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The Political Economy of Public Investment when Population is Aging – A Panel Cointegration Analysis

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The Political Economy of Public Investment when Population is Aging – A Panel Cointegration Analysis

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Philipp Jäger and Torsten Schmidt¹

The Political Economy of Public Investment when Population is Aging – A Panel Cointegration Analysis

Abstract

Time preferences vary by age. Notably, according to experimental studies, senior citizens tend to discount future payoffs more heavily than working-age individuals. Based on these findings, we hypothesize that demographic change has contributed to the cut-back in government-financed investment that many advanced economies experienced over the last four decades. We demonstrate for a panel of 13 OECD countries between 1971 and 2007 that the share of elderly voters and public investment rates are cointegrated, indicating a long-run relationship between them. Estimating this cointegration relationship via pooled dynamic OLS (D-OLS) and fully modified OLS (FM-OLS) we find a negative and significant effect of population aging on public investment. Moreover, the estimation of an error correction model reveals long-run Granger causality running exclusively from aging to investment. Our results are robust to the inclusion of additional control variables typically considered in literature on the determinants of public investment.

JEL Classification: H54, D72, J11, J14, D91

Keywords: Public investment; population aging; panel cointegration

May 2015

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1 Introduction

Public investment as a share of GDP has been constantly declining in most advanced economies for more than four decades. This seems puzzling since the overall empirical evidence indicates that the economic returns to public capital are markedly positive (for an overview see Bom and Ligthart, 2014). Based on these high returns some economists (e.g. IMF 2014) have recently advocated additional public investment spending to foster economic growth. However, we argue in this paper that the growing share of elderly voters has contributed to the decline in public investment rates. Thus, raising public investment levels might become increasingly difficult in greying democracies.

The argument that demographic change will shift voting power towards older generations and therefore affects the composition of government spending, dates at least back to Preston (1984). In the last decades, empirical studies have demonstrated that population aging in fact influences government spending components including social welfare, health and education (Harris et al., 2001; Krieger and Ruhose, 2013; Poterba, 1997; Sanz and Velázquez, 2007; Shelton, 2008; Tepe and Vanhuysse, 2009). Furthermore, evidence based on survey data reveals age-specific preferences regarding welfare and educational spending (Boeri et al., 2001; Bonoli and Häusermann, 2009; Cattaneo and Wolter, 2009; Sørensen, 2013). However, existing studies typically focus on different functions of government spending, but do not distinguish between consumption and investment. This paper aims to fill this gap and investigates the effect of aging on public investment as a whole regardless of the specific beneficiary or function of investment.

So far, the most common explanation for the decline in public investment rests on fiscal pressures (Vuchelen and Caekelbergh, 2010). The few existing studies¹ on the determinants of public investment usually find public investment to be negatively correlated with the budget deficit or the debt level “reflecting the political reality that it is easier to cut-back or postpone investment spending than it is to cut current expenditures” (Oxley and Martin, 1991, p. 161). Based on this line of reasoning, public investment has been declining because it’s a flexible budget item and voters seem to be rather insensitive to cuts in times of fiscal pressure, given the limited visibility and more diffuse character of investment spending.

However, public investment did not recover once fiscal stress eased in the 1990s (Vuchelen and Caekelbergh, 2010) rendering the fiscal pressure narrative incomplete. Privatization and the emergence of Public Private Partnerships (PPP) also offer only limited explanatory power for the constant decline in government-financed investment (Heinemann, 2006; Mehrotra and Väililä, 2006). The definition of public investment provides the main reason for the negligible role of privatization. Public investment is commonly measured by gross fixed capital formation of the general government, which does not include investments made by state-owned enterprises that are commercially run (Gonzalez Alegre et al., 2008). Therefore, privatization of these companies does not affect public investment.² PPP on the other hand is a rather new phenomenon and still accounts for less than one tenth of public investment spending in developed countries (IMF, 2014). Hence, the factors behind the steady decline in public investment are yet not fully understood.

In this paper, we propose an additional explanation for the declining share of public investment that is related to ongoing demographic change. The underlying theory rests upon a simple political economy mechanism following the voter group decision model developed by Craig and Inman (1986). We stipulate that the elderly, on average, prefer a smaller investment to GDP ratio compared to working-age individuals. This claim is based on experimental studies investigating age-specific time preferences, which find the elderly to discount future payoffs more heavily than working-age people. Since greying societies are characterized by a steadily growing share of elderly

¹These studies include Bacchiocchi et al. (2011), Haan et al. (1996), Heinemann (2006), Keman, (2010), Mehrotra and Väililä (2006), Sturm and Haan (1998) as well as Turrini, (2004).

²Enterprises are considered commercial, if their sales revenues cover more than 50% of the actual production costs. Some public utilities have been reclassified from the general government to the private sector in the mid '90s (Balassone and Franco, 2000), however the findings of Heinemann (2006) suggest a minor impact of privatization overall.

people, the voting power of senior citizens rises and thus their preferences should become increasingly influential. Therefore, we hypothesize that population aging reduces the overall demand for public investment, and hence the availability of public investment as a measure to enhance economic growth during the demographic transition.

