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The incidence of the Argentine pension system on income distribution: A cross-sectional analysis

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Abstract This article examines the impact of the Argentine pension system on income distribution through a cross-sectional analysis. It considers contributory, non-contributory, and semi-contributory benefits, along with the different financing sources. Using data from the 2017–2018 *Encuesta Nacional de Gasto de los Hogares*, an accounting incidence approach and various inequality indices are applied to assess the redistributive effect under different scenarios and assumptions. The analysis also decomposes redistribution into intra- and inter-age effects. Results indicate that the pension system had a progressive impact on income distribution during the period analysed, although the extent of this effect varies across the scenarios.

Keywords Argentina, income redistribution, pension scheme, retirement, social security financing, social security scheme

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Introduction

From the point of view of the individual, a pension system has two basic objectives. The first is to smooth consumption over the life cycle, i.e. to allow individuals to shift consumption from their productive years to their retirement years to maintain a relatively constant standard of living over time. The second is to insure against longevity risk, understood as uncertainty about the length of an individual's life. Such uncertainty refers to the possibility that individuals will survive longer than expected, which would imply an underestimation of the savings needed to support themselves in retirement (Barr and Diamond, 2006).

From a public policy perspective, there may be two additional objectives to be pursued through the pension system. The first is poverty alleviation, which consists of guaranteeing the retired population a minimum income threshold that allows them to meet their basic needs. The second is income redistribution, both intergenerational and intragenerational (Barr and Diamond, 2006). In this sense, while the first refers to the redistribution of income between different generations, the second refers to the redistribution of income between members of the same generation (Barr and Diamond, 2006). Likewise, these objectives should be pursued in a way that does not negatively affect economic growth.

With respect to the redistributive objective, there has been much controversy about the impact of defined benefit pension systems on income redistribution. For instance, their regressive effects have been emphasized, as they would tend to favour higher-income individuals who have a longer life expectancy and enter the labour market at an older age. In this sense, it is argued that individuals with greater resources will receive benefits for a longer time after having contributed fewer years (Friedman, 1962; Friedman and Cohen, 1972; Holzmann, 1998; Schwarz, 2006). However, authors such as Beattie and McGillivray (1995) and Barr (2002) recognize that although perverse redistribution from the poor to the rich is inevitable in a system that seeks to diversify the longevity risk of an entire population, lower-income individuals reaching retirement age would receive higher replacement rates due to the lump-sum components of pension benefits and the existence of universal minimum benefits, thus highlighting the progressivity of pensions.

Recent empirical studies for Argentina provide evidence supporting the relevance of these longevity-driven regressive mechanisms. Bramajo and Grushka (2020) show that mortality risks among adults aged 65 or older vary systematically with pension income, with higher-income beneficiaries experiencing lower mortality probabilities and receiving benefits for longer periods. Consistently, Argentina's National Social Security Administration (ANSES, 2024) confirms persistent longevity differentials across income groups, finding that each doubling of pension income is associated with an increase of approximately 0.6–0.8 years in life expectancy at age 65. These findings indicate that socio-economic differences in longevity remain significant in Argentina.

To know whether a pension system has a progressive or regressive impact on the income distribution in a country, it is necessary to conduct an incidence analysis. Numerous incidence analyses of fiscal policies carried out in different countries can be cited (Sung and Park, 2011; Wang, Caminada and Goudswaard, 2012; Djindjić, 2014; Hwang, 2016; Lustig, 2017). However, to the best of our knowledge, there are no incidence analyses that isolate the redistributive effect of the pension system at a given point in time (both its sources of financing and its benefits), disregarding the rest of the fiscal intervention programmes. In some countries, analyses of the impact of social security systems (including the pension system) have been carried out. In this sense, the work of Oshio (2002), Bruckmeier and Schwengler (2010), He and Sato (2013) and Cai and Yue (2020) can be mentioned. The work of Oshio (2002) stands out among the others because it evaluates the redistributive effect from three perspectives: i) the distribution of income within the same age group, ii) the distribution of income between different age groups, and iii) the distribution of lifetime income, which is highly relevant when dealing with a public programme that involves different generations. Carrying out this type of exercise and comparing the results of different countries is extremely useful, not only to answer the theoretical question about the progressivity or regressivity of pension systems, but also to investigate the causes of the lower efficiency of fiscal policy in progressively affecting income distribution in Latin American countries (Martorano, 2018) compared to European Union countries (Goñi, López and Servén, 2011).

In this context, this article seeks to answer the following research question: to what extent does the Argentine pension system reduce income inequality, and through which channels – intergenerational or intragenerational redistribution – does it operate most effectively? Given the characteristics of the Argentine system – particularly its high coverage achieved through moratoria (pension regularization programmes introduced in 2005 that allowed individuals with incomplete contribution histories to access pensions by regularizing or "buying back" missing years of contributions) (Geri, de Santis and Moscoso, 2019) and non-contributory benefits that have significantly expanded coverage, especially among low-income groups (Arza, 2012; Mesa-Lago, Cruz Saco and Gil, 2021) – our hypothesis is that most of the redistributive effect is explained by intragenerational redistribution among the elderly, rather than redistribution between working and retired individuals. This hypothesis is closely linked to the demographic and labour-market conditions that shape the functioning of the pension system and determine the relative importance of each redistributive channel.

Argentina is entering a phase of sustained population ageing. The share of individuals aged 65 or older increased from 10.6 per cent in 2010 to 12.42 per cent in 2024 and is projected to reach 16.4 per cent by 2040 (World Bank, 2024; INDEC, 2025).[1] Consistent with this trend, the old-age dependency ratio – defined as the population aged 65 or older relative to the working-age population (aged 15–64) – rose to 18.82 per cent in 2024 and is expected to approach 25 per cent by 2040 (World Bank, 2024; INDEC, 2025). These dynamics suggest growing pressure on a pay-as-you-go (PAYG) system already characterized by structural imbalances.

Labour-market conditions further shape the functioning of the pension system. Labour informality constitutes a central structural constraint for the contributory PAYG scheme in Argentina. According to EDIL (2024), the share of informal salaried workers reached 36.4 per cent in the second quarter of 2024, a level consistent with the historical range of 32–36 per cent observed since the early 2000s. Informality is particularly prevalent among young workers (around 57 per cent) and among women (approximately 37 per cent), reducing contribution density and increasing the incidence of incomplete careers. These structural features help explain both the persistent coverage gaps in old-age protection and the continued relevance of pension moratoria and non-contributory benefits in expanding access.

Against this background, this article contributes to the literature by presenting the first incidence analysis of the Argentine pension system that decomposes the redistributive effect into inter-age, intra-age, and incompleteness components. Drawing on Oshio's (2002) analytical framework, the study tailors the approach to the specific institutional configuration of Argentina – a country characterized by widespread labour informality, high coverage achieved through moratoria and non-contributory benefits, and a complex mix of contributory and tax-based financing. These features offer a unique opportunity to explore the redistributive implications of pension systems under conditions markedly different from those of high-income countries, and to examine how non-contributory instruments shape both intra- and intergenerational equity.

In addition, the article conducts a comprehensive set of sensitivity analyses, varying key assumptions such as the classification of pension income (as a government transfer or deferred income), the inclusion of the pension deficit as part of the redistributive mechanism, and the scope of income definitions. These exercises not only assess the robustness of the main findings but also highlight how conceptual choices in incidence analysis can substantially affect the interpretation of redistributive outcomes.

1. World Bank data retrieved from the World Bank databank.

This article is structured as follows. The next section describes the general regime of the Argentine pension system, detailing its sources of financing and eligibility conditions for access to benefits. In turn, we present the methodological approach of the incidence analysis of the Argentine pension system, and then describe the database used. We define operationally the different types of income included in the analysis by means of mathematical equations, before explaining how the inter- and intra-age redistributive effects are measured and the characteristics of the concentration indices used. We explain which parameters were varied in the sensitivity analysis before presenting the results obtained for each proposed scenario, as well as the results obtained from the sensitivity analysis, and then offer our concluding remarks.

The Argentine Integrated Pension System

The general system of the Argentine Integrated Pension System (SIPA) is organized on a pay-as-you-go (PAYG) basis with defined benefits. It provides benefits for old age, disability and survivorship. All persons over 18 years of age who are employees or who individually or jointly engage in a lucrative activity are compulsorily insured (Ley N° 24.241, 1993). SIPA benefits are primarily contributory and are financed predominantly by contributions from employees and their employers. Specifically, workers' contributions amount to 11 per cent of gross wages (Ley N°24.241, 1993) and their employers' contributions should be 16 per cent of gross wages (Ley N° 25.453, 2001; Ley N° 25.565, 2002). The personal contribution of the self-employed is 27 per cent of the presumed income in the case of the self-employed (Ley N° 24.241, 1993) and 11 per cent of the gross income in the case of workers registered in the simplified regime (Ley N° 24.977, 1998).

SIPA also receives funding from general taxes. Most of the non-contributory financing comes from consumption taxes, namely: i) 11 per cent of the VAT collection, ii) 28.69 per cent of the liquid fuels collection, iii) 100 per cent of the additional tax on cigarettes, vi) 70 per cent of the simplified regime tax and vii) 100 per cent of the tax on banking transactions, known as the cheque tax (Ley N° 23.966, 1991; Ley N° 26.028, 2005; Ley N° 24.625, 1995; Ley N° 27.430, 2017).

Table 1 summarizes, for each SIPA pension benefit (contributory, semi-contributory and non-contributory), the conditions of access and the criteria for determining the average benefit and/or average assets (in US dollars – USD) received by its beneficiaries for 2018.

The contributory pension benefits provided by the SIPA to the elderly are: i) *Prestación Básica Universal* (PBU), ii) *Prestación Compensatoria* (PC), iii) *Prestación Adicional por Permanencia* (PAP), and iv) *Pensión por Fallecimiento* (PF). The PBU benefit is a fixed amount that is subject to periodic updating. This benefit is

Table 1. *Retirement benefits under the SIPA*

Contributory benefits				
Benefit	**Requirements**		**% wage**	**Benefit amount**
PBU	*Men* 65 years	*Women* 60 years	-	-
PC=PAP	30 years of service with contributions. Qualify for PBU, not receiving disability retirement.		45%[1]	Average earnings 10 years before retirement.
PF	Being a widow, widower or cohabitant of the deceased.		70%	Average remuneration 5 years before termination.
Semi-contributory benefits				
	Requirements		**Average benefit with moratoria**	**Average benefit without moratoria**
	Men	*Women*		
Moratoria Previsional 2004	65 years	60 years	USD 409.83	USD 720.84
	Debts for contributions prior to September 1993			
Moratoria Previsional 2014	*Men*	*Women*	USD 409.83	USD 720.84
	65 years	60 years		
	Debts for contributions prior to December 2003.			
Moratoria Previsional 2016	Only women aged 60 years or older but younger than 65 years. Debts for contributions prior to December 2003.		USD 409.83	USD 720.84
Non-contributory benefits				
	Requirements			**Average benefit**
PNCV	70 years of age or older. Not receiving, neither the applicant nor their spouse, any other benefit from any pension scheme. Not possessing resources that allow for their own subsistence and that of their cohabiting family group. In the case of a marriage, old-age pension will only be processed in favor of one spouse.			USD 282.74
PUAM	65 years of age or older. Not receiving or being entitled to any benefit from any pension scheme. Maintaining residence in the country once the pension has been applied for.			USD 318.80

Note: In the case of 30 years of contributions.

Sources: Author's elaboration based on Leyes 25.865 (2004); 26.970 (2014); 27.260 (2016), Decreto 432/97; and Boletín estadístico de la seguridad social (ANSES, 2017).

intended for members who have reached the minimum retirement age for each sex (age 60 for women or age 65 for men) and have completed 30 years of contributory service, with the possibility of compensating for each missing year of service with two years of additional age. Since the amount of the PBU is the same for all beneficiaries, regardless of their income during the active phase, its existence contributes to the solidarity of the system, since both those who earn a minimum wage and those who reach a high salary have access to the same basic benefit (MTEySS, 2011).

In contrast, the PC and the PAP operate as a single benefit,[2] access to which requires qualification for the PBU and no disability retirement. The amount of both benefits is equal to 1.5 per cent of the average updated salary received during the ten years prior to retirement, multiplied by each year of service up to a maximum of 35 years; i.e. they are equal to 45 per cent of the pre-retirement salary assuming 30 years of contributions. Like the PBU, these benefits are indexed according to the evolution of inflation and wages (Ley N° 27.426, 2017). Finally, the PF asset is equal to 70 per cent of the pre-retirement salary and requires being a widow, widower or partner of the deceased worker (Ley N° 24.241, 1993).

With regard to semi-contributory pension benefits, these are those that originated with the implementation in 2005 of the *Plan de Inclusión Previsional*, hereafter PIP (Ley N° 25.994, 2004; Decreto N° 1454/5, 2005). The objective of this plan was to achieve the inclusion in the pension system of people who were not eligible for retirement benefits (Bravo Almonacid, 2013). The two pillars of the PIP were the *Prestación Previsional Anticipada* (PPA) and the *Moratorias Previsionales* (MP). The former allowed access to a pension benefit for unemployed older adults who, after accrediting 30 years of contributions, were five years short of the retirement age (age 60 for men or age 55 for women). The MP provide for two types of pension debt regularization schemes, the *Moratoria Previsional Permanente* (MPP) and the *Moratorias Previsionales Excepcionales* (MPE). The MPP is a continuous scheme that was reopened in January 2004 and started with the enactment of Ley N° 25.865 (2004). This law updated the special regime for the regulation of the social security obligations of self-employed persons established by Ley N° 24.476 (1995). In 2004, this regime was made permanent for self-employed workers who owed contributions prior to September 1993 (Decreto N° 164/04, 2004). The MPE, carried out between January 2005 and April 2007, established that any person who met the age requirements could access the moratorium and enjoy the benefits of a social security benefit, regardless of whether he or she had contributed to the system. As a result, this measure extended the coverage to the entire adult population of

2. Between 1994 and 2008, the PAP benefit was lower than the PC benefit to encourage workers to leave the public PAYG system and transfer their contributions to a private pension fund manager. The abolition of the funded subsystem in 2008 equalized the replacement rates of the two benefits.

retirement age, without specifying any requirements regarding their work history (Madera, 2012).

Finally, with regard to non-contributory benefits for the elderly, the SIPA grants: i) the *Pensión no Contributiva por Vejez* (PNCV) (Decreto N° 432/97, 1997) and ii) the *Pensión Universal para el Adulto Mayor* (PUAM) (Ley N° 27.260, 2016). In particular, the PUAM is intended for all persons aged 65 or older who are not covered by old-age pension retirement or unemployment insurance. Its creation was aimed at maintaining the pension coverage standards achieved by the moratorium plans and consists of a monthly payment equal to 80 per cent of the minimum pension benefit, which in turn is equal to 82 per cent of the minimum living and mobile wage as of December 2017. As of December 2017, the average PUAM benefit was USD 318.80, while the average non-contributory old-age pension benefit was USD 282.74, which is why the PUAM is expected to gradually replace the PNCV (the requirements are more flexible and the benefit is higher).

Regarding benefit adequacy and coverage during the period analysed, Calabria and Gaiada (2019) report that pension coverage among individuals of retirement age reached 89.6 per cent in 2018 (rising to 95.1 per cent when including those who remained employed). More than three-quarters of beneficiaries received the minimum pension, while fewer than 4 per cent obtained more than three times the minimum. They also estimate average replacement rates of approximately 60–67 per cent for contributory pensions, around 30–35 per cent for semi-contributory benefits, and roughly 25 per cent for non-contributory schemes and the PUAM. Consistent with this general pattern, Durán, Geri and González (2021) find that replacement rates for a typical contributory pension were around 45–52 per cent of pre-retirement earnings, with minimum pensions representing a substantially lower proportion of average wages. Semi-contributory and non-contributory benefits were even less adequate. Moreover, when compared with basic consumption baskets estimated specifically for older adults, transfers such as the PUAM may fall short of ensuring consumption smoothing or preventing income deprivation in old age.

These country-specific patterns align with broader regional trends. Using comparable survey data, Gaiada, Calabria and Guinsburg (2022) show that pension coverage in Latin America increased from 67.4 per cent to 87.1 per cent between 1996 and 2015, driven largely by the expansion of non-contributory and semi-contributory schemes. Over the same period, the average replacement rate in the region rose from 37.4 per cent to 51.7 per cent. This comparative evidence underscores that Argentina's high coverage levels – exceeding 90 per cent by 2015 – reflect the impact of pension moratoria and other inclusion mechanisms.

From a macro-fiscal perspective, the public pension scheme operates with a persistent financing gap. In 2018, the SIPA recorded a deficit of about

2.7 per cent of GDP, after considering both contributory and semi-contributory components (Bour, Susmel and Urbiztondo, 2019). This shortfall is financed through the general revenues and earmarked taxes previously described, underscoring the system's structural dependence on tax-based funding rather than on social security contributions alone.

Incidence analysis: Methodological approach

Redistribution can be analysed through both cross-sectional and life-cycle perspectives. This study adopts a cross-sectional incidence approach, which is particularly suited to evaluating the redistributive impact of the pension system at a specific point in time. This perspective enables the analysis of how the current structure of the system reallocates income across individuals and age groups, based on observed income, transfers, and financing flows. In contrast, life-cycle approaches aim to assess redistribution across an individual's lifetime, often incorporating expected benefits, contribution histories, and survival probabilities. Both approaches are methodologically valid and offer complementary insights: whereas life-cycle models are well suited to studying long-term actuarial equity and lifetime redistribution, cross-sectional analyses are particularly informative for assessing the immediate distributive effects of the system as it operates in practice.

The objective of any incidence analysis of an existing public expenditure programme is to compare the current distribution of income (given the existence of the programme) with a counterfactual distribution (in the absence of such a programme), taking into account the distributional impact of both the expenditure and its financing sources. Since this study focuses on the current pension system in Argentina (rather than, for example, analysing a hypothetical reform), it is an ex-post incidence analysis. The income observed in the presence of the programme is divided into income that is due to the existence of the programme and income that is not due to the existence of the programme. In contrast, the counterfactual income is the income that would result if the programme did not exist. According to Gasparini, Cicowiez and Sosa Escudero (2012), there are several reasons why this counterfactual income may differ from the income that does not originate in the existence of the programme. This is because there are behavioural effects that occur when the government decides to intervene or not to intervene in the market: individuals respond to the incentives created by the existence or non-existence of a programme.

Thus, for example, the existence of the pension system – as currently defined by the rules that created and regulate it – generates certain incentives in economic agents. If the pension system did not exist or if its rules were different, the incentives would likewise be different, and economic agents would probably react differently to such incentives. In particular, it is likely that the savings-consumption and labour supply patterns of the agents would have been different to smooth their intertemporal consumption, affecting the values of income during both the active and passive stages.

Determining how individuals would react in the absence of the pension system is complex. One approach that attempts to account for behavioural effects is the general equilibrium approach, which assumes specific functional forms for both individual utility and firm production functions, and a government budget constraint. The disadvantage of these models, however, is that they necessarily start from very stylized (and restrictive) assumptions about the preferences of individuals, the degree of heterogeneity among them, the way in which capital and labour determine output in an economy, and the way in which the pension system operates, among others. Following Moncarz (2015), for example, a weakness of these models is the assumption of full rationality, which has been the subject of much controversy, especially regarding long-term decisions such as those related to social security. Finally, one of the most attractive arguments to justify the existence of pension programmes is the myopic behaviour of individuals (Samuelson, 1975).

Since a full treatment of the effects of the absence of a pension system on individuals' labour supply and saving patterns would significantly increase the complexity of the analysis, an accounting approach is adopted. This means that behavioural effects are ignored, implicitly assuming that individuals currently receiving a pension would not have changed their behaviour during their working lives in the absence of a pension system. According to Lustig (2018), the main drawback of this approach is that it tends to overestimate the redistributive effects of the system, since individuals who have retired from the labour market and receive income only from the pension system would receive zero income in the absence of the system. This is because it is assumed that they would not have changed their behaviour during the active phase to ensure a minimum level of income in old age. However, this weakness of the accounting approach can be overcome (at least partially) by considering pension income as deferred income rather than as a pure transfer from the government (Lustig, 2018).

Indeed, on the one hand, if pension income is considered as a transfer from the government, the contributions that finance such pensions should be considered as taxes, and the application of the accounting approach will lead to assigning arbitrarily low pre-system income values to families composed of elderly adults of retirement age. On the other hand, the treatment of pensions as deferred income is based on the fact that active workers may consider pension contributions not as taxes but as compulsory savings. Therefore, it is expected that even in the absence of the pension system, such savings would have existed on a voluntary basis, making it possible to generate a floor of income for the retirement period.

Although the weak link that the Argentine pension system presents between the value of individual contributions and pension rights would make it more natural to consider pensions as government transfers, in the following analysis two scenarios will be analysed. First, we consider a benchmark scenario, in which pension income is considered as a transfer in its entirety. Second, we consider an alternative scenario (which we will call the hybrid scenario), in which it is assumed that part of the retirement income and pensions correspond to transfers, while another part corresponds to deferred savings. Following Lustig (2018), in the hybrid scenario, the part of the retirement income or pension considered as a transfer is calculated as ωY_p, where ω is the ratio between the pension system's contribution deficit and total expenditure on benefits, and Y_p is the amount of the benefit. Similarly, different variants of the hybrid scenario are simulated, assuming different levels of contributory deficits, to assess their impact on redistribution. These simulations constitute a key contribution of this article, as they allow for a detailed sensitivity analysis that highlights how conceptual choices regarding pension income classification and financing assumptions can substantially alter the measured redistributive effect.

For the benchmark scenario, the starting point is the definition of *pre-system income*, which requires estimating household income in the complete absence of the pension system, i.e. excluding pensions from their observed income and adding to it any payments made to finance the system. In contrast, *post-system income* would be the income observed in reality when the pension system is in operation, i.e. *pre-system income* minus taxes and contributions plus pension benefits. For the hybrid scenario, *pre-system income* implies subtracting from observed income only the part of pensions valued as transfers (leaving the part considered as deferred income). The *post-system income* is exactly the same as in the basic scenario: observed income.

Finally, it should be clarified that although it would be expected that, in the absence of the system, changes in the behaviour of families or individuals could alter the amount to be paid in taxes not affected by the financing of the pension system (which would affect the pre-system income), following the logic of the accounting criterion adopted in this work, it is assumed that the amount of tax payments made by families and companies not affected by the financing of the pension system will not be affected by its absence and, therefore, will not have any effect on the *pre-system income* of families or individuals. Thus, the analysis to be carried out allows us to isolate (under the assumptions adopted) the distributional impact of the taxes that finance the SIPA from the rest of the taxes.[3]

3. Although the amounts to be paid as a fourth category of income tax are estimated (without this tax having affected the financing of the pension system during the period analysed), the only purpose of this calculation is to allow a correct estimation of the pre-system income, considering the available data referring to the net income of the households.

Data base

To conduct the incidence study, the databases corresponding to the *Encuesta Nacional de Gasto de los Hogares* (ENGHo), conducted between November 2017 and November 2018 in Argentina, were used. The ENGHo collects income and expenditure data from approximately 45,000 households located in urban centres with 2,000 inhabitants or more throughout the country, in order to know the variation of consumption habits according to the different seasons of the year. In this sense, the bases of persons, households and consumption were unified, thus having information corresponding to 21,530 households. Likewise, the variables were weighted with the weights present in the databases. In this sense, the ENGHo uses an expansion factor for households, which makes it possible to obtain all the estimates of interest in the survey. Initially, this factor corresponds to the factor of the dwelling to which it belongs, which results from the sampling design and includes the product of the inverse of the inclusion probabilities of the dwelling (ENGHO, 2019). Thus, the use of weights makes it possible to extrapolate the results to a universe composed of 12,632,469 households.

Further, given that the survey was conducted over a year, income and expenses were deflated by taking into account inflation in each quarter, expressing all monetary units at the values of the two-month period November 2017–December 2017. This makes the values comparable. In addition, all amounts were expressed in US dollars according to the average rate between November and December 2017 (USD 17.75).

It is worth mentioning that since we are working with a household survey, our observation unit is the household and not the individual; therefore, the observed income is Total Household Income (THI). For the purposes of this article, we define THI as the sum of labour income, both from the main occupation and from other occupations, and non-labour income, that is, taking into account pensions, inter-household transfers, rents, subsidies and own consumption. However, according to Gasparini (1998), the use of THI ignores the fact that the level of well-being depends not only on family income but also on the composition of the household. Thus, given that people live in households where the budget is shared, the standard of living of a person is related to the total income of the household and its composition (Gasparini, Cicowiez and Sosa Escudero, 2012). For this reason, the analysis is complemented by using household income per capita (HIP) as a measure of well-being, which is the most widely used indicator of individual well-being in poverty and inequality studies (Gasparini, Cicowiez and Sosa Escudero, 2012). However, to replicate Oshio's (2002) analysis, the squared coefficient of variation (SCV) will be computed using the THI, while the rest of the indicators will be computed for the HIP.

Operational definition of the different types of income in each scenario

Since the contributions to the pension system are calculated based on gross income, and considering that the income of dependent employees declared in the ENGHo corresponds to net income, it was necessary to estimate the value of gross income to determine more precisely the value of contributions to the system.

The challenge lies in the validity of the fourth category of income tax (see below) in the period under study. In fact, in order to convert the income declared by individuals in the surveys (net income) into gross income, it must be taken into account that the net income is the result of deducting from the gross legal wage not only the contributions that finance the different social security subsystems (which represent a fixed percentage of the gross wage) and other elements that may or may not be related to the gross legal wage (such as compulsory insurance and trade union dues, for example), but also, where applicable, the payment of the fourth category income tax.

When calculating the fourth category of income tax, to obtain the net taxable income (NTI), the following are first deducted from the gross statutory wage (GSW): i) social security contributions, ii) payments for union dues, iii) the corresponding personal deductions, and iv) a fixed amount called non-taxable income. The amount of tax payable based on the value of the NTI is determined by taking into account the income range or bracket in which this value falls.[4] Each income bracket is associated with a fixed amount to be paid and a tax rate applied to the difference between the NTI and the lower limit of the corresponding income bracket. It is worth mentioning that, given the progressive nature of the tax, the higher the range in which the taxpayer's NTI falls, the higher both the fixed amount to be paid and the tax rate applicable to the difference.

Thus, the net wage (NW) declared by formal employees in a dependent relationship is expressed in equation (1).

$$\mathrm{NW} = NTI - FA_c - g_c\mathrm{S} + NTM \tag{1}$$

Where NTI is the net taxable income, FA_c is the fixed amount that depends on the bracket c in which the NTI falls, g_c is the rate applicable to the corresponding

4. The fourth category of income tax, which was in force at the time of the study, adopted the so-called "graduated progressivity" system. It divides net taxable income into brackets or steps (defined in the legislation) and applies a different rate to each bracket (which increases as one moves from one bracket to the next). In practice, this is equivalent to charging a fixed amount depending on the bracket in which the GNSI falls and applying a tax rate to the difference between the GNSI and the lower limit of the corresponding bracket (which also increases as we move from a given bracket to the next higher one).

bracket c, S is the surplus to which the rate g_c is applied, and NTM is the non-taxable minimum that is added back after being subtracted from the GSW to calculate the NTI.

