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Research article

Pension expenditures and old-age poverty in OECD countries

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Abstract: Using data from 37 OECD countries during the 1980–2020 period, this paper seeks to quantify the potential role and impacts of pension expenditures for reducing old-age poverty (66–75 and 76-and-over age groups). To that end, we applied panel data techniques controlling for key macroeconomic and demographic variables. Our results suggest that pension expenditure shows a linear and nonlinear impact on old-age poverty rates. The estimated elasticities increase (in absolute value) with the ratio of public pension expenditure to GDP, being higher (in absolute value) for the 76-and-over than for the 66–75 age group, suggesting a larger impact of pension expenditure on poverty for the oldest population group. When evaluating how each explanatory variable explains the poverty rates, we found that the macroeconomic situation is the main driver of its reduction, followed by public pension expenditure. Applying the panel dynamic threshold model, we also endogenously identified a common threshold of 7.21 in the ratio of public pension expenditures to GDP in the relationship between this variable and the poverty rate of both age groups. These results hold across various sensitivity analyses testing for structural breaks and heterogeneous relationships in the baseline empirical model and for the possibility of multiple thresholds in the threshold model. Our findings have significant implications for policymakers and researchers interested in social welfare and economic policy, providing evidence supporting the role of pension expenditures in reducing poverty among the elderly.

Keywords: pension expenditures; poverty; population aging; panel threshold models; OECD countries

JEL Codes: I32, H55, J11, J14, H63

1. Introduction

Pension reforms have been the focus of the policy agenda for the Organisation for Economic Co-operation and Development (OECD) countries for several decades since, on average, public spending on pensions has accounted for 7% of their gross domestic product (GDP). Also, this value is expected to significantly rise over the coming decades in response to population aging[1]. Pension policies are usually very controversial since they imply making long-term decisions under considerable short-run political pressures.

There is a growing concern about the future, particularly regarding the trend of lower pensions for today's workers compared to past generations, which could increase poverty rates in OECD countries. Reforms are essential to ensure both pension adequacy and the financial sustainability of pension systems. The aging population is putting increasing strain on these systems. The approach to these reforms will vary by nation, with the structure of the pension system and the investment choices made by individuals and pension funds playing crucial roles (see, e.g., OECD, 2023).

The journey to age equality is one of the primary purposes of the 2030 Agenda for Sustainable Development Goals (SDGs). The recent G20 meeting identified aging as a relevant factor in the world's economic and developmental growth, stressing the protection of older human rights by safeguarding access to healthcare services, decent work, and lifelong learning opportunities, as they are often overlooked and underappreciated. The United Nations (2019) warned about the necessity of promoting the economic, social, and political inclusion of older people since inequalities and poverty often deepen in old age.

Even though poverty risks in most OECD economies have moved from older to younger groups since the mid-1990s (OECD, 2019), the relative old-age (65+) poverty rate[2] is higher than the total population (13.5% versus 11.8%, respectively). The justification is the very high old-age poverty rates registered by some countries, such as Australia, Latvia, Estonia, Lithuania, and the United States.

According to OECD (2019), older age groups (76 and over) are characterized by significantly higher poverty rates in OECD countries. Some arguments have been put forward to explain this feature. First, pension benefits are often price-indexed[3]; second, an important percentage of the 76-and-over age group is female (distinguished by lower pensions and higher life expectancy); and third, some older people have not been covered during their whole working period.

Additionally, it is important to mention that income declines further with age in the old age group. In particular, those over 75 receive significantly lower pensions than the 66–75-year-old group in all OECD economies, with an average of 14 percentage points of difference. Nevertheless, for non-OECD G20 countries, the scenario is the opposite.

Several factors have undermined pension system solvency for OECD countries, necessitating innovative solutions. In addition to the behavior of demographic variables, the current low economic growth and the low interest rates for savers and financial service providers who provide life insurance are crucial to understanding why pension funds and life insurers adopt riskier investment options,

[1] Razin and Schwemmer (2021) provided a macroeconomic framework to understand the effects of population aging on policies relating to welfare-state policy.

[2] It is defined as income below half the domestic median equalized household income.

[3] Most OECD economies index their pensions to prices. This price indexation can benefit governments with budget constraints. However, this situation increases the risk of fueling pensioner poverty since earnings diminish over time.

looking for higher yields. This scenario challenges many countries' current and future retirement income (OECD, 2019).

Private pensions have achieved a more significant role due to recent pension reforms in OECD countries[4]. In particular, many countries that started as an entirely public system have moved to mixed systems over the years. For instance, economies such as Mexico, Poland, the Slovak Republic, Sweden, and Hungary substituted part of the public pension for mandatory private pensions to compensate for retirement income. Private pension programs have also been promoted in the United Kingdom, Iceland, and Switzerland since the 1980s. Due to the low level of public pensions, particularly for middle and high earners, and the need to cover future retirement income, voluntary private pensions have become crucial to nations such as the United States, Ireland, Canada, and Japan, among others. In most of these countries, private pensions represent more than 40% of coverage. The main argument for this trend is that a wide range of mandatory replacement rates across OECD countries[5] are voluntary, and private savings are necessary to guarantee elderly living standards.

The increasing shift from public pension schemes to a mix of public and private pension provision in the OECD economies and a change from the defined benefit (DB) to the defined contribution (DC) programs (Barr and Diamond, 2009; OECD, 2015 or Orenstein, 2011) has an impact on the expansion of income inequality and elderly poverty. According to Hwang (2016), Wang et al. (2012), Pestieau (1992), Behrendt (2000), Goudswaard and Caminada (2010), and Been et al. (2017), public pensions are expected to involve more equalized income distribution between individuals than private plans, which present a link between the contributions paid and the benefits received. This statement is based on the fact that public programs are more associated with income-related funding and flat-rate benefits, or benefits based on working time rather than past earnings, meaning that lower-income individuals are the primary beneficiaries. According to the OECD (2015), several economies have transitioned from DB to DC plans, recognizing that these plans do not facilitate interpersonal redistribution[6]. The first study to analyze this topic with panel data was by van Vliet et al. (2012). This study was followed by other empirical research that considered cross-sectional data, such as those by Weller (2004) and Milligan (2008). Van Vliet et al. (2012) argued that a higher share of private pensions does not necessarily lead to increased poverty among the elderly. These findings, which are robust across several econometric specifications, stand in contrast to theoretical expectations and previous analyses. In contrast, Been et al. (2017) provided empirical evidence of higher levels of elderly poverty and income inequality in European countries, a consequence of the higher relative weight of private pensions on pension provision.

The most common indicator for computing economic poverty in all nations, both rich and poor, is a conditioned measure to calculate economic well-being or resources relative to needs. The vast majority of the elderly in rich nations depend on cash incomes to satisfy their basic needs. In other words, family economic transfers are not as crucial for the elderly in rich economies compared to

[4] Private pensions can be voluntary or mandatory. Besides, the definition of the private scheme may differ between nations and across time.

[5] According to Martin and Whitehouse (2008), the replacement rate (pension received in retirement relative to earnings when working) from mandatory schemes for an average earner with a full career ranges from 31% in the United Kingdom to 96% in Greece.

[6] There are some exceptions, such as in Sweden and Denmark, where the DC programs involve the possibility of sharing collective risk.

developing countries (Smeeding and Williamson, 2001). For this reason, annual income is the most popular measure of economic resources to avoid poverty in rich nations, better than consumption (Johnson and Smeeding, 2013).

In recent decades, most European Union (EU) governments have focused on significant pension reforms to guarantee system sustainability (see European Commission, 2010 and Zaidi et al., 2006 for an overview of these reforms). The primary purpose of the pension scheme is to alleviate poverty and provide retirement income within a fiscal constraint. According to the standard approach, the financial success of pension reform is associated with a reduction in future spending (Schneider, 2009). Nonetheless, the World Bank claims that "pension systems need to provide adequate, affordable, sustainable, and robust benefits". The problem is that there is no consensus on the best method to determine whether pension entitlements will remain adequate. A unique indicator to measure pension adequacy is difficult (Grech, 2013).

For instance, the EU Member States have considered several instruments, such as the at-risk-of-poverty rate of the elderly, their median relative income, and the aggregate replacement ratio, among others, to ensure adequate retirement incomes for all and to allow living standards after retirement.

Among the OECD countries, there are several different retirement schemes. In particular, there are first-tier programs in which past earnings are not relevant to determine the retirement income and are related to a minimum standard of living for the elderly; see, for instance, residence-based basic, targeted, contribution-based basic, and minimum. Another critical piece in the architecture of the pension system associated with smoothing consumption between retirement and work life is the mandatory earnings-related components. This second link is considered the second-tier program and is based on defined benefits, points, notionally defined contributions, and funded-defined contributions. Finally, the third-tier scheme is composed of voluntary private frameworks[7].