In order to analyze our hypothesis empirically and accounting for non-stationarity in the data, we employ panel cointegration and error correction techniques using data from 13 OECD countries for the period from 1971 to 2007. The cointegration tests indicate a long-run relationship between public investment and population aging. Subsequently, we estimate this cointegration relationship applying pooled dynamic OLS (D-OLS) and fully modified OLS (FM-OLS). Consistent with our hypothesis we find a negative relationship between public investment and population aging. The estimated coefficient is statistically as well as economically significant. Moreover, results from an error correction model suggest long-run Granger causality running exclusively from aging to public investment. The use of a reduced-form model is justified, because cointegration relationships are invariant to model extensions (Herzer et al., 2012; Lütkepohl, 2007). Adding control variables typically considered in the literature indeed leaves our results broadly unchanged. Furthermore, our conclusion neither depends on the choice of the estimator nor on the lead and lag specification and is robust to the exclusion of Japan and Germany, the only two sample countries where population is expected to decline in the close future.

The outline of the paper is as follows: In the next section we introduce and discuss the theoretical foundation of the link between demographic change and the decline in public investment using a simple group decision model. In section 3 we introduce the data as well as the empirical approach and present our results in section 4. Section 5 concludes.

2 Background and Hypothesis

Typically, public investment affects the individual's well-being mostly in the long-run. Constructing durable public goods, e.g. streets and buildings, is time-consuming, but the economic lifetime of these assets is most likely rather long. Hence, time preferences matter for the assessment of public investment projects. Those people who discount the future more heavily will be less prone to appreciate additional public investment spending, holding everything else constant.

Empirical evidence derived from experimental studies shows that time preferences vary with age (Harrison et al., 2002; Read and Read, 2004). In these experiments, participants are forced to make a range of simple inter-temporal choices. Most often, the two available options are, either receiving a fixed amount of money sooner or more of it later. The responses to different sets of payout options enable the authors to infer individual discount rates. Based on such an experiment involving 268 Danes between 19 and 75 years of age, Harrison et al. (2002) find that retired individuals discount future payoffs more than non-retired people. Similarly, the study conducted by Read and Read (2004), including 123 British people between 19 and 89 years, suggests higher annual discount rates of elderly individuals. Interestingly, Read and Read (2004) find the differences in annual discount rates between senior citizens and working-age people especially pronounced if the second payoff option lies in the distant future (10 years after the experiment). A possible explanation for these results is the decline in survival probability associated with rising age (Sozou and Seymour, 2003). Based on this argument, elderly people strongly prefer immediate consumption, since future consumption prospects are uncertain. A related argument states that the expected utility of future consumption decreases with age since people anticipate a deterioration of their mental and physical capacity (Trostel and Taylor, 2001).

The size and composition of public spending is determined within the political process. Voter preferences should crucially affect the structure of public spending in democratically organized societies, since politicians are likely to increase their election probability by responding to the perceived interest of the majority of their citizens. In

this context, the preferences of the median voter often serve as a reference point. However, we depart from the median voter model and thus from using median age as main explanatory variable, because the demand for public investment might not be strictly declining with age. In fact, the results of Read and Read (2004) indicate a U-shape relationship between age and the discount rate with elderly individuals discounting the most.

Therefore, we instead motivate our empirical analysis using the voter group decision model developed by Craig and Inman (1986). The authors model the final outcome of political decision making as the weighted average of the outcome favored by each voter group. Voter groups can be defined along different dimensions. Craig and Inman (1986) split voters based on their income. Following our hypothesis we assume age to be the crucial socio-economic cleavage and distinguish two groups: elderly people and working-age individuals. In our case the model looks like the following:

$$I^* = \pi^{work} * I^{work} + \pi^{old} * I^{old} \quad (1)$$

where I^* is the observed level of public investment determined in the political process. I^{work} and I^{old} indicate the demand for public investment for working-age and elderly individuals respectively. π^{work} and π^{old} specify the political strength of each group, whereas $\pi^{work} + \pi^{old} = 1$. Similar to Sanz and Velázquez (2007) we assume that the political weight of a demographic group is based on the relative size of a group to the total voting population.³ As argued above, we hypothesize that elderly people demand less public investment than working-age individuals ($I^{old} < I^{work}$) based on different time preferences. Thus, holding everything else constant, an increase in π^{old} results in a decline of I^* in our model.

Two limitations apply within this theoretical framework. First, population aging partially follows from a continuous rise in life expectancy. Therefore, today's citizens are likely to have different time preferences compared to their same-age counterparts four decades ago. Thus, demographic change can also lead to an increase of I^{work} and I^{old} . Moreover, elderly's preferences might not be entirely driven by self-interest, but by altruistic sentiments towards their offspring or the society as a whole. Hence, the question whether a growing share of elderly voters has contributed to the decline in public investment has to be solved empirically

However, even if the data is consistent with our theoretical model, our results could be driven by a potentially negative relationship between population aging and public investment returns. The underlying argument is the following: besides the increase in life expectancy, the ongoing demographic transition is characterized by a decrease in fertility rates. Hence, population aging could also be associated with population shrinkage and thus less need for public capital. The statistical relationship between the share of elderly voters and public investment might therefore be spurious, since decreasing investment returns and not the shift in voting power would be responsible for the decline in public investment rates.