It should be noted that the calculation of the NW expressed in equation (1) assumes that the taxpayer, at the time of calculating the NTI, does not make any type of deduction for personal deductions, union dues or family expenses. In fact, taking these elements into account requires information on certain personal characteristics of the taxpayer that is difficult (if not impossible) to obtain from the ENGHO, such as whether the taxpayer is a member of a trade union, whether the income of the taxpayer's spouse or common-law spouse is less than the non-taxable minimum, whether the children living in the household are the taxpayer's legitimate children and, if they are between the ages of 18 and 21, whether they are attending college.

Equation (2) shows how the excess over net income is calculated.

$$\mathrm{NW} - NTM + FA_c = NTI - g_c(NTI - LB_c) \tag{2}$$

Where LB_c is the lower bound of the net profit interval c. Rearranging the terms and clearing the NTI yields equation (3), which expresses the net profit before taxes as a function of the observed net wage, the non-taxable minimum, and the fixed amount, rate, and lower bound of the profit category interval.

$$\frac{NW - NTM + FA_c - g_c LB_c}{(1 - g_c)} = NTI \tag{3}$$

After defining equation (3) and knowing the different possible values of FA_c, g_c y LB_c during the 2017 tax period, a categorical variable with ten categories was constructed, where category 0 corresponds to people who do not pay income (have a net salary less than or equal to the non-taxable minimum), category 1 corresponds to people who are in the first net income interval, category 2 corresponds to people who are in the second net income interval, and so on. Once this variable is constructed, the net income before taxes for each person is obtained and then the gross wage (GW) is defined from it according to equation (4).

$$\mathrm{GW} = \frac{(NTI + NTM)}{\left(1 - \tau_{ap}\right)} \tag{4}$$

Where τ_{ap} is the contribution rate corresponding to the employee and is the sum of 3 per cent to health insurance, 3 per cent to the National Institute of Social Services for Retired Persons and Pensioners (INSSJyP) and 11 per cent to the SIPA.

Once the gross salary was obtained, it was considered that in the pre-system situation (without the pension system), workers receive a net salary according to equation (5), where only $\tau_{ap} = 6$ *per cent* is deducted.

$$\mathrm{NW} = \mathrm{GW}\left(1 - \tau_{ap}\right) \tag{5}$$

Benchmark scenario

The hypothetical income of each household before the pension system is obtained by deducting from the THI the income from retirement and pensions, returning to them at the same time the amounts paid as contribution rates of formal workers and the fraction destined to finance the SIPA from VAT and taxes on the consumption of diesel, other liquid fuels and cigarettes of all households. Equation (6) summarizes the calculation of pre-system income:

$$PSI_h = GW_h\tau_a + NW_h + VAT_h + tfuel_h + tdiesel_h + tkerosene_h + tcigarettes_h - R\&P_h \tag{6}$$

The variable PSI_h represents the pre-system income of the household, which will depend, among other factors, on the activity category, employment status and consumption level of each individual member of the household. For active individuals employed in a formal dependent relationship, the employee contribution rate (τ_a = 11 per cent in the general regime) is added back to household income, where GW_h denotes the gross wage calculated above and NW_h the household's net wage (reported in the survey). For inactive individuals receiving retirement or contributory pensions, pension income ($R\&P_h$) is deducted from household income. Finally, the fraction of the amount paid as value added tax that is destined to the pension system (VAT_h) must be added to the pre-system income, as well as that corresponding to other consumption taxes such as the tax on diesel oil $tdiesel_h$, fuel $tfuel_h$, kerosene $(tkerosene_h)$ and cigarettes $tcigarettes_h$. In such regard, the statutory VAT rate amounts to 21 per cent.

Equations (7) to (11) show the amounts to be returned to households for each type of tax:

$$VAT_h = \left[\frac{Total\ Expenditure_h}{(1 + vat)} \times vat\right] \times 0.11 \tag{7}$$

$$tfuel_h = \left[0,379 \times q_{fuel}\right] \times 0,\ 2869 \tag{8}$$

$$tdiesel_h = [0,234 \times q_{diesel}] \times 0,\ 2869 \tag{9}$$

$$tkerosene_h = [0,234 \times q_{kerosene}] \times 0,\ 2869 \tag{10}$$

$$tcigarettes_h = \frac{Cigarette\ Expenditure}{(1 + vat + 0.07)} \times 0.07 \tag{11}$$

Equation (7) indicates that for each household h with positive consumption, its gross consumption is calculated, then it is multiplied by the effective VAT rate to obtain what is paid for this tax. Finally, 11 per cent is applied, which is what is intended to finance the pension system.

In equations (8) to (10), to obtain the amount that finances the system, the amounts consumed are multiplied by the corresponding fixed amount of each tax, and finally 0.2869 is applied, which is destined to the pension system.

Finally, for those households that consume cigarettes, their gross expenditure is calculated and the 7 per cent corresponding to the cigarette tax is applied (equation (11)).

Hybrid scenario

For the hybrid scenario, the calculation of the pre-system income of each household is similar to that of the benchmark scenario, except that only the part of the pensions that are considered transfers (and not all of these pensions) are removed from income, as well as the contributions that are considered deferred savings. Equation (12) formally defines the PSI_h in this scenario.

$$PSI_h = GW_h\tau_a + NW_h + VAT_h + tfuel_h + tdiesel_h + tkerosene_h + tcigarettes_h - (1 - \omega)R\&P_h \tag{12}$$

Where ω is the fraction of the pension that is considered as a transfer from the government.

Measurement of inter- and intra-age redistributive effects

Once the income of each household was obtained with and without the pension system, different inequality indices (Gini, Theil, and Atkinson) were applied to both the pre-system and post-system per capita incomes to determine the redistributive effect. To compute the intra- and inter-group redistributive effects, following Oshio (2002), the squared coefficient of variation (SCV) was applied to the household's total pre-system and post-system income which, in addition to satisfying the Dalton-Pigou properties, is scale invariant, and symmetry, it also satisfies the property of additive decomposability (which Gini and Atkinson do not) and, as an advantage over Theil, allows us to dimension an effect called the "incompleteness effect", which, as will be explained later, is relevant in the analysis of the incidence of a pension system.

To carry out the analysis, households were divided into six age groups according to the age of the head of the household at 10-year intervals. Once the groups were defined, equations (13) and (14) were applied to obtain the intra- and inter-age redistribution effects according to the SCV.

$$SCV = \frac{V}{\mu^2} \tag{13}$$

$$SCV = \sum_{t=1}^{K} \omega_i(\mu_i - \mu)^2 + \sum_{i=1}^{K} \omega_i V_i \tag{14}$$

Similarly, following Oshio (2002), the change in income distribution caused by the pension system can be expressed according to equation (15).

$$\frac{SCV^* - SCV}{SCV} = \frac{1}{V}\sum_{t=1}^{n=6} \omega_i\left[(\mu_i^* - \mu^*)^2 - (\mu_i - \mu)^2\right] + \frac{1}{V}\sum_{t=1}^{n=6} \omega_i(V_i^* - V_i) + \frac{\mu^2 - \mu^{*2^*}}{\mu^{*2}}\frac{V^*}{V} \tag{15}$$

Focusing on the left-hand side of equation (15), the term SCV^* indicates the value of the coefficient applied to the income distribution after the pension system, while the SCV refers to the value of the coefficient for the counterfactual of no pension system (pre-system income). Therefore, a positive value would imply an increase in inequality, since the SCV^* would be greater than the SCV. On the contrary, a negative value would imply an improvement in income redistribution. Analysing the right-hand side of equation (15), the first term expresses the extent to which the gap between the average of each group and that of the population as a whole is reduced, which is called the "inter-group redistribution effect". The second term shows the degree to which the variance of each group as a whole is reduced, the "within-group redistribution effect". Finally, the third term reflects the so-called "incompleteness effect" of income redistribution. Regarding this last effect, Oshio (2005) points out that "the social security system is usually incomplete in terms of income redistribution, since its benefits are financed not only by workers' own contributions but also by employers' contributions and government subsidies" (Oshio, 2005, p. 94). This incompleteness effect reflects the fact that the lack of accurate information to reallocate the totality of the components that finance the pension system will result in the average pre-system income being lower than the average post-system income.

This is because, to construct the pre-system income, all the income corresponding to retirement and pensions would be removed without restoring all the components that finance it. In the specific case of this work, this incompleteness effect may arise, among other sources of financing not considered, from the fact that employer contributions are not taken into account in the base scenario (due to the lack of detailed information to take them into account). In this sense, it is recognized that in standard incidence analyses it is usually assumed that there is tax shifting; this implies that both suppliers and demanders can adapt to the existence of the tax by shifting it according to their different elasticities, making it fall on workers (Gómez Sabaini, Santiere and Rossignolo, 2002). Although the analysis carried out in the base scenario of the study by Oshio (2002) considers only the contributions of workers, the various sensitivity scenarios proposed in this article make certain assumptions regarding this and other issues, all of which are explained below.

The inclusion and explicit decomposition of the incompleteness effect represents an important contribution of this study. Most incidence analyses, particularly those applied to pension systems, tend to either omit this dimension or subsume it implicitly within the residual of other effects. However, failing to account for the mismatch between the removal of benefits and the full restoration of their financing sources can lead to a systematic overestimation of redistribution. By quantifying this incompleteness, the analysis highlights a key limitation of the accounting approach and underscores the importance of transparency in the construction of counterfactual incomes. In contexts such as Argentina, where detailed data on employer contributions and tax incidence are often unavailable, making this effect explicit allows for a more accurate interpretation of the redistributive performance of the pension system and provides a cautionary note for comparative studies that rely solely on observed income transfers.

Sensitivity analysis: Employer contributions, self-employed and moratoria

To analyse the robustness of the results and to extend the analysis, a sensitivity analysis was carried out by applying comparative statics with respect to the benchmark scenario, taking into account situations not considered in the previous scenarios.

First, to overcome the drawbacks associated with the incompleteness effect and to be able to evaluate how the results change with the inclusion of employer contributions with respect to the base case, a sensitivity analysis is performed that, under certain assumptions, includes employer contributions as a source of financing. In this sense, from the perspective of capital, employer contributions are part of the labour costs of firms and, as such, should be included in an analysis of the functional (rather than personal) distribution of income (Marshal, 1984). Nevertheless, the assumption proposed by Lustig (2018) is used, according to which it is assumed that employers pass on contributions to workers in the form of lower wages, so that in the absence of the pension system, gross wages are assumed to increase in the same proportion as employer contributions. Although this assumption is questionable, since employer contributions are by law a cost to the employer and there is no guarantee that these funds would be transferred to the employee in their absence, different transfer values of employer contributions are simulated.

In this sense, if employers pass on the employer's costs to the worker, they can only do so before setting the gross wage, since they cannot legally pass on these costs once the gross wage has been set. Thus, one way to pass on these costs is to set the gross wage below the market wage (the wage that would prevail in the absence of the pension system). Equation (16.1) shows how the employer's contribution τ_{co}, which can vary in an interval [0, 0.127][5] (i.e. the employer can pass on none or all of the burden), could be transferred to the statutory gross wage (SGW), starting from a market gross wage (GMW) from which the employer's contributions that finance the pension system have been deducted.

$$GMW - \tau_{co}\mathrm{SGW} = \mathrm{SGW} \tag{16.1}$$

$$GMW = \mathrm{SGW}(1 + \tau_{co}) \tag{16.2}$$

Thus, in the sensitivity section, we consider different τ_{co} transfer scenarios in the interval [0, 0.127] and then obtain the net income in the absence of the pension system according to equation (17), analogous to the benchmark case, where $\tau_{ap} = 6\%$.

$$\mathrm{NW} = \mathrm{GMW}\left(1 - \tau_{ap}\right) \tag{17}$$

However, as the ENGHo does not distinguish between formal and informal self-employed, in the baseline scenario only dependent employees paying pension

5. The statutory employer contribution rate (16 per cent) has been modified several times. In 2017–2018, the current rate was approximately 12.7 per cent (ANSES, 2011).

contributions were considered formal. To broaden the analysis, a sensitivity scenario that includes formal self-employed workers is also considered.

Finally, although pension moratoria are semi-contributory benefits that are part of the pension system and a mechanism for income redistribution, a scenario without moratoria is considered, which is implemented operationally by eliminating pension income below the minimum pension benefit.

Results

Table 2 summarizes the main descriptive household statistics for the sample analysed (weighted values). For example, about 57 per cent of the household heads are men, while the remaining 42 per cent are women, and the average age of the household heads is about 51 years old. It can also be seen that the average THI is practically three times higher than the HIP, which is to be expected when correcting for the number of persons in the household.

Among households headed by elderly persons of retirement age (age 65 for men or age 60 for women), 91.3 per cent receive retirement and/or pension income. This income represents 41.47 per cent of the total income of elderly households. Similarly, according to data from the National Social Security Administration (ANSES), of the 6,852,090 SIPA beneficiaries in 2018, 52.79 per cent entered the system from the pension moratorium programme (ANSES, 2018).

Table 2. *Descriptive statistics of households*

Variables	Minimum	Q1	Median	Average	Q3	Maximum
Age of the head of household	16	37	48.5	50.5	63	98
Total household income (USD)	0	712	1,169.05	1,514.30	1,888.64	121,848.74
Household income per capita (USD)	0	248.18	417.99	591.02	723.41	44,949
Gender						
Men				*Women*		
57%				42%		
Proportion of households with heads aged 65+ receiving retirement or pension: 91.3%						
Proportion of ITH explained by retirement or pension: 41.47%						

Note: All values are weighted values.
Source: Authors' elaboration based on data from ENGHo (2019).

Benchmark scenario: Pensions as government transfers

Figure 1 shows the density functions and the Lorenz curve for the distribution of per capita income before and after the pension system.

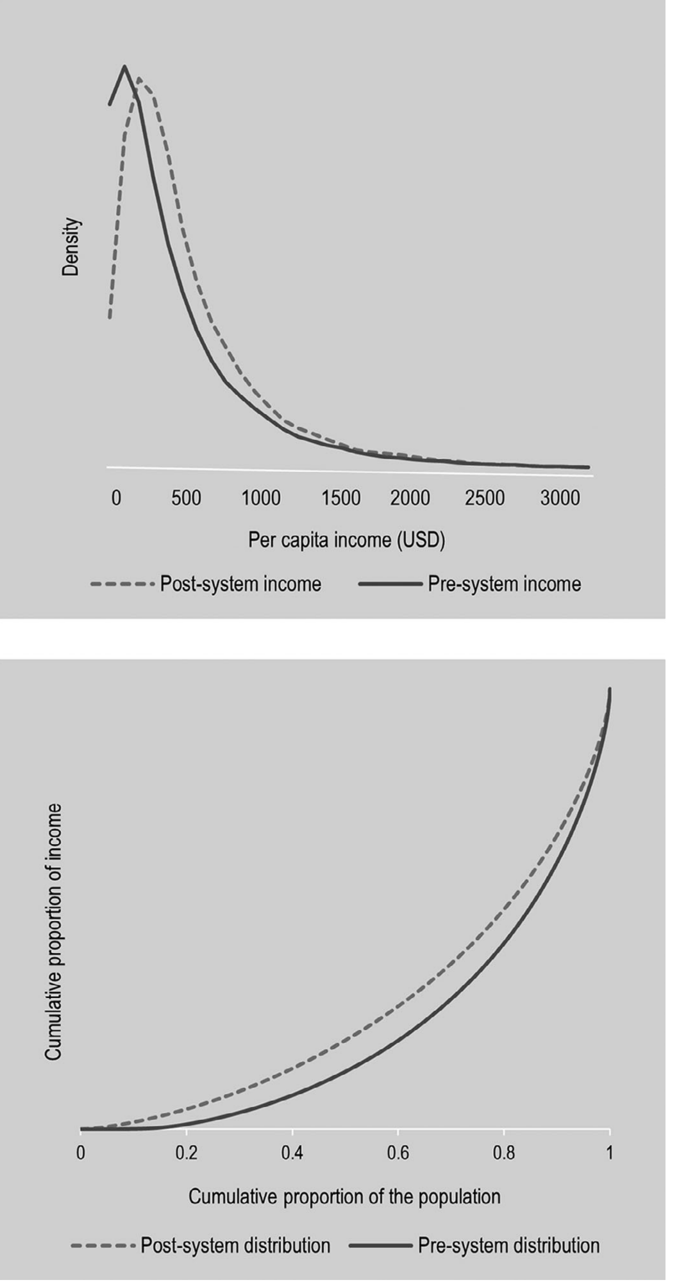
As shown, the pre-system income density function is slightly shifted to the left, indicating that there is a higher proportion of individuals with lower incomes. In turn, the area between the equal distribution line and the Lorenz curve would be larger in the absence of the pension system. Moreover, it is found that the post-system income distribution dominates the pre-system income distribution in the Lorenz sense. This provides robustness to inequality comparisons since the ordering is maintained regardless of the inequality index used. This implies that only the magnitude of the change in inequality may be affected by the index used, but not the sign of the comparison (Gasparini et al., 2012). Indeed, as shown in Table 3, regardless of the indicator used, it can be concluded that the income distribution after the system is less unequal than before the system.

Next, repeating the analysis carried out by Oshio (2002), Table 4 shows the results of decomposing the SCV and applying it to the different age groups defined. In the pre-system situation, it is observed that the average income tends to increase until the age group 40-49 and then gradually decreases in the older age groups. Similarly, the highest SCV values correspond to the older segment of the population. This suggests that as people age, their ability to generate income decreases, but very differently for individuals in the same age group, which would imply a greater degree of inequality. Table 4 also shows that the SCV of the total population is 1,799, of which 1,685 (about 94 per cent) can be explained by inequality within the age group.

If we consider the incidence of the pension system, we can see that the value of the total SCV decreases from 1,799 to 1,311, which, as explained above, means that the pension system contributes to the correction of income inequalities. In this sense, it is observed that the greatest decrease in the SCV occurs for the age groups aged 60 or older. In addition, the average income of those younger than 60 years of age decreases in the post-pension system situation and increases for the elderly, which indicates that income transfers from the working population to the retired population are taking place.

Table 4 also shows the decomposition of the redistributive effect of the total population according to equation (15). The redistributive effect is 27.1 per cent overall, 4.9 per cent between age groups and 5.3 per cent within age groups. However, it is the incompleteness effect that accounts for most of the redistributive effect, about 16.9 per cent. Such an effect is consistent with the gap between average pre-system income (USD 1,363.97) and post-system income (USD 1,514.30), which is caused by the lack of data to accurately impute and construct all types of income.

Figure 1. *Benchmark scenario. Density plot and Lorenz curve of pre- and post-system per capita incomes*



Source: Authors' elaboration based on data from ENGHo (2019).

Table 3. *Benchmark scenario. Inequality indices for different income concepts (HIP)*

	Pre-system income	Post-system income	Reduction in inequality
Gini	0.549	0.443	-19.31%
Theil	0.590	0.387	-34.41%
Atkinson ($\varepsilon = 0.5$)	0.282	0.167	-40.78%
Atkinson ($\varepsilon = 1$)	0.577	0.304	-47.31%
Atkinson ($\varepsilon = 1.5$)	0.844	0.443	-47.51%

Source: Authors' elaboration based on data from ENGHo (2019).

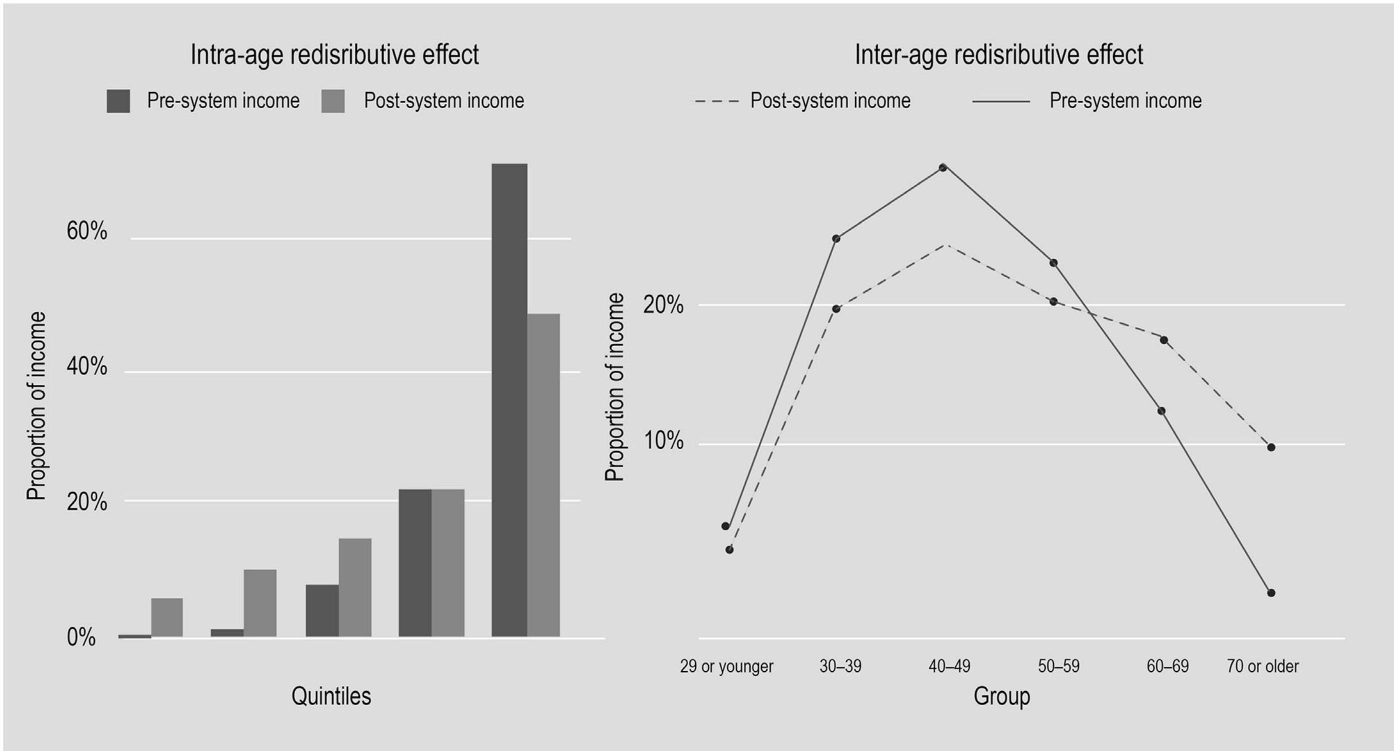
Table 4. *Benchmark scenario. Redistributive effect by age group (THI)*

Age groups	Average pre-system income (USD)	Average post-system income (USD)	SCV pre-system	SCV post-system
29 years or younger	1,114.64	1,043.44	0.565	0.522
30–39 years	1,584.38	1,472.56	0.646	0.633
40–49 years	1,822.67	1,703.07	0.8	0.669
50–59 years	1,663.34	1,669.59	1.389	1.25
60–69 years	1.198.15	1,676.53	6.23	3.13
70 years or older	424.69	1,258.43	3.72	0.633
Total	1,363.97	1,514.30	1.799	1.311
Inter-age income inequality pre-system				0.113
Intra-age income inequality pre-system				1.685
Inter-age income inequality post-system				0.021
Intra-age income inequality post-system				1.289
SCV reduction rate				-0.271
Inter-age redistributive effect				-0.049
Intra-age redistributive effect				-0.053
Incompleteness effect				-0.169

Source: Authors' elaboration based on data from ENGHo (2019).

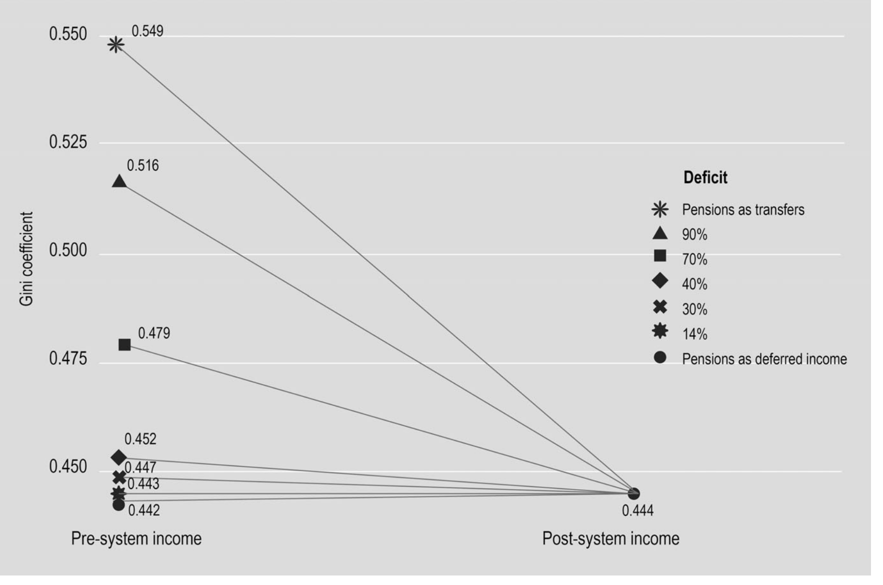
To analyse the intra-age redistribution effect in more detail, the groups aged 60-69 and aged 70 or older were considered and divided into quintiles according to the level of income before and after the pension system. Thus, for each age group considered,

Figure 2. *Benchmark scenario. Redistributive effects of the pension system*



Source: Authors' elaboration based on data from ENGHo (2019).

Figure 3. *Redistributive effect of the pension system for the hybrid scenario considering different levels of deficit (HIP)*



Source: Authors' elaboration based on data from ENGHo (2019).

the first quintile includes the 20 per cent of households with the lowest income, while quintile 5 includes the 20 per cent of households with the highest income. The income share of each quintile in relation to total income was then calculated. The left panel in Figure 2 shows the main results, where it can be observed that participation in the pension system of the three lowest income groups improves their income, while the participation of higher income groups reduces their total income. Thus, the pension system redistributes resources from high-income pensioners to those with lower incomes.

To analyse the inter-age effect, the participation of each age group in relation to total income was calculated. The right panel in Figure 2 shows that in the situation before the pension system, there is a pronounced drop in the income share of the older age groups. This situation is attenuated when the effects of the pension system are taken into account. Thus, there is a decrease in the income share of groups aged 59 or younger and a corresponding increase for groups aged 60 or older. This effect may also be related to the smoothing of consumption, which is one of the main objectives of pension systems.

