070*
shr_30_49	-0,606	-0,605	-0,404	-0,411	-0,836	-0,950	-0,972	-0,962
	-7,190	*-7,090*	*-2,520*	*-2,520*	*-3,200*	*-3,760*	*-5,960*	*-6,340*
shr_50_64	-0,075	-0,064	-0,433	-0,319	-0,125	0,292	0,014	0,011
	-0,440	*-0,390*	*-2,730*	*-2,010*	*-0,290*	*0,560*	*0,090*	*0,070*
shr_65_74	1,841	1,940	2,075	1,698	0,768	0,587	1,463	1,437
	5,820	*6,040*	*3,700*	*2,850*	*0,820*	*0,550*	*4,270*	*4,330*
shr_75plus	**-0,655**	**-1,255**	**3,086**	**-3,341**	**3,572**	**-2,112**	**-4,259**	**-4,171**
	-0,690	***-1,330***	***0,980***	***-1,320***	***0,840***	***-0,500***	***-2,830***	***-2,860***
square75	**-5,364**	**-0,548**	**-67,600**	**21,039**	**-69,571**	**-11,489**	**23,476**	**23,035**
	-0,650	***-0,070***	***-1,520***	***0,600***	***-1,540***	***-0,240***	***1,890***	***1,880***
cuk8094	**-0,027**	**-**	**-0,055**	**-**	**-0,075**	**-**	**0,004**	**-**
	-2,100	-	***-3,610***	-	***-2,900***	-	***0,350***	-
output gap	0,640	0,640	0,413	0,415	0,773	0,755	0,357	0,357
	5,790	*5,800*	*3,680*	*3,830*	*3,480*	*3,320*	*5,200*	*5,200*
or	-0,023	-0,023	-0,001	0,003	-0,066	-0,060	-0,016	-0,016
	-3,450	*-3,360*	*-0,120*	*0,370*	*-4,650*	*-4,150*	*-2,430*	*-2,410*
N	630	630	164	164	134	134	283	283
Centered R^2	0,363	0.356	0,345	0.278	0,538	0,500	0,536	0,535

t-statistics in italic. 2SLS indicates that all regressions have been performed using two-stage least squares. HAC means that t-statistics have been computed from White/Newey_West estimates of heteroskedasticity and autocorrelation consistent standard errors.