We think diminishing investment returns following from a fall in resident population are only a minor problem for our analysis. First of all, within our sample period, population increases almost steadily over time. According to UN projections, this trend is expected to continue at least until 2025 for all sample countries except for Germany and Japan. Hence, population shrinkage is no likely scenario for most sample countries in the medium term. Similarly, absolute employment as an additional proxy for investment returns kept increasing over time in all countries except for Japan since the late 90s. We account for the distinctiveness of Japan and Germany by excluding them as a robustness check. Moreover, we also control for population and population growth in an alternative specification.

³Sanz and Velázquez (2007) relied on simple population shares instead.

3 Empirical strategy and data

3.1 Econometric model

In order to analyze the nexus between public investment and population aging we apply the following model:

$$\ln(\text{public investment})_{it} = a_i + b_i \text{trend} + \beta \ln(\text{elderly voter share})_{it} + \varepsilon_{it} \quad (2)$$

where i and t denote the country and time dimension. Following the standard in the literature, *public investment* is operationalized by gross fixed capital formation of the general government as a share of GDP. The *elderly voter share* is defined as proportion of the population aged 65 and older to total voting age population (population older than 20) and captures political strength of the elderly (π^{old} from equation (1)). We choose 65 as threshold age because it has been the most frequent legal pension age for males in our sample. A similar variable, using 70 instead of 65 as threshold age, has recently been employed by Krieger and Ruhose (2013) to study the impact of population aging on family benefits and educational expenditures. Related empirical studies have operationalized population aging using old-age dependency ratios (Shelton, 2008; Tepe and Vanhuyse, 2009) or the share of elderly to total population (Harris et al., 2001; Poterba, 1997; Sanz and Velázquez, 2007). However, we think that these indicators are less appropriate for our purposes, because they capture relative voting power of elderly citizens only indirectly.

In addition, we include country fixed effects a_i to capture country-specific factors that are rather stable over time, e.g., geography and the political system. As common in the applied cointegration literature (Herwartz and Theilen, 2014; Herzer and Nunnenkamp, 2013; Malinen, 2012), we also add deterministic country-specific time trends to account for only imperfectly observable factors that change over time, e.g. the remaining life expectancy of citizens, the degree of privatization and private public partnerships, the price level of public investment and the stock of public capital which might affect investment returns. Including time trends is also in line with previous work on the determinants of public investment (Bacchiocchi et al., 2011; Mehrotra and Väililä, 2006).

The estimated relationship between population aging and public investment might be spurious if both variables are non-stationary and not cointegrated. Under these assumptions the error term ε_{it} becomes non-stationary potentially resulting in distorted t-statistics that falsely indicate statistical significance (Entorf, 1997; Granger and Newbold, 1974; Phillips, 1986). If *public investment* and the *elderly voter share* by contrast are cointegrated (which means both variables are non-stationary but a linear combination of them is stationary) some favorable properties apply such as the robustness of panel cointegration estimators against omitted variables and reverse causation (Herzer et al., 2012).

As shown by Phillips and Durlauf (1986) as well as Engle and Granger (1987) time series cointegration relationships can be estimated consistently even though the regressor and the error term are correlated. While, this result does not carry over to the estimation of panel data using pooled OLS, two panel cointegration estimators, the D-OLS estimator suggested by Kao and Chiang (2001) as well as Mark and Sul (2003) and the FM-OLS estimator developed by Phillips and Moon (1999), have been demonstrated to provide consistent and normally distributed coefficient estimates under these circumstances (Choi, 2006; Kao and Chiang, 2001). Therefore, under the assumption of cointegration, we obtain a consistent estimate of β using D-OLS or FM-OLS even if the *elderly voter share* is endogenous. This is in contrast to standard regression analysis where a correlation between the error term and the explanatory variables results in inconsistent parameter estimates. Moreover, with increasing sample size, parameter estimates of cointegration relationships converge faster to their true values compared to estimates based on stationary variables (Stock, 1987). This so called super-consistency property should result in a rather precise estimate of β .

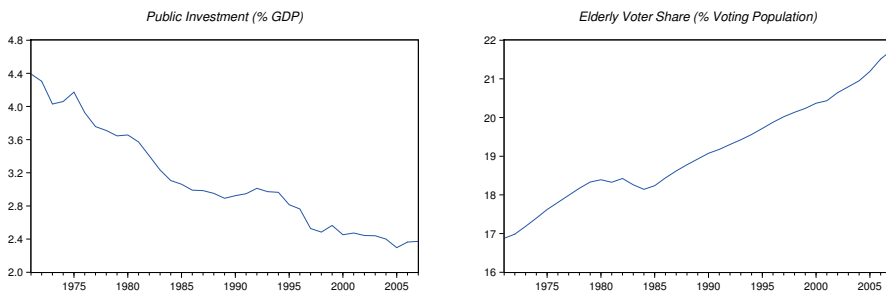
3.2 Data

In this study we use data for a balanced panel of 13 OECD countries⁴ for the period of 1971 to 2007 yielding 481 observations. The data selection is based on data availability as well as three additional deliberations. First of all, we only consider democracies because political decision making differs from authoritarian societies. Secondly, to keep our sample homogenous we focus on aging societies only.⁵ Last but not least, we exclude the period after 2007 to avoid distortions resulting from the deep drop in GDP as well as massive fiscal stimulus packages following the financial crisis. Although the sample mean of public investment has increased substantially in 2009, it returned to pre-crisis levels in 2011. Therefore, we believe the period between 2008 and 2011 is exceptional.