Deferred pension scenario

This section analyses a hybrid scenario in which only part of the pensions are considered as government transfers, while the rest is valued as deferred income. Figure 3 shows the Gini coefficients for the two types of income under different assumptions regarding the pension system deficit. It should be clarified that the deficit percentages refer to the share ω of pension benefits not financed by contributions. It can be observed that the smaller the deficit, the smaller the redistributive effect of pensions. This is logical, since a smaller deficit implies that a larger share of pensions is considered as deferred income. Likewise, from a deficit of 14 per cent, i.e. considering 86 per cent of pensions as deferred income, the system becomes regressive. In the specific case of Argentina, in 2018 about 40 per cent of social security benefits were financed by non-contributory sources from taxes. Although the social security system is larger than the pension system, this percentage can be taken as an approximation due to the lack of data. This means that the 40 per cent deficit scenario in Figure 3 would best represent the situation in Argentina, which implies moving from a Gini of 0.4518 to a Gini of 0.4435 (a reduction in inequality of only 1.84 per cent compared to 19.31 per cent in the baseline scenario).

Sensitivity analysis

In this section, we show the results of comparative statics with respect to the base case by considering four scenarios under different assumptions. Given that in many cases the person legally responsible for a tax is not necessarily the one who bears its burden, it is interesting to evaluate what would happen if employers were to pass on employer contributions to some extent in the form of lower gross wages.

Table 5 presents the results, where as τ_c increases, the incompleteness effect becomes increasingly smaller, which is partly explained by the fact that as τ_c increases, the average *post-system income* becomes proportionally smaller relative to the *pre-system income*. In addition, when analysing the THI, it is observed that increases in τ_c lead to a decrease in the redistributive effect measured by the SCV. In contrast, if we compare it with the HIP (measured with the Gini coefficient), the opposite occurs. One explanation for this discrepancy is that, as τ_c increases, the variance of pre-system income increases, but this is more than compensated for by the increase in *pre-system income*, generating a decrease in the SCV of *pre-system income* and, therefore, an increase in the redistributive effect.

Of course, since the ENGHo does not ask about the contributions of self-employed workers, the analysis carried out so far only considers employees. This is because the ENGHo does not allow us to distinguish between registered self-employed workers (who pay contributions) and non-registered employees. Therefore, in order to include them in the analysis, the operational definition used by Jiménez (2011) was applied, according to which formal self-employment is considered to extend to those qualified self-employed workers and those non-professional workers with some qualification, provided that their labour income is in the last 6 deciles of the distribution and that they carry out their activities in establishments with more than five employees. This made it possible to include 705 households in the analysis (which would be 518,017 in weighted terms), assuming that they carry out activities as contributors under the simplified regime (known as *monotributo*). The reduction in inequality, as measured by the Gini coefficient, is 19.58 per cent in this case, slightly higher than in the base case.

Finally, to evaluate the scenario without semi-contributory benefits, pension income was eliminated for those households with pension income (per capita per elderly person) below the minimum benefit (USD 408.26). In this scenario, the reduction in the Gini coefficient was 12.76 per cent, which is much lower than the reduction estimated for the base case.

Discussion

Given the results of this study, it is relevant to analyse to what extent they differ (or agree) with other similar papers in the literature on the incidence of pension systems and, more importantly, what are the advantages and limitations of the approach adopted.

Comparing our results with those obtained by Oshio (2002), that author finds that pension programmes in Japan generate a decrease in SCV of 32.44 per cent with an inter-age effect of 17.59 per cent, an intra-age effect of 0.49 per cent and an incompleteness effect of 14.36 per cent. That is, Oshio also finds a high redistributive effect, largely explained by a high incompleteness effect. This may be due, as in our baseline scenario, to the fact that employer contributions are not considered a source of financing for the pension system. Furthermore, the author does not consider non-contributory sources of financing either. The big difference is in the proportion of the total effect that is explained by the intergenerational effect: for Oshio it represents 54 per cent of the total, while for us it represents barely 20 per cent. Even in the scenario in which we assume that employer contributions are fully transferred, reducing the incompleteness effect, the effect that assumes greater weight is the intra-age effect rather than the inter-age effect. This would imply that in Argentina the pension system contributes more to intragenerational redistribution than to intergenerational redistribution, which may be largely due to the Pension Inclusion Plan.

Table 5. *Benchmark scenario. Simulation of shift of employer contributions*

Tao= 1%	Pre-system income	Post-system income	Redistributive effect
Average (ITF)	1,363,470	1,514,300	-
SCV	1.791	1.311	-26.82%
Inter-effect	-	-	-4.90%
Intra-effect	-	-	-4.86%
Incompleteness effect	-	-	-17.08%
Average (IPF)	495,249	591,019	-
Gini	0.5484	0.4430	-19.22%
Tao= 3%	Pre-system income	Post-system income	Redistributive effect
Average (ITF)	1,378,580	1,514,300	
SCV	1.776	1.311	-26.20%
Inter-effect	-	-	-5.00%
Intra-effect	-	-	-5.97%
Incompleteness effect	-	-	-15.24%
Average (IPF)	500,701	591,019	-
Gini	0.5492	0.4430	-19.34%
Tao= 5%	Pre-system income	Post-system income	Redistributive effect
Average (ITF)	1,393,669	1,514,300	-
SCV	1.762	1.311	-25.63%
Inter-effect	-	-	-5.11%
Intra-effect	-	-	-7.09%
Incompleteness effect	-	-	-13.43%
Average (IPF)	506,147	591,019	-
Gini	0.5501	0.4430	-19.47%
Tao= 7%	Pre-system income	Post-system income	Redistributive effect
Average (ITF)	1,408,740	1,514,300	-
SCV	1.749	1.311	-25.05%
Inter-effect	-	-	-5.21%
Intra-effect	-	-	-8.19%
Incompleteness effect	-	-	-11.65%
Average (IPF)	511,584	591,019	-
Gini	0.5510	0.4430	-19.60%
Tao= 9%	Pre-system income	Post-system income	Redistributive effect

(Continued)

Table 5. *Benchmark scenario. Simulation of shift of employer contributions -* Continued

Tao= 1%	Pre-system income	Post-system income	Redistributive effect
Average (ITF)	1,423,786	1,514,300	-
SCV	1.736	1.311	-24.50%
Inter-effect	-	-	-5.31%
Intra-effect	-	-	-9.28%
Incompleteness effect	-	-	-9.91%
Average (IPF)	517,014	591,019	-
Gini	0.5519	0.4430	-19.73%
Tao= 11%	Pre-system income	Post-system income	Redistributive effect
Average (ITF)	1,438,817	1,514,300	-
SCV	1.724	1.311	-23.96%
Inter-effect	-	-	-5.41%
Intra-effect	-	-	-10.36%
Incompleteness effect	-	-	-8.19%
Average (IPF)	522,438	591,019	-
Gini	0.5528	0.4430	-19.86%
Tao= 12.7%	Pre-system income	Post-system income	Redistributive effect
Average (ITF)	1,451,578	1,514,300	-
SCV	1.714	1.311	-23.52%
Inter-effect	-	-	-5.49%
Intra-effect	-	-	-11.27%
Incompleteness effect	-	-	-6.75%
Average (IPF)	527,043	591,019	-
Gini	0.5535	0.4430	-19.96%

Source: Authors' elaboration based on data from ENGHo (2019).

Unfortunately, to the best of our knowledge, there are no other studies that use the SCV to analyse the impact of the pension system on income distribution. However, there are studies that use other indicators that do not allow the same decomposition, so the same wealth of information cannot be obtained, but only the sign of the impact. In this line, Sung and Park (2011) conclude that the contribution of the Republic of Korea public pensions to the reduction of inequality was small in 2007 (-0.6 per cent). For his part, Hwang (2016) concludes that Republic of Korea public pensions had effects that exacerbated income inequality in old age between 1998 and 2010 and even counteracted the palliative effects of public assistance. However, this may be because the Republic of Korea pension system did not include all workers until 2006. Indeed, the author admits that those who have public pensions are those with a quality work history, so the overall coverage of public pensions is low and is concentrated among older people and those with a higher socio-economic level. These results contrast with those found for Argentina, where, regardless of the inequality index used, the pension system contributes to reducing per capita income inequality.

Other authors who find a progressive effect of pensions are Cai and Yue (2020), who conclude that social security transfers in the People's Republic of China reduce the Gini coefficient from 0.460 to 0.436. However, they warn that when breaking down the Gini coefficient and calculating the marginal effects of each element of social security, formal sector and urban pensions are the most unequal income drivers. This could be explained by the fact that China's pension system, like Argentina's, is fragmented in several programmes.

Elsewhere, works such as that of Djindjić (2014) that study the impact of the tax system, conclude, for example, that in Serbia taxes substantially reduce inequality, with state pensions being one of the components that contribute most to such reduction. Wang, Caminada and Goudswaard (2012) reached a similar conclusion when they studied the impact of the tax system in 28 countries, concluding that pension programmes contribute most to the reduction of the Gini coefficient.

Finally, and no less relevant, Lustig (2017) finds that, on average for the 17 Latin American countries, fiscal policy as a whole reduces the Gini coefficient by 14.35 per cent, considering pensions as deferred income. In this sense, Argentina is the country where inequality is reduced the most, with a reduction in the Gini of 37.1 per cent. Furthermore, if pensions are considered as direct transfers, as in our base scenario, the average Gini in Latin America is reduced by 15.62 per cent and in Argentina by 40.71 per cent. Furthermore, when comparing both scenarios, Lustig finds that in Argentina contributory pensions generate a drop in the Gini well above the Latin American average, although below that of the United States of America and the European Union.

Before discussing the methodological limitations of this study, it is also important to consider the macroeconomic context in which the results were obtained. Although all monetary variables in this study were deflated to constant prices to avoid inflation distortions and ensure comparability across households, high-inflation periods may generate intergenerational redistributive effects (Doepke and Schneider, 2006; Meh and Terajima, 2011). In the Argentine context, this mechanism can be observed in the performance of pensions and wages during 2017–2018. Pension benefits increased by 28.4 per cent while inflation reached 47.6 per cent, leading to a significant real decline in their purchasing power (Resolución N° 176-E/2017; Resolución N° 242/2018;

INDEC, 2018). Registered workers also experienced real income losses, as the RIPTE index (average taxable wage of registered employees) rose by 31.18 per cent over the same period – only slightly above the mobility adjustment. Thus, both groups were adversely affected by inflation, but pensioners faced a marginally deeper erosion due to lagged indexation.

As regards the limitations of the analysis, one aspect to be considered is the lack of any comparison of the relative efficiency of the pension system in redistributing income with other components of the social security system or tax system. Here, some works suggest that social spending (which includes the pension system) has contributed significantly in relative terms to reducing inequality in Latin America during the first decade of the 2000s, even when controlled for other relevant factors (Martorano, 2018).

However, there is a common problem in household surveys related to the possible systematic underreporting of monetary income by recipients. This concern is based on the discrepancies between the gross income reported in household surveys and the values estimated by national accounts or other secondary sources (Salvia and Donza, 1999). Similarly, there is evidence that the degree of underreporting in pension surveys is lower than for other income sources (Sanz, 2019).

Another point to bear in mind is the fact that we do not take into account the existence of other programmes, such as the National Institute of Social Services for Retired Persons and Pensioners (INSSJyP), which saves on out-of-pocket health expenses for the elderly.

Finally, while this study evaluates redistribution at a single point in time, it is worth noting that part of the debate on the regressivity of defined benefit systems refers to life-cycle patterns that cannot be captured in a cross-sectional incidence analysis. Issues such as differential life expectancy, contribution histories and duration of benefit receipt – discussed in the literature cited in the introduction – require longitudinal information to be properly assessed. In this framework, our findings show that, in the short run, the Argentine pension system generates a predominantly progressive intragenerational effect, largely driven by minimum benefit guarantees and high coverage among older adults. Whether these short-run redistributive patterns persist or reverse over individuals' lifetimes remains an open question and is beyond the scope of our data; a comprehensive life-cycle incidence analysis is therefore a promising avenue for future research.

Conclusion

The results indicate that the pension system had a progressive impact on income redistribution during the period analysed. Thus, for the base scenario, considering per capita income, the Gini coefficient of pre-system income was estimated at 0.549, which is reduced to 0.443 when the effects of the pension system are taken into account, implying a decrease of 19.31 per cent in this coefficient.

However, the calculation of the SCV for total household income shows a decrease from 1.799 to 1.311 (27.1 per cent). When this coefficient is broken down, 5.3 per cent is due to the intra-age effect, which implies the redistribution of resources from high-income pensioners to those with lower incomes. In turn, 4.9 per cent is due to the inter-age effect, i.e. the redistribution of income from the working population to the inactive population. Finally, most of the effect is due to the incompleteness effect (16.9 per cent), which is caused by the lack of information to fully impute all the components that finance the system. However, if we consider the shift of contributions, we find that in the case of a complete transfer (12.7 per cent), the incompleteness effect is reduced to 6.75 per cent, the intra-age effect to 11.27 per cent and the inter-age effect to 5.49 per cent. Similarly, the reduction in inequality is 23.52 per cent.

In the alternative hybrid scenario, where only part of the pensions are considered as transfers and the rest are considered as deferred savings, the redistributive effects of the pension system are significantly reduced. This is because incomes in the absence of the system are reduced to a lesser extent and therefore the calculated inequality is lower than in the base case.

These findings support the hypothesis that, in the Argentine context, redistribution occurs predominantly through intragenerational mechanisms, which represents one of the central empirical contributions of this study. The structure of the system – marked by moratoria, non-contributory benefits, and high informality – favours transfers within the elderly population rather than across generations. This distinguishes Argentina from many high-income countries, where intergenerational redistribution tends to dominate.

Methodologically, the study contributes by explicitly decomposing the redistributive impact into intra-age, inter-age, and incompleteness effects, and by introducing a detailed sensitivity analysis across multiple scenarios. The results highlight that the measurement of redistribution is highly sensitive to conceptual choices, such as the classification of pension income and the treatment of the system's financing sources. Ignoring these dimensions can lead to over- or underestimation of the equity effects of pension systems.

From a policy perspective, the findings suggest that in middle-income countries with inclusive but fragmented pension systems, intragenerational redistribution may be more effective in reducing inequality than traditional intergenerational transfers. At the same time, the results underline the importance of improving data availability and transparency regarding pension financing to enable more accurate evaluations of redistributive performance.

More specifically, the evidence points to several areas where policy improvements could enhance the equity and efficiency of the system. First, the strong role of intragenerational redistribution highlights the need to consolidate and strengthen non-contributory and semi-contributory programmes targeted to low-income elderly populations, especially in contexts of widespread informal employment. Recognizing this dynamic is essential for policy makers, as it underscores that the system's redistributive performance depends primarily on the design and sustainability of these mechanisms. Second, the significant weight of the incompleteness effect underscores the urgency of increasing the transparency and traceability of pension financing sources, particularly employer contributions and general revenues. A better understanding of who effectively bears the cost of pension financing would allow for more informed debates about fairness and sustainability.

Finally, the sensitivity of redistributive outcomes to conceptual definitions – such as the classification of pension income as a transfer versus deferred income – suggests that policy evaluations should be cautious in interpreting results based on narrow accounting assumptions. Promoting methodological pluralism and requiring robustness checks in fiscal incidence analysis could contribute to more balanced policy discussions in the field of social protection.

In terms of future research, new studies could explore how redistributive patterns vary by gender, region, or employment history, particularly in contexts marked by labour informality and segmentation. In addition, the development of more frequent and consistent indicators based on data sources with greater temporal coverage – beyond traditional household surveys such as the EPH – would make it possible to monitor redistributive performance over time and improve the responsiveness of social protection policy. Such efforts would not only enrich academic understanding of redistributive systems in middle-income countries but also support the design of more equitable and sustainable pension reforms.

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Universal health coverage and maternal mortality in Morocco: An ARDL bounds analysis

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Abstract Universal Health Coverage (UHC) represents a critical pathway toward improved maternal health outcomes, yet causal evidence remains limited in Middle East and North Africa (MENA) countries. This study assesses the impact of UHC expansion on maternal mortality in Morocco using Autoregressive Distributed Lag (ARDL) bounds testing with annual data from 2000 to 2023. We estimate a health production function controlling for education, fertility, urbanization, and public health expenditure. Results show that UHC expansion significantly reduces maternal mortality, with a long-run elasticity of –0.179 ($p < 0.001$). Based on the estimated elasticities, UHC alone explains roughly one quarter of the observed decline in maternal mortality, while combined with public health spending, these two factors account for nearly half of the overall reduction. The findings highlight strong complementarity between financial protection and health investment, providing new econometric evidence of UHC's effectiveness in the MENA region and informing strategies toward Sustainable Development Goal achievement. The study acknowledges methodological limitations including non-significant education and fertility coefficients that diverge from the established literature, likely reflecting ARDL-specific estimation properties and Morocco's institutional context.

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Introduction

Universal Health Coverage (UHC) remains a global health policy priority, yet progress has stagnated. Approximately 8 million preventable deaths occur annually from poor-quality health care rather than lack of access (UNICEF, 2024). Financial protection has deteriorated, with 2 billion people facing health-related financial hardship (WHO EMRO, 2024). The findings of a 2025 World Health Organization (WHO) systemic analysis reveals that maternal mortality declined by only 40 per cent globally between 2000 and 2023, with progress stagnating since 2015. Achieving the 2030 Sustainable Development Goal (SDG) target of 70 deaths per 100,000 live births by 2030 requires an unprecedented 15 per cent annual reduction rate (Cresswell et al., 2025).

The theoretical foundation for understanding UHC's health impact has evolved from Grossman's (1972) health capital model toward frameworks recognizing health as emerging from complex production processes. Contemporary health economics emphasizes how institutional arrangements – particularly health care financing and system organization – shape health production (Or, 2000). However, establishing causal relationships between UHC reforms and health outcomes presents methodological challenges, particularly regarding endogeneity where poor health outcomes may motivate reforms rather than reforms driving improvements (Reich et al., 2016).

Recent evidence challenges conventional UHC assumptions. Analysis of 4.1 million births from 60 low- and middle-income countries demonstrates that UHC expansion benefited poorer populations in early stages but became less so as overall coverage increased (Hone et al., 2024). UHC effects depend on design features, baseline health system capacity, and implementation quality rather than formal reform adoption alone (Kruk et al., 2018).

Morocco offers an opportunity for examining UHC's impact on maternal mortality. The country's systematic health coverage expansion from 13.5 per cent in 2000 to 87 per cent in 2023 (see Table 2), documented by Dougadir, Radah and Elhia (2025) as a deliberate institutional reform trajectory, coincided with maternal health improvements. Maternal mortality declined from 271 to 70 per 100,000 live births – a 74 per cent reduction compared to the global 40 per cent average (Souza et al., 2024). Skilled birth attendance increased from 40 per cent to 87 per cent, institutional delivery rates rose from 31 per cent to 84 per cent, and antenatal care coverage improved from 32 per cent to 88 per cent. This transformation occurred within Morocco's dual institutional framework: Compulsory Health Insurance (*Assurance maladie obligatoire* – AMO) for formal sector workers and retirees, and Medical Assistance Scheme (*Régime d'assistance médicale* – RAMED) offering free care for the poorest quintile. This dual approach reflects international best practices for middle-income countries with substantial informal economies (Frenk et al., 2006).

Our analysis covers 2000–2023, capturing the evolution from the dual AMO-RAMED system to the 2022–2023 generalization that replaced RAMED with AMO-TADAMON and extended coverage to self-employed workers through AMO *Travailleurs non salariés* (AMO-TNS). However, the brief post-reform observation window limits assessment of these recent structural changes' long-term effects.

The Middle East and North Africa (MENA) region remains underrepresented in the international UHC impact literature. This geographical gap limits policy guidance specifically relevant to MENA contexts, where countries face unique combinations of challenges including political instability, economic volatility, distinctive epidemiological profiles, and oil-dependent economic structures. Rigorous time-series econometric studies establishing causal UHC effects in this region are scarce (Stenberg et al., 2017).

Understanding the temporal dynamics of UHC effects – distinguishing between immediate impacts and longer-term equilibrium relationships – requires sophisticated analytical approaches. Traditional cross-sectional and short panel data approaches cannot adequately address these complexities. Recent methodological advances in Autoregressive Distributed Lag (ARDL) applications provide solutions for data-constrained developing country contexts, particularly given their capacity to handle mixed integration orders, structural breaks, and modest sample sizes.

This study examines whether UHC expansion causally reduces maternal mortality in Morocco while quantifying both short-run and long-run effect magnitudes. We employ comprehensive ARDL bounds testing combined with Error Correction Modelling, building on recent research demonstrating statistically significant UHC effects using similar approaches (Konca, 2024). Our analytical framework distinguishes immediate policy impacts from sustained effects while controlling for concurrent socio-economic changes.

The study contributes to health economics literature in three ways. First, we provide the first application of ARDL methodology to maternal mortality and UHC analysis in the MENA region. Second, we quantify UHC's specific contribution to Morocco's maternal mortality reduction, demonstrating that coverage expansion accounts for approximately 25 per cent of the total decline. Third, we demonstrate complementarity between UHC coverage and public health expenditure, suggesting balanced investment strategies optimize outcomes. Morocco's systematic approach to coverage expansion offers lessons for other middle-income countries pursuing health system reforms.

The research questions are: 1) Does UHC expansion causally reduce maternal mortality in Morocco, and what is the magnitude of both short-run and long-run effects? 2) How do UHC effects interact with public health expenditure? 3) What are the dynamic adjustment mechanisms governing convergence to long-run equilibrium? 4) How stable are these relationships across different time periods?

The remainder of this article proceeds as follows. The next section presents our empirical literature on UHC impact evaluation. Thereafter we set out our ARDL methodology and data sources and report our econometric results. We then discuss the findings and quantify UHC's contribution before offering concluding recommendations.

Literature review

Health economic theory has evolved from Grossman's (1972) health capital model toward frameworks recognizing health as emerging from complex production processes combining medical inputs, environmental factors, and institutional arrangements. The health production function H = f (M, E, L, I) positions UHC within institutional factors (I) that fundamentally shape resource allocation and access patterns (Or, 2000). Contemporary research emphasizes the quality–coverage nexus, demonstrating that high-quality health systems could prevent half of maternal and newborn deaths, yet quality improvements require deliberate policy interventions across multiple system levels (WHO, 2024). This quality dimension challenges traditional UHC evaluation focusing primarily on coverage metrics, while highlighting that worldwide, more preventable deaths – an estimated 8 million annually – occur from poor-quality health care than from lack of access to care (UNICEF, 2024).

The existing literature faces fundamental methodological constraints limiting causal inference (Table 1). Most studies employ cross-sectional approaches unable to address reverse causality where poor health outcomes motivate reforms rather than reforms driving improvements (Wagstaff et al., 2015). Omitted variable bias persists when unobserved factors simultaneously influence UHC expansion and health outcomes. The temporal dimension remains inadequately addressed in current research designs. Understanding immediate versus sustained UHC effects requires analytical approaches capable of capturing dynamic adjustment processes.

ARDL methodology provides solutions for data-constrained developing country contexts (El Baouchari and Raouf, 2024). The approach accommodates mixed integration orders without requiring pre-testing for unit roots, handles structural

Table 1. *Key UHC impact studies: Methodological overview*

Study	Country/ Region	Sample	Method	UHC Effect
Hone et al. (2024)	60 LMICs	4.1 M births	Panel analysis	Positive early, diminishing for poor
Konca (2024)	Türkiye	1974-2018	ARDL	+0.43% life expectancy (p<0.01)
Dowou et al. (2023)	Sub-Saharan Africa	Multi-country	Systematic review	Mixed, weak financing
Dinga et al. (2024)	SADC countries	Panel data	Fixed effects	Government spending reduces MMR
Moreno-Serra and Smith (2012)	Global	Cross-country	Instrumental variables	Positive population health
Spaan et al. (2012)	Africa and Asia	Systematic review	Meta-analysis	Heterogeneous effects

Source: Authors' elaboration.

Table 2. *Variable definitions and data sources*

Symbol	Description	Unit	Source
MMt	Maternal mortality rate	Deaths per 100,000 live births	WHO Global Health Observatory, World Bank WDI
EDUt	Primary education completion rate	Percentage (%)	World Bank World Development Indicators
TFt	Total fertility rate	Births per woman	World Bank World Development Indicators
URBt	Urbanization rate	Percentage (%)	World Bank Urbanization Database
DPSt	Public health expenditure per capita	USD (constant 2015)	WHO Global Health Expenditure Database, Morocco MoH
CSUt	Universal health coverage rate	Percentage (%)	Morocco ANAM, High Commission for Planning

Source: Authors' elaboration.

breaks effectively, and performs well with modest sample sizes (Pesaran, Shin and Smith, 2001).

Recent validation emerges from Türkiye's comprehensive UHC analysis. Konca (2024) demonstrates statistically significant effects on life expectancy in both the short run (p = 0.001) and long run (p = 0.011), with health expenditure showing positive long-run effects (p = 0.001) and an error correction term of -0.58 suggesting moderate adjustment speed. The error correction mechanism falls within the theoretically acceptable range, indicating stable convergence to long-run equilibrium.

Mathematical modelling approaches, including optimal control theory applied to epidemic interventions (Elhia, Chokri and Alkama, 2021; Boujallalm, Elhia and Balatif, 2021; Boujallalm, Balatif and Elhia, 2021; Elhia et al., 2014), complement econometric methods by providing dynamic optimization perspectives. While such models excel at prospective policy design, they require extensive parametric structure often unavailable for maternal mortality determinants. ARDL provides robust reduced form estimates without full structural specification, accommodating mixed integration orders and modest sample sizes typical of developing country contexts.

Maternal mortality serves as an ideal indicator for UHC effectiveness assessment. Deaths are largely preventable with appropriate intervention, making mortality rates directly responsive to access and quality improvements. Maternal services span the care continuum from prevention through emergency intervention, providing comprehensive health system assessment. Outcomes reflect equity dimensions central to UHC goals, as mortality risks concentrate among socio-economically disadvantaged populations (Miller et al., 2016).

Global assessments reveal the remaining challenges. The maternal mortality ratio declined 40 per cent between 2000 and 2023, representing a low 2.2 per cent annual reduction – far below the 15 per cent required for 2030 targets (WHO et al., 2025). WHO systematic analysis of 2009–2020 maternal deaths demonstrates that quality emerges as the critical mediating factor determining whether health-care contact translates into survival (Cresswell et al., 2025).

Financial protection measurement presents additional analytical complexity that influences UHC evaluation methodology. Recent research suggests catastrophic health spending may be systematically underestimated for service-specific analysis, particularly among poorer households, when aggregate rather than service-specific thresholds apply (Ataguba et al., 2024). This methodological insight carries profound implications for understanding UHC effectiveness and supports comprehensive analytical approaches that capture both direct health effects and broader welfare implications. The intersection of coverage expansion with financial protection mechanisms becomes particularly relevant for understanding how institutional arrangements translate into measurable population health improvements.