Guided by these research and policy needs, our paper aims to examine the effects of pension expenditures on poverty, controlling for key macroeconomic and demographic variables using a sample of 37 OECD countries during the period 1980–2020. Our results suggest that the estimated elasticities increase (in absolute value) with the ratio of public pension expenditure to GDP, being higher (in absolute value) for the 76-and-over age group than for the 66–75 age group, suggesting that the impact of pension expenditure on poverty is larger for the oldest population group. Moreover, applying the panel dynamic threshold model, we also endogenously identify a common threshold of 7.21 in the ratio of public pension expenditures to GDP in the relationship between this variable and the poverty rate of both old-age groups.

The paper is structured as follows: The following section reviews the main theoretical and empirical studies on this topic. Section 3 presents the data and explanatory variables used in the empirical model. Section 4 briefly describes several econometric methodologies implemented to make our results more robust. Section 5 offers interpretations of the empirical results. Finally, Section 6 offers some concluding remarks and suggests potential avenues for future research.

2. Literature review

The adequacy of the pension system on elderly poverty protection is one of the main aspects that have been analyzed by academics and policymakers because old-age pension schemes are the most

[7] See Table 4.1 in OECD (2019) for more details and a description of these mandatory and voluntary retirement schemes.

significant component of social protection and seek several social goals (Scholz et al., 2006; Crawford and O'Dea, 2012; Knoef et al., 2016). Examining pension adequacy refers to the degree of poverty alleviation and consumption smoothing (see Barr and Diamond, 2006; Palacios and Sluchynsky, 2006; Abatemarco, 2009; or Caucatt et al., 2013, to name a few). In particular, Draxler and Mortensen (2009) emphasized that adequacy can be measured considering three dimensions: intragenerational redistribution, intergenerational redistribution, and lifetime income smoothing.

Although a replacement rate of 70% is usually applied as the rule of thumb for an adequate pension (Haveman et al., 2007), authors such as Chybalski and Marcinkiewicz (2016) contended that it is not a perfect instrument since it only focuses on consumption smoothing. For this reason, several other instruments have been applied (poverty line 40%, 50%, or 60% median relative income and at-risk-of-poverty rate, among others).

According to the International Labour Organisation (ILO), statutory pension programs must warrant adequate benefits, meaning a decent standard of living for the rest of their lives. For this reason, one of the main purposes of this study is to examine whether the OECD pension system provides enough protection against elderly poverty.

The adequacy of pension schemes is relevant not only because the aging of the population implies unsustainability for paying pensions for longer but also because it involves a crucial reduction in governments' earning capacity to pay those pensions. As birth rates decline, the expected revenues from salary-related pension contributions in both public and private pension[8] programs decrease. Moreover, migration is not able to change this trend worldwide. For these reasons, most economies struggle to face these demographic problems and overcome economic recessions.

Developed countries are characterized by low participation rates in old-age workers; this, together with higher longevity and low effective retirement age generates high dependency ratios, calling into question the social security systems (Gruber and Wise, 1999). Besides, higher unemployment rates translate into low contributions. This feature can cause severe consequences for today's and future pension system's sustainability. For many OECD nations, lower contributions may damage the pay-as-you-go (PAYG) finances and, in the future, may lead to a reduction in elderly income. Also, it is usually associated with higher reliance on non-contributory benefits, aggravating the fiscal pressure from the aging problem (OECD, 2019). Analyzing the British pension system, Carroll (2008) highlighted that the new pension regime will lead to more poverty, which is exacerbated by smaller family sizes, the decline in marriage, and more lone-person households, among other economic factors.

Most economies fight old-age poverty by combining two-income maintenance strategies. On the one hand, social insurance financed through worker contributions (Bismarckian approach) or taxes focuses more on poverty mitigation (Beveridgean approach). This option includes a universal, pay-as-you-go, defined-benefit, social-retirement scheme. In both the Germanic-Bismark and the British-Beveridgian social retirement systems, income replacement is afforded for all participants. The benefits design can be organized into two or more tiers: a lower tier with a higher replacement rate for lower lifetime earners, and a high minimum benefit offered for higher tiers. Additionally, the upper tier, which is more related to contributions, can pay out benefits at a much lower fraction of working-time earnings for high earners (Smeeding and Williamson, 2001). In most economies, social insurance is the higher source of income for the elderly. On the other hand, social assistance

[8] See Pedersen (2004) and De Deken (2013) for an extensive discussion on the classification of public and private pensions.

is a non-contributory system. Most economies combine the social retirement system with some form of social assistance. These programs are more focused on low-income elders and couples with an income-related safety net. According to the World Bank (2012), social assistance programs should be implemented as a valuable instrument to reach social protection and to alleviate poverty, social exclusion, and inequality. OECD (2019) supports that contributory basic pensions and minimum pensions should remain tools in preventing and mitigating old-age poverty.

Most empirical studies have concentrated on the link between social protection and poverty. However, there is less literature available on the specific relationship between pensions and poverty. Notable references in this area include Dethier et al. (2011), Faye (2007), Stewart and Yermo (2009), and Long and Pfau (2009).

Analyzing whether social protection benefits should be assigned to all or only to those with some specific criteria, Cruz-Martínez (2019) corroborated that universal social pensions are economically viable and efficient instruments to smooth income poverty. According to ILO (2014), social protection is considered a relevant tool to reduce poverty and inequality since one of the main purposes is to promote economic growth through improving health, productivity, domestic demand, and structural changes. Among economists who support the positive impact of social protection on poverty, one of the main arguments is associated with transfer coverage. This concept refers to the percentage of income/consumption that is provided by the social protection systems (Fiszbein et al., 2013, 2014).

The field of poverty alleviation is rich with diverse opinions and perspectives, although there is yet to be a consensus on the best method to alleviate income poverty. Mkandawire (2005) highlighted the effectiveness of the universalistic system in the 1960s and 1970s, while the rise of neoliberal social policy has brought targeting to the forefront as the most efficient means of poverty reduction[9]. When examining universal protection systems with means-tested programs, the former is usually associated with higher budgets and transfers that show flatter income distributions (see, e.g., Kidd, 2015, Danson et al., 2012).

Nevertheless, studies such as Dethier et al. (2011) show that old-age poverty can persist even in a scenario characterized by minimum pensions. They point out that even with high pension benefits and coverage levels, this does not necessarily imply poverty alleviation. In particular, the authors stress that the reduction in absolute poverty due to universal minimum pensions ranges from 2% for Brazil to 24% for Costa Rica, meaning that the impact depends on the country analyzed. Similarly, Ayala et al. (2021) ensured that most of the benefits associated with minimum pensions are insufficient to cover poverty risk in terms of incidence and intensity.

According to Förster et al. (2007), in developed countries, the average income of the elderly is significantly lower compared with other age groups, and their poverty rate is much higher. Despite the generous transfers for old people, including minimum pensions, in the OECD countries, elderly poverty remains an important societal problem. This situation can be confusing since these economies are characterized by universal pension, which guarantees levels above the poverty line. Three arguments could explain this behavior: family composition, take-up issues, and the fact that some elderly people keep unprofitable assets (Dethier et al., 2011).

Empirical evidence shows different stances about the relationship between pension expenditures and poverty. Smeeding and Williamson (2001) claimed that the result is not conclusive. These authors

[9] The main arguments to justify that targeting is the best alternative for social provision compared to universalism are explained in detail in Mkandawire (2005).

established that well-targeted spending schemes and adequate minimum benefits contribute to lower poverty rates in rich countries. Higher levels of public social spending are associated with low poverty levels and income inequality in this case. Nevertheless, expenditures that are not well targeted, unreasonable spending, and programs with low minimum benefits cannot reduce elderly poverty rates.

Using the Survey of Health, Ageing and Retirement in Europe, Fonseca et al. (2014) offered weak evidence about retirement as a determinant to protect against poverty in European countries.

By 2005, European Union countries experienced pension reforms shifting from earnings-related defined-benefit (DB) PAYG structure toward other pension models, such as personal accounts or notional-defined contribution (NDG) schemes. However, Zaidi et al. (2006) underscored that in most cases, these reforms were conducted for fiscal sustainability purposes rather than for prioritizing elderly poverty and income adequacy. In particular, they predicted that when the benefit ratio declines, the at-risk-poverty rates among the 65+ will increase substantially for EU25 during the period 2025–2050. These authors emphasize that, since people aged 76 and over depend more on national pensions, the reduction in generosity will translate into higher risk-at-poverty rates[10]. Analyzing the impact of pension reforms in ten major European countries since the 1990s, Grech (2013) found that even in nations where generosity has decreased, poverty has been alleviated, particularly in economies in which minimum pensions improved.

In addition, analyzing public and private schemes, a more significant public share in the pension provision aids in reducing poverty among old-age people (Oshio and Shimizutani, 2005; Milligan, 2008). Similarly, higher spending on public pensions leads to lower income inequality and poverty among the elderly in 17 European countries from 1995 to 2008 (Been et al., 2017).