Data on gross fixed capital formation of the general government as well as on GDP stems from the AMECO database of the European Commission. Both variables are measured in nominal terms. According to the 1993 System of National Accounts (SNA 93), which is the relevant version for our sample, capital formation is defined as the net acquisition of assets such as equipment or buildings that are used in production for more than one year. Maintenance is classified as consumption unless it improves the performance, increases the capacity or prolongs the expected working lives of these assets. In contrast to the most recent SNA version from 2008, public investment spending in our study does not include expenditures on research and development as well as on arms. Public investment data for the US in the SNA 93 version stems from the OECD. The aging variable has been constructed using OECD data, too.

Based on the intuition of the voter group decision model, we expect public investment to vary with the share of elderly people to the voting population. Figure 1 plots the average public investment to GDP ratio and the average elderly voter share for the 13 OECD countries in our sample. It clearly indicates a negative correlation. The average increase in the elderly voter share by approximately 5 percentage points presumably underestimates the true shift in voting power, because elderly people tend to vote more frequently than working-age individuals and this age-specific participation gap seems to have increased over time.⁶ The country-specific figures, provided in the appendix (Figure 2, Figure 3), demonstrate that the opposing trends in public investment rates and elderly voting shares can be observed in all sample countries. Summary statistics are presented in Table 8 in the appendix.

Figure 1: **Cross-sectional average of Public Investment and the Elderly Voter Share, 1971-2007**



⁴ Austria, Belgium, Canada, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Sweden, United Kingdom, United States.

⁵In practice, this resulted in the exclusion of Ireland. In fact, Ireland combines growing public investment with a decreasing share of elderly voters, therefore rather supports our argument. However, since the relationship between age and the discount factor might be non-monotonic (Read and Read, 2004) and hence a decrease in the elderly voter share might not increase public investment in the same way as an increase in the elderly voter share decreases public investment, we decided to exclude Ireland.

⁶We calculated age-specific voter turnout rates in national election for all European countries in our sample except for Italy based on the four rounds of the European Social Survey between 2002 and 2008 (ESS, 2002). On average, elderly voter turnout rates exceed working-age turnout rates by approximately 6 percentage points. Moreover, the gap increases over time, which is in line with the findings by Smets (2010). Similar results have been reported for the US (File, 2014), Canada (Barnes and Virgint, 2010) and Japan (Takao, 2009). Based on the ESS dataset, elderly voter turnout is higher in all European countries considered except for Belgium where voting is mandatory.

The existing empirical evidence suggests that population aging and the composition of public spending are related. Specifically, aging has been shown to affect spending on education, health, social security and defense. This is important because public investment also includes spending on those items. Table 1 demonstrates that in the last year of our sample (2007), on average, one third of public capital formation was devoted to economic affairs which mainly includes transportation infrastructure. Theoretically, it is not obvious which age group benefits most from transportation infrastructure. In contrast, the 15.7 percent of public investment spent on education can be assumed to be less beneficial for the elderly (Harris et al., 2001; Inman, 1978; Poterba, 1997), whereas the 14.4 percent attributed to health, defense and social welfare expenditure has been shown to be favored by senior citizens (Sanz and Velázquez, 2007). Hence, the composition of public investment in 2007 does not clearly favor one specific age group. Data on investment for specific functions is not consistently available before 1995 for most sample countries with the exception of the US. However, the composition of public investment has remained more or less stable since then and can therefore be assumed to be rather constant. This argument is supported by the composition of public investment in the US, which changed only modestly since 1970.

Table 1: **Public Investment by function, cross-sectional average, 2007**

Function	Investment share
Economic affairs	33.95%
Education	15.71%
General public services	11.11%
Recreation, culture and religion	7.46%
Housing and community amenities	7.42%
Health	6.78%
Environment protection	5.71%
Public order and safety	4.22%
Social protection	3.84%
Defence	3.81%

Sample except Canada. US data: Investment on "Environment protection" is included in "Housing and community amenities".

4 Empirical results

Since the variable properties greatly affect the estimation strategy we conduct panel unit root and cointegration tests prior to the estimation.

4.1 Panel unit root tests

In order to achieve robust results and taking the relatively low power of unit root tests into account, we analyze the time series properties of gross fixed capital formation and the elderly voter share by employing the following three slightly different "first generation" panel unit root tests (PURT): the LLC-PURT proposed by Levin et al. (2002), the IPS-PURT suggested by Im et al. (2003) and the Fisher-augmented Dickey Fuller-PURT (Fisher-ADF) following Maddala and Wu (1999). All employed tests are derived from augmented Dickey Fuller regressions, thus a rejection of the null hypothesis suggests stationarity at least for a fraction of the panel. We choose these type of tests, because they generally outperform panel stationary tests (Hlouskova and Wagner, 2006).

However, first generation PURTs may be misleading under cross-sectional dependence. We address this issue by subtracting cross-sectional averages for each observation, a procedure suggested by Levin et al. (2002) and often applied in empirical research (e.g. Canning and Pedroni, 2008; Schmidt and Vosen, 2013).⁷ Furthermore,

⁷Other approaches proposed in the literature include the estimation of common factors using principle component analysis (Bai and Ng, 2004). Given that the precision based on few cross-sections (less than 20) is limited (Sul, 2009) we refrain from using PCA.

we make use of the cross-sectionally augmented IPS-PURT (CIPS) developed by Pesaran (2007) which accounts for the potential cross-sectional dependence by adding cross-section averages to the ADF regressions. Since both variables seem to be heavily trending, we include a trend in the level equations.