Morocco's systematic UHC implementation represents one of the MENA region's most documented experiences, with the dual approach combining mandatory insurance for formal workers with targeted free care for the poor, reflecting international best practices for middle-income countries with significant informal economies (Frenk et al., 2006). Key milestones include the 2005 launch of AMO covering formal sector employees, RAMED's scale-up in

2012 providing free care for the poorest quintile, and then the 2021 Royal Initiative toward Universal Coverage integrating both schemes. Combined with concurrent primary care investments and provider training programmes, Morocco's experience provides a natural experiment for understanding how well-designed reforms translate into measurable health improvements.

Post-COVID-19 pandemic evidence further emphasizes the critical importance of learning from successful UHC experiences to inform recovery strategies and accelerate progress toward the United Nations Sustainable Development Goals 2030 targets. Building resilient health systems requires context-based health worker redistribution, evidence-based task-shifting policies, and results-based financing mechanisms, with high political commitment and multi-sectoral collaboration being fundamental for success (Debie et al., 2024). The pandemic derailed progress across multiple health targets, particularly in low- and middle-income countries, highlighting the urgency of understanding which institutional arrangements and implementation strategies can deliver sustained health improvements (Yin et al., 2024). Recent systematic reviews of barriers to maternal health access in Africa show that the most significant obstacles include transportation barriers, economic constraints, cultural beliefs, and poor quality of care – factors that well-designed UHC systems should theoretically address (Dahab and Sakellariou, 2020).

The convergence of these theoretical insights, methodological advances, and empirical evidence reveals three fundamental gaps that constrain current understanding of UHC's causal impact on population health outcomes. First, no existing research applies ARDL bounds testing with Error Correction Modelling to maternal mortality and UHC evaluation in the MENA region, despite demonstrated advantages for handling health system data constraints. Second, MENA countries remain severely underrepresented in UHC impact literature despite rapid health reforms. Third, most studies examine short-term UHC effects while long-run equilibrium relationships remain largely unexplored. Our study addresses these deficiencies by employing ARDL techniques to capture both short-run dynamics and long-run relationships, focusing on maternal health as a comprehensive system performance indicator, and providing UHC impact evidence from the MENA region.

Methodology

Theoretical framework and model specification

Our empirical strategy builds on the health production function framework, specifically adapted for time-series analysis of institutional reforms and informed by recent theoretical developments in maternal health determinants (Souza et al., 2024). We conceptualize health outcomes as emerging from complex production processes where health care inputs, socio-economic determinants, and institutional factors interact to determine population health. This theoretical foundation guides our specification of health outcomes as a function of these key determinants:

$$H_t = f(HC_t, SE_t, I_t, \varepsilon_t)$$

where H_t represents health outcomes at time t, HC_t denotes health care inputs, SE_t encompasses socio-economic determinants, I_t captures institutional factors (particularly UHC), and ε_t represents the stochastic error term. Translating this conceptual framework into an estimable econometric model, we employ a double-log specification that facilitates elasticity interpretation while addressing heteroscedasticity concerns common in health economics applications:

$$\ln(MMt) = \alpha_0 + \alpha_1 \ln(EDUt) + \alpha_2 \ln(TFt) + \alpha_3 \ln(URBt) + \alpha_4 \ln(DPSt) + \alpha_5 \ln(CSUt) + \varepsilon t$$

Our variable selection reflects careful consideration of both theoretical foundations and empirical constraints, with each variable capturing key dimensions of the health production process as detailed in Table 2.

This double-log specification allows direct elasticity interpretation while capturing potential non-linear effects and facilitating comparison of relative impacts across different determinants of maternal mortality.

ARDL bounds testing framework and advantages

The Autoregressive Distributed Lag (ARDL) bounds testing approach (Pesaran, Shin and Smith, 2001) offers compelling advantages for health economics applications in developing countries. Unlike traditional cointegration methods requiring uniform integration orders, ARDL accommodates variables that are I(0), I(1), or mutually cointegrated without pre-testing for unit roots. This methodological flexibility is particularly valuable given our modest sample size and the mixed integration properties commonly encountered in health and economic data.

The approach illustrates superior small sample properties, performing well with limited observations unlike Johansen cointegration methods that require large samples of reliable inference. The inclusion of lagged dependent variables helps address simultaneity bias that may arise from reverse causality – a particularly relevant concern when health outcomes potentially influence health policy decisions. Additionally, ARDL exhibits greater robustness to structural breaks compared to traditional cointegration approaches, making it especially suitable for analysing policy interventions such as UHC expansion where institutional changes may alter underlying relationships.

Building on recent successful applications in health economics research (El Baouchari and Raouf, 2024; Konca, 2024; Okamoto et al., 2024), our general ARDL(p,q_1,q_2,q_3,q_4,q_5) specification takes the form:

$$\begin{aligned}\Delta\ln(MM\_t) = {} & \alpha_0 + \alpha_1 t + \Sigma_{i=1}{}^{p}\,\beta_i\,\Delta\ln(MM\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_1\}}\,\gamma_{1i}\,\Delta\ln(EDU\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_2\}}\,\gamma_{2i}\,\Delta\ln(TF\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_3\}}\,\gamma_{3i}\,\Delta\ln(URB\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_4\}}\,\gamma_{4i}\,\Delta\ln(DPS\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_5\}}\,\gamma_{5i}\,\Delta\ln(CSU\_\{t-i\}) \\ & + \delta_1\ln(MM\_\{t-1\}) + \delta_2\ln(EDU\_\{t-1\}) \\ & + \delta_3\ln(TF\_\{t-1\}) + \delta_4\ln(URB\_\{t-1\}) \\ & + \delta_5\ln(DPS\_\{t-1\}) + \delta_6\ln(CSU\_\{t-1\}) + \varepsilon\_t\end{aligned}$$

Optimal lag structure selection employs information criteria (AIC, SIC, HQ) across all possible combinations, with maximum lags constrained by sample size considerations (maximum = 2, following the established rule of maximum lags $< T/10$ for $n = 24$).

Cointegration testing

We employ annual data spanning 2000–2023, capturing Morocco's complete UHC reform cycle from 13.5 per cent to 87 per cent coverage. This time span aligns with ARDL methodology requirements, which perform adequately with samples of 20–30 observations when strong cointegrating relationships exist (Pesaran, Shin and Smith, 2001). Recent health economics applications in comparable contexts employ similar sample sizes (El Baouchari and Raouf, 2024; Konca, 2024).

The bounds test examines the joint significance of lagged level variables to determine whether long-run equilibrium relationships exist among our variables. We test the null hypothesis H_0: $\delta_1 = \delta_2 = \delta_3 = \delta_4 = \delta_5 = \delta_6 = 0$ (no cointegration) against the alternative H_1: at least one $\delta_i \neq 0$ (cointegration exists). Decision rules follow Pesaran, Shin and Smith (2001): if the F-statistic exceeds the upper I(1) bound, we reject the null hypothesis and confirm cointegration; if the F-statistic falls below the lower I(0) bound, we conclude no cointegration exists; otherwise, the results remain inconclusive. Critical values incorporate adjustments for sample size and regressor characteristics, ensuring robust inference despite our modest sample constraints.

Error Correction Model and dynamic analysis

Upon confirming cointegration through the bounds test, we estimate the Error Correction Model (ECM) to analyse short-run dynamics and adjustment mechanisms following recent best practices in health economics applications. The ECM specification allows us to distinguish between immediate policy impacts and longer-term equilibrium effects, providing crucial insights into the temporal dynamics of UHC reform impacts on maternal mortality outcomes.

The Error Correction Model takes the following form:

$$\begin{aligned}\Delta\ln(MM\_t) = {} & \lambda_0 + \Sigma_{i=1}{}^{p-1}\ \lambda_i\ \Delta\ln(MM\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_1-1\}}\ \psi_{1i}\ \Delta\ln(EDU\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_2-1\}}\ \psi_{2i}\ \Delta\ln(TF\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_3-1\}}\ \psi_{3i}\ \Delta\ln(URB\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_4-1\}}\ \psi_{4i}\ \Delta\ln(DPS\_\{t-i\}) \\ & + \Sigma_{i=0}{}^{\{q_5-1\}}\ \psi_{5i}\ \Delta\ln(CSU\_\{t-i\}) + \varphi_1 ECT\_\{t-1\} + \upsilon\_t\end{aligned}$$

where, ECT_{t-1} represents the error correction term lagged one period. The coefficient φ_1 must be negative and statistically significant for a valid error correction mechanism, indicating the speed of adjustment toward long-run equilibrium following short-run disturbances. Long-run elasticities are calculated from the cointegrating relationship as:

$$\theta_i = -\delta_{i+1}/\delta_1 \qquad \text{for } i = 1, 2, 3, 4, 5$$

Representing the long-run elasticities of maternal mortality with respect to each explanatory variable. Short-run effects emerge from the coefficients on differenced variables in the ECM, capturing immediate impacts of changes in explanatory variables on maternal mortality changes. The adjustment speed is measured by the absolute value of the error correction coefficient φ_1, indicating the proportion of disequilibrium corrected each period.

Data validation and quality assurance

Our data validation framework employs rigorous procedures to ensure reliability, temporal consistency, and analytical integrity throughout the estimation period. Given the critical importance of data quality for robust econometric inference, we implement comprehensive validation protocols that address potential measurement concerns while maintaining international statistical standards.

Cross-verification procedures systematically compare data across multiple authoritative sources to identify and resolve potential inconsistencies. This multi-source validation approach enables detection of outliers, measurement errors, and methodological changes that might compromise temporal consistency. Where minor discrepancies emerge between sources, we prioritize WHO and World Bank data for international comparability while incorporating country-specific refinements from national statistical agencies where appropriate.

Temporal consistency analysis examines each data series for structural breaks, outliers, and implausible values that might reflect changes in measurement methodology or data collection procedures. We employ statistical outlier detection methods alongside expert knowledge of Morocco's health system evolution to distinguish genuine policy effects from measurement artifacts. For the limited instances of missing observations (less than 2 per cent of total data points), we implement cubic spline interpolation procedures that preserve underlying trends while maintaining data integrity and ensuring smooth transitions between observed values.

Data transformation procedures include logarithmic transformations for all variables to facilitate elasticity interpretation while addressing heteroscedasticity concerns. We verify that all series exhibit positive values throughout the analysis period, ensuring appropriate logarithmic transformation. Unit root properties are carefully examined to confirm the appropriateness of our ARDL methodology, with particular attention to potential structural breaks coinciding with major policy reforms.

Robustness framework and diagnostic testing

Our comprehensive robustness framework addresses potential methodological concerns through multiple analytical approaches designed to ensure the reliability and generalizability of our findings. Given the importance of robust inference in policy-relevant research, we implement extensive diagnostic testing, sensitivity analysis, and alternative estimation procedures that collectively validate our core empirical results across different methodological assumptions and specifications.

Diagnostic testing encompasses comprehensive residual analysis including Jarque-Bera tests for normality, Breusch-Godfrey LM tests for serial correlation up to three lags, Breusch-Pagan-Godfrey tests for heteroscedasticity, and Ramsey RESET tests for specification adequacy. Parameter stability analysis employs CUSUM and CUSUM-squared tests alongside breakpoint unit root tests that allow for endogenous structural breaks, addressing concerns about parameter constancy during Morocco's evolving policy regime. These stability tests prove particularly important given the significant policy changes occurring during our estimation period.

Sensitivity analysis examines alternative lag structures selected by different information criteria (AIC and HQ alongside our SIC-selected baseline), alternative functional forms testing robustness to logarithmic transformation assumptions, and variable transformations using level variables for key regressors to assess sensitivity to functional form choices. Bootstrap procedures with 1,000 replications generate robust critical values adapted to our sample size while providing bootstrap confidence intervals for long-run coefficient estimates to address potential small-sample bias concerns.

Alternative estimation methods include Fully Modified OLS (FMOLS) and Dynamic OLS (DOLS) as robustness checks for long-run coefficient estimates, providing additional approaches to address endogeneity and serial correlation concerns. Given potential endogeneity concerns inherent in health policy analysis, we employ HAC (Heteroskedasticity and Autocorrelation Consistent) standard errors using the Newey-West procedure with automatic lag selection, ensuring robust inference in the presence of autocorrelation and heteroscedasticity while maintaining the reliability of our statistical inferences for policy interpretation.

Results

Descriptive statistics and unit root analysis

The descriptive statistics reveal substantial variation across all variables, with UHC coverage demonstrating the highest coefficient of variation (0.676), followed by public health spending (0.457) and maternal mortality (0.437). This considerable variation provides sufficient identification power for robust econometric estimation, whereas the systematic trends observed in key variables suggest the potential for meaningful cointegrating relationships. The patterns align with global trends in maternal mortality reduction, although Morocco's progress of 74 per cent reduction has been notably more rapid than the global average of 40 per cent over similar time frames (WHO et al., 2025).

Correlation analysis reveals theoretically consistent relationships, with strong negative correlations between maternal mortality and education (-0.958), urbanization (-0.942), public health spending (-0.970), and UHC coverage (-0.823). The high correlations among explanatory variables (ranging from 0.569 to 0.958) reflect Morocco's integrated development trajectory, where improvements in socio-economic indicators have progressed simultaneously, supporting the reliability of our subsequent econometric analysis.

The unit root tests reveal a mixed integration order pattern that is ideal for ARDL analysis. Three variables (LMM, LEDU, LTF) are integrated of order one I(1), while three variables (LURB, LDPS, LCSU) are stationary I(0). This combination validates our methodological choice, as traditional cointegration methods require all variables to be I(1), whereas ARDL can accommodate mixed integration structures without compromising analytical rigor. This finding is consistent with recent applications of ARDL methodology in health economics research and confirms the appropriateness of our empirical strategy (Table 3 and Table 4).

ARDL model selection and estimation

Model selection and specification justification. The automatic model selection procedure systematically evaluated 486 different ARDL specifications across all possible lag combinations, constrained by sample size considerations. The selection process employed multiple information criteria to ensure robust specification choice while maintaining statistical reliability (Table 5).

Table 3. *Descriptive statistics (2000–2023)*

Variable	Mean	Std. Dev.	Minimum	Maximum	CV
MM (per 100,000)	142.54	62.31	70.00	271.00	0.437
EDU (%)	85.58	13.27	58.70	104.70	0.155
TF (births/woman)	2.51	0.136	2.267	2.796	0.054
URB (%)	58.97	3.87	53.40	67.30	0.065
DPS (USD per capita)	52.57	24.06	13.86	86.69	0.457
CSU	0.37	0.25	0.135	0.87	0.676

Note: CV denotes coefficient of variation. MM = maternal mortality rate, EDU = primary education completion rate, TF = total fertility rate, URB = urbanization rate, DPS = public health spending per capita, CSU = universal health coverage rate (proportion, 0–1 scale).

Source: Authors' elaboration.

Table 4. *Augmented Dickey-Fuller Unit Root Test results*

Variable	Level	First Difference	Order of Integration
LMM	-2.440 (0.352)	-4.508** (0.009)	I(1)
LEDU	-1.876 (0.661)	-3.708** (0.046)	I(1)
LTF	-2.255 (0.438)	-17.402*** (0.000)	I(1)
LURB	-6.730*** (0.000)	-	I(0)
LDPS	-4.272** (0.014)	-	I(0)
LCSU	-4.360** (0.012)	-	I(0)

Notes: ***, *, * denote significance at 1%, 5%, and 10% levels respectively. P-values in parentheses. All tests include constant and trend. Lag length selected by SIC with maximum lags = 5.

Source: Authors' elaboration.

Table 5. *Model selection results (Top 5 models by SIC)*

Model	LogL	AIC	SIC*	HQ	Specification
168	60.09346325	-4.281223	-3.636517	-4.129350	ARDL(2,0,2,2,1,0)
11	64.01829528	-4.365299	-3.571814	-4.178378	ARDL(2,2,2,1,2,1)
167	60.56089249	-4.232808	-3.538508	-4.069252	ARDL(2,0,2,2,1,1)
2	65.00796687	-4.364360	-3.521282	-4.165756	ARDL(2,2,2,2,2,1)
165	60.18668121	-4.198789	-3.504489	-4.035233	ARDL(2,0,2,2,2,0)

Source: Authors' elaboration.

The selection of ARDL(2,0,2,2,1,0) over competing specifications requires careful justification given the marginal differences in information criteria (Table 6). While ARDL(2,2,2,1,2,1) achieves marginally superior AIC (-4.281 vs -4.365), the SIC criterion appropriately favours our selected model due to its stronger penalty for additional parameters. With 24 observations, the 17-parameter specification of the second-ranked model would consume 71 per cent of available degrees of freedom, risking severe overfitting and unstable parameter estimates. SIC applies a penalty of k×ln(n) compared to AIC's k×2, making it particularly appropriate for small samples where parsimony is crucial. The selected specification achieves optimal balance by capturing 99.83 per cent of maternal mortality variation while maintaining statistical reliability through appropriate parameterization with 13 parameters. The negligible improvement in fit (0.05% R^2 difference) does not justify the notable increase in model complexity.

ARDL estimation results. The estimation results offer exceptional explanatory power ($R^2 = 99.83\%$) and reveal several key findings that support our theoretical framework. Most importantly, the significant negative coefficient on UHC coverage (LCSU: -0.482, $p = 0.0018$) provides strong evidence that UHC expansion reduces maternal mortality in the short run. Both contemporary and lagged public health expenditure demonstrate negative, significant effects, suggesting that health spending improvements generate sustained benefits. The autoregressive structure reveals strong persistence in maternal mortality changes through lagged dependent variables, consistent with the gradual nature of health system transformations.

Cointegration analysis and long-run relationships

Bounds testing for cointegration. Both the F-statistic (13.782) and t-statistic (-8.177) substantially exceed their respective critical values at the 1 per cent

Table 6. *ARDL(2,0,2,2,1,0) estimation results with HAC standard errors*

Variable	Coefficient	HAC Std. Error	t-Statistic	p-value
Short-run Dynamics				
LMM(-1)	2.1647***	0.4480	4.8323	0.0009
LMM(-2)	-3.8549***	0.7410	-5.2024	0.0006
LEDU	0.2255	0.2442	0.9236	0.3797
LTF	-1.0130	0.6562	-1.5439	0.1570
LTF(-1)	0.0434	0.3937	0.1102	0.9147
LTF(-2)	0.5389***	0.1736	3.1044	0.0126
LURB	14.5812***	2.9489	4.9445	0.0008
LURB(-1)	-18.0634***	4.4813	-4.0308	0.0030
LURB(-2)	-6.8052	5.1134	-1.3309	0.2160
LDPS	-0.2192***	0.0628	-3.4911	0.0068
LDPS(-1)	-0.2371**	0.1404	-1.6890	0.1251
LCSU	-0.4823***	0.1104	-4.3690	0.0018
Constant	55.5574***	14.8296	3.7466	0.0046
Model Diagnostics				
R^2	0.9983	Adjusted R^2	0.9960	
F-statistic	434.56***	DW statistic	3.192	

Notes: ***, *, * denote significance at 1%, 5%, and 10% levels respectively. HAC standard errors employ the Newey-West procedure with automatic lag selection to address potential autocorrelation and heteroscedasticity.
Source: Authors' elaboration.

significance level, providing compelling evidence for cointegration among the variables. This confirms the existence of a stable long-run equilibrium relationship, validating the ARDL approach and justifying the estimation of long-run coefficients. The strength of our cointegration results establishes the foundation for meaningful causal inference regarding UHC effects on maternal health outcomes (Table 7).

Long-run elasticity estimates. The long-run elasticity estimates provide crucial insights into the sustainable impacts of health system determinants on maternal mortality. Most importantly, the UHC elasticity of -0.1793 indicates that each 1 per cent increase in UHC coverage reduces maternal mortality by approximately 0.18 per cent in long-run equilibrium ($p < 0.001$). This considerable effect size shows that Morocco's UHC expansion from 13.5 per cent

Table 7. *ARDL bounds test results*

Test Statistic	Value	Critical Values (1% level)	Decision
F-statistic	13.782***	I(0): 4.537, I(1): 6.370	Strong Cointegration
t-statistic	-8.177***	-	Confirmed

Notes: Critical values from Pesaran et al. (2001). Case III: unrestricted intercept and no trend. Both statistics substantially exceed critical values at the 1% significance level.

Source: Authors' elaboration.

Table 8. *Long-run coefficient estimates*

Variable	Coefficient	Std. Error	t-Statistic	p-value	95% CI
LEDU	0.0883	0.0748	1.1206	0.2780	[-0.071, 0.248]
LTF(-1)	-0.1601	0.2869	-0.5582	0.5840	[-0.767, 0.447]
LURB(-1)	-3.8240***	0.9127	-4.1899	0.0000	[-5.737, -1.911]
LDPS(-1)	-0.1698***	0.0332	-5.1160	0.0001	[-0.240, -0.099]
LCSU	-0.1793*	0.0398	-4.5055	0.0003	[-0.263, -0.096]

Notes: ***, *, * denote significance at 1%, 5%, and 10% levels respectively. All coefficients represent long-run elasticities.

Source: Authors' elaboration.

to 87 per cent coverage contributed significantly to the observed maternal mortality reduction through this specific channel.

The public health spending elasticity (-0.1698, $p < 0.001$) reveals remarkably similar magnitude effects, suggesting strong complementarity between coverage expansion and resource allocation. The urbanization elasticity (-3.824, $p < 0.001$) indicates substantial mortality reductions associated with urban residence, reflecting superior health care infrastructure and emergency obstetric services in urban areas. Notably, education and fertility demonstrate non-significant long-run effects, a methodological finding discussed in detail below in the section on methodological limitations (Table 8).

Error correction model

The estimated error-correction coefficient (–2.69, $p < 0.001$) indicates an exceptionally rapid adjustment process following deviations from the long-run equilibrium (Table 9). Although its magnitude exceeds the conventional range observed in macroeconomic time-series models, such large negative values are

Table 9. *Error correction model results*

Variable	Coefficient	Std. Error	t-Statistic	p-value
ECT (-1)	-2.6902*	0.3281	-8.1978	0.0000
ΔLCSU	-0.4823***	0.1104	-4.3690	0.0018
ΔLDPS	-0.2192***	0.0628	-3.4911	0.0068
ΔLURB	14.5812***	2.9489	4.9445	0.0008
R^2	0.9363	Adjusted R^2	0.9044	

Note: Only key coefficients shown for brevity. Complete results available upon request.
Source: Authors' elaboration.

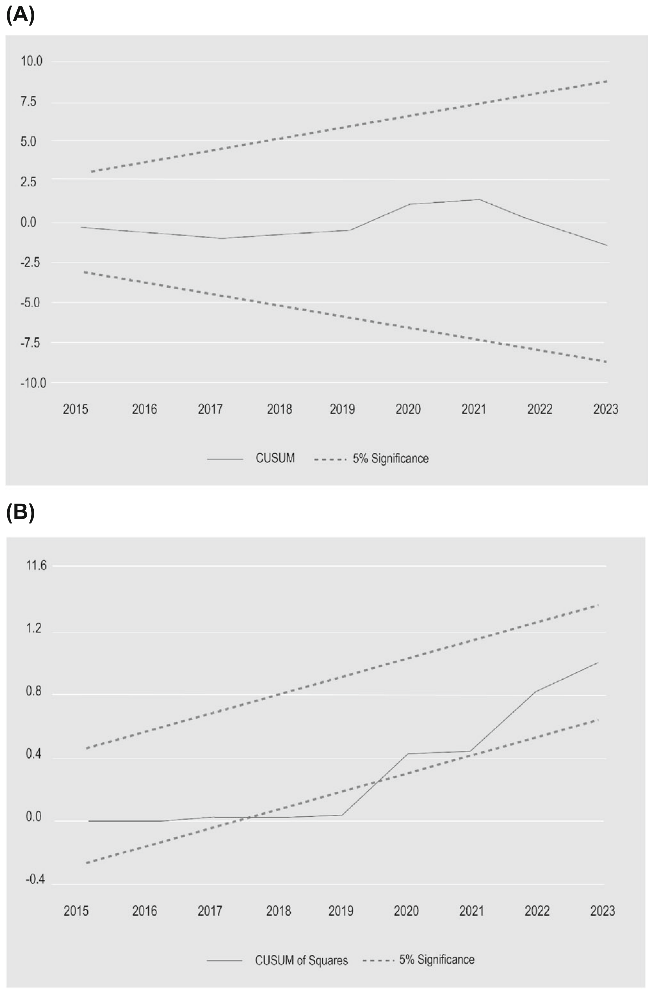
theoretically plausible when structural reforms abruptly remove access barriers, producing short-term overshooting before stabilization. According to institutional discontinuity theory (North, 1990; Acemoglu and Robinson, 2012), such regime shifts trigger immediate behavioural and institutional adjustments rather than gradual convergence. Structural-break tests confirm two major discontinuities – in 2012 (RAMED launch, t = −3.85) and 2020 (reform acceleration, t = −2.87) – that coincide with sharp reductions in maternal mortality (−8.2 per cent per year during 2012–2017 versus −3.4 per cent pre-reform) and a 235 per cent surge in institutional deliveries. These dynamics illustrate that Morocco's UHC reforms generated rapid, front-loaded effects followed by stabilization toward a new long-run equilibrium. HAC standard errors and parameter-stability diagnostics (Figure 1; Table 11) support specification robustness.

Parameter stability and causality analysis

Parameter stability: CUSUM tests. The CUSUM stability tests evaluate parameter constancy across the latter portion of our sample period (2015–2023), following the standard practice of reserving the initial two-thirds of observations (2000–2014) for parameter initialization in recursive testing procedures. The CUSUM test (Figure 1, Panel A) shows that recursive residuals remain largely within the 5 per cent significance bounds throughout the testing period, with a slight deviation near the upper-bound around 2020–2021. This minor instability coincides with the COVID-19 pandemic period and accelerated UHC reforms, suggesting temporary parameter variation during this exceptional period.

The CUSUM-squared test (Figure 1, Panel B) reveals more pronounced instability, with the statistic exceeding the 5 per cent confidence bands after 2020. This indicates some variance instability in the later period, likely reflecting structural changes in health system dynamics during the pandemic and

Figure 1. *CUSUM and CUSUM-Squared Stability Tests*



Source: Authors' elaboration.

intensified reform implementation. Despite these temporary deviations, the overall pattern supports parameter stability for most of the testing period, validating our long-run coefficient estimates while acknowledging some structural evolution in recent years.

Granger causality analysis. The Granger causality tests indicate no significant predictive relationships between past values of UHC coverage and future

Table 10. *VAR Granger causality test results*

Null Hypothesis	Chi-sq	df	p-value	Decision
LCSU does not Granger cause LMM	1.072	2	0.585	Cannot reject H_0
LMM does not Granger cause LCSU	2.883	2	0.237	Cannot reject H_0

Source: Authors' elaboration.