Martin and Whitehouse (2008) maintained that, thanks to the new OECD income distribution due to pension reforms, older people's poverty rates have shown a reduction compared to the poverty rate of the population as a whole, except for seven countries (see Förster and Mira d'Ercole, 2005). However, these authors also highlight that cutting pensions in nations such as Germany, Japan, Italy, Poland, or the Slovak Republic leads to lower relative retirement incomes, which can result in a greater risk of elderly poverty, especially for low-income workers. At the micro-level, Engelhardt and Gruber (2006) computed the elasticity of poverty in the United States and identified that social security generosity highly contributed to elderly poverty during the 1968–2001 period.

Smeeding and Williamson (2001) and Turner et al. (2017) highlighted that a relatively inexpensive and well-targeted intervention can slow down the upward trend in elderly risk of poverty. More specifically, a longevity insurance benefit[11] as part of Social Security has contributed to reducing the poverty rate for people aged 75 and older in Ireland and Poland, since its poverty rate has achieved lower levels than those aged 65 and older. Other studies such as Milevsky (2005), Webb et al. (2007), and Iwry and Turner (2009) have corroborated that this longevity insurance annuity is a very appropriate option to provide insurance protection at an affordable cost for the most vulnerable age group.

Applying several measures of pension system's generosity, Lefèbvre and Pestieau (2006) underscored the correlations between pensions and poverty. Nonetheless, these authors did not control for other variables that can capture structural differences between economies. To overcome this drawback,

[10] These projections are computed based on the assumption that national benefit reduction is not compensated by working longer or increasing income with private pensions.

[11] A longevity insurance annuity is defined as a deferred annuity that starts receiving payments at an advanced age at which an essential part of the birth cohort would have died (Turner et al., 2017).

Jacques et al. (2021) assessed the impact of public pension spending on poverty among old age in 27 European countries, for the period 1995–2014, controlling for demographics and macroeconomic controls. The authors identified a robust nonlinear relationship between these two variables.

Another strand of the literature displays evidence that pension schemes do not guarantee protection against old-age poverty (see Zaidi and Grech (2007) and Kawinski et al. (2012), among others). Studies such as van Vliet et al. (2012) did not find significant evidence to support a relationship between private pension schemes to alleviate poverty rates among the elderly in some OECD countries during 1995–2007. Apart from no robust link between the share of spending in private pensions and old-age poverty, they did not detect an association between elderly inequality and private pension expenditure. Bednarczyk (2018) revealed that the Polish pension scheme does not provide enough protection against income poverty for future pensioners. In particular, this author suggested that a growing number of people do not fulfill the minimum insurance period required to receive the guaranteed minimum pension, bringing to light the unsustainable pension system and the increased risk of elderly poverty. This article exhibits some arguments for a greater risk of elderly poverty rates in the future, such as the flexible forms of employment that do not allow pension entitlements, women who have lower salaries and shorter professional working time, or people who have been characterized by low income along with their professional careers.

Arnold and Farinha Rodrigues (2015) emphasized that pension expenditure is the least efficient redistribution tool for Portugal with a defined-benefit pension system. They underline that this is explained by the fact that the higher the earnings during working age, the higher the pension received in retirement. However, this study claims that in addition to contributory pensions, the means-tested non-contributory pension scheme was especially effective in reducing poverty rates among pensioners aged 65 and above.

Most recently, Zhu et al. (2023) investigated how the higher Old Age Living Allowance (OALA) affected the institutionalized life course of older adults in Hong Kong, finding that higher OALA significantly increased income-based and expenditure-based poverty. Further, using data from the China Health and Retirement Longitudinal Study (CHARLS) (2018), Li et al. (2023) applied a heteroskedastic probit model with two steps least-squares and instrumental variables estimators to examine whether pensions have a beneficial impact on lowering relative poverty. According to their findings, pensions have observable anti-poverty benefits, significantly reducing relative poverty at the regional level.

Finally, some studies have examined the association between retirement on an old-age pension and subjective economic well-being. Shin et al. (2015) assessed the degree to which the South Korean government's 2008 introduction of the Basic Old-Age Pension (BOAP) program has improved the financial security of the elderly using a panel data analysis using data from the 2008–2010 Korea Welfare Panel Study. The findings indicate that the BOAP improves the affordability of basic subsistence products like heating and wholesome meals, especially for the elderly, which benefits older people's financial welfare. The effects, however, are restricted to these few results; other significant measures, such as general financial well-being, do not change. For its part, Palomäki (2019) used the longitudinal component of the EU-SILC 2010–2013, which includes 73,614 person-years and 26,680 persons aged at least 55, to investigate the relationship between retirement on an old-age pension and subjective economic well-being. Individual linear fixed effect regression models and retirement transition and retirement route dummies are used to analyze the data. The findings indicate that a slight decline in subjective economic well-being is linked to retirement in old age in general. A closer look, however,

shows that retirement is unquestionably associated with adverse outcomes for those retiring from various labor market positions. In addition, using a variety of subjective well-being metrics, Pak (2020) assesses the effectiveness of South Korea's 2014 social pension reform, one of the most significant social welfare expansions in the country's history. In order to isolate the causal effects of the benefit increase attributable to the reform, the author estimates a matched difference-in-differences model using data from the Korean Longitudinal Study of Ageing. According to this author's findings, recipients' financial happiness increased by an average of 4.8%–5.7% due to the reform; the correlation was higher for pensioners, seniors over 70, and those at the bottom of the wealth distribution. Lastly, Kuitto et al. (2023) presented a comparative empirical evaluation of the relationship between the generosity of public pension payments and the occurrence and severity of old-age poverty in 14 advanced OECD welfare states between 1980 and 2010. Using an instrumental-variable method, these authors show that pension systems and earnings-related schemes effectively lower the probability of old-age poverty, although a closer look reveals that retirement is unquestionably associated with adverse outcomes for those retiring from various labor market statutes.

3. Data

Our analysis covers 37 OECD countries during the period 1980–2020. In particular, our sample compromises data for Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Slovak Republic, Spain, Sweden, Switzerland, Turkey, United Kingdom, United States, Chile, Estonia, Israel, Slovenia, Latvia, Lithuania, and Costa Rica.

The poverty rate (*poverty*) is defined as the proportion of individuals over 65 with a disposable income lower than 60% of the median domestic disposable equivalent income of the total population. A higher value of this indicator means a higher at-risk-of-poverty rate among the elderly. Relative poverty is commonly computed by the OECD, the European Commission, and by other international groups (see, e.g., Förster, 1993; Hagenaars et al., 1998; and Ramprakash, 1994). Relative measures of poverty rates are preferred to absolute ones (see Bourguignon, 2003; Cantillon, 2011; and van Vliet et al., 2012; among others).

One of the main purposes of this study is to assess whether the generosity of public pension schemes can contribute to the poverty rates for the OECD countries during the period 1980–2020. For this goal, we include pension spending (*pension*) in our analysis[12]. It is important to mention that it covers income for persons retired from the labor market or incomes for people who have reached a "standard" pensionable age or fulfilled the necessary contributory conditions. This variable incorporates early retirement pensions and reflects expenditures on services such as daycare, home help, rehabilitation, and residential care.

Following van Vliet (2010) and Caminada et al. (2012)'s suggestions, we incorporate control macroeconomic and demographic factors such as real per capita GDP (*gdp*) (expressed in constant

[12] All cash expenditures (including lump-sum payments) on old-age and survivors' pensions as a percentage of GDP. Old-age cash benefits provide an income for persons retired from the labor market or guarantee income when a person has reached a "standard" pensionable age or fulfilled the necessary contributory requirements. Old age also includes social expenditure on services for the elderly people, services such as daycare and rehabilitation services, home-help services, and other benefits in kind. It also includes expenditure on the provision of residential care in an institution.

2010 euros); the public debt-to-GDP ratio (*debt*); the ratio of total government spending to GDP (*govspending*)[13]; the Gini index (*gini*)[14]; and the dependency ratio (*dependency*), computed as the ratio between individuals aged 65 and over and the working-age population (15–64)[15] [16]. The definitions and sources of the variables are presented in Table 1.

Table 1. Explanatory variables and data sources.

Variable	Description	Source
Poverty66	Poverty rate after taxes and transfers for the 66–75 age group	OECD Statistics
Poverty76	Poverty rate after taxes and transfers for 76-and-over age group	OECD Statistics
Pension	All cash expenditures on old-age and survivor pensions as a percentage of gross domestic product	OECD Statistics
Gini	Gini index measuring the extent to which the distribution of income among individuals or households within an economy deviates from a perfectly equal distribution	World Income Inequality Database
GDP	Per capita gross domestic product at constant 2010 U.S. dollars	World Development Indicators (World Bank)
Dependency	Dependency ratio computed as the ratio between individuals aged 65 and over and the working-age population (15–64)	World Development Indicators (World Bank)
Govspending	General government spending	World Development Indicators (World Bank)
Debt	The ratio of public debt to gross domestic product	World Development Indicators (World Bank)
Unemployment	The unemployment rate expressed as total unemployment as a percentage of the total labor force	World Development Indicators (World Bank)
Life_expectancy	Life expectancy at birth, total (years)	World Development Indicators (World Bank)

[13] Cash payments are incurred for government operating activities in providing goods and services. It includes compensation of employees (such as wages and salaries), interest and subsidies, grants, social benefits, and other expenses such as rent and dividends.