The PURTs presented in Table 2 suggest that *public investment* and the *elderly voter share* are both non-stationary in log-levels, but stationary in first differences. Therefore, we conclude *public investment* and the *elderly voter share* to be integrated of order 1 and proceed by testing for cointegration.

Table 2: Panel unit root tests

Variables	Linear trend	Unadjusted data				Demeaned data		
		LLC	IPS	Fisher-ADF	CIPS	LLC	IPS	Fisher-ADF
<i>Log-Levels</i>								
Elderly Voter Share	yes	-0.19	-0.64	43.97*	-1.78	-1.34	1.17	25.94
Public Investment	yes	-1.89*	-0.50	33.83	-2.55	-1.53	-0.57	32.73
<i>First differences</i>								
Elderly Voter Share	no	-5.04**	-7.87**	121.90**	-2.64**	-5.89**	-7.83**	119.95**
Public Investment	no	-18.70**	-17.26**	270.26**	-5.39**	-19.70**	-18.17**	285.37**

Note: All variables in logs. All ADF regressions include a constant. Lags have been selected using the Schwartz criterion starting from 5 lags. ** (*) Indicates significance at the 1% (5%) level.

4.2 Panel cointegration tests

To test for a common trend of *public investment* and the *elderly voter share*, we apply two panel cointegration tests, one residual-based test developed by Pedroni (2004, 1999) as well as the panel version of a system equation test proposed by Johansen (1991). The idea of the Pedroni test is to estimate equation (2) and test whether the residuals $\hat{\epsilon}_t$ are stationary. Pedroni (2004, 1999) suggests several test-statistics. We employ Pedroni's panel cointegration test based on the group and panel ADF statistics, because they outperform all other panel cointegration tests considered in Wagner and Hlouskova (2010).

To test the robustness of our results we also apply the Fisher-type panel version of the Johansen cointegration test. Since the test statistics of residual-based cointegration tests are affected by normalization, exchanging the dependent and independent variable might alter the conclusion if the sample is small. The system approach developed by Johansen treats all variables as endogenous, thus does not suffer from this problem. Again, we demean the data to mitigate the impact of cross-sectional dependence. Table 3 shows that the null hypothesis of no cointegration is rejected for all panel cointegration test specifications except for the group ADF-t-statistic using the unadjusted dataset. Moreover, the Johansen test indicates that only one cointegration relationship exists, supporting the results of the PURTs. We conclude that *public investment* and the *elderly voter share* are cointegrated and therefore share a long-run relationship.

Table 3: Panel cointegration tests

	Unadjusted data		Demeaned data	
<i>Pedroni (1999,2004)</i>				
Panel t-statistics (ADF)	-2.53**		-3.31**	
Group t-statistics (ADF)	-1.28		-1.82*	
	Cointegration rank			
<i>Johansen</i>	$r \leq 0$	$r \leq 1$	$r \leq 0$	$r \leq 1$
Fisher trace	69.57**	31.77	48.16**	26.36

Note: All variables in logs. Lags for the Pedroni tests have been selected using the Schwartz criterion starting from 5 lags. One lag has been selected for the Johansen Test. Cointegration relationship includes constant and trend. ** (*) Indicates significance at the 1% (5%) level.

4.3 Estimation

Since *public investment* and the *elderly voter share* are cointegrated, we can estimate β from equation (2) consistently using pooled D-OLS and FM-OLS. Both panel cointegration estimators allow for potential endogeneity as well as serial correlation. We apply the pooled version of these estimators because they considerably outperform their group-mean equivalents in simulations (Wagner and Hlouskova, 2010). D-OLS accounts for the correlation between the explanatory variable and the error term by including leads and lags of the explanatory variable in first differences. In order to estimate D-OLS, we augment equation (2) by *elderly voter shares* in first differences as well as leads and lags of *elderly voter shares* in first differences. In contrast, FM-OLS corrects for endogeneity and serial correlation using estimated adjustment factors. To keep comparisons valid and preserve degrees of freedom, we use one lead and one lag for all D-OLS specifications. However, the results are robust to different lag and lead specifications.

We continue using cross-sectionally demeaned data to alleviate cross-sectional dependence, which is in line with earlier research (Herzer and Nunnenkamp, 2013). As a benchmark, we also report OLS estimates, even though OLS might suffer from a non-negligible bias and distorted t-statistics (Kao et al., 1999; Kao and Chiang, 2001).

In line with our theoretical argument, we find a negative and significant effect of the *elderly voter share* on *public investment* in all specifications. As demonstrated in Table 4 the estimated coefficients suggest that a 1% increase in the *elderly voter share* is associated with an average decrease in *public investment* rates of 0.64% – 0.97%. To illustrate the magnitude of these coefficients, we exploit the evolution of sample means over time. During our period of investigation, the mean of the *elderly voter share* rose by 29%. At the same time, the cross-sectional average of *public investment* as a ratio to GDP decreased by 46%. Based on a coefficient of -0.8, population aging would explain around half of the decline in the sample mean of *public investment*. This implies a rather sizeable effect of aging. In the following, we test the robustness of our results using cross-sectionally demeaned data.