Table 11. *Comprehensive diagnostic test results*

Test	Statistic	p-value	Conclusion
Residual Diagnostics			
Jarque-Bera Normality	0.196	0.907	Normal residuals
Breusch-Godfrey LM(3)	5.818	0.121	No serial correlation
Breusch-Pagan Heteroscedasticity	0.300	0.972	Homoscedastic
Ramsey RESET	1.473	0.179	Correct specification
Parameter Stability			
CUSUM	Within	—	Stable
CUSUM-squared	Exceeds	—	Post-2020
Structural Break Tests			
Breakpoint test – LMM (2020)	-2.873	0.0003	Break
Breakpoint test – LCSU (2012)	-3.848	0.0021	Break

Source: Authors' elaboration.

maternal mortality changes, and vice versa. This finding complements rather than contradicts our ARDL cointegration results. While the ARDL demonstrates strong contemporaneous and long-run equilibrium relationships, the absence of Granger causality suggests that UHC effects operate primarily through immediate access mechanisms rather than lagged predictive relationships. This pattern is consistent with rapid access improvements following coverage expansion, where health benefits materialize quickly rather than requiring extended implementation periods (Table 10).

Diagnostic testing and robustness analysis

Model validation. The diagnostic tests validate our model specification. The Breusch-Godfrey test reveals serial correlation up to lag 3; however, the Newey-West HAC correction applied in our estimation procedure ensures robust standard errors and valid hypothesis testing. Additional tests confirm residual normality and homoscedasticity. Parameter stability analysis (CUSUM) demonstrates coefficient constancy, while structural break tests formally identify discontinuities in 2012 (RAMED launch, t = -3.85) and 2020 (reform acceleration, t = -2.87), validating the institutional step-functions underlying our rapid adjustment dynamics. These diagnostic results collectively support the reliability of our ARDL framework for examining the relationship between universal health coverage and maternal mortality (Table 11).

Robustness analysis. The robustness analysis demonstrates remarkable stability of UHC coefficient estimates across alternative specifications, with values ranging from -0.1721 to -0.1846 – representing variation of only 7.3 per cent. All estimates remain statistically significant at conventional levels, confirming that our primary findings do not depend on specific methodological choices or specification assumptions. The bootstrap confidence intervals provide additional assurance against small-sample bias concerns, while alternative cointegration estimators (FMOLS and DOLS) yield virtually identical results, further validating our analytical approach (Table 12).

Discussion

Principal findings and global significance

This analysis provides robust econometric evidence that Universal Health Coverage expansion reduces maternal mortality in Morocco through both immediate and sustained mechanisms. The long-run elasticity estimate of -0.179 per cent demonstrates that each 1 per cent increase in UHC coverage reduces maternal

Table 12. *Sensitivity analysis results*

Model Specification	UHC Long-run Coefficient	95% CI	p-value
Baseline ARDL	-0.1793	[-0.263, -0.096]	0.0003
Alternative lag selection (AIC)	-0.1846	[-0.281, -0.088]	0.0008
Alternative lag selection (HQ)	-0.1721	[-0.254, -0.090]	0.0005
Bootstrap (1000 replications)	-0.1793	[-0.267, -0.091]	0.0002
FMOLS estimator	-0.1834	[-0.276, -0.091]	0.0004
DOLS estimator	-0.1765	[-0.269, -0.084]	0.0007

Source: Authors' elaboration.

mortality by approximately 0.18 per cent in equilibrium, representing a quantification of UHC effects in the MENA region. This notable effect size indicates that Morocco's UHC expansion from 13.5 per cent to 87 per cent coverage contributed substantially to the observed maternal mortality reduction through this specific channel.

These findings gain importance given global UHC progress has stagnated since 2015, while achieving SDG maternal mortality targets requires unprecedented annual reduction rates of 15 per cent (WHO et al., 2025). The convergence of strong cointegration evidence (F-statistic = 13.782), valid error correction mechanisms, and substantial coefficient stability across methodological approaches (variation < 7.3%) establishes compelling evidence for causal relationships rather than mere association. This methodological rigor addresses persistent challenges in health system reform evaluation, particularly regarding endogeneity concerns that have limited previous research in developing country contexts.

Dynamic adjustment and complementarity mechanisms

The error correction mechanism reveals rapid adjustment dynamics characteristic of institutional reforms addressing multiple access barriers simultaneously. Morocco's dual-programme expansion (AMO + RAMED) generated convergence within approximately 4.5 months, reflecting maternal mortality's responsiveness to emergency obstetric care where timely facility access produces immediate survival gains.

The short-run UHC elasticity (-0.482) exceeding the long-run magnitude (-0.179) documents initial utilization surges followed by capacity-constrained moderation, underscoring that coverage expansion requires parallel infrastructure investment. The comparable elasticities for UHC coverage (-0.179) and health expenditure (-0.170) demonstrate strong complementarity, indicating that optimal resource allocation demands balanced investment in financial protection and system capacity.

Methodological limitations

Our ARDL analysis reveals non-significant coefficients for education and fertility, apparently contradicting the extensive literature demonstrating these factors as powerful maternal mortality determinants (Miller et al., 2016). This divergence reflects methodological rather than substantive issues. Three factors explain this pattern.

First, ARDL's lagged structure may capture education and fertility effects indirectly through correlated variables (UHC, health spending, urbanization), particularly given Morocco's integrated development trajectory where socio-economic improvements progressed simultaneously. Second, Morocco's universal free maternal care since 2008 – providing all pregnancy monitoring, delivery, and diagnostic services free in public facilities regardless of insurance status – may have attenuated education's direct marginal effect by removing financial barriers traditionally linked to educational inequalities. Third, our dependent variable is estimated by the United Nations using a model incorporating fertility and skilled attendance as explanatory variables, creating generated variable bias where estimation model inputs correlate with our regressors (Pagan, 1984). These factors suggest results that reflect ARDL estimation properties and institutional context rather than the absence of education-fertility effects. Policy makers should continue prioritizing girls' education and family planning as essential maternal mortality reduction components.

All variables except UHC coverage derive from international estimates rather than observed national data, with MMR estimates incorporating substantial uncertainty. Our 24-year series, while appropriate for ARDL methodology, provides limited degrees of freedom constraining model complexity. Implicit multicollinearity among slowly evolving variables (UHC, health spending, urbanization) may reduce coefficient precision without introducing bias, making inference fragile for individual regressors while maintaining overall model reliability.

Our demand-focused framework does not directly capture supply-side limitations. Recent evidence documents health workforce shortages and territorial inequalities (Bouzaidi and Ragbi, 2024), emergency care system deficiencies (CESE, 2023), and quality gaps requiring deliberate improvement initiatives (Elomrani et al., 2024). These constraints moderate UHC effectiveness, suggesting our elasticity estimate (-0.179) may represent a lower bound on achievable effects with concurrent supply strengthening. The documented complementarity between UHC and public health expenditure confirms that optimal policy requires balanced demand-side and supply-side investment.

Finally, Morocco's institutional specificities – universal free maternal care, dual AMO-RAMED structure, specific epidemiological context – limit direct extrapolation to other MENA countries. Our findings reflect Morocco's gradual dual-track expansion through 2021 with limited observation of the recent unified system, whose long-term effects require additional years of data for rigorous assessment.

Methodological contributions and external validity

This study addresses methodological challenges that have constrained UHC impact research in developing countries. Our ARDL application with $n = 24$ observations demonstrates rigorous time-series analysis remains feasible despite data limitations common in low- and middle-income countries. The ARDL framework's capacity to accommodate mixed integration orders, handle structural breaks, and address simultaneity bias through lagged dependent variables is particularly valuable for health system reform evaluation.

Our comprehensive robustness analysis across alternative specifications, information criteria, and estimation methods validates core findings while demonstrating methodological sophistication required for policy credibility. These findings provide econometric evidence of UHC's impact in the MENA region, addressing a critical deficiency in global health economics literature. Morocco's dual economy structure, middle-income status, and gradual reform approach characterize many countries pursuing UHC, suggesting broader applicability across similar contexts. Recent systematic reviews confirm mixed UHC effects globally (Hone et al., 2024), while our substantial elasticity estimates suggest well-designed reforms can achieve significant health gains even in complex institutional environments. These methodological and empirical contributions generate several actionable insights for health policy design.

Policy implications and strategic guidance

The results yield four inter-related policy lessons with relevance beyond the Moroccan context. First, integrated investment strategies prove essential: coverage expansion and public health spending are mutually reinforcing, requiring simultaneous attention to financial protection and service quality. Second, adaptive implementation approaches demonstrate effectiveness: rapid error-correction dynamics indicate that UHC effects emerge quickly, permitting front-loaded implementation with real-time adjustments. Third, policy makers should calibrate expectations: recognizing initial over-reaction and subsequent moderation maintains political momentum once early gains taper. Finally, enhanced monitoring systems enable the identification of the high elasticity of maternal mortality to UHC shocks, which justifies its inclusion as a sentinel indicator in performance dashboards, enabling timely feedback loops for adaptive governance.

Study limitations and research directions

Future research should extend these findings through panel ARDL applications combining multiple MENA countries, enabling larger samples and cross-country heterogeneity analysis. Granular investigation of transmission mechanisms –

antenatal care utilization, institutional delivery, emergency obstetric access – would clarify causal pathways. Comprehensive UHC assessment requires examining effects on child health, infectious disease control, and broader health system responsiveness. Household survey analysis would enable distributional effects investigation, assessing whether benefits reach intended populations. Finally, as Morocco's unified Compulsory Health Insurance (AMO) matures, comparative analysis of dual versus single-tier systems would inform institutional design choices for middle-income countries.

Conclusion

This study provides an econometric analysis of Universal Health Coverage's impact on maternal health outcomes in the Middle East and North Africa region, demonstrating that UHC expansion significantly reduces maternal mortality through both immediate and sustained mechanisms. Our findings contribute fundamentally to global understanding of UHC effectiveness when progress toward universal coverage has stagnated and innovative approaches are urgently needed for achieving the 2030 SDG targets, while acknowledging methodological limitations that inform the need for cautious interpretation.

Our application of ARDL bounds testing with Error Correction Modelling to UHC impact evaluation represents methodological advancement, successfully addressing mixed integration orders, structural breaks, and endogeneity concerns that have limited previous research in developing country contexts. The successful application with 24 observations confirms this approach's viability for data-constrained environments, providing a template for future research while maintaining methodological rigor required for policy credibility. While the error correction coefficient of -2.69 exceeds the typical range observed in financial markets, it reflects the institutional nature of health system reforms where policy changes generate immediate access improvements that temporarily exceed sustainable long-run gains before equilibrating toward stable relationships. While the distinction between short-run over-reaction and long-run equilibrium effects reveals complex adjustment processes involving immediate access improvements and gradual capacity equilibration.

Our findings address the four research questions directly. First, UHC expansion causally reduces maternal mortality (long-run: -0.179, $p < 0.001$; short-run: -0.482), with initial impacts exceeding sustained effects by factor 2.7. Second, UHC (-0.179) and health expenditure (-0.170) effects demonstrate strong complementarity, indicating that coverage expansion and quality improvement require simultaneous implementation. Third, adjustment dynamics (ECT = -2.69) indicate a rapid yet transitory overshooting response driven by institutional discontinuities in 2012 (RAMED) and 2020 (reform acceleration), validating step-function adjustment mechanisms. Fourth, robust cointegration (F = 13.782) and parameter stability confirm relationship stability. These results reveal that dual programme UHC approaches addressing formal and informal sectors simultaneously, with parallel infrastructure investment, optimize maternal health outcomes in middle-income contexts.

These findings provide robust empirical support for accelerated UHC investment, demonstrating that well-designed reforms achieve health improvements necessary for meeting global targets. The near-equal elasticities for UHC coverage (-0.179 per cent) and public health spending (-0.170 per cent) offer clear guidance for optimal resource allocation, indicating that coverage expansion and quality improvement should be pursued simultaneously rather than sequentially. This complementarity insight challenges conventional approaches that often prioritize single dimensions and supports integrated strategies addressing both financial protection and system capacity. As a pioneering UHC impact study in the MENA region, this analysis fills an important limitation in global health economics literature while demonstrating methodological approaches that enable other countries to conduct similar evaluations despite data constraints.

Morocco's experience provides valuable lessons for countries pursuing health system reforms, particularly those with similar dual economy structures and gradual implementation approaches. The findings contribute to understanding global maternal mortality challenges, demonstrating that meaningful progress remains achievable through appropriate institutional reforms even as global progress has stagnated since 2015. With maternal mortality requiring 15 per cent annual reductions to meet 2030 SDG targets, understanding effective interventions becomes increasingly important for evidence-based policy making. The study simultaneously advances health production function theory by explicitly incorporating institutional variables while demonstrating the importance of dynamic adjustment processes, with evidence for strong complementarity between UHC coverage and public expenditure contributing to understanding optimal health system design.

Future research should extend these findings through panel ARDL applications enabling larger sample sizes, cross-country comparative studies examining design feature effectiveness, and granular analysis of transmission mechanisms linking UHC expansion to health improvements. Understanding differential effects across population subgroups and examining UHC efficiency relative to alternative interventions represent additional priorities.

Universal Health Coverage represents both a moral imperative and an effective development strategy. Morocco's experience provides evidence that middle-income countries can implement transformative health system reforms despite resource constraints and institutional complexity. While global progress toward UHC remains off track, Morocco's systematic approach shows that rapid advancement remains possible with appropriate policy design, adequate financing, and sustained political commitment.

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Intersectionality and structural barriers in the last-mile delivery of social protection: Evidence concerning women's access to social protection in Gujarat, India

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Abstract Using an intersectional lens, this mixed-methods study examines how structural barriers in last-mile delivery shape women's access to social protection in cattle-rearing households in urban Gujarat, India, during and after the COVID-19 pandemic. Drawing on survey data from 427 households, interviews with frontline workers, key informants, and a focus group discussion, the study shows that ostensibly neutral last-mile implementation mechanisms can reproduce intersectional exclusions. Digital information dissemination

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and selective door-to-door outreach tend to disadvantage women without smartphone access and women from historically marginalized caste communities. Administrative requirements linked to identity documentation further constrain access by imposing high opportunity costs and reinforcing women's dependence on male household members. The findings point to the need for information dissemination beyond digital platforms, simplified and flexible administrative processes, and the use of linking organizations to mediate last-mile access.

Keywords coverage, ICT, India, social exclusion, social security administration, social structure, takeup, women

Introduction

Social protection refers to the policies and programmes that protect people against risks and vulnerabilities linked to life-cycle phases, mitigate the impact of shocks, and support people facing chronic incapacities to secure basic livelihoods and necessities (Adato and Hoddinott, 2008; Beveridge, 1942; Norton, Conway and Foster, 2001; Schneider, Mokomane and Graham, 2016). In line with the recent literature, we understand social protection through a rights-based approach, which frames recipients as active social actors entitled to protection, rather than as passive beneficiaries (Piron, 2004; Hickey, 2007). During the COVID-19 pandemic, several countries expanded social protection measures; however, significant implementation gaps persisted, particularly for marginalized groups who were formally entitled to social protection (Devereux, 2021). Many countries have lacked comprehensive strategies to ensure that social protection reaches populations facing intersecting disadvantages. Informal women workers (those without employment contracts, employment benefits, or representation) in urban settings are often overlooked in the design and implementation of social protection, despite being disproportionately affected by crises such as the COVID-19 pandemic, as existing measures have failed to adequately respond to their needs (Patel, 2022). Despite growing recognition of last-mile implementation gaps, far less attention has been paid to how such failures are produced through intersecting social vulnerabilities and recipients' everyday interactions with institutions.

Gender, caste and inequalities in access to social protection

The gendered effects of the COVID-19 pandemic on women in India are well-documented. The pandemic exacerbated preexisting gender inequalities (Rawat, 2025). Women were more likely to experience domestic violence, lose livelihoods, find it difficult to rejoin the labour force, and have an increased burden of unpaid work (Rukmini, 2020a; Rukmini, 2020b; Sanyal and Shrivastava, 2021; Deshpande, 2022). Their loss of livelihood led to higher food insecurity, poverty, and indebtedness for women (Agarwal, 2021). Despite these increased vulnerabilities, a significant gap has persisted between formal entitlement and actual access to social protection, with many women unable to obtain cash transfers and food assistance despite meeting eligibility criteria (Pande et al., 2020). The impact of the pandemic on women was not acknowledged in the government's annual budget, and many women-specific schemes faced budget cuts (Gupta, Ghosh and Bindal, 2021).

Caste inequities were also exacerbated during the pandemic in India (Siddiqui, 2021; Acharya and Christopher, 2022). Caste is a form of hierarchical social stratification that has historically been constructed through ideologies of purity and pollution, and graded inequality (Mudliar and Koontz, 2018). Many people from marginalized castes who live in urban areas were stigmatized as being "super-spreaders" of the COVID-19 virus, and the social distancing rules added to their ostracization (Siddiqui, 2021). The National Commission for Dalit Human Rights found significant gaps in social protection for marginalized castes and little awareness of announced entitlements, such as increased coverage under the national health insurance scheme and emergency cash transfers during the pandemic (NCDHR, 2020, p. 3).

These patterns are not unique to India. Comparable evidence across a range of low- and middle-income country contexts points to persistent gaps between formal entitlement and actual access to social protection, especially for women and socially marginalized groups (Arza, 2015; Camilletti, 2020; Lund, 2006; Kapindu, 2011; Ouma and Adésínà, 2018; Rocon et al., 2016; Santos, 2016; Staab, 2020; Zohir et al., 2008). Across these contexts, scholarship consistently identifies gender and caste as central axes through which experiences of crisis and access to social protection are structured (Adhikari et al., 2014; Cookson et al., 2023; Sabarwal, Sinha and Buvinic, 2011; Ulrichs, 2016; Holmes and Scott, 2016; Pankaj, 2019), especially in settings characterized by high levels of informality. We therefore make the methodological choice to focus on gender and caste in this article.[1]

1. While gender and caste constitute the primary analytical axes in this article, we recognize that access to social protection is also shaped by class position, marital status, and literacy. We briefly examine how these factors condition women's institutional positioning and mediate interactions with last-mile delivery systems, particularly in navigating documentation requirements, digital platforms, and bureaucratic procedures. While other axes of inequality, such as migration status or disability, are also relevant, they fall beyond the analytical scope of this study.

While much of the literature recognizes that barriers to social protection emerge through everyday interactions between institutions and recipients, these interactions are often examined along a single axis of disadvantage. This study moves beyond single-axis analyses by providing empirical evidence on how gender and caste intersect to shape access to social protection at the "last mile".

Analytical framework: Intersectionality and structural barriers in social protection access

Intersectionality provides a useful analytical lens that understands discrimination as the result of interlocking systems of oppression and domination (Crenshaw, 1989). As systems of power operate simultaneously rather than independently, disadvantage cannot be fully understood through single-axis analyses. Crenshaw (1989) introduced the term intersectionality to highlight how the experiences of white women define the experience of sex discrimination, while black men define the experience of racial discrimination, thus erasing the experiences of black women. "This tendency is perpetuated by a single-axis framework that is dominant in antidiscrimination law, and that is also reflected in feminist theory and antiracist politics" (Crenshaw, 1989, p. 139). Similarly, in India, feminist movements have tended to ignore caste as an important intersection, while anti-caste movements have typically overlooked gender (Siddiqui, 2020). When gender inequity intersects with other factors, such as caste-based discrimination, it further limits women's access to resources and opportunities (Thompson, Rohwerder and Mukherjee, 2023).

Identifying intersecting axes of disadvantage does not, on its own, explain how these power relations become consequential in everyday encounters with the State. Institutions responsible for the last-mile delivery of social protection are often formally framed as "neutral" toward vulnerability, treating disadvantaged groups as homogeneous categories, without factoring in the role of power structures and social positions within these vulnerabilities (Dutta, Agarwal and Sivakami, 2020).[2] However, it is important to note that social protection systems are also shaped by underlying normative assumptions and institutional value judgements (Bender and Kaltenborn, 2025), including ideas about responsibility, deservingness, and equity, which are mediated through socio-cultural contexts and power imbalances (Cookson and Barrantes, 2024). As a result, ostensibly neutral institutional designs and implementation processes can reproduce existing intersectional disadvantages through their operation at the last mile.

2. Specifically, Dutta, Agarwal and Sivakami (2020) describe how analyses of the impact of the COVID–19 pandemic in India have often treated all vulnerable groups as homogeneous entities without factoring in the role of power structures and social positions within these vulnerabilities.

Recent scholarship, therefore, highlights the analytical value of intersectional perspectives in social protection research, particularly for understanding how inequalities emerge at the intersections of policy design, implementation, and social structure, and for informing more responsive programme design (Frey, Thomas and Alajääskö, 2024). Building on this work, this study uses an intersectional lens to understand the structural barriers that shape access to social protection at the last mile.

Drawing on Kidd (2017), we conceptualize structural barriers as system-level constraints embedded in the design and implementation of social protection programmes. These barriers are not necessarily explicitly discriminatory, yet they interact with existing social relations and vulnerabilities, such as gender, caste and marital status, to shape who can access entitlements in practice. Existing global literature identifies structural barriers, including complex documentation requirements (Silliman Bhattacharjee, 2023; Hossain, 2011), burdensome administrative procedures (Martins et al., 2024; Pellissery, 2005), unequal modes of information dissemination, geographical inaccessibility (Rodrigues, Almeida and Fausto, 2021; Saprii et al., 2015), and inadequate supporting infrastructure to access social protection (Chen, Chindarkar and Xiao, 2019; Goldblatt, 2009; Kulkarni et al., 2022). Such structural barriers have also been identified in the Indian context. For example, Drèze, Khera and Somanchi (2021) show how discrepancies in demographic information between a woman's Aadhaar card and other required documents can prevent her from accessing social protection. While such barriers are increasingly well documented, more research is needed to clarify how their effects vary across intersecting social locations. Where multiple disadvantages intersect, such constraints can reinforce exclusion from social protection despite formal eligibility (Kidd, 2017). To operationalize this analytical framework, the study examines how intersecting identities and structural barriers shape access to social protection within a specific empirical setting, cattle-rearing households in urban Gujarat. The following section introduces the study context and explains the selection of cattle-rearing households in urban Gujarat as a critical site for examining last-mile social protection delivery.

Study context

India provides a critical context for examining last-mile social protection delivery, given the scale of informality, the centrality of women's unpaid and informal labour, and the reliance of social protection systems on documentation-based eligibility. In India, there are 420 million informal workers, many of whom were negatively impacted by the COVID-19 pandemic (Bacil and Soyer, 2020). Over

80 million households are engaged in cattle-rearing, the majority of them being smallholders. Approximately 70 per cent of human labour in India's dairy industry is performed by women (Gulati and Juneja, 2023). During the pandemic, supply chains in the dairy industry were disrupted, with difficulties experienced in transporting dairy products, increased inventory costs and a decline in revenue (Das, Sivaram and Thejesh, 2021). Women from cattle-rearing households were chosen for this study as they face a disproportionate share of the challenges posed by shocks in the dairy industry, including the COVID-19 pandemic, food shocks, and recurrent health emergencies in the household. In particular, this study examines challenges in accessing food security schemes and health insurance among women from cattle-rearing families in Gandhinagar and Ahmedabad, Gujarat, as pandemic-induced shifts in information dissemination and administrative practices continue to shape access to social protection. Food security schemes and health insurance were selected for this study because vulnerable informal-sector workers were at high risk of falling into poverty due to out-of-pocket health and food expenses during the pandemic, and these schemes sought to provide protection against these shocks. While empirically grounded in urban Gujarat, this study contributes to wider debates on access to social protection in the Global South.

Methods

Our research employs an exploratory qualitative analysis complemented by descriptive statistics to examine social protection access for women residing in cattle-rearing households in Gandhinagar and Ahmedabad, Gujarat, and their related challenges. For food security schemes, we focused on the Public Distribution System, a food assistance programme through which staple food grains are provided to vulnerable households at subsidized prices. During the pandemic, the Public Distribution System was expanded, through the Pradhan Mantri Garib Kalyan Anna Yojana, to offer free additional grains beyond the usual entitlements. Although eligibility for both of these schemes is determined at the household level, the quantity of grain provided typically corresponds to household size, making it beneficial to register all household members on the ration card.[3] For health insurance, we focused on the Pradhan Mantri Jan Arogya Yojana (PM-JAY), often described as the world's largest public health insurance programme. At the time of data collection, PM-JAY had recently absorbed the Gujarat state-provided health insurance scheme, Mukhyamantri Amrutam

3. Ration cards are a form of identity proof through which eligible people can purchase subsidized food grains from the Public Distribution System. There is typically one ration card per household that lists all eligible household members.

Yojana, and therefore, data were collected on access to both health insurance schemes.[4]

The data for this study were collected between August 2022 and February 2023, shortly after the third wave of the COVID-19 pandemic in Gujarat. Four different data types were collected:

- Semi-structured individual interviews with 24 frontline workers including community health workers,[5] ration-shop workers, and dairy cooperative heads were designed to explore personal experiences and insights regarding the last-mile implementation of social protection, and challenges faced in delivering social protection to women from cattle-rearing communities, especially women from marginalized caste communities.
- Key informant interviews with five experts possessing extensive experience in gender and social protection, and cattle-rearing communities, to validate and expand upon the emerging patterns.
- One focus group discussion involving eight women affiliated with the Pethapur Mahila Dudh Utpadak Sahakari Mandali, a dairy cooperative run entirely by women. The participants were sampled to be representative of the different caste groups in the ward.
- Descriptive statistics from a cross-sectional structured survey using stratified random sampling to women from 427 smallholder cattle-rearing households (who own ten or fewer cattle) in Gandhinagar and Ahmedabad.

For the qualitative research, we purposively sampled respondents for their potential to provide in-depth insights into intersectional access to social protection. The qualitative interviews and focus group discussions were conducted in either English or Gujarati, and the data produced were translated into English. The interviews and focus group discussions were transcribed and subsequently coded using MAXQDA for the thematic analysis in English and Gujarati separately. Emerging codes and themes were then systematically extracted and analysed using an intersectional framework, focusing on gender, caste, structural barriers, and access to social protection. After the thematic analysis, we compared the coding in English and Gujarati to ensure that information was not lost in the process of translation.

4. Eligibility for both schemes is determined at the household level, while each household member must hold a valid health insurance card to access entitlements. Such household-based targeting, while administratively convenient, risks overlooking unequal decision-making power within households, a concern that becomes salient in the implementation failures examined in this study.
5. By the term "community health worker", we mean the ASHA (Accredited Social Health Activists) and Anganwadi (primary nutrition and early childhood care) workers. Only women are recruited for these roles. Community health workers are envisioned to act as "link workers", bridging communities to health and social protection services (Sapri et al., 2015).