[14] The Gini index provides a synthetic measure of inequality. It ranges from 0 (in the case of perfect equality) to 1 (a situation in which one person captures all resources in an economy). This index has been widely used for inequality research.

[15] The unemployment rate and life expectancy at birth were also initially considered as potential explanatory variables in a general statistical model, but, as they were found statistically insignificant in the regression results, they were eliminated following the general specific approach (Hendry, 1995, ch. 9). The same situation occurred with the longevity variable suggested by an anonymous referee to incorporate its effect on the burden of public pension plans and its consequent effect on poverty.

[16] As an anonymous referee pointed out, our empirical findings should be interpreted cautiously because omitted variable bias could occur if other relevant factors, such as social safety nets, labor market dynamics, or health spending, are not included in the model.

According to Barro (2000), *govspending* and *debt* are explanatory factors that should be taken into account to reflect economic maturity and the corresponding impact on income distribution, since the poverty rate is measured relative to the median income. On the other hand, it is well-known that poverty rates are conditioned by the median income among nations and time. To overcome this issue, we have included economic factors such as the public debt-to-GDP ratio, government spending, and unemployment rates as proxies to control the economic cycle. Additionally, we have added time- and country-fixed effects.

To produce a data matrix without missing values, we apply two complementary procedures: the technique of multiple imputations developed by King et al. (2001), which permits the approximation of missing data and allows us to obtain better estimates, and the simultaneous nearest-neighbor predictors proposed by Fernandez-Rodriguez et al. (1999), which infer omitted values from patterns detected in other simultaneous time series.

Since the appropriate econometric treatment of a model depends crucially on the pattern of stationarity and non-stationarity of the variables under study, as a preliminary step before carrying out the estimation, we perform a variety of unit root tests in panel datasets. Specifically, we use the Levin et al. (2002), Harris and Tzavalis (1999), Breitung (2000), Im et al. (2003), and Fisher-type (Choi, 2001) tests. The results of these tests (not shown here to save space but available from the authors upon request) decisively reject the null hypothesis of a unit root for the (logarithm) of *poverty*, *pension,* and *gdp,* as well as for *dependency* and *govspending* [indicating that they are stationary in levels, i.e., I(0)], while they do not reject the null for *gini* and *debt* [suggesting that these variables can be treated as first-difference stationary, i.e., I(1)].

4. Econometric methodologies

In this section, we briefly describe several econometric methodologies that we use to achieve robust results.

4.1. Baseline empirical model

In the first step, we consider three basic panel regression methods: the fixed-effects (FE) method, the random effects (RE) model, and the pooled ordinary least square (POLS) method to estimate the following baseline empirical model:

$$lpoverty_{it} = \alpha_i + \delta_1 lpension_{it} + \delta_2\, lpension^2{}_{it} + \delta_3 \Delta gini_{it} + \delta_4 lgdp_{it} + \delta_5 dependency_{it} + \delta_6 \Delta debt_{it} + \delta_7 govspending_{it} + \xi_{it} \quad (1)$$

Note that, given that our dependent variable is stationary, we transform the non-stationary (*gini* and *debt*) variables into stationary variables taking their first difference. Notice also that we introduce a quadratic pension term to approximate the possible nonlinear relationship between public pension expenditure and the poverty rate.

In order to determine the empirical relevance of each of these potential methods for our panel data, we make use of several statistical tests. In particular, we test FE versus RE using the Hausman (1978) test statistic to test for non-correlation between the unobserved effect and the regressors. Additionally, to choose between pooled-OLS and RE, we use Breusch and Pagan (1980)'s Lagrange multiplier test to

test for the presence of an unobserved effect. Finally, we use the F test for fixed effects to test whether all unobservable individual effects are zero, to discriminate between pooled-OLS and RE.

Since pension expenditures can respond to changes in poverty rates, we decided to control for the possible endogeneity problem. To that end, we treat pension expenditures and the square of pension expenditures as two endogenous variables. In particular, we estimate the so-called fixed effects instrumental variables or the fixed effects two-steps least-squares technique (FE-IV or FE-2SLS) and the random effects instrumental variables or the random effects two-steps least-squares procedure (RE-IV or RE-2SLS). Hausman's (1978) test is implemented to discriminate between them.

4.2. Threshold model

In a second step, we study whether the poverty rates in the OECD countries behave in a different manner above or below a particular pension expenditure threshold. For this purpose, we implement a dynamic panel threshold model to estimate the asymmetric relationship between these two variables, controlling for fixed effects and the macroeconomic and demographic factors used previously.

The threshold model can be expressed as:

$$y_{it} = \alpha_i + X_{it}(q_{it} < \tau)\delta_1 + X_{it}(q_{it} \geq \tau)\delta_2 + e_{it} \quad (2)$$

In which y_{it} is the dependent variable (which, in this case, corresponds to the 66–75 and the 76 and over age group poverty rates), q_{it} is the threshold variable, τ is the threshold parameter, which divides the Equation into two regimes, α_i are the nation-specific intercepts, and e_{it} is the error term.

An alternative for the above Equation is:

$$y_{it} = \alpha_i + X_{it}(q_{it}, \tau)\delta + e_{it} \quad (3)$$

where $X_{it}(q_{it}, \tau) = \begin{cases} X_{it} I(q_{it} < \tau) \\ X_{it} I(q_{it} \geq \tau) \end{cases}$ and I(.) is an indicator function.

As a common practice in empirical research in this area, instead of searching over the whole sample, the range is limited within the interval $(\underline{\tau}, \overline{\tau})$ as quantiles of the threshold variable. To compute the $\tau's$ estimator, we select the value that minimizes the residual sum of squares (RSS):

$$\hat{\tau} = \arg\min S_1(\tau) \quad (4)$$

Then, the ordinary least squares estimator of δ is:

$$\hat{\delta} = \{X^*(\tau)' X^*(\tau)\}^{-1} \{X^*(\tau)' y^*\} \quad (5)$$

where y^* *and* X^* are within-group deviations.

When the threshold parameter is unknown, the nuisance parameter problem appears due to the fact that estimation and inference are more complex because the τ estimator's distribution is nonstandard. Hansen (1999) solves this concern by proving that $\hat{\tau}$ is a consistent estimator for τ.

Additionally, we apply a simple test to clarify whether the coefficients estimated are statistically different across regimes. We compute the following test to verify if there exists an asymmetric relationship between poverty rates and pension expenditures indicating the presence of a different behavior above and below the estimated threshold:

$$H_0: \delta_1 = \delta_2 \qquad H_1 : \delta_1 \neq \delta_2 \quad (6)$$

In this case, the null hypothesis captures the linear model. Therefore, if we reject this hypothesis, the threshold model is preferred. The corresponding F statistic is computed as:

$$F_1 = \frac{(S_0 - S_1)}{\hat{\sigma}^2} \tag{7}$$

In which S_0 is the RSS of the linear model, and the F statistic is based on bootstrap iterations, crucial to determine the significance of the estimated threshold.

5. Empirical results

5.1. Empirical results from the baseline empirical model

Table 2 reports the results obtained using the three basic panel regression methods described above for the 37 OECD countries during the 1980–2020 period for poverty rates after taxes and transfers of the age groups of 66–75 and 76 and over. Based on the specification tests, the RE model is the most appropriate for poverty rates in both age groups.

Table 2. Parameter estimates for the empirical model (both age groups 66–75 and 76 and over poverty rates).