Table 4: Estimation results

Explanatory variable	Unadjusted data			Demeaned data		
	D-OLS (1)	FM-OLS (2)	OLS (3)	D-OLS (4)	FM-OLS (5)	OLS (6)
<i>Elderly Voter Share</i>	-0.71* (-2.14)	-0.75** (-2.74)	-0.64 (-3.18)	-0.89** (-2.86)	-0.97** (-3.58)	-0.88 (-4.41)
N	442	468	481	442	468	481

Note: The dependent variable is public investment. All variables in logs. t-statistics in parentheses. All estimations include country-specific constants and trends. D-OLS with one lead and one lag. D-OLS and FM-OLS covariance estimated using the sandwich estimator allowing for heterogeneous variances. ** (*) Indicates significance at the 1% (5%) level.

4.4 Robustness checks

We start by augmenting our empirical model with the following variables suggested in the literature on the determinants of public investment: the debt-to-GDP ratio, real GDP per capita and total population.⁸ The debt-to-GDP ratio should account for the usual “fiscal stress” argument raised in the public investment literature. GDP per capita is a rough measure of average income, which might affect the citizen’s demand for public investment. Population, as argued at the end of chapter 2, is intended to proxy returns to public investment. We include all controls either in log-levels or in growth rates. Further information on the construction of the control variables and summary statistics can be found in the appendix.

⁸Population or population growth has been included by Heinemann (2006) & Sturm and Haan (1998); GDP or GDP growth by Bacchiocchi et al. (2011) & Heinemann (2006) & Mehrotra and Väilä (2006) & Turrini (2004).

Given that panel cointegration estimates should be robust to omitted variable bias, we expect only small changes in β . The results shown in Table 5 confirm this expectation. The coefficients on the *elderly voter share* are similar to the estimates provided in Table 4 and remain significant at conventional significance levels. In line with previous considerations population and population growth seem to be positively correlated with *public investment*. In contrast, the level and growth rates of government debt as well as GDP per capita tend to have opposing effects on *public investment*. The results are similar, if control variables are simultaneously included in log-levels and growth rates.

Table 5: Estimates with control variables

Explanatory variables	D-OLS (1)	FM-OLS (2)	D-OLS (3)	FM-OLS (4)
Log-levels				
<i>Elderly Voter Share</i>	-0.74* (-2.01)	-0.73** (-2.88)	-1.07**(-3.05)	-0.97**(-3.61)
<i>Government Debt</i>	-0.12 (-1.71)	-0.12** (-3.49)		
<i>GDP per capita</i>	0.50 (1.32)	0.75** (3.53)		
<i>Total Population</i>	4.33** (2.86)	3.27** (3.80)		
Growth rates (coefficient*100)				
<i>Government Debt</i>			0.51* (2.08)	0.18 (1.77)
<i>GDP per capita</i>			-0.57 (-0.42)	-1.81** (-3.12)
<i>Total Population</i>			3.46 (0.58)	4.00 (1.02)
N	428	454	415	441

Note: The dependent variable is public investment in log-levels. 14 observations are missing due to incomplete data on public debt levels. All variables cross-sectionally demeaned. t-statistics in parentheses. D-OLS with one lead and one lag. All estimations include country-specific constants and trends. D-OLS and FM-OLS covariance estimated using the sandwich estimator allowing for heterogeneous variances. ** (*) Indicates significance at the 1% (5%) level.

To assess the robustness of our results, we also re-estimated our model without Japan and Germany, where, unlike the other sample countries, resident population already declined after 2007 or is expected to decline in the close future. Hence, the cut-back in public investment in both countries might be in fact driven by declining investment returns instead of increasing elderly voting power. The estimates presented in Table 6 show that the negative and significant correlation between *public investment* and the *elderly voter shares* also prevails in our reduced sample, indicating that falling investment returns following from a decline in future population are only a minor problem for our analysis. We also estimate the D-OLS model using different lag and lead length (see Table 9 in the appendix). Varying the number of leads and lags does not affect our basic results. In fact, all coefficients remain significant at least at the 5% level.

Table 6: **Sample without Japan and Germany**

Explanatory variable	D-OLS	FM-OLS
<i>Elderly Voter Share:</i>		
<i>Sample without Japan</i>	-0.91** (-2.85)	-0.93** (-3.37)
<i>Sample without Germany</i>	-0.98* (-2.50)	-1.07** (-3.05)
<i>Sample without Japan and Germany</i>	-0.97* (-2.37)	-0.99** (-2.71)

Note: The dependent variable is public investment. All variables in logs and cross-sectionally demeaned. t-statistics in parentheses. D-OLS with one lead and one lag. All estimations include country-specific constants and trends. D-OLS and FM-OLS covariance estimated using the sandwich estimator allowing for heterogeneous variances.