Participants for the quantitative cross-sectional survey were recruited via a stratified random sampling method. Madhur Dairy provided a list of wards[6] with dairy cooperatives in Gandhinagar. Wards were randomly selected from this list, and cattle-rearing households were chosen at random when enumerators visited these wards. Given that a list of wards with dairy cooperatives could not be obtained for Ahmedabad, the Indian Institute of Public Health, Gandhinagar, facilitated access to a list of wards in Ahmedabad and their respective cattle-rearing populations. Wards were randomly selected from this list. Data were collected through face-to-face interviews using KoboCollect on enumerators' smartphones. The data were then cleaned and analysed through Stata 14.2.

Participants were provided with information about the study, and written consent was obtained. The study was conducted after obtaining appropriate institutional review board approvals. We have utilized pseudonyms for participants to ensure privacy and confidentiality. To protect participants and reduce deductive disclosure, local stakeholders recommended redacting the caste name and instead using the umbrella term "marginalized caste community".

Results

The results show how structural barriers embedded in the last-mile delivery of social protection shape access in intersectional ways. Seemingly neutral implementation practices interact with gendered, caste-based, and marital-status inequalities to produce patterned exclusions. While these barriers may affect many people seeking social protection, women from cattle-rearing households experience them more acutely due to intersecting vulnerabilities.

Structural barriers in accessing information about social protection

Digital dissemination and intersectional exclusions in information access. WhatsApp is a common communication channel through which community health workers, who are typically women, and ration-shop workers can disseminate information on health and social protection. The shift to using digital platforms is a relatively new phenomenon that became popular during the COVID-19 pandemic due to the strict lockdown and social distancing regulations. Community health workers create WhatsApp groups that usually only include women in the ward. These groups are used to share information on health practices, such as mask-wearing during the pandemic, as well as details about new social protection schemes and how to access them. Only women are

6. A ward is a basic administrative unit in Indian cities.

typically included in these groups, as the health information is primarily relevant to pregnant and breastfeeding women, as well as parents of young children.

Most community health workers acknowledged that not all women have access to smartphones.[7] In such cases, information is reportedly sent to husbands, based on the assumption that they will relay messages to their wives. As a result, when digital platforms are used, women's access to information about women-targeted social protection schemes and health interventions is mediated through their husbands. Attempts to contact women directly are limited, as women often provide their husbands' phone numbers because they primarily use the same device. Information is therefore missed when husbands are unavailable or lack sufficient prepaid mobile phone credit.

These constraints intersect with caste-based exclusions. Women from marginalized caste communities often face compounded barriers to digital access, and in our interviews, they were also less likely to be proactively contacted or included in WhatsApp groups by community health workers. In our interviews, several community health workers reported avoiding making phone calls or digital outreach to women from marginalized caste communities altogether, often citing anticipated disinterest or non-responsiveness. As one community health worker explained:

> "There is a WhatsApp group for women, especially mothers, so that we can disseminate information to them … Many of the mothers are on WhatsApp. Those who can afford phones are on WhatsApp, even though they mostly use their husbands' phones. People from marginalized caste communities are not interested anyway, so it does not matter if they are not on WhatsApp" (Community Health Worker – 3).

This quote illustrates the common misconception among community health workers and local government officials that women from marginalized caste communities are simply not interested in social protection programmes and related information. This perception conflates lack of access with lack of interest, obscuring the structural and social barriers that limit women's inclusion in digital information channels. As conceptualized earlier, these patterns illustrate how apparently neutral dissemination mechanisms, such as WhatsApp groups and phone-based outreach, operate as structural barriers when they intersect with gendered resource constraints, caste-based assumptions, and frontline discretion.

7. In this article, we use "smartphone" to refer to internet-enabled mobile devices that support app-based platforms such as WhatsApp, while "mobile phone" includes basic phones limited to voice calls and SMS.

Constraints associated with door-to-door dissemination. After the relaxation of pandemic-era social distancing rules, some community health workers and local government officials preferred to revert to door-to-door visits to disseminate information directly to beneficiaries. While door-to-door dissemination is often presented as a corrective to digital exclusion, our findings show that its effectiveness is also shaped by a combination of frontline practices and broader caste- and gender-based constraints.

Our survey data indicate that information delivered to women does not automatically translate into access, as decisions about uptake are often negotiated within households. When visiting households, community health workers primarily engage with the women. This is due to their preference for speaking woman-to-woman, often discussing topics such as health and nutrition, which women tend to be more responsible for in their households. Another reason is that men are usually away at work when visits occur. However, data from our structured survey (n=427) indicate that men unilaterally make decisions about accessing social protection in approximately 53 per cent of households. Women unilaterally make these decisions in only about 12 per cent of households, while decisions are jointly made in approximately 34 per cent of cases. Men's decision-making authority is thus likely to influence health outcomes for the entire household and has direct implications for access to social protection, including health insurance during times of health crises. Consequently, whether a household accesses social protection often depends on whether men are informed by women, and whether they value the scheme and act on that information. These dissemination gaps, therefore, matter not only because information fails to reach women, but because women's ability to act on information is mediated by gendered intra-household power relations. Where decision-making authority is primarily held by men, dissemination failures may interact with existing intra-household power dynamics to limit women's effective access to social protection.

Another set of constraints relates to how frontline workers organize and undertake door-to-door dissemination under conditions of limited incentives and institutional support. Ration-shop workers are responsible for ensuring that beneficiaries receive the grain to which they are entitled under the Public Distribution System and for informing residents about new entitlements available under specific schemes. During the COVID-19 pandemic, for example, the government introduced the Pradhan Mantri Garib Kalyan Anna Yojana, through which eligible beneficiaries could obtain additional rations. Several frontline workers noted that the absence of formal incentives, additional remuneration, or institutional support for outreach beyond routine duties led them to rely on informal social networks and ad hoc arrangements to disseminate information about new entitlements. Most ration-shop workers and community health workers therefore relied on word of mouth through women's social networks. Some instructed women beneficiaries collecting rations, particularly those who were well-connected within the community, to spread the information among their neighbours. In other cases, ration-shop workers made direct phone calls to well-known women and asked them to pass the message on to others. In a few wards, ration-shop workers paid local women a small amount of money to visit households and convey details of the new scheme. However, reliance on these informal dissemination channels tended to privilege women who were already socially connected within the ward. Women from marginalized caste communities were often less embedded in these networks or lived in geographically peripheral areas within the ward, increasing the likelihood that they would be missed by such forms of outreach. As a result, information dissemination remained uneven, even within the same ward. Those responsible for last-mile implementation could not credibly confirm whether these methods consistently reached socio-economically marginalized households or women residing at the geographic and social margins.

Door-to-door dissemination practices are further shaped by caste-based settlement patterns and safety concerns reported by frontline workers, which affect whom they are able or willing to reach through door-to-door visits. It is often the case that marginalized caste communities live in the peripheries due to historical segregation. This geographical barrier acts to maintain an intersectional injustice. Community health workers express a certain sense of hesitation in disseminating information to people from marginalized caste communities, even when there is a sizeable proportion of marginalized caste community households in their ward of jurisdiction. Community health workers typically exclude the houses of women who live in the peripheries, i.e. where the marginalized caste households are found, citing geographical barriers:

> "There is difficulty in reaching some houses in marginalized caste settlements, especially when we have to go alone to geographically remote areas within our ward. We cannot go there easily as there is reluctance amongst the people who live there. Additionally, they live in far-off places in the ward. If something bad were to happen to us there, no one would even come to our rescue" (Community health worker – 6).

A few community health workers described feeling unsafe during door-to-door visits and reported experiencing inappropriate comments from some men in the areas they served, which reduced their willingness to personally disseminate information in certain neighbourhoods. Health workers also perceived women in marginalized caste communities as less receptive during door-to-door outreach. Several community health workers explained their reluctance to engage more actively in these areas by referring to concerns about hostility or a perceived lack of interest among residents. However, as reflected in interviews as well as in focus group discussions, such perceptions are shaped by long-standing histories of exclusion and unequal power relations. Distrust and suspicion toward community health workers and government representatives emerged as responses to past experiences of stigmatization and disregard, particularly among marginalized caste communities. Rather than indicating disinterest, limited engagement often reflected these broader socio-historical dynamics, which continue to shape interactions at the citizen-administration interface. As one community health worker explained:

> "We asked them why they do not answer the door and why they do not allow us to enter the house. They feel a sense of fear and suspicion in their minds when they see people from different communities. To be honest, we also feel that when we see them. Since those women do not like it if we go to them, we stopped going. If they ever approach us, we will give them information about the schemes" (Community Health Worker – 4).

This excerpt illustrates how interactions between community health workers, who in our study sites were predominantly from relatively privileged castes, and women from marginalized caste communities are shaped by long-standing histories of exclusion, stigma, and asymmetric power relations. Rather than reflecting disinterest or refusal on the part of marginalized women, such encounters point to deeply rooted distrust toward state representatives that has accumulated over time. In turn, frontline workers' responses to this distrust influence who is reached through door-to-door dissemination and who is not.

Importantly, these dynamics unfold within institutional arrangements that formally treat dissemination as a neutral and routine task, while leaving considerable discretion to frontline workers in deciding how, where, and to whom information is conveyed. In practice, this discretion allows historical social hierarchies and everyday safety concerns to shape outreach decisions, resulting in the systematic exclusion of women living at the social and geographic margins. Given that women from marginalized caste communities are more likely to be omitted from door-to-door visits, they often do not receive information about available schemes, eligibility criteria, or procedural steps, making access to health insurance and food security programmes substantially more difficult despite formal entitlement.

In a small number of cases, community health workers sought to address these barriers by partnering with local government officials, dairy cooperative leaders, or members of women's collectives during door-to-door visits. As these intermediaries were already trusted within the community, their involvement helped ease communication and reduce suspicion, suggesting that leveraging existing social networks can partially mitigate exclusions produced at the citizen-administration interface.

Administrative barriers and documentation-related exclusion

A central structural barrier to accessing social protection in the study sites relates to documentation and the associated administrative processes required to prove and maintain eligibility. Social conventions and traditional family structures dictate that women change their surname after marriage, which in turn makes it mandatory for them to update this information across multiple official documents. Still, many women are often not aware of this rule, and changing the documents is a cumbersome process that requires providing several supplementary documents and undertaking multiple visits to the responsible government office. For example, women and their family members must submit affidavits confirming the woman's legal marriage, obtain approval from local governance authorities, and provide proof of their marital status. Women must then rely on community health workers or local government agents to upload this information so that identity documents, most notably Aadhaar cards, can be updated. After updating their Aadhaar card, all other social protection documentation, such as ration cards and health insurance cards, needs to be updated.

These documentation requirements constitute a structural barrier, as access depends not only on formal eligibility but on navigating complex administrative sequences that are unevenly accessible across households and social groups. In practice, the administrative burden required to update documentation often leads to the updating of documents not being considered a priority in the household that women are married into. As a participant in the focus group discussion described:

> "Men think that getting a health insurance card is too much of a hassle. When men hear of such schemes, they avail of them, but often only for themselves. It is such a hassle for them to avail for themselves that they do not want to put this effort in for the women in their family, especially for their wives who they recently married" (Focus Group Discussion Participant).

These administrative burdens are experienced unevenly, becoming more prohibitive at the intersection of gender, caste and class. While this lack of documentation affects many women, it is even more pronounced in marginalized caste communities. As many men in these communities are daily-wage labourers, this limits their ability to take a day off (and also to lose wages) in order to go to the district-level *Jan Seva Kendra*[8] multiple times to arrange for documents for themselves, or their wives. The opportunity cost of lost wages, combined with travel time and bureaucratic delays, significantly constrains the ability of households to complete required updates. In several cases, respondents also reported being ill-treated or turned away when attempting to update documents, reflecting the persistence of caste-based prejudice within administrative spaces.

Literacy further shapes how these structural barriers are navigated. Sixty per cent of women in the survey could not read or write, limiting their ability to independently manage documentation processes that involve form-filling, digital interfaces, and interactions with officials. As a result, women, particularly, daughters-in-law,[9] remain dependent on husbands or other male relatives to complete administrative procedures, reinforcing gendered dependencies in access to social protection.

Exclusion is further compounded by the interdependence of documentation across schemes. For instance, to obtain a health insurance[10] card, one needs to show a ration card as proof. In many cases, some documentation may present the surname prior to marriage, while other documents have the post-marriage surname. In such cases, women are unable to prove their eligibility despite being formally entitled, resulting in the denial of social protection entitlements.

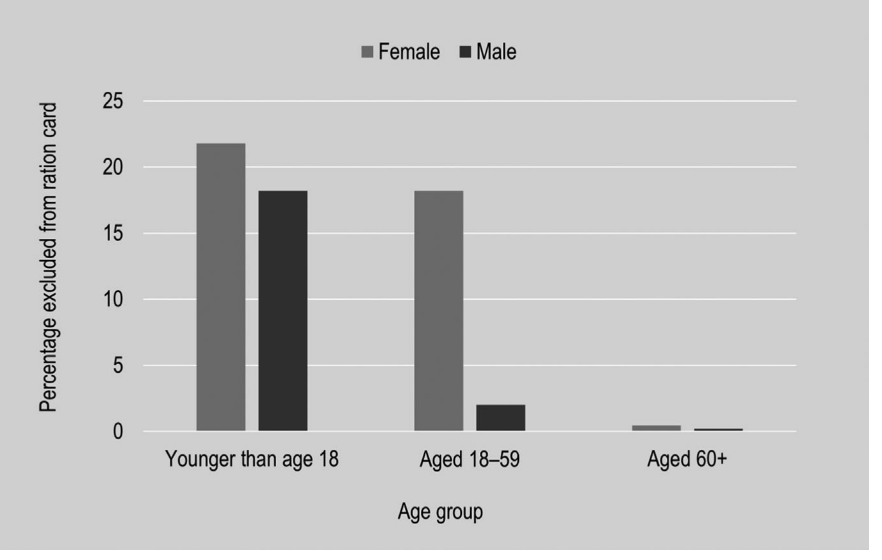
As shown in Figure 1, women and children were disproportionately omitted from the household ration card. In 18 per cent of the households surveyed in this study, women aged 18–59 were reported to be absent from the household ration card. This age bracket typically corresponded to the daughter-in-law of the house. Twenty per cent of the households we surveyed reported that the children's names were absent from the household ration card. When this was brought up during the key informant interviews and expert interviews, it was clarified that for children to have documentation, both their parents must first have valid documents. As many mothers do not have updated documents, their children also do not have documents and, as a result, cannot access Public Distribution System entitlements.

8. The term Jan Seva Kendra translates to "Public Service Center". In this context, it is run by the government and aims to act as a one-stop hub where people can apply for documents, such as ration cards, birth and marriage certificates, and social protection, without having to visit multiple government offices.

9. We use the term daughter-in-law (instead of wife) as most of the households included in our study are multigenerational families, with at least two adult women: the mothers-in-law and the daughters-in-law. While neither are typically the central decision-maker in the household, daughters-in-law are responsible for most of the household chores, including cooking.

10. The health insurance card needed to access entitlements under the PM-JAY scheme is called the Ayushman Card.

Figure 1. *Gendered exclusion from household ration cards across age groups (%)*



Source: Authors' elaboration.

Frontline workers do encourage women to update their surnames in all their documentation to access maternal and child-focused social protection schemes. However, as highlighted in the expert interviews, this is often limited to those within the social network of the community health worker and, thus, those who possess some social capital, leaving women from marginalized caste communities, once again, excluded from such support. Currently, collective organizations such as women's self-help groups and dairy cooperatives are especially proactive in facilitating marginalized women's access to social protection by providing information, helping update documents, and completing forms to enrol in various social protection schemes. They act as bridges, given that they know, on the one hand, the women and their families and, on the other, they are familiar with government administration and know how to access social protection. Nevertheless, there is a stark lack of formal, state-backed alternative processes that allow women without proper documentation to benefit from social protection – or to easily obtain the necessary documentation in the first instance.

Discussion

While well-designed social protection has the potential to reduce out-of-pocket expenditure during health shocks and protect women's food security, particularly in times of crisis such as the COVID-19 pandemic, our findings illustrate how ostensibly neutral last-mile institutions can interact with gendered and caste-based inequalities in ways that contribute to uneven access to social protection. Drawing on an intersectional approach, this study examines how gender and caste influence women's access to social protection. Intersectionality highlights how multiple axes of power and disadvantage interact, producing barriers that cannot be understood through single-axis analyses alone (Crenshaw, 1989). By applying this lens to last-mile implementation, we show how the last-mile delivery of social protection systems can nonetheless exclude marginalized women when intersecting social hierarchies are rendered invisible in everyday implementation processes.

An important result of this study is that two interrelated last-mile implementation mechanisms, digital dissemination of information and selective door-to-door outreach, systematically limit women's access to social protection. Women from marginalized caste communities were less likely to receive information about social protection because dissemination increasingly relied on digital platforms to which many women lack access, or on door-to-door outreach by community health workers who often did not visit marginalized neighbourhoods, citing geographical constraints, safety concerns, or perceived disinterest. The switch to digital platforms has failed to acknowledge existing structural inequities and consider how many marginalized women do not have access to mobile phones. According to the Mobile Gender Gap Report, only 26 per cent of women in India have smartphones (Shanahan, 2022), most of them likely from privileged backgrounds. There continues to be a gender gap in mobile phone ownership and usage in India, especially in low-income settings, and women tend to experience digital barriers more acutely than men (Nagpal and Bamezai, 2022). Our results concur with similar studies that show how the digital divide has prevented women who would benefit most from health insurance and food security programmes from accessing these, especially at a time when timely and unrestricted access was critical, such as during the COVID-19 pandemic (Biswal and Rath, 2022). Comparative evidence from other Global South contexts, particularly Kenya and Indonesia, similarly demonstrates that the use of digital and technological platforms in social protection delivery can generate new forms of exclusion when access to public information is mediated solely through digital platforms (Onyango and Ondiek, 2023; Dwi Wahyunengseh et al., 2020). Although the pandemic has subsided, digital modes of information dissemination continue to be used, suggesting that this form of exclusion is not temporary but is now embedded in routine implementation practices. Our results resonate with institutional analyses of social protection delivery, which suggest that where responsibilities for outreach and information dissemination are fragmented and accountability mechanisms are limited, frontline actors exercise considerable discretion in implementation (Bender and Agbaam, 2025; Cameron and Shah, 2014; Munge et al., 2019). This is not simply the result of logistical constraints; it arises from seemingly neutral implementation arrangements that fail to account for structural inequalities in accessing information, thereby translating existing intersectional social disadvantage into uneven access to social protection.

Previous studies highlight how people from marginalized caste communities experience systematic exclusion in accessing information about social protection and in availing of social protection schemes (Chander, 2019; Saha, 2019; Patel, Das and Das, 2018). This exclusion is not incidental and stems partly from the reluctance of privileged-caste community health workers to engage with marginalized caste communities (Roalkvam, 2014). Our study further explores this to show that community health workers tend to justify excluding women from marginalized caste communities from WhatsApp groups and door-to-door dissemination by attributing to these women perceived hostility or disinterest. Such interpretations overlook how the distrust and suspicion that women from marginalized caste communities feel toward community health workers or local government representatives are rooted in long histories of exclusion and stigmatization. This distrust likely intensified during the COVID-19 pandemic amid stigmatizing narratives (Siddiqui, 2021). This strained relationship has widened the distance between marginalized communities and frontline implementers and continues to limit effective dissemination of information concerning social protection.

Similar forms of selective engagement and exclusion by frontline workers have been documented in other contexts, including discrimination against homeless populations and individuals with low socio-economic status in Brazil's health and social protection systems (Albuquerque et al., 2019; Martins et al., 2024), as well as frontline workers' prioritization of communities located closer to administrative centres in Colombia (Castañeda et al., 2005) and constrained outreach to remote and vulnerable households in Pakistan due to financial, transportation, and security-related barriers (Kumar et al., 2023). In this sense, our findings support Sepúlveda Carmona's (2017) argument that social protection systems that appear neutral to existing socio-cultural norms can inadvertently perpetuate, or even deepen, marginalization.

Additionally, we find that in cases where door-to-door dissemination does occur, community health workers often direct information on social protection to women in the household. However, decisions regarding uptake are typically made by men in the household and are influenced by prevailing gender and caste norms. Prior studies similarly document how intra-household power dynamics condition women's ability to translate information into access, with men in some households fearing social stigma from women's economic empowerment, while in others supporting women's increased autonomy (Farnworth, Ravichandran, and Galiè, 2023; Surendran-Padmaja, Khed and Krishna, 2023). Therefore, women's access to social protection is shaped not only by programme design but by the interaction between structural barriers in last-mile implementation and intersecting gender- and caste-based power relations.

A third important result of this study concerns the bureaucratic barriers created by post-marriage surname changes, which frequently prevent women from accessing health insurance and food security schemes when their updated identities remain undocumented. Our findings corroborate earlier research showing how the lack of necessary documentation is a structural barrier that prevents women from accessing social protection (Drèze, Khera and Somanchi, 2021; Falcao, Sachin and Painkra, 2019). Similar dynamics have been documented in Maharashtra, India, where women are excluded from social protection because their post-marriage identities are not recognized by administrative systems (Kalyan Shankar, Deulgaonkar and Sahni, 2024). We extend this literature by showing how these documentation barriers are not gender-neutral in their effects but intersect with caste to deepen exclusion. Women in marginalized caste communities often face limited access to information about post-marriage documentation requirements, reflecting broader inequalities in literacy, information dissemination, and bureaucratic design. Given that administrative procedures are not tailored to accommodate these constraints, women frequently depend on other household members to navigate them. For households reliant on informal or daily-wage labour, the opportunity costs of repeated visits to government offices further compound these barriers. These findings align with international evidence showing how complex documentation and bureaucratic requirements can intensify exclusion for disadvantaged groups. In South Africa, applicants for disability grants must navigate multiple government departments, substantially increasing barriers to access (Goldblatt, 2009), while in Nepal, the absence of citizenship documentation or birth certificates has been identified as a key reason for exclusion from the universal social pension (Speck, 2021). As a result, such seemingly neutral administrative requirements translate into unequal access outcomes at the last mile. Our results also show that additional assistance in updating surnames on documents to receive social protection is often limited to those in the social network of the community health worker, thus furthering the systemic exclusion of marginalized caste communities who lack the social capital to access social protection. The absence of effective grievance redressal or social accountability mechanisms further limits women's ability to contest such exclusion.

The critical limitation of this study is that the number of women from marginalized caste communities included in the sample for qualitative interviews is low. This reflects both the reluctance of women from marginalized caste communities to participate in the study, as well as the fact that very few women from marginalized castes are working in capacities responsible for the implementation of social protection. We tried to overcome this barrier by including women from marginalized castes in the focus group discussion as well as by conducting expert interviews with NGO workers who worked with women from marginalized castes and district-level government officials who were more sensitive to the challenges faced by women from marginalized castes. In this regard, we recommend further research in this area, particularly in-depth participant-centred qualitative studies with women from marginalized castes, conducted after establishing better trust and rapport, as indicated in other studies that recommend best practices for conducting fieldwork with marginalized communities (Mosavel et al., 2011; Potnis and Gala, 2020).

Conclusion: Policy recommendations

Our findings show that exclusion from social protection is produced at the citizen-administration interface through routine last-mile implementation practices that interact with gendered and caste-based hierarchies. Digital information dissemination that assumes smartphone access, selective door-to-door outreach shaped by frontline workers' discretion and power asymmetries, and strict documentation and record-matching requirements together function as structural barriers that disproportionately constrain marginalized women's ability to access social protection despite formal eligibility. In this way, exclusion is reproduced within ostensibly neutral institutions at the last mile. Addressing these patterns requires policy responses that focus on implementation design and everyday administrative practice. Our results allow us to derive the following policy recommendations:

First, training and sensitization programmes for community health workers and local government representatives should address entrenched social hierarchies and the mutual distrust that can exist between those responsible for implementing social protection and marginalized communities. At the same time, high workloads, inadequate remuneration, and limited training can reinforce frontline workers' selective outreach and reliance on existing social networks, thereby reproducing social hierarchies in last-mile implementation. Therefore, addressing exclusion requires structural improvements in working conditions, incentives, and institutional support for frontline workers; this recommendation concurs with existing studies that emphasize the importance of adequate financial incentives for community health workers (Sarin et al., 2016; Pani et al., 2022). Second, greater representation of women from marginalized caste communities among community health workers could help address information gaps by reducing caste-based barriers to outreach. However, as previous research cautions, recruitment alone does not automatically dismantle caste hierarchies, as marginalized caste health workers may continue to face discrimination or restricted mobility in privileged-caste neighbourhoods (Sharma, Webster and Bhattacharyya, 2014). Complementary strategies are therefore required. These include deliberately using trusted social networks within marginalized caste communities, such as women's self-help groups, dairy cooperatives, and other linking organizations, as channels for disseminating information through face-to-face interactions. Such approaches can help counter both caste-based exclusion and the limitations of digital dissemination (Thomas et al., 2022; George et al., 2023). Third, the growing reliance of social protection programmes on strict documentation requirements and exact record matching across administrative systems needs urgent review from a gender and equity perspective. This calls for flexible administrative procedures, including allowing alternative forms of proof of identity, simplifying correction processes for name mismatches, and training frontline officials to resolve documentation issues rather than obstruct access to social protection. Comparable social protection programmes have addressed documentation gaps by accepting locally verified alternatives, such as voter cards in Uganda and clinic cards, affidavits, or school records in South Africa, often combined with temporary enrolment while formal identity documents are corrected (Brook, Jones and Merttens, 2014; Goldblatt, 2005). More broadly, shifting toward more universal social protection schemes, drawing on subnational evidence such as the universalization of the Public Distribution System in Tamil Nadu (Swaminathan, 2008), may reduce exclusion errors and reliance on discretionary gatekeeping.

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Socioeconomic mortality gaps and contributory pensions in Uruguay

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Abstract In this article, we study how differential mortality at the socioeconomic level affects inequality within social security in Uruguay, a country with a relatively low incidence of informality and unemployment in the Latin American region. Lower-income workers meet pension access requirements to a lesser extent and are more likely to retire later than higher-income workers. Also, they tend to die younger. Using administrative records, we simulate work histories adjusted by survival probabilities by socioeconomic level. This research found that differential mortality is less important than differences in contribution patterns in Uruguay. The bottleneck is access to benefits at the usual retirement ages.

Keywords ageing, equal treatment, pension scheme, retirement, social policy, social security planning, Uruguay

Introduction

Contributory pension systems pursue multiple objectives, including providing insurance against longevity risk, smoothing consumption over the life cycle, reducing poverty in old age, redistributing income, and mitigating inequality (Barr and Diamond, 2010; Holzmann and Hinz, 2005). From the perspective of individuals, two objectives are particularly central. First, pension systems provide insurance against the risk of outliving individual savings by providing annuitized income streams that ensure financial security in old age. Second, they facilitate intertemporal consumption smoothing, partially correcting for myopia and

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behavioural or informational frictions that may otherwise lead individuals to undersave for retirement.