	Lpoverty66			Lpoverty76		
	FE	RE	OLS	FE	RE	OLS
$lpension_{it}$	0.3819***	**0.3888*** **	−0.5760***	0.4160***	**0.3745** **	−0.4183***
	(0.1531)	**(0.1501)**	(0.1823)	(0.1717)	**(0.1671)**	(0.1321)
$lpension^2{}_{it}$	−0.2529***	**−0.2568*** **	0.1458***	−0.4074***	**−0.3753*** **	0.0433
	(0.0554)	**(0.0528)**	(0.0540)	(0.0622)	**(0.0587)**	(0.0418)
$\Delta gini_{it}$	0.0514***	**0.0544*** **	0.0499***	0.0385***	**0.0416*** **	0.0297***
	(0.0085)	**(0.0077)**	(0.0052)	(0.0095)	**(0.0085)**	(0.0049)
$lgdp_{it}$	0.4386***	**0.1848**	−0.1778***	−0.3257*	**−0.2203***	−0.1659***
	(0.1550)	**(0.1134)**	(0.0421)	(0.1739)	**(0.1228)**	(0.0430)
$\Delta dependency_{it}$	3.5256***	**3.3250*** **	3.4830***	8.3722***	**7.6499*** **	4.9080***
	(0.8964)	**(0.8460)**	(0.6056)	(1.0055)	**(0.9376)**	(0.6035)
$govspending_{it}$	−0.0141***	**−0.0152*** **	−0.0294***	0.0085*	**0.0050**	−0.0272***
	(0.0045)	**(0.0044)**	(0.0040)	(0.0050)	**(0.0048)**	(0.0039)
$\Delta debt_{it}$	−0.0026**	**−0.0025** **	0.0015*	−0.0056***	**−0.0048*** **	0.0005
	(0.0011)	**(0.0010)**	(0.0008)	(0.0013)	**(0.0012)**	(0.0007)
Constant	−7.8415***	**−5.3036*** **	−0.9084	−0.7403	**−1.6506**	−0.4099
	(1.5886)	**(1.2045)**	(0.6247)	(1.7818)	**(1.3099)**	(0.6308)
Country FE	Yes	**Yes**	No	Yes	**Yes**	No
Year FE	Yes	**Yes**	Yes	Yes	**Yes**	Yes
R^2 overall	0.0975	**0.1865**	0.3416	0.1069	**0.1269**	0.2444
R^2 within	0.2566	**0.2538**		0.2425	**0.2414**	
R^2 between	0.1072	**0.1865**		0.0894	**0.1202**	
BIC	1200.71	**1158.63**	2478.12	1433.47	**1328.91**	2500.92
AIC	979.23	**956.38**	2256.65	1211.99	**1145.22**	2279.45

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	Lpoverty66			Lpoverty76		
	FE	RE	OLS	FE	RE	OLS
Breusch and Pagan test (POLS vs RE)	6133.89 [0.0000]			5439.46 [0.0000]		
F test for fixed effects (POLS vs FE)	65.43 [0.0000]			48.35 [0.0000]		
Haussman test (FE vs RE)	19.08 [0.9996]			10.56 [0.8533]		

Notes: In the ordinary brackets below the parameter estimates are the corresponding *z*-statistics, computed using White's (1980) heteroskedasticity robust standard errors. The associated p-values are in the square brackets below the specification tests. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. In bold type, we highlight the relevant estimation method for each case according to the specification tests.

Controlling for the possible endogeneity of pension expenditures, the Hausman test identifies that the best method is RE-2SLS for the 66–75 age group poverty rate and FE-2SLS for the 76-and-over poverty rate (Table 3). Therefore, we discuss only the results based on these estimation methods that are found to be the relevant ones for our sample.

Table 3. Parameter estimates for the empirical model controlling for endogeneity (both age groups 66–75 and 76-and-over poverty rates).

	Lpoverty66		Lpoverty76	
	FE-2SLS	**RE-2SLS**	**FE-2SLS**	RE-2SLS
$lpension_{it}$	0.5633***	**0.5303*** **	**0.5940*** **	0.5280***
	(0.1609)	**(0.1585)**	**(0.1780)**	(0.1756)
$lpension^2{}_{it}$	−0.3509***	**−0.3145*** **	**−0.5049*** **	−0.4339***
	(0.0591)	**(0.0568)**	**(0.0654)**	(0.0627)
$\Delta gini_{it}$	0.0301***	**0.0285*** **	**0.0273*** **	0.0197**
	(0.0080)	**(0.0072)**	**(0.0088)**	(0.0078)
$lgdp_{it}$	−0.4165***	**−0.3523*** **	**−0.9879*** **	−0.7140***
	(0.0968)	**(0.0859)**	**(0.1071)**	(0.0926)
$\Delta dependency_{it}$	1.8240**	**1.3879** **	**6.2400*** **	4.6662***
	(0.7939)	**(0.74288)**	**(0.8782)**	(0.8125)
$govspending_{it}$	−0.0188***	**−0.0202*** **	**0.0032**	−0.0009
	(0.0045)	**(0.0044)**	**(0.0050)**	(0.0049)
$\Delta debt_{it}$	−0.0023**	**−0.0021** **	**−0.0047*** **	−0.0038***
	(0.0011)	**(0.0011)**	**(0.0012)**	(0.0011)
Constant		**1.0375**		4.5471***
		(0.8740)		(0.9498)
Country FE	Yes	**Yes**	**Yes**	Yes
Year FE	Yes	**Yes**	**Yes**	Yes
R^2 overall	0.1928	**0.1813**	**0.2031**	0.0898
R^2 within	0.1912	**0.1924**	**0.2000**	0.2001

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	Lpoverty66		Lpoverty76	
	FE-2SLS	**RE-2SLS**	**FE-2SLS**	RE-2SLS
R^2 between	0.1838	**0.2216**	**0.0905**	0.1006
BIC	1000.93	**989.65**	**1203.55**	1245.69
AIC	966.55	**958.23**	**1169.17**	1124.26
Haussman test	−10.58		28.21	
(FE-2SLS vs RE-2SLS)			[0.0002]	
Under identification test	LM statistic: 975.152		LM statistic: 975.152	
	χ^{2i} p-value: 0.0000		χ^{2i} p-value: 0.0000	
Weak identification test	Cragg-Donald Wald F statistic:		Cragg-Donald Wald F statistic:	
	1.7e+04		1.7e+04	
	10% maximal IV size: 7.03		10% maximal IV size: 7.03	
	15% maximal IV size: 4.58		15% maximal IV size: 4.58	
	20% maximal IV size: 3.95		20% maximal IV size: 3.95	
	25% maximal IV size: 3.36		25% maximal IV size: 3.36	
Sargan statistic	0.589		0.589	
	p-value: 0.9875		p-value: 0.9875	

Notes: In the ordinary brackets below the parameter estimates are the corresponding *z*-statistics, computed using White's (1980) heteroskedasticity robust standard errors. The associated p-values are in the square brackets below the specification tests. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. In bold type, we highlight the relevant estimation method for each case according to the specification tests. Sargan-Hansen test is a test of overidentifying restrictions. The joint null hypothesis is that the instruments are valid (i.e., uncorrelated with the error term) and that the excluded instruments are correctly excluded from the estimation Equation.

As shown in Table 3, public pension expenditure and its square are highly significant (at the 1% significant level), suggesting a strong nonlinear relationship between both age groups' poverty rates and public pension expenditure. Our results are in line with Jacques et al. (2021).

As expected, the Gini coefficient, usually applied to measure income inequalities, offers a direct significant association with poverty rates in both age groups, since individuals who are characterized by low income relative to others are likely to live below the poverty line.

Turning to the case of real per capita GDP, not surprisingly, we found that it contributes to alleviating poverty rates in both age groups.

When the pension expenditure growth is not associated with the increase in the number of retirees, a negative consequence can be generated in the elderly's income. For this reason, to control for the population aging, we also included the dependency ratio for the OECD countries during the analyzed period. The results suggest that the higher the dependency ratio is, the higher both poverty rates are[17].

With regard to the public debt-to-GDP ratio as a proxy for the economic cycle, an indirect relationship is identified by the data. This result suggests that a higher involvement of the state through

[17] Controlling for other variables, we also find evidence that the enhancement of life expectancy jeopardizes the pension system's sustainability, compromising poverty rates. These additional results are not shown here to save space, but they are available from the authors upon request.

more generous public programs implies a higher debt-to-GDP ratio, but, simultaneously, it translates into lower poverty rates.

Concerning government spending, our results suggest that this variable is fundamental and highly statistically significant to reduce poverty among the 66–75-year-old group. This finding aligns with those by Smeeding (2006) and Smeeding and Williamson (2001). Interestingly, we find a positive association between government spending and the poverty rate for the age group older than 75, indicating that different schemes for income maintenance in old age produce very different anti-poverty results and the need for better-targeted spending.

It is noticeable that the values of the objective functions (the Akaike information criterion and the Bayesian information criterion, AIC and BIC, respectively) of the RE-2SLS for 66–75 poverty rates and FE-2SLS for the age group of 76 are lower than the values of the objective function of the OLS, FE, and RE estimation, justifying the appropriate use of instrumental variables to correct for the presence of endogenous regressors.

Finally, as shown in Table 3, the estimated results from the standard tests of validity suggest that the instruments we use satisfy both relevance and exogeneity conditions.