** (*) Indicates significance at the 1% (5%) level.

4.5 Long-run Granger causality

So far we have assumed that causality runs from population aging to public investment. However, our previous results only indicate a long-run relationship between two variables with no information on the direction of causality. Therefore, we employ an error correction model to test the response of each variable to a deviation from the long-run equilibrium. As standard in the literature we measure the deviation by the residual:

$$\hat{\varepsilon}_{it} = \ln(\text{public investment})_{it} - \hat{a}_i - \hat{b}_i \text{trend} - \hat{\beta} \ln(\text{elderly voter share})_{it}$$

obtained from the cointegration regression. In order to test for long-run Granger causality we estimate the following two equations and examine the significance of α_1 and α_2 using a t-test, respectively. The lagged first differences of *public investment* and *elderly voter shares* account for potential short-run effects.

$$\Delta \ln(\text{public investment})_{it} = c_{1i} + \alpha_1 \hat{\varepsilon}_{it-1} + \sum_{j=1}^5 \delta_{11} \Delta \ln(\text{public investment})_{it-j} + \sum_{j=1}^5 \delta_{21} \Delta \ln(\text{elderly voter share})_{it-j} + u_{1it} \quad (3)$$

$$\Delta \ln(\text{elderly voter share})_{it} = c_{2i} + \alpha_2 \hat{\varepsilon}_{it-1} + \sum_{j=1}^5 \delta_{12} \Delta \ln(\text{public investment})_{it-j} + \sum_{j=1}^5 \delta_{22} \Delta \ln(\text{elderly voter share})_{it-j} + u_{2it} \quad (4)$$

A significant α_1 would suggest long-run Granger causality running from the *elderly voter share* to *public investment*, while a significant α_2 signals long-run causality from *public investment* to the *elderly voter share* (e.g. Bronzini and Piselli, 2009; Herzer et al., 2012). Granger causality is not restricted to be unidirectional, in fact, both variables might be actually endogenous. Since all variables in the model are stationary we can estimate (3) and (4) using OLS (Bronzini and Piselli, 2009). Following common practice, we start from a general model with 5-lags and successively eliminate insignificant short-run dynamics (Herzer et al., 2012; Herzer and Nunnenkamp, 2013), though the main results do not depend on the lag length. Table 7 shows that α_1 but not α_2 is statistically significant from zero suggesting that long-run Granger causality exclusively runs from the *elderly voter share* to *public investment*, and not vice versa, supporting the argument that population aging has contributed to the decline in public investment.

Table 7: Error correction model

	Dependent variable: Public Investment	Dependent variable: Elderly Voter Share
	α_1	α_2
D-OLS residual	-0.34** (-8.90)	-0.01 (-1.20)
FM-OLS residual	-0.35** (-8.83)	-0.01 (-1.24)

Note: Insignificant short-run dynamics have been removed based on their t-value starting from 5 lags until all lagged differences are significant at the 10% level. t-statistics in parentheses. Residuals from the D-OLS estimation with one lead and one lag. ** Indicates significance at the 1% level.

5 Conclusion

The share of public investment declined almost steadily in most advanced economies since the seventies. The prevalent explanation for the cut-back in government financed investment centers on fiscal pressures caused by a rising level of public debt. Due to strong opposition against spending cuts for public consumption, governments consolidate by reducing public investment. In this paper, we suggest an additional explanation for the declining investment share. Based on a voter group decision model developed by Craig and Inman (1986), we argue that the ongoing demographic transition has decreased public investment. Experimental research reveals that senior individuals value future payoffs less than working-age people. Therefore, a raising fraction of elderly voters is likely to cause a reduction in the overall demand for durable public goods and hence public investment.

We demonstrate for a panel of 13 OECD countries between 1971 and 2007 that the share of elderly voters and public investment rates are indeed cointegrated and negatively correlated. Moreover, long-run Granger causality exclusively runs from aging to investment. Thus, our results suggest that public investment tends to decline in greying societies, even though returns to public capital are considerably positive, simply because the aging electorate does demand less and less investment spending. This finding is in line with previous empirical evidence on the negative impact of population aging on educational expenditures (Harris et al., 2001; Poterba, 1997). Alternative funding sources including Public Private Partnerships and public debt, or reforms of the political system, e.g. the introduction of demeny voting, could help to alleviate the growing opposition to public investment spending.

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7 Appendix

7.1 Variable information and summary statistics

Table 8: Summary statistics

Country	Public investment		Elderly voter share		Public debt		GDP per capita		Population	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Austria	3.18	1.38	20.40	0.84	49.71	17.12	22.14	5.15	7.78	0.25
Belgium	2.78	1.20	20.46	1.19	98.81	26.95	21.96	4.61	10.03	0.25
Canada	2.92	0.48	15.32	1.53	70.63	18.72	21.45	4.05	27.33	3.33
Denmark	2.30	0.76	19.84	0.67	48.31	23.45	29.38	5.97	5.19	0.13
Finland	3.16	0.45	17.88	2.00	27.23	18.34	21.14	5.31	4.97	0.2
France	3.26	0.25	20.44	1.22	43.32	16.58	21.52	3.97	57.93	3.14
Germany*	2.97	1.15	20.17	1.51	42.52	16.59	21.54	4.31	79.49	2.27
Italy	2.71	0.46	19.36	3.28	85.25	26.09	19.42	4.09	56.29	1.29
Japan	5.12	0.78	16.96	4.72	84.61	51.8	22.13	5.6	121.05	6.54
Netherlands	3.59	0.58	17.25	0.84	60.45	13.78	23.76	5.02	14.9	0.97
Sweden	3.62	0.74	22.42	1.15	48.36	15.37	24.39	4.82	8.56	0.32
United Kingdom	2.40	1.13	20.92	0.56	48.01	9.82	21.2	5.65	57.56	1.5
United States	2.33	0.16	17.02	0.61	53.91	10.61	26.06	5.86	251.16	28.76

Public investment and public debt in % of GDP. GDP per capita in thousand 2005 Euro. Population in million residents. *German series have been linked using growth-rates of West Germany.