Beyond these insurance functions, governments frequently incorporate explicit redistributive mechanisms aimed at reducing poverty and inequality. Pension formulas often generate higher implicit returns for low-income workers through minimum benefits, progressive replacement rates, or contribution subsidies. These design features seek to compensate for unequal saving capabilities across individuals and to promote redistribution within the social security system.

However, the distributive capacity of pension systems may be affected by systematic differences in mortality across population groups (Auerbach et al., 2017; Boado-Penas, Habermann and Naka, 2020; Breyer and Hupfeld, 2010; Goda, Shoven and Slavov, 2011; Nelissen, 1999; Olivera, 2019; Shi and Kolk, 2022). Given that pensions are typically paid as lifelong annuities with limited withdrawal flexibility, individuals who live longer receive benefits for a longer period. If life expectancy were identical across individuals, this feature would not alter redistribution outcomes. In practice, however, mortality varies substantially across demographic and socioeconomic groups. One well-established dimension is gender, which has generated extensive debate regarding whether pension systems should differentiate benefits using gender-specific mortality tables (IWPR, 2000; James, Cox Edwards and Wong, 2003; Bertranou, 2006).

An equally important and increasingly studied dimension concerns socioeconomic mortality gaps, defined as systematic differences in mortality rates or life expectancy across socioeconomic groups. A large body of evidence shows that individuals with lower socioeconomic status – measured by income, education, or occupation – experience higher mortality risks and shorter life expectancy than more advantaged individuals (Chetty et al., 2016; Bound et al., 2014; Duggan, Gillingham and Greenlees, 2008; Gerdtham and Johannesson, 2004; Olshansky et al., 2012; Shi and Kolk, 2022). These differences imply that individuals from lower socioeconomic groups face shorter periods of benefit collection and a higher probability of death before the retirement age, potentially weakening the redistributive objectives of contributory pension systems.

This raises a central question for pension policy design: how do socioeconomic mortality gaps affect inequality within social security systems? In particular, differential survival may offset or even reverse the progressive redistribution embedded in pension rules. Several studies find that accounting for socioeconomic differences in mortality substantially reduces the progressivity of pension systems and, in some cases, generates regressive outcomes (Auerbach et al., 2017; Boado-Penas, Habermann and Naka, 2020; Breyer and Hupfeld, 2010; Goda, Shoven and Slavov, 2011; Nelissen, 1999; Olivera, 2019; Shi and Kolk, 2022). Evidence also indicates that mortality differentials are typically larger among men, reinforcing distributional concerns in male pension outcomes (Auerbach et al., 2017; Breyer and Hupfeld, 2010; Duggan, Gillingham and Greenlees, 2008; Goda, Shoven and Slavov, 2011; Koskinen and Martelin, 1994; Sánchez-Romero and Prskawetz, 2017).

Empirical evidence on socioeconomic mortality differentials is extensive in high-income countries. Using administrative records from the United States of America, Chetty et al. (2016) document a strong positive gradient between income and life expectancy, with widening gaps over time and larger disparities among men. Similarly, Duggan, Gillingham and Greenlees (2008) find substantial survival differences by lifetime income among Social Security beneficiaries. Evidence from European countries confirms comparable patterns: Gerdtham and Johannesson (2004) show a significant negative relationship between mortality risk and income in Sweden, while studies based on education gradients reveal persistent and sometimes widening longevity differences across socioeconomic groups (Bound et al., 2014; Olshansky et al., 2012).

A growing literature evaluates how these mortality differences affect pension redistribution. Auerbach et al. (2017) show that increasing differential mortality reduces the progressivity of the United States of America's Social Security system across cohorts. Boado-Penas, Habermann and Naka (2020) find that uniform annuity divisors in the Swedish notional account system favour highly educated individuals with longer life expectancy. Similar redistributive reversals are documented for Germany (Breyer and Hupfeld, 2010) and the Netherlands (Nelissen, 1999). Using simulations, Goda, Shoven and Slavov (2011) demonstrate that differential mortality reduces effective replacement-rate progressivity, particularly for men. At the European level, Olivera (2019) shows that mortality differences increase pension wealth inequality, although their relative importance varies across countries.

While evidence on mortality inequality is abundant for more developed economies, research for Latin America remains limited. Recent contributions have begun documenting differences in life expectancy by income around retirement age (Edwards, Soto and Zurita, 2020; Valderrama, 2024), yet little is known about how these mortality patterns interact with pension redistribution in the region.

In addition to mortality differences, inequality within pension systems is strongly influenced by heterogeneous contribution histories. In Latin America – and Uruguay in particular – lower-income workers experience higher informality, unemployment, and lower contribution densities throughout their working lives (Bucheli et al., 2005; Bucheli, Forteza and Rossi, 2006; Bucheli, Forteza and Rossi, 2010; Forteza et al., 2009; Forteza et al., 2011; Lagomarsino and Lanzilotta, 2004; Lavalleja, Rossi and Tenenbaum, 2019). Simulation-based studies for Uruguay show that incomplete contribution histories significantly shape access to benefits and system progressivity (Forteza and Ourens, 2012; Forteza and Mussio, 2012; Forteza and Rossi, 2013). However, these analyses generally abstract from socioeconomic differences in mortality.

This article contributes to the literature by analysing how differential mortality across socioeconomic groups affects inequality within the contributory social security system for men, measured through the distribution of the present value of pension benefits net of lifetime contributions. Focusing on Uruguay – a country combining European-like demographic dynamics with Latin American labour market characteristics – the analysis evaluates whether socioeconomic mortality gaps meaningfully alter redistribution outcomes within the pension system. By jointly considering mortality differences and contribution heterogeneity, the paper provides new evidence on the relative importance of longevity inequality compared with labour market inequalities in shaping within-system pension inequality.

To address this question, the article evaluates inequality within the social security system by analysing the present value of pension benefits net of lifetime contributions, following the approach of Auerbach et al. (2017), Forteza and Mussio (2012), and Goda, Shoven and Slavov (2011). Given that complete administrative work histories are not available in Uruguay, the analysis relies on simulated life-cycle employment trajectories for representative male workers across socioeconomic groups. The simulations incorporate income profiles, contribution histories, and mortality rates, and pension entitlements are computed under the rules introduced in October 2023. Inequality within the system is measured using the Gini RSV index of the present value of benefits net of contributions. The analysis compares a baseline scenario with differential mortality across socioeconomic groups to alternative scenarios that isolate the role of mortality differences and contribution patterns.

The results show that differential mortality across income groups has only a modest effect on inequality within the Uruguayan pension system. If mortality differences were eliminated, the Gini index of net pension benefits would decrease by 0.16 percentage points. In contrast, equalizing contribution patterns across socioeconomic groups would reduce inequality by 2.06 percentage points, highlighting the much larger role of labour market trajectories in shaping pension inequality.

The Uruguayan social security system

Uruguay operates a contributory multi-pillar pension system. The current institutional framework was originally established by the 1995 reform and subsequently modified by Law 20,130, approved in 2023. The reform introduced significant changes to eligibility conditions, retirement age, and benefit calculation rules in order to address demographic pressures and improve the long-term sustainability of the system. It also harmonized the rules across most pension subsystems, establishing a common framework for all workers, with some exceptions.

The system is financed through worker contributions (15 per cent of the contribution wage), employer contributions (7.5 per cent of the contribution wage), earmarked taxes, and general revenues. General revenues cover any deficit in the pay-as-you-go pillar and finance redistributive benefits such as the solidarity supplement. A ceiling on pensionable wages limits the maximum benefits generated in the pay-as-you-go pillar. In addition, a variable solidarity supplement increases the combined benefits received from both pillars for low-income retirees. As of January 2026, the solidarity supplement is calculated as 440 United States dollars (USD) minus 33 per cent of the full pension, ensuring that lower-income workers receive a higher rate of return on their contributions than higher-income workers.

The allocation of contributions between the two pillars depends on earning levels. Ten per cent of nominal wages up to approximately USD 3,500 (January 2026 values) finances the pay-as-you-go pillar, together with all employer contributions, earmarked taxes, and general revenues. The remaining worker contributions finance the individual capitalization pillar: 5 per cent of wages up to USD 3,500 and 15 per cent of wages between USD 3,500 and USD 6,800 are accumulated in individual retirement accounts managed by pension fund administrators. The accumulated balances and the returns generated finance retirement benefits in the funded pillar. There are no mandatory contributions for wages over USD 6,800. Upon retirement, these balances are typically converted into life annuities. Currently, life annuities are provided only by the State Insurance Bank (*Banco de Seguros del Estado* – BSE) using uniform mortality tables. Differential mortality across socioeconomic groups implies that individuals with higher life expectancy receive benefits for longer periods, which may generate redistribution across groups.

One of the central changes introduced by Law 20,130 is the gradual increase in the statutory retirement age from 60 to 65 years, depending on the year of birth, with some exceptions. Workers born in 1973 or later face a retirement age of 65. However, individuals with long contribution histories may retire earlier, at age 63 with at least 38 years of contributions or age 64 with at least 35 years of contributions.

The reform also modified the minimum contribution requirements. To qualify for a retirement pension at age 65, individuals must accumulate at least 30 years of contributions. The minimum retirement age increases as the contribution period decreases, reaching age 70 with at least 15 years of contributions. In addition, benefits from the individual capitalization pillar can be converted into a life annuity starting at age 65, regardless of the number of years of contributions accumulated.

Another important change concerns the benefit calculation formula. The replacement rate of the public pillar depends on both the number of years of contributions and the retirement age, thereby providing incentives for longer working lives. Workers who retire later or accumulate longer contribution histories receive higher replacement rates. Under the new rules, the replacement rate for a worker retiring at age 65 with 30 years of contributions is 45 per cent of the average wage calculated over the best 20 years of earnings.

Upon retirement, workers receive a combined benefit from both pillars. The public pillar provides a defined benefit financed through current contributions, while the capitalization pillar provides additional retirement income through an annuity or programmed withdrawals based on the accumulated balance in the individual account.

Uruguay maintains detailed administrative records of employment histories through the Social Insurance Bank (*Banco de Previsión Social* – BPS). These records have been systematically collected since the mid-1990s and allow the system to track contribution histories and calculate pension benefits with relatively high precision. The availability of these administrative data has facilitated empirical research on labour market trajectories, contribution densities, and pension outcomes.[1]

Methodology

Simulation strategy

This study simulates work-lives using observed administrative records and external information. The simulated lives represent male individuals from three socioeconomic groups defined by formal income levels observed in the administrative data. For each group, we simulate 1,000 individuals whose working lives are modeled according to the group-specific probabilities of contributing to the system, wage levels, and expected survival rates. Details on the adjustment of the simulations to the observed data are provided in the Appendix.

Individuals are assumed to retire as soon as they meet the eligibility criteria, that is, once they have accumulated the minimum required years of contributions and reached the statutory retirement age. Retirement benefits are calculated according to the rules in force since 2023. This framework allows us to compute the expected present value of pension benefits received, lifetime contributions paid,

1. Data is available upon request to the Social Insurance Bank (*Banco de Previsión Social* – BPS).

Table 1. *Characteristics of the scenarios*

Scenarios/Features	Base	Common mortality	Common contribution
Income	Specific to the group	Specific to the group	Specific to the group
Contribution	Specific to the group	Specific to the group	Common
Survival	Specific to the group	Common	Specific to the group

Source: Author's computations.

and the Gini RSV index, which together define the baseline scenario representing the current institutional setting.

In addition, we construct alternative scenarios in which key variables are either specific to each socioeconomic group or equalized across all simulated individuals, as summarized in Table 1. For instance, in the "Common Mortality" scenario, all individuals face the same survival rates. In the "Common Contribution" scenario, the probability of contributing is identical across groups. When equalizing attributes, we adopt the characteristics of the high-income group, which exhibits higher contribution probabilities and longer life expectancy.

For each scenario, we simulate individuals' contribution lives, expected contributions, pension entitlements, and the present value of benefits net of contributions, and we compute the Gini RSV index. Comparing the Gini index in the "Common Mortality" scenario with the "Baseline" scenario allows us to evaluate the effect of differential mortality on inequality within the system. Similarly, comparing the "Common Contribution" scenario with the "Baseline" scenario quantifies the impact of differences in contribution patterns.

Each methodological element is described in detail below.

Data source

The analysis uses administrative records from the BPS of Uruguay, comprising 352,197 male formal workers born between 1915 and 1986 who recorded at least one month of positive remuneration between April 1996 and April 2015. Monthly employment records are available from April 1996, when systematic recording of employment histories began, until April 2015.

The dataset includes personal characteristics: gender, date of birth, and date of death (if applicable); employment details: branch of activity, type of task performed, type of contract (monthly or day labourer), and remuneration; and company information: size of the company and branch of activity in which companies operate. We also have information on the timing of contributions and the taxable earnings.

Information is also available on benefits received during working life, including sickness, maternity, and unemployment benefits, as well as retirement information such as the retirement date, pension amount, and date of death when applicable.

Socioeconomic classification

Individuals are classified into three socioeconomic groups based on terciles of the formal income distribution within their birth cohort. The classification uses the individual's average monthly formal earnings over the entire observation period, adequately adjusted. Once assigned, individuals remain in the same socioeconomic group throughout the simulations.

Potential endogeneity between income and mortality may arise when income is measured over a short period (Auerbach et al., 2017), since adverse health shocks may simultaneously reduce earnings and life expectancy. To mitigate this concern, we measure income over a long observation window. Moreover, since the objective of the study is to evaluate lifetime inequality in the net pension benefits, this potential endogeneity has limited implications for the analysis.

Survival by socioeconomic group

In 2025, a process of social dialogue was initiated in Uruguay with the aim of discussing potential measures to improve the performance of the country's social security. One of the issues highlighted in these discussions was differential mortality by socioeconomic groups, measured through income levels. At present, complete mortality tables are not available. However, some point estimates of life expectancy exist for retirees and for the general population at specific ages. The available evidence indicates that differential mortality is present in Uruguay, that it declines with age, and that it is more pronounced for men than for women at any given age. It also shows that differences are larger between extreme income groups than between adjacent ones. For example, the difference in life expectancy at age 65 between men in the first and the tenth income deciles is estimated to be around six years. These gaps become smaller when income groups are defined more narrowly. This underscores the need to develop complete official mortality tables by socioeconomic group.

Although the administrative records contain information on deaths occurring during the observation period, these data do not allow us to estimate mortality rates by age, cohort, and socioeconomic group with sufficient precision.

Table 2. *The proportion of persons with each level of education by income group*

Educational level/Tercile	Tercile 1 (%)	Tercile 2 (%)	Tercile 3 (%)	Total (%)
Incomplete Primary	1.4	0.0	0.0	0.5
Completed Primary	16.8	2.2	2.3	7.5
Completed Lower Secondary	22.4	23.3	6.8	16.4
Completed Secondary School	44.0	40.4	17.8	30.1
Post-Secondary Education	15.4	34.1	73.1	45.5
Total	100.0	100.0	100.0	100.0

Notes: Post-secondary education does not necessarily imply university studies or completion of the highest level attained. There are only observations with formal education.

Source: Author's computations based on ECH 2015, INE.

Instead, we use survival projections by educational level for Uruguay, published by the Wittgenstein Centre for Demography and Global Human Capital (2018).[2] These projections provide survival probabilities for males by age and educational level for the period 2015 to 2020. The educational levels considered are no education, incomplete primary education, complete primary education, the first level of secondary education, complete secondary education, and post-secondary education (which does not necessarily imply tertiary education or completion of the highest level).

From these projections, we compute survival probabilities conditional on being alive at age 20. Survival probabilities for intermediate ages are obtained through interpolation.

To map educational categories into the income-based socioeconomic groups, we rely on data from the 2015 Continuous Household Survey (ECH) conducted by the National Institute of Statistics (2015). Male individuals aged 35 or older are classified into income terciles according to their formal earnings, and the distribution of educational attainment within each tercile is calculated. Table 2 reports the resulting correspondence between education levels and income groups.

Each simulated individual is then randomly assigned an educational level consistent with the distribution observed within their income group.

Simulation of work lives

We simulate 1,000 male individuals in each socioeconomic group, beginning at age 20. Individuals are assumed to enter the labour market in 2015 and initially start their careers without a formal contribution. For each month until retirement or until

2. See Wittgenstein Centre. *Wittgenstein Centre Data Explorer Version 2.0 (Beta)*, 2018.

reaching age 100, contribution status (whether they contribute or not) and earnings are simulated based on the BPS data.

Following Bucheli, Forteza and Rossi (2006), the probability of contributing in a given month depends on the individual's age and their contributory status in the previous month. These probabilities are estimated from observed frequencies within each socioeconomic group. We perform Monte Carlo simulations to generate the complete contribution history for each individual.

Formally, let $c_{t,e,z}^{i,g}$ denote the probability that individual i in socioeconomic group g contributes in month t at age e, conditional on the previous month's contribution status z, where $z = \{contribute, , did\ notcontribute\}$. For each month we generate a random draw from a uniform distribution in the interval $[0, 1]$, sim_t. An individual contributes in month t if: $sim_t \geq \left(1 - c_{t,e,z}^{i,g}\right)$; otherwise, they do not contribute.

Monthly formal wage data from the administrative records are deflated using the average wage index. Workers are classified by age and socioeconomic group, excluding observations with zero wages. Within each group, wage deciles are computed.

Each simulated individual i is randomly assigned a wage decile d within their socioeconomic group g as a permanent characteristic. Monthly wages are then randomly drawn conditional on age, socioeconomic group, and wage decile. If the individual does not contribute in a given month, their wage is zero.

Both wage trajectories and contribution densities differ across socioeconomic groups. However, conditional on belonging to a given group and decile, the wage path is simulated independently from the realization of contribution status in each period.

Individuals are assumed not to experience disability. They continue working until they become eligible for retirement or reach age 100. Once eligibility conditions are satisfied, pension benefits are calculated according to the 2023 regulations and the baseline assumptions.

The present value of net pension benefits and the Gini index

For each individual, we compute the present value of the expected pension benefits net of lifetime contributions:

$$PV = Benefits - Contributions =$$
$$\sum_{t=e}^{100} \frac{S_{e_o}(t)}{(1+\rho)^{t-e_0}} B_e(t) - \sum_{t=e_o}^{\min(e-1;\,100)} \frac{S_{e_o}(t)}{(1+\rho)^{t-e_o}} C(t)$$

Where $B_e(t)$ y $C(t)$ are the actual flows of benefits received and contributions paid monthly from the initial age e_0 to 100 years when the worker retires at the age e; ρ is the discount rate equal to the interest rate of the average pension funds since 1996, equivalent to 6 per cent real annual; and $S_{e_0}(t)$ is the probability of being alive at t given the individual was alive at e_0.

Individuals pay contributions from the start of their formal working life until retirement or until the age of 100, whichever comes first. Pension benefits are received from retirement age until age 100. Benefits include payments from both the public pillar and the individual capitalization pillar. Contributions include worker and employer contributions, assuming that employer contributions are ultimately borne by workers through lower wages (Hamermesh and Rees, 1993; Gruber, 1999; Brown, Coronado and Fullerton, 2009).

Since net benefits may be negative, particularly for individuals with lower survival probabilities, inequality is measured using the Gini RSV index, which allows for negative values (Ostosiewicz and Vernizzi, 2017; Raffinetti, Siletti and Vernizzi, 2015, 2017).

Sensibility analysis

First, we analyse an alternative scenario in which all individuals convert their individual account balances into an annuity at age 65, regardless of their number of years of contributions. Finally, given that approximately 85 per cent of individuals aged 65 or older receive some pension benefit in Uruguay (Rofman and Oliveri, 2012), we simulate an additional scenario with more lenient contribution requirements in which both survival probabilities and contribution densities are equalized across socioeconomic groups.

Results

Differences in access by socioeconomic group and setting

First, we analyse differences in pension access across scenarios and socioeconomic groups. In the simulations, individuals claim the common pension as soon as they become eligible and continue working until they qualify or reach age 100. Individuals who do not qualify for a common pension are those who fail to accumulate at least 15 years of contributions between the ages of 70 and 100. It is also important to note that, if individuals are allowed to claim an annuity from the capitalization pillar at age 65 regardless of their contribution history, no individual would remain without some form of pension benefit. However, in many cases, the resulting pensions would likely be very small. These assumptions help explain why the simulated figures differ from observed retirement outcomes, where roughly 20 per cent of the population does not qualify for a contributory pension.

The proportion of individuals who fail to retire varies markedly across socioeconomic groups. In the baseline scenario – where individuals retain their observed characteristics in terms of income, contribution patterns, and mortality – around 11 per cent of individuals in the lowest socioeconomic group fail to obtain any benefit, a situation that does not occur in the other socioeconomic groups. For approximately a tenth of individuals with lower socioeconomic status, participation in the social security system represents a net loss: they contribute for several years but ultimately fail to qualify for a pension benefit. By construction, this result disappears in the Common Contribution scenario.

Among individuals who meet the minimum eligibility requirements, the age at which they do so varies substantially across socioeconomic groups, as shown in Figure 1. About 85 per cent of individuals in the highest income group retire at age 65 or younger. Slightly less than 10 per cent meet the requirements at age 63, implying an accumulation of 38 years or more of contributions. More than 25 per cent qualify at age 64, having accumulated between 35 and 37 years of contribution. Among the intermediate-income group, only a small fraction can claim a pension before age 65. Around 70 per cent meet the requirements by age 66, and about 95 per cent do so by age 67. In both of the two highest-income groups, all individuals who qualify for retirement meet the requirements before age 70.

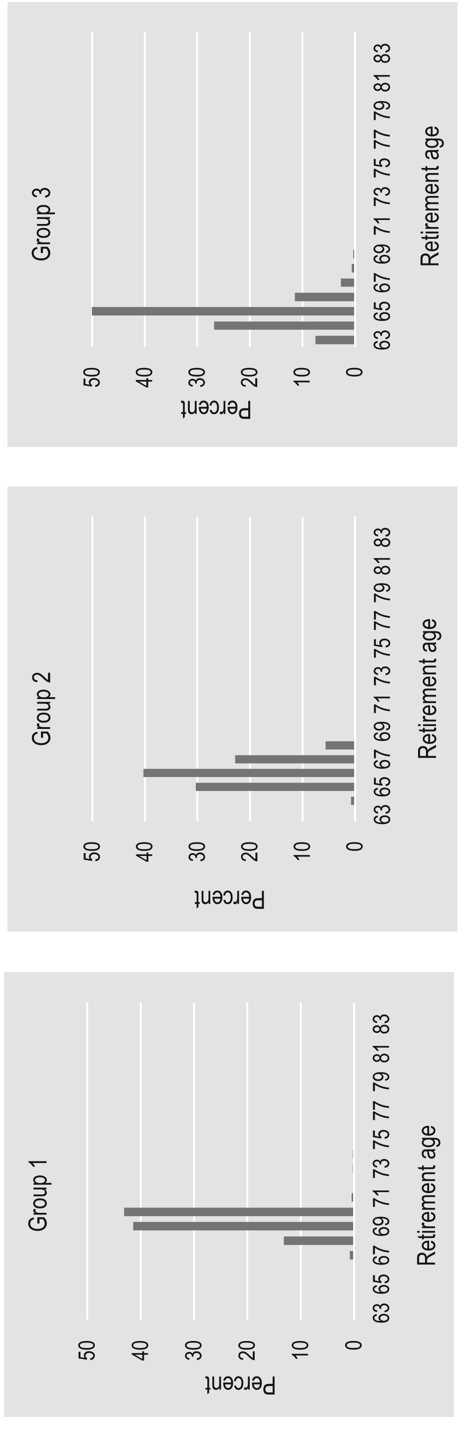
The lowest-income group shows substantially greater dispersion than the higher-income groups. No individuals in this group meet the requirements before age 66, and about 90 per cent do so between the ages of 68 and 70. Moreover, some individuals in the lowest income-group meet the minimum requirements at much older ages – up to age 84.

Regarding the alternative scenarios, there are no significant changes in access to pensions or in the retirement age under the common mortality scenario, which applies the survival rates of the highest-income group to all individuals. In contrast, in the equalized contributions scenario, where the probabilities of contributing at each age is set equal to the highest-income group, access to benefits becomes similar across all individuals.

Inequality in different scenarios

The inequality of benefits net of expected contributions does not substantially change between the baseline and common mortality scenarios. In the "Common Mortality" scenario, access to different benefits and retirement age does not

Figure 1. *Individuals meeting minimum requirements by age and socioeconomic group in the baseline scenario (%)*



Source: Author's computations based on simulated work lives.

change significantly. In this sense, the fall in inequality is only 0.16 percentage points in the point value (see Table 3). In this scenario, the low contribution density means these individuals cannot substantially benefit from improved longevity.

Concerning the "Common Contribution" scenario, we observe a drop in inequality of 2.06 percentage points. In this case, the ages at which individuals meet minimum requirements in the lower-income group vary substantially, and the percentage of individuals who do not manage to access benefits. In this sense, all the variables are very similar to those of the high-income group. Higher contributions allow individuals to retire earlier and access higher-quality benefits.

The analysis of these results shows the greater importance of individuals' contribution density in obtaining lower inequality. The crucial thing is to be able to access benefits more equally. In countries with a lower incidence of informality and minor differences in contributions by income group, differential mortality is likely to be more critical. In countries with more significant variability in ethnic, racial, or geographic conditions of formal workers, differential mortality could play a more relevant role.

Sensitivity analysis

The choice of purchasing an annuity at age 65, regardless of years of contribution

Suppose all individuals choose to purchase an annuity from their individual account balances at age 65, while maintaining the other assumptions of the base scenario. Under this assumption, the results are similar in terms of the ages at which retirement eligibility is achieved. However, the proportion of individuals in the lowest-income group who do not receive any benefit disappears completely. Even if they do not meet the minimum contribution requirements

Table 3. *GiniRSV index of present value of net benefits by scenario*

Scenarios	Gini RSV
Base	0.5011
Common mortality	0.4995
Common contribution	0.4805

Notes: Produced using RStudio's Gini_RSV package. The Gini correction made by Raffinetti, Siletti and Vernizzi (2015, 2017) admits negative values in the variable of interest.

Source: Author's computations.

for the general scheme, they can still access an annuity at age 65 through the individual capitalization pillar, although the benefit is likely to be small due to the relatively low accumulated contributions.

Nevertheless, individuals with lower incomes still tend to access meaningful benefits at later ages. It is important to note that the solidarity supplement only applies when individuals meet the eligibility requirements of the general scheme. As a result, the overall patterns of inequality remain broadly unchanged compared to the baseline analysis, although Gini RSV indices are slightly higher in all cases. The differences observed across the alternative scenarios also remain.

Lenient control of contribution requirements

Mechanisms that recognize years of contributions outside formal employment records could facilitate access to retirement benefits, particularly for individuals from lower-income groups (Colombo, 2013). However, there is insufficient information to explicitly model these mechanisms. To approximate their potential effects, we construct an additional scenario that combines more lenient contribution requirements with differential mortality.