Given the nonlinear relationship detected, we follow the suggestion of Jacques et al. (2021) and calculate the elasticity of the poverty rate to public pension expenditures as follows:

$$\varepsilon_{it}(lpension) = \frac{\partial lpoverty_{it}}{\partial lpension_{it}} = \hat{\delta}_1 + 2\hat{\delta}_2 lpension_{it} \tag{8}$$

where $\hat{\delta}_1$and $\hat{\delta}_2$are the estimated coefficients in Table 3 for $lpension_{it}$ and $lpension^2{}_{it}$, respectively. Note that the elasticity of poverty to public pension spending is not constant but depends on the level of public pension expenditures ($lpension_{it}$). Figures 1 and 2 plot the estimated elasticity of the poverty rate for the 66–75 and 76-and-over age groups, respectively. As can be seen, they are both significantly different from zero and increase (in absolute value) with the ratio of public pension expenditure to GDP. Regarding the 66–75 age group (Figure 1), the estimated elasticity gradually increases from −0.43 to −0.69, with −0.56 as its average value. Turning to the case of the 76-and-over age group (Figure 2), the estimated elasticity rises smoothly from −0.94 to −1.36, with an average value of −1.16. Note also that both figures suggest that elasticity finally stabilizes for a pension expenditure/GDP ratio of 7.8%. Therefore, the estimated elasticity is higher (in absolute value) for the 76-and-over than for the 66–75 age group, suggesting that the impact of pension expenditure on poverty is larger for the oldest population group.

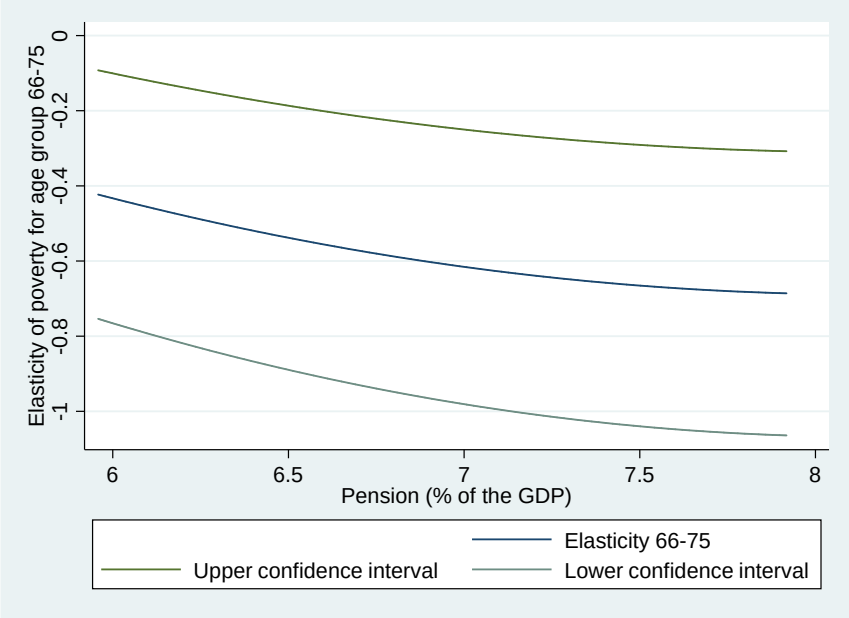

Figure 1. Elasticity of poverty rate for the age group 66–75 to public pension expenditures.

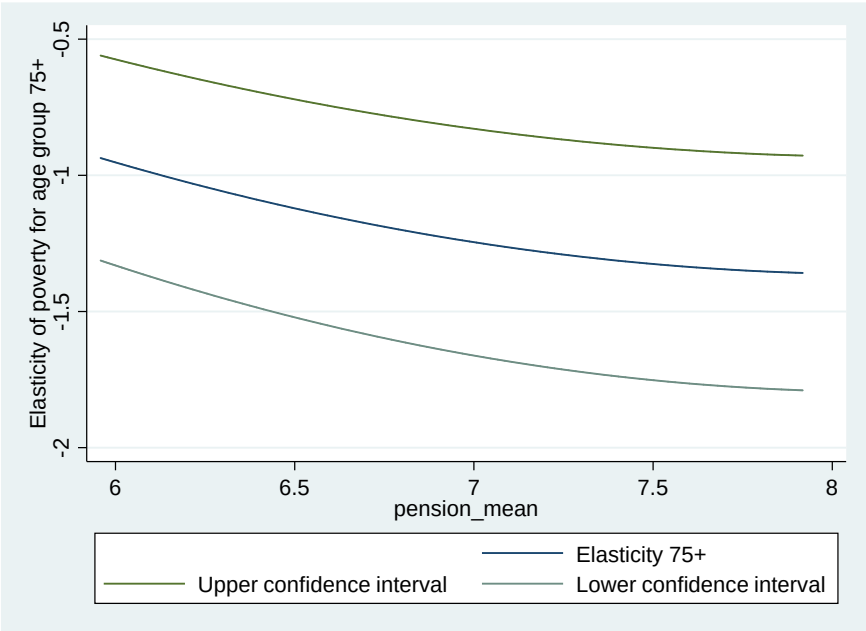

Figure 2. Elasticity of poverty rate for the age group of 76 and over to public pension expenditures.

Lastly, to gauge the predictive power of our basic model and assess how each explanatory variable contributes to the explanation of the dependent variable, we perform stochastic dynamic simulations. Table 4 reports the results for both poverty rate indicators under study. Column 2 represents the actual values of the dependent variables averaged over the period under study, while column 3 shows the averaged predicted values. The remaining columns present the contribution of the explanatory variables. As can be seen, our model delivers a high level of forecast accuracy. Moreover, our results

suggest that the macroeconomic situation is the main driver of the reduction of the poverty rate for both age groups under study (contributing to explaining 56.83% of the poverty rate for the 66–75 age group and 77.30% for the age group of 76 and over), with increasing inequality and public debt accumulation playing a much less important role. In absolute value terms, public pension expenditure (in levels and quadratic terms) explains 32.30% of the poverty rate for the 66–75 age group and 21.71% for the age group of 76 and over. Finally, as for government spending, while this measure of the financial sustainability of public programs represents 10.75% of the absolute variation in 56.83% of the poverty rate for the 66–75 age group, it only accounts for 0.83% of the poverty rate for the age group of 76 and over.

Table 4. Predictive power and relative contributions of the explanatory variables.

	Actual	Predicted	Contributions (%)						
	rate	rate	$lpension$	$lpension^2$	$\Delta gini$	$lgdp$	$\Delta dependency$	$govspending$	$\Delta debt$
Lpoverty 66	−2.2314	−2.4257	14.65	17.65	0.02	56.83	0.05	10.75	0.03
Lpoverty 76	−1.8317	−1.9484	7.96	13.74	0.01	77.30	0.12	0.83	0.04

Notes: Contributions are expressed in absolute values.

5.2. Empirical results from the threshold model

Based on the previous results, we estimate a dynamic panel threshold model controlling for fixed effects and the common macroeconomic factors to shed further light on the impact of pension expenditures on poverty rates for both age groups[18]. The estimated results are shown in Table 5. As can be seen, these results corroborate a nonlinear relationship between pension expenditures and poverty rates regardless of the measure employed (65–75 or 76-and-over age group poverty rates)[19]. The upper part of Table 4 offers the estimated pension threshold, the corresponding p-value, and the 95% confidence interval. The estimated pension threshold for old age groups under study is 1.9756, which is statistically significant at a 10% and 5% significance level for 65–75 and 76-and-over age groups. This threshold corresponds to a 7.21 ratio of public pension expenditures to GDP. Recall that $\hat{\delta}_1$ and $\hat{\delta}_2$ represent the estimated marginal effects of the pension expenditure on poverty rates when the pension expenditure is below and above the estimated threshold value, respectively. Therefore, the estimated impact on the reduction in poverty rate is higher in magnitude for both age groups when pension expenditure is above its threshold (−0.83 versus -0.64 for the 65–75 age group and −0.92 versus −0.72 for the 76-and-over age group). Note also that these estimated marginal effects are higher in absolute value for the 76-and-over age group. Turning to the effect of the macroeconomic and demographic variables on poverty rates (in the lower part of this table), in general terms, we detect a similar pattern to those previously obtained for both age groups from the baseline empirical model. A

[18] As this methodology requires a balanced panel, our estimations are based on the 2002–2020 period, given the data availability for all macroeconomic and demographic variables.

[19] We can conclude this result since we have enough significant evidence to reject the null hypothesis of a single model. For this reason, the dynamic threshold model is more appropriate.

positive and significant relationship is found for both the dependency ratio and economic activity. However, a negative relationship is identified for the public debt-to-GDP ratio. As for the effect of income inequality, the estimated coefficient suggests a positive relationship between the Gini index and the poverty rate for the 65–75 age group; the opposite occurs for the 76-and-over age group (although in the latter case, the estimated coefficient is not statistically significant). Finally, concerning government spending, this alternative estimation method does not detect a statistically significant impact of this variable on poverty rates.

Table 5. Results of dynamic panel threshold model (age groups 66–75 and 76-and-over poverty rates).