Population, debt and GDP data is taken from the AMECO database. We divide nominal debt by nominal GDP to obtain the debt to GDP ratios and deflate GDP using the GDP deflator provided by the AMECO database. For the debt to GDP ratio 14 observations are missing (Canada 1971-74, Netherlands 1971-74, France 1971-76).

Figure 2: Public Investment (% GDP), 1971-2007

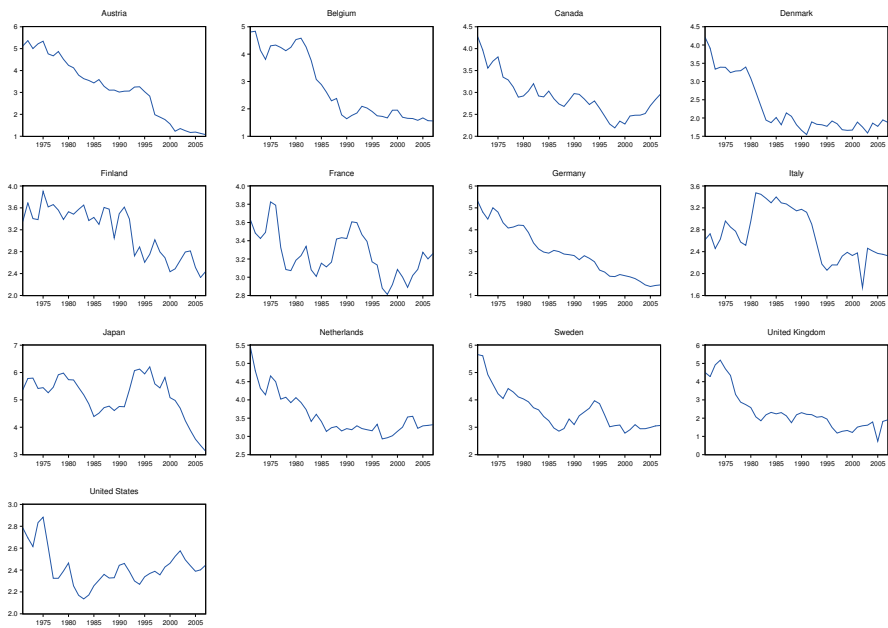
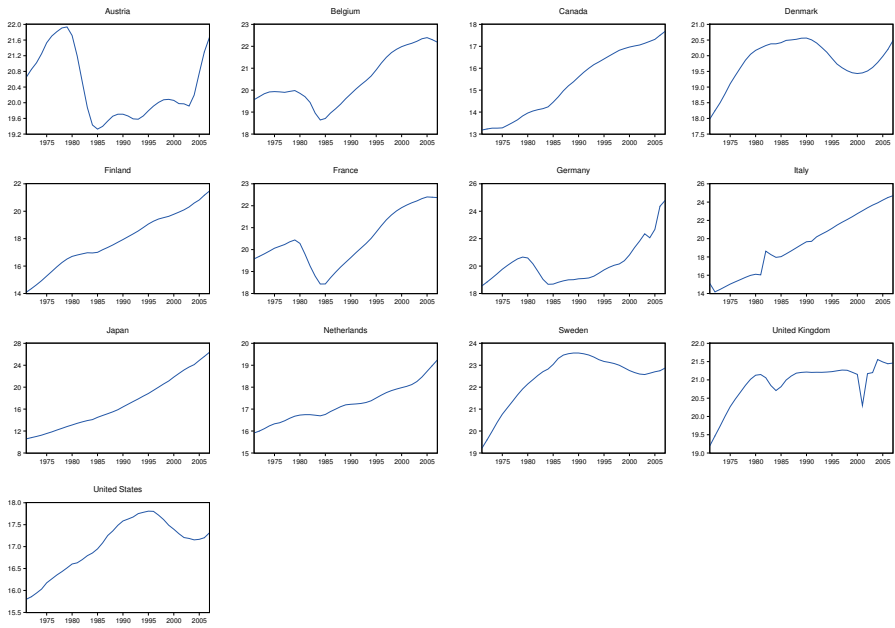


Figure 3: Elderly Voter Share (% Voting Population), 1971-2007



7.2 Additional robust check

Table 9: D-OLS estimates with different lead and lag length

Explanatory variable	leads=0,lags=1 (1)	leads=1,lags=0 (2)	leads=1,lags=2 (3)	leads=2,lags=1 (4)	heterogenous (max 5) (5)
<i>Elderly voter share</i>	-0.97** (-3.15)	-0.83** (-2.74)	-0.95** (-2.85)	-0.75* (-2.29)	-1.03* (-2.54)
N	455	455	429	429	406

Note: The dependent variable is public investment. All variables in logs and cross-sectionally demeaned. t-statistics in parentheses. Estimation includes country-specific constants and trends. Heterogenous: Country specific Lag and Lead selection based on Schwartz criterion with a maximum number of 5. D-OLS covariance estimated using the sandwich estimator allowing for heterogeneous variances.** (*) Indicates significance at the 1% (5%) level.