	Lpoverty66	Lpoverty76
Threshold estimates ($\hat{\tau}$)	1.9756	1.9756
Significance of threshold p-value	0.0614	0.0488
95% confidence interval	[1.9683, 1.9804]	[1.9659, 1.9804]
Impact of threshold variable on interest rate:		
$\hat{\delta}_1$	−0.6414***	−0.7272***
	(0.1504)	(0.1562)
$\hat{\delta}_2$	−0.8313***	−0.9251***
	(0.1476)	(0.1532)
Impact of control variables on interest rate:		
$\Delta gini_{it}$	0.0221*	−0.0068
	(0.0132)	(0.0137)
$lgdp_{it}$	1.0500***	0.5077***
	(0.1846)	(0.1917)
$\Delta dependency_{it}$	0.1210**	2.4354***
	(0.0929)	(1.0310)
$govspending_{it}$	0.0019	0.0044
	(0.0056)	(0.0058)
$\Delta debt_{it}$	−0.0053***	−0.0079***
	(0.0012)	(0.0013)
Constant	−12.3450***	−5.7260***
	(1.9842)	(2.0604)
R^2 within	0.1000	0.0288
R^2 between	0.0160	0.0140
R^2 overall	0.3056	0.3056

Notes: In the ordinary brackets below the parameter estimates are the corresponding *z*-statistics, computed using White's (1980) heteroskedasticity robust standard errors. In the square brackets below the specification tests are the associated p-values. *, **, and *** indicate significance at 10%, 5%, and 1% respectively.

5.3. Sensitivity analysis[20]

A variety of sensitivity analyses were conducted to enhance the robustness of our findings. In particular, we test for structural breaks and heterogeneous relationships in the baseline empirical model and explore the possibility of multiple thresholds in the threshold model.

5.3.1. Structural breaks

The potential for structural breaks or changes in the relationship over time is a concern, given the extensive study period from 1980 to 2020. Therefore, we conduct tests for structural breaks in the baseline empirical model to strengthen the robustness of our findings.

We have implemented a method developed by Ditzen et al. (2021) and Karavias et al. (2021) for detecting and dating multiple structural breaks in panel data. This method can efficiently identify an unknown number of breaks at unknown break dates. Based on asymptotically valid tests, these authors developed a new Stata command (xtbreak) that can identify breaks, ascertain their number and location, and provide break date confidence intervals. We specifically look at testing hypothesis H_0: no breaks against H_1: s breaks, where s is specified by the researcher.

Table 6. Test for multiple breaks according to Karavias et al. (2021).

		Bai and Perron critical values		
	Test statistic	1% critical value	5% critical value	10% critical value
H_0: no breaks				
H_1: 3 breaks				
SupW(tau)	2.02	2.77	2.46	2.29
Estimated break points:	2005, 2008, and 2011			
H_0: no breaks				
H_1: 2 breaks				
SupW(tau)	1.99	3.12	2.71	2.52
Estimated break points:	2008 and 2011			
H_0: no breaks				
H_1: 1 breaks				
SupW(tau)	2.20	3.82	3.12	2.81
Estimated break points:	2009			

Note: SupW(tau) refers to the sup-Wald test for the existence of a structural break. The null hypothesis is H_0, and the alternative hypothesis is H_1.

As seen in Table 6, we initially endogenously identified three potential structural breaks (located in the years 2005, 2008, and 2011) within the period under study. However, we cannot reject the null hypothesis (that there are no breaks). To further validate our findings, we conducted successive tests for the absence of breaks with two break dates (located in 2009 and 2011) and a single break date (located in 2009). Once again, the results obtained indicate that we cannot reject the null hypothesis of no break. This underscores the importance of our research, as we can confidently conclude that

[20] We would like to express our gratitude to an anonymous reviewer for suggesting that we conduct these robustness tests.

there is no significant change in the relationship estimated using the baseline empirical model, thereby adding significant value to the field.

5.3.3. Heterogeneous relationships

Since the paper covers a wide range of OECD countries, using a uniform model for all these diverse countries may overlook important country-specific factors, such as differences in pension systems, welfare policies, or economic structures. To gain further insights into the relation between pension expenditures and old-age poverty and to obtain additional evidence on the robustness of our results, we look for clusters in our sample to group countries with the same characteristics, maximizing commonalities and minimizing differences.

The more notable the differences between the groups (i.e., the more prominent the inter-cluster distances) and the larger the similarities within a group (i.e., the smaller the intra-cluster distances), the more distinct the clustering is. Then, we apply k-medians clustering to find a partition in which countries within each cluster are as close to each other and as far from the countries in other clusters as possible[21]. Each cluster is defined by its cluster center or centroid, the point at which the Canberra distance from all the countries is minimized (see Jain, 2010)[22]. By repeating the algorithm with a different set of randomly selected initial centroids and determining, among the various local minima, the one that creates graphs where each point in a group is closer to a point in the neighboring groups, we overcome the two possible limitations of the k-medians method (the selection of the number of clusters and the dependence of the results on the initial partition)[23]. Applying this methodology, the 37 countries in our sample are distributed into four groups (k = 4) with similar characteristics. The classification of the countries belonging to each group is listed in Table 7, and in Figure 3, the reader can find a mapping with the countries belonging to each group. As observed, the grouping formed includes countries with various regimes and types of welfare states discussed in the literature (e.g., Esping-Andersen, 1990; and Arts and Gelissen, 2002). This indicates that k-medians clustering, which considers unobserved heterogeneity, can reveal a more accurate classification than the traditional ones that may change over time and be influenced by ongoing reforms.

Table 7. Composition of the groups of countries detected by the k-medians cluster analysis.

	Countries
Group 1	Israel, Japan, Latvia, Lithuania, United Kingdom, and United States.
Group 2	Australia, Chile, Czech Republic, Estonia, Finland, Korea, Luxembourg, Norway, and Slovak Republic.
Group 3	Austria, Canada, Denmark, France, Germany, Hungary, Iceland, Ireland, Netherlands, New Zealand, Poland, Portugal, Slovenia, Spain, Sweden, Switzerland, and Turkey.
Group 4	Belgium, Costa Rica, Greece, Italy, and Mexico.

[21] As Cardot et al. (2012) showed, k-medians are more robust to outliers than k-means.

[22] The Canberra distance between two points *a* and *b* in *n* dimensional space is calculated as the sum of the absolute differences in each dimension, weighted by the inverse of the sum of absolute value: $\sum_{i=1}^{n} \frac{|a_i - b_i|}{|a_i| + |b_i|}$. Simulation results reveal that the application of Canberra distance yields more accurate load estimates than other distance functions such as the Euclidean or the Manhattan distances (Al-Wakeel et al., 2017).

[23] A detailed account of this process is not shown here to save space, but they are available from the authors upon request.

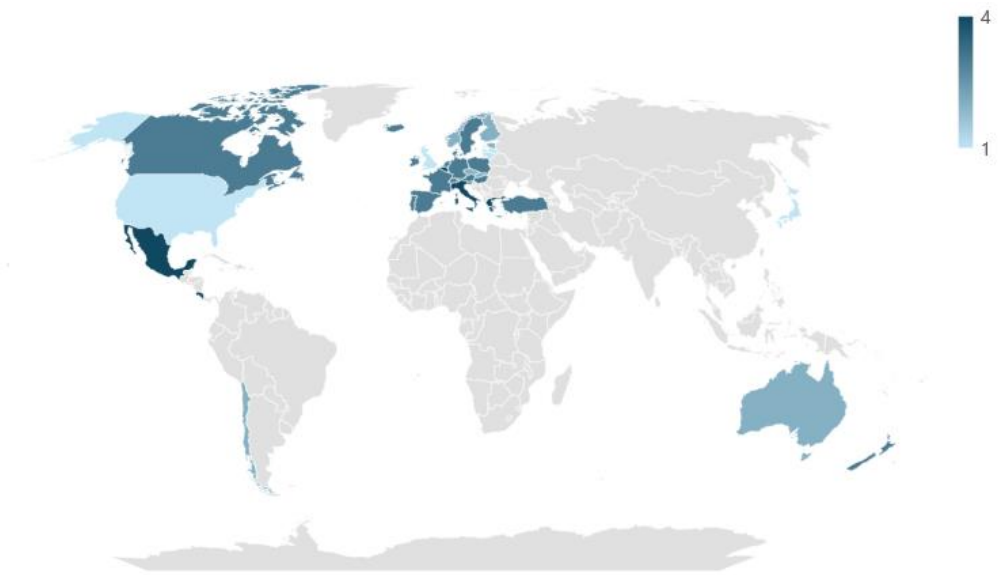

Figure 3. Groups of countries detected by the k-medians cluster analysis.

Once again, the Hausman test identifies the best method as the RE-2SLS for the poverty rate of the 66–75 age group and the FE-2SLS for the 76-and-over age group. Table 8 presents the results obtained by applying these estimation methods. As shown, the results suggest a nonlinear relationship between pension expenditures and old-age poverty rates for all country groups considered; this conclusion is valid for both age groups. While the magnitude of the effect differs, the estimated sign is the same for all groups. It is important to note that, except for group 4, the parameters fall within the 95% confidence interval we estimated for the total number of countries, suggesting that our results are based on a relatively homogeneous sample.

Table 8. Parameter estimates for the empirical model controlling for endogeneity (both age group 66–75 and 76-and-over poverty rates): Estimation by groups.

	Lpoverty66				Lpoverty76			
	Group 1	Group 2	Group 3	Group 4	Group 1	Group 2	Group 3	Group 4
$lpension_{it}$	0.5791***	0.5973***	0.5823***	5.6742***	0.4627**	0.4354**	0.0722**	4.4614***
	(0.2079)	(0.1680)	(0.1780)	(1.8746)	(0.2169)	(0.2205)	(0.1926)	(0.6690)
$lpension^2{}_{it}$	−0.4429***	−0.2136***	−0.2086***	−1.5436***	−0.2268***	−0.5714***	−0.2182**	−1.1265***
	(0.0999)	(0.0331)	(0.0985)	(0.2400)	(0.0782)	(0.1204)	(0.0861)	(0.1673)
$\Delta gini_{it}$	0.0014***	0.0725***	0.0361***	0.0469**	0.0201***	−0.0058	0.0724***	0.0721***
	(0.0176)	(0.0199)	(0.0184)	(0.0088)	(0.0073)	(0.0166)	(0.0112)	(0.0119)
$lgdp_{it}$	−0.9061***	−0.1183***	−0.2810***	−0.6963***	−0.3922***	−0.5077***	−0.0687***	0.1719
	(0.3778)	(0.0723)	(0.0587)	(0.0674)	(0.0795)	(0.1674)	(0.0896)	(0.1792)
$\Delta dependency_{it}$	9.0289***	3.7564	3.7248***	1.9240***	4.3297***	8.7212***	−4.6236***	−1.0456
	(2.2589)	(3.2458)	(0.0672)	(0.7381)	(1.0812)	(1.5905)	(1.4696)	(1.3782)

Continued on next page

	Lpoverty66				Lpoverty76			
	Group 1	Group 2	Group 3	Group 4	Group 1	Group 2	Group 3	Group 4
$govspending_{it}$	−0.0614***	−0.0032***	−0.0032***	−0.0745	−0.0641***	−0.0410***	−0.0280***	−0.0272***
	(0.0112)	(0.0152)	(0.258)	(0.0022)	(0.0063)	(0.0105)	(0.0062)	(0.0075)
$\Delta debt_{it}$	0.0017	−0.0097***	−0.0138***	0.0025***	−0.0042	0.0020	−0.0019	−0.0048***
	(0.0035)	(0.0029)	(0.0091)	(0.0038)	(0.0048)	(0.0036)	(0.0050)	(0.0017)
Constant	7.5598*	−5.7806***	−5.2716***	−4.6420***	2.2655**	3.5175**	−4.1692***	0.1201***
	(4.1926)	(2.4337)	(3.9840)	(3.9401)	(0.9828)	(1.5351)	(1.4327)	(1.1566)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: In the ordinary brackets below the parameter estimates are the corresponding *z*-statistics, computed using White's (1980) heteroskedasticity robust standard errors. In the square brackets below the specification tests are the associated p-values. *, **, and *** indicate significance at 10%, 5%, and 1%, respectively. Based on the Hausman test results, we apply the RE-2SLS for the poverty rate of the 66–75 age group and the FE-2SLS for the 76-and-over age group.

5.3.3. Multiple thresholds

The threshold model assumes that the relationship between pension expenditures and poverty changes abruptly at a given point in time. However, identifying a single threshold could oversimplify the relationship. That is why we adopt a more flexible approach, allowing for the possibility of multiple thresholds that could capture the complexity of the relationship more accurately. In particular, we test the possibility of multiple-threshold models. For this purpose, we apply a further test in which the null hypothesis captures a single-threshold model; in the alternative hypothesis, the double-threshold model is considered as follows:

$$y_{it} = \alpha_i + X_{it}(q_{it} < \tau_1)\delta_1 + X_{it}(\tau_1 \leq q_{it} < \tau_2)\delta_2 + X_{it}(q_{it} \geq \tau_2)\delta_3 + e_{it} \quad (9)$$

where τ_1 and τ_2 are the two threshold values and δ_1, δ_2 and δ_3 capture the three impacts for each regime.

Table 9 offers the estimation results. We compare the linear and the single-threshold models and compare the single-threshold with the double-threshold model. In both age groups, we clearly reject the null hypothesis of the linear model. However, we rule out the existence of a double-threshold model since we cannot reject the null hypothesis between the single and double models, further supporting our previous results.

Table 9. Multiple-threshold models and their impact on poverty rate.

Lpoverty66 Threshold estimator (level=95):			
Model	Threshold	Lower	Upper
Th_1	1.9756	1.9683	1.9804
Th_21	1.9756	1.9683	1.9804
Th_22	2.3046	2.2880	2.3151

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Threshold-effect test:							
Threshold	RSS	MSE	F-stat	Prob	Crit10	Crit5	Crit1
Single	44.3856	0.0822	38.41	0.0614	34.4394	40.0684	55.8758
Double	42.6060	0.0789	22.56	0.2938	32.7085	37.6947	49.1222
Lpoverty76							
Threshold estimator (level=95):							
Model		Threshold		Lower		Upper	
Th_1		1.9756		1.9689		1.9804	
Th_21		1.9756		1.9689		1.9804	
Th_22		2.1994		2.1833		2.2029	
Threshold-effect test:							
Threshold	RSS	MSE	F-stat	Prob	Crit10	Crit5	Crit1
Single	47.8582	0.0886	38.74	0.0488	32.2925	39.5421	53.4082
Double	45.5958	0.0844	26.79	0.1615	29.9388	35.6755	49.4871

Notes: In the threshold-effect test, we show the RSS, the mean squared error (MSE), the F statistic (F-stat), the probability value of the F statistic (Prob), and the critical values at 10%, 5%, and 1% significance levels (Crit10, Crit5, and Crit1, respectively). Th_21 and Th_22 refer to the two estimators in the double-threshold model. *Single* considers that the null hypothesis is the linear model, and the alternative hypothesis is the single-threshold model; *Double* captures a single-threshold model in the null hypothesis and a double-threshold model in the alternative hypothesis.

6. Concluding remarks

This paper addresses the ongoing debate regarding the effectiveness of public pension systems in providing adequate protection against income poverty for future pensioners. Given that old-age pension schemes are a crucial component of social protection, we have examined whether these systems sufficiently meet the needs of the elderly. In particular, we have attempted to shed light on the relationship between elderly poverty rates and pension expenditures by applying panel data techniques to data from 37 OECD countries during a period of almost 40 years (1980–2020), controlling for demographic and macroeconomic factors.

Our main empirical results suggest that pension expenditure shows a linear and nonlinear impact on old-age poverty rates. The estimated elasticities increase (in absolute value) with the ratio of public pension expenditure to GDP, being higher (in absolute value) for the 76-and-over age group than for the 66–75 age group, suggesting that the impact of pension expenditure on poverty is larger for the oldest population group. When evaluating how each explanatory variable explains the poverty rates, we find that the macroeconomic situation is the main driver of its reduction, followed by public pension expenditure. Finally, applying the panel dynamic threshold model, we endogenously identify a common threshold of 7.21 in the ratio of public pension expenditures to GDP in the relationship between this variable and the poverty rate of both age groups. The robustness of our results is confirmed by a sensitivity analysis that involves testing structural breaks and heterogeneous relationships in the baseline empirical model and exploring the possibility of multiple thresholds in the threshold model.

The findings of this study significantly contribute to the discourse on age, income, financial satisfaction, and retirement. They carry substantial policy implications, revealing that pension

expenditures play a crucial role in reducing poverty rates among the elderly. Moreover, our estimates propose that focusing on the age group of 76 and older could lead to an even more pronounced relative impact on poverty reduction, underscoring the importance of this research in shaping future policies.

Given the premise of this paper, namely that public spending is flexible, it would be worth examining, as a natural extension, how intragenerational redistribution, intergenerational redistribution, and lifetime income stabilization create a partisan and institutional problem. Rapid economic development, industrialization, and urbanization have posed significant challenges to the sustainability and solvency of OECD pension expenditures. This also underscores the complexity of reducing old-age poverty in OECD countries, a task that requires a combination of public pension funds, individual voluntary plans, and company-funded pension plans. Another interesting direction for future studies would be to investigate how societal change influences pension systems' sustainability. The increasing number of elderly people living alone is not just a personal issue but one with significant effects on the national economy, exacerbating poverty and living expenses at a time when the workforce is shrinking, and the number of elderly people in need of pension support is rising[24]. These are the areas that our future research agenda aims to explore, with the potential to significantly impact social welfare and economic policy.

Use of AI tools declaration

The authors declare they have not used Artificial Intelligence (AI) tools in the creation of this article.

Author contributions

All authors contributed to the study's conception, design, Methodology, formal Analysis, validation and writing.

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Conflict of interest

All authors declare no conflicts of interest in this paper.

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