In this scenario, survival rates and contribution probabilities are assumed to be equal across socioeconomic groups. Under these assumptions, the Gini RSV index decreases by 2.69 percentage points relative to the baseline scenario, and by 0.63 percentage points compared to the equalized contribution scenario, reinforcing the main finding.

Discussion and conclusions

This article examines how socioeconomic differences in mortality interact with labour market inequalities to shape redistribution within the contributory pension system in Uruguay. Using administrative records and simulated life-cycle employment trajectories for male workers across income groups, we estimate the present value of pension benefits net of lifetime contributions and measure inequality through the Gini RSV index. The analysis compares a baseline scenario reflecting observed heterogeneity in mortality and contribution patterns with counterfactual scenarios that isolate the effects of each factor.

The results show that differential mortality across socioeconomic groups has a limited effect on inequality within the Uruguayan pension system. Eliminating mortality differences across groups reduces the Gini RSV index by only 0.16 percentage points. In contrast, equalizing contribution patterns across socioeconomic groups leads to a substantially larger reduction in inequality, highlighting the dominant role of labour market trajectories in shaping pension outcomes. These findings indicate that the primary bottleneck in the system is not differential longevity but unequal access to pension eligibility conditions driven by differences in contribution lives.

The simulations also reveal important differences in pension access across socioeconomic groups. Individuals in higher-income groups are significantly more likely to meet minimum eligibility requirements and tend to qualify for retirement earlier. Lower-income workers, by contrast, accumulate contributions more slowly and are therefore more likely to reach retirement age without meeting the minimum requirements. In some cases, individuals in the lowest-income group contribute to the system for several years but fail to qualify for a pension benefit, implying negative net returns from participation in the contributory system. This pattern reinforces the central role of contribution density in determining both access to benefits and the timing of retirement.

These results contribute to the broader literature on socioeconomic mortality gaps and pension redistribution. While studies for high-income countries often find that differential mortality substantially reduces the progressivity of pension systems (Auerbach et al., 2017; Boado-Penas, Habermann and Naka, 2020; Breyer and Hupfeld, 2010; Goda, Shoven and Slavov, 2011), our findings suggest that this mechanism plays a smaller role in Uruguay. This difference likely reflects the institutional and labour market context of many Latin American countries, where incomplete contribution histories and labour market informality remain key determinants of pension eligibility. As a result, inequalities arising from labour market trajectories dominate the distributional outcomes of the pension system.

From a policy perspective, these findings suggest that addressing disparities in pension outcomes requires focusing primarily on improving access to benefits. Policies that increase contribution density among lower-income workers – such as measures to reduce informality, facilitate contribution continuity, or recognize contribution periods outside formal employment – are likely to have a great impact on pension inequality. Similarly, mechanisms that strengthen redistributive components of the system could mitigate the negative returns experienced by individuals who contribute but fail to meet eligibility thresholds.

The sensitivity analysis supports these conclusions. Allowing individuals to access annuities from the individual capitalization pillar at age 65 eliminates the share of individuals who receive no benefit but does not substantially alter the distribution of net pension wealth, since the resulting pensions are often very small. Likewise, simulations with more lenient contribution requirements reduce inequality but do not change the relative importance of contribution histories compared with mortality differences.

Despite these contributions, the analysis has several limitations. First, complete mortality tables by socioeconomic status are not available for Uruguay, requiring the use of survival projections by educational attainment as a proxy. While this approach is consistent with existing evidence on socioeconomic mortality gradients, more precise mortality estimates would improve future analyses. Second, the study relies on simulated employment trajectories because full lifetime administrative histories are unavailable. Although the simulation strategy uses detailed administrative data to reproduce observed contribution patterns, future work using longer administrative panels would allow more precise estimates of pension outcomes. Finally, the analysis focuses exclusively on male workers.

Future research could extend the analysis along several dimensions. Estimating mortality tables by income or education using linked administrative and demographic data would allow a more precise evaluation of longevity inequality within the pension system. Incorporating female labour market trajectories would provide a more comprehensive assessment of pension redistribution. Finally, comparative studies across Latin American countries could help determine whether the relative importance of mortality and contribution patterns observed in Uruguay generalizes to other institutional contexts.

In sum, the evidence presented in this article suggests that in Uruguay, the main source of inequality within the contributory pension system arises from unequal labour market trajectories across socioeconomic groups. Policies aimed at improving contribution continuity and facilitating access to pension benefits among lower-income workers are therefore likely to be the most effective in reducing pension inequality.

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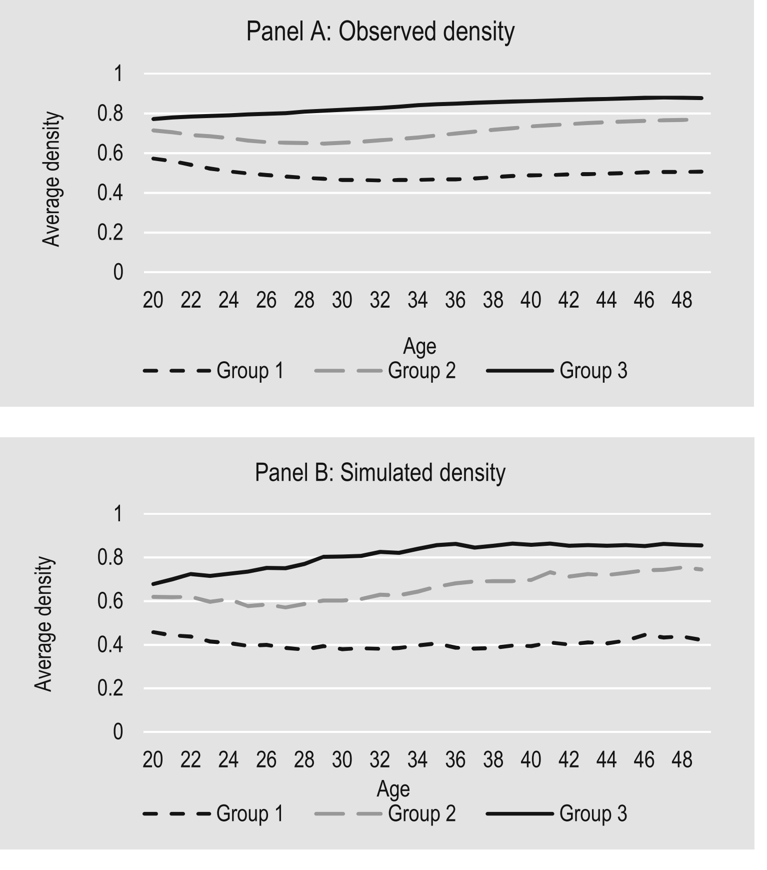
Appendix:
Adjustment of simulations and external data

This appendix presents the adjustment of the simulations to the data.

First, we compare observed and simulated contribution densities and income.

Figure A.1. shows the average contribution densities by age and socioeconomic group observed (Panel A) and simulated (Panel B). The patterns are similar, although the simulated densities are lower at initial ages and for the lower income group.

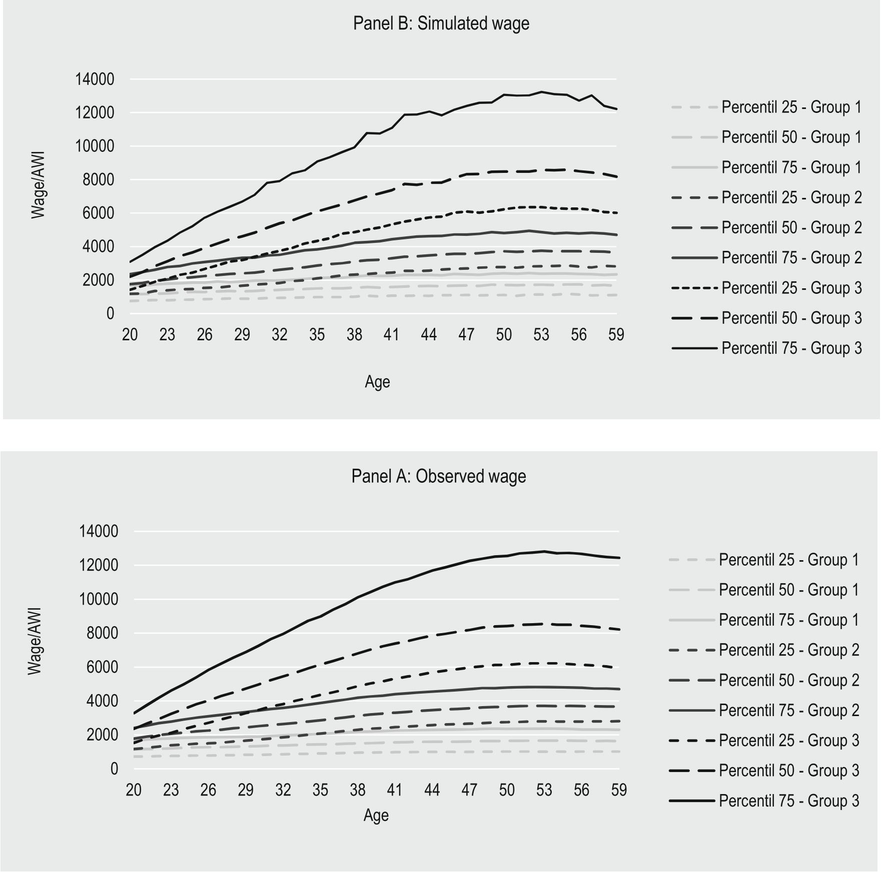
Figure A.1. *Observed and simulated contribution density by socioeconomic group and age*



Source: Author's computations.

Figure A.2 shows the ratio of remuneration to the average wage index by age and income group. The observed (Panel A) and simulated (Panel B) age profiles are very similar.

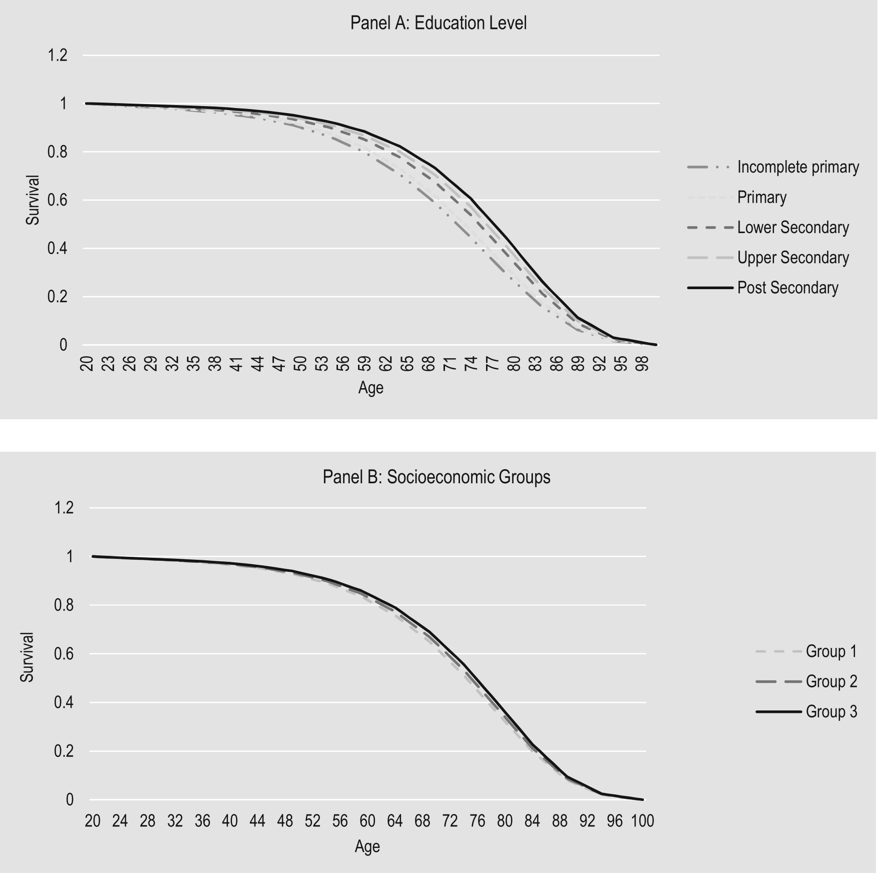
Figure A.2. *Ratio of remuneration to AWI by age and socioeconomic group (selected percentiles)*



Source: Author's computations.

Figure A.3 shows survival by age for different educational levels (Panel A) and by socioeconomic groups (Panel B), conditional on being alive at 20 years of age.

Figure A.3. *Survival by age, education level and socioeconomic group for males alive at 20*



Source: Author's computations.

A more marked difference is observed in survival by educational levels since the socioeconomic strata group individuals of different educational levels, and we compute survival rates for formal workers; perhaps the dispersion is more significant if we consider the whole population. There are no individuals with "No education" among formal workers. These data are valuable as a first approximation but highlight the need to delve into survival data for socioeconomic groups based on income levels.

117

The targeting efficiency of rural social assistance programmes and income inequality: Evidence from the 2019–2021 Guangdong Thousand-Village Survey

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Abstract Utilizing data from the 2019–2021 Guangdong Thousand-Village Survey, this study analyses the targeting efficiency of the contemporary social assistance programmes and empirically investigates their effects on income inequality among rural households. We show that social assistance helps alleviate the relative income deprivation of impoverished rural households. However, many have no access to social assistance, with the under-coverage rate being approximately 80 percent. The inclusion error of social assistance further exacerbates income inequality among rural households. Therefore, it is important to improve coordination among different rural social assistance programmes, pay more attention to the invisible poor, and ensure that all eligible households share assistance.

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Keywords benefit administration, China, coverage, social assistance, standard of living, poverty

Introduction

In 2020, China achieved the milestone of eradicating rural poverty by lifting all 98.99 million impoverished rural residents out of poverty. As a part of the basic social security system, social assistance played a crucial role in meeting the needs of vulnerable groups. To build a classified and tiered social assistance system, it is crucial to have a deep understanding of the impact of social assistance on increased earnings and income inequality.

Given the diversity of social assistance needs, it is important to establish a dynamic tracking system for recipients. In many countries, social assistance programmes are designed for the poorest groups, while in others the focus lies on unemployed migrants, children, women, or the homeless. Policy makers employ a range of methods to identify target groups, including proxy means testing, community-based targeting, hybrid targeting (Follett and Henderson, 2023), and geographic microtargeting (Smythe and Blumenstock, 2022). However, previous studies have documented prevalent coverage errors across various types of social assistance programmes (Li and Walker, 2017; Hümbelin, 2019). These include both in-kind assistance programmes, such as food subsidies and cash transfer mechanisms that are part of China's means-tested Minimum Living Standard Guarantee (Dibao) programme (Li and Walker, 2018).[1] Similar implementation challenges have been observed in housing welfare schemes (Blundell, Fry and Walker, 1988) and service-oriented interventions, notably Kosovo's hybrid targeting approach for child poverty alleviation (Roelen and Gassmann, 2011). Therefore, it is worthwhile to explore the causes of coverage errors in social assistance.

Research indicates that effective social assistance programmes can mitigate long-term societal costs by preventing poverty traps and associated negative externalities (Niño-Zarazúa, 2019). Some studies have found that, even with coverage errors, poverty-oriented assistance programmes are still more beneficial than universal programmes (Rema and Olken, 2018; Grosh and Leite, 2009). While social assistance programmes are designed to reduce income inequality, coverage errors in their implementation risk creating a double burden: not only do such errors lead to inefficient allocation of public resources, but they may

1. In English, also called the Minimum Livelihood Guarantee or Minimum Living Security, Dibao is considered the world's largest cash-based social assistance programme.

inadvertently perpetuate the very disparities these policies aim to address. Coverage errors simultaneously create dual inefficiencies across different contexts. In Russia, where a large portion of assistance benefits fails to reach intended recipients (Andreeva, Bychkov and Feoktistova, 2021), misallocated resources deprive other eligible groups of needed support.

Social assistance programmes can help stimulate consumption among low-income households (Deaton, 1992), but their effectiveness is often constrained by implementation challenges. Using Panel Study of Income Dynamics (PSID) data from 1978 to 1992, Blundell and Pistaferri (2003) found that while the United States' food assistance programme increased household consumption, it fell short in providing adequate protection against financial shocks. Similar patterns are observed in medical and educational assistance programmes (Gertler and Gruber, 2002). While governments provide targeted assistance, benefit amounts often prove insufficient during earnings fluctuations. Furthermore, programmes requiring meeting employment conditions may trap recipients in unstable jobs, offsetting benefits (Sod-Erdene et al., 2019), thereby discouraging some eligible families from applying.

From a theoretical perspective, relative poverty emphasizes the structural position within income distribution, while absolute poverty focuses on the sufficiency of basic survival needs. However, in China, the two concepts are intrinsically linked. China successfully eradicated extreme poverty by 2020, marking historic improvements in living standards. Yet income gaps and uneven public service distribution persist, which may still trigger a "sense of relative deprivation" among some families through social comparison, thereby affecting their welfare perceptions and social stability. Meanwhile mis-allocation of assistance (such as allocating resources to ineligible households) may affect relative poverty through two pathways: one directly alters the income distribution, while the other influences subjective deprivation through social comparison psychology. Both pathways differ from the reduction of absolute poverty but share equally significant policy implications. The 7th National Population Census data reveals that over 500 million rural residents, including nearly 300 million migrant workers in cities, still face relatively low incomes. Sustained efforts must prioritize safeguarding vulnerable groups' basic needs to boost rural earnings and domestic consumption. Accordingly, the aim of this article is to better illuminate the mechanism of the impact of social assistance on relative poverty, broadening the dimensions of policy evaluation and providing evidence for constructing a multidimensional poverty governance system in the "post-poverty alleviation era".

Guangdong, a benchmark region for urban-rural development gaps, is commonly referred to as "where the wealthiest and poorest areas in China coexist", highlighting the urgency of addressing regional income inequality. This makes it an optimal microcosm for studying the governance of relative poverty in contemporary China. Accordingly, this study employs relative deprivation theory with 2019–2021 Guangdong Thousand-Village Survey data[2] to address two issues: i) What is the targeting efficiency of rural social assistance programmes, and what forms of coverage errors manifest in its implementation? ii) Does social assistance provision effectively enhance household income and reduce income inequality among the rural poor?

Theoretical framework and hypotheses

Theoretical framework

Social comparison theory posits that individuals derive self-worth through perceived social status (Festinger, 1954; Tajfel, 1970). This process drives systematic comparisons between in-group and out-group positioning to negotiate status mobility. Relative deprivation emerges when such comparisons reveal intra-group disparities, particularly when disadvantaged members perceive deprivation of entitlements. Empirical evidence confirms that these perceptions induce psychological distress and constrain opportunity structures (Mummendey et al., 1999). Coverage errors in social assistance systems directly shape individuals' social status perceptions. When groups lack critical resources, members experience heightened relative deprivation – a psychological state that exacerbates intra-group inequality (Yitzhaki, 1979). This deprivation further correlates with systemic risks including violent crime, public health deterioration, and social instability. Methodologically, we adopt Kakwani's (1977) generalized Gini coefficient framework, which quantifies relative deprivation through expected deprivation values. This approach builds on Yitzhaki's foundational linkage between deprivation theory and inequality metrics, enabling precise measurement of rural households' deprivation levels.

Social assistance is a redistribution mechanism that benefits vulnerable groups (Fiszbein, Kanbur and Yemtsov, 2014; Gough et al., 1997). Many empirical studies have explored the relationship between social assistance and welfare outcomes, poverty, and social inequality in developing countries (Baird et al., 2013; Bastagli et al., 2019).

Social assistance programmes have demonstrated measurable impacts across five key dimensions: poverty alleviation (Angelucci and Attanasio, 2009; Gertler, Martinez and Rubio-Codina, 2012; Skoufias and Di Maro, 2008; Skoufias, Unar and Gonzalez de Cossio, 2013), educational attainment (Baird, McIntosh and Özler, 2011; Barrera-Osorio et al., 2011; Benhassine et al., 2015; Filmer and

2. The dataset can be requested from the website of Survey Data Center, Jinan University.

Schady, 2011; Macours, Schady and Vakis, 2012), health enhancement (Attanasio, Oppedisano and Vera-Hernández, 2015; Barham and Maluccio, 2009; Behrman and Parker, 2013; Fernald and Hidrobo, 2011; Fernald, Gertler and Neufeld, 2008; Miller, Tsoka and Reichert, 2011; Ramírez-Silva et al., 2013), asset accumulation (Covarrubias, Davis and Winters, 2012; Maluccio, 2020; Masino and Niño-Zarazúa, 2020; Todd, Winters and Hertz, 2020), and labour market participation (Alzúa, Cruces and Ripani, 2013; Ardington, Case and Hosegood, 2009; Asfaw et al., 2014; Attanasio et al., 2010; Barrientos and Villa, 2015).

While social assistance programmes can reduce poverty (Stewart, Huerta and Sader, 2009), previous studies reveal critical limitations in their long-term effectiveness. Studies using supplemental poverty measures show that cash and food assistance temporarily lower child poverty rates (Pac et al., 2017). However, three structural issues weaken these benefits: i) programmes often miss eligible families due to targeting errors, ii) benefit amounts are too low to make a meaningful difference, and iii) strict eligibility rules exclude vulnerable groups (Valencia Lomelí, 2008; Teichman, 2008; see also Riphahn, 2001). This is particularly evident in emerging economies, where child-focused programmes frequently fail to lift families out of poverty sustainably. Short-term assistance may unintentionally create dependency. In China, Dibao recipients reported lower income expectations and life satisfaction, suggesting temporary assistance harms long-term motivation (Gao, Zhai and Garfinkel, 2010). Similarly, time-limited cash transfers often prevent people from gaining stable jobs or reintegrating into society (Oosterhoff and Yunus, 2022; Bargain and Donni, 2012). These findings highlight a key dilemma: while social assistance meets urgent basic needs, it rarely addresses the root causes of poverty. Effective reforms must combine immediate support with skills training and job opportunities.

Social assistance serves as a vital tool for promoting common prosperity, particularly in addressing persistent rural poverty, particularly given China's persistent rural poverty challenges. As the foundational social security mechanism, it serves as a critical instrument for income redistribution. Previous studies confirm its positive effects on improving living standards for low-income households (Weitoft et al., 2008). However, comparative studies indicate nuanced outcomes: China's new rural pension system demonstrates stronger poverty reduction effects for elderly rural residents compared to urban Dibao programmes (Li, 2018). Recent years have witnessed declining enrollment rates in both urban and rural Dibao systems. A notable challenge emerges in households with incomes slightly above eligibility thresholds – many still experience poverty due to high medical and education costs, so-called "expenditure-driven poverty". Current assistance programmes often fail to address these families' specific needs, reducing their anti-poverty effectiveness.

Therefore, it is essential to accurately identify the needs of different groups and implement a variety of social assistance programmes, such as medical and educational support, to enhance the effectiveness of social assistance.

Hypotheses

On the basis of the above, a first hypothesis is proposed.

H1. Social assistance programmes play a positive role in reducing income inequality among rural households.

In China, the rural Dibao programme has improved the living standards of many rural impoverished households, playing a crucial role in income redistribution. Utilizing data from the National Bureau of Statistics between 2007 and 2018, several studies have measured the Gini coefficient to assess the programme's impact on income disparities. The results show that the rural Dibao has contributed to about a 4 per cent reduction in income inequality – an improvement, albeit still below international benchmarks (Pu and Xu, 2021). Nevertheless, it has proven effective in narrowing rural income gaps. However, several challenges persist, including low benefit levels, insufficient fiscal investment, and pronounced regional economic disparities. Compared with social insurance schemes, social assistance programmes have shown a greater capacity to moderate income inequality. Yet, due to exclusion errors in rural social assistance, the programme's reach remains limited. For instance, in Guangdong province, fewer than 5 percent of rural residents received Dibao benefits, and the programme's coverage rate in 2012 was lower than the national average (Xie and Yang, 2015).

Given the above analysis, the following hypothesis is proposed.

H2. Inadequate coverage and exclusion errors in social assistance programmes have played a significant role in deepening income inequality in rural areas.

Data source and method

Data sources

Our analysis draws on the Guangdong Thousand-Village Survey (2019–2021), an ongoing provincial panel study initiated and conducted by Jinan University since 2018. The survey selected samples from 30 counties and 119 administrative villages using the probability proportional to size (PPS) sampling method. The questionnaire covered a range of topics, including land transfers, poverty alleviation, social assistance, environmental sanitation, and rural governance. Given that the 2018 baseline survey focused on collecting foundational information about regions and respondents, its questionnaire content was significantly different from subsequent tracking surveys. While the survey design from 2019 to 2021 remained consistent and included more detailed assessments of social assistance, the analysis in this article utilizes panel data from 2019 to 2021, with a total sample size of 8,728 households.

Variable selection and descriptive statistics

Variable selection.

• Explained variables. Income inequality is commonly assessed using measures such as the Gini coefficient (or Gini index), the Lorenz curve, the Theil index, and the Atkinson index. In this study, we adopt the Kakwani index to capture the degree of relative deprivation among low-income rural households and conduct regression analysis based on this measure.Traditional inequality metrics such as the Gini coefficient primarily measure inter-group income disparities but fail to account for intra-group individual positioning. To address this limitation, scholars introduced relative deprivation theory – the psychological state resulting from when individuals lack resources possessed by their peers (Runciman, 1961). Following this framework, Yitzhaki (1979) mathematically linked relative deprivation to Gini coefficient calculations. Kakwani (1984) advanced this framework through stochastic modeling. The model assumes that individual income (y) and the average income of a group Y (μ) are continuous random variables with a probability density function. A group Y consists of n samples, their income is y_i respectively. Furthermore, Kakwani introduces the index of relative deprivation for individual households to indicate the extent of income inequality. The formula is as follows:

$$RD(y, yi) = 1/n\mu_Y\left[\sum_{j=i+1}^{n}(yi - yi)\right] = Y_i^+\left[\left(\mu_{y_i}^+ - y_i\right)/\mu_Y\right] \qquad (1)$$

As shown in formula (1), μ_Y represents the average income of all surveyed households in the group. μ_{y_i} refers to the average income of households whose income exceeds y_i in group Y. Y_{i}^+ represents the percentage of households whose income exceeds y_i in group Y. The index of relative deprivation in income ranges from 0 to 1. We use this method to compare an individual's income to those with higher income levels. This comparison allows us to measure the relative deprivation in income for individual households, with a range between 0 and 1.

• Focal variables. The focal variables include social assistance programmes, such as the Dibao, medical assistance, and educational assistance, all of which are dummy variables. Households that have previously obtained support from any assistance programmes are coded as 1, while others are coded as 0.

• Control variables. At the household level, the variables include age (continuous variable), gender (dummy variable, male = 1), education level of household heads (continuous variable), household head engages in agricultural activities (dummy variable) and family size (continuous variable). At the village level, the variables include the number of registered residents, the distance from villages and towns, population served by per health worker and the number of village primary schools.

Descriptive statistics of variables.

• Basic characteristics of rural households. As shown in Table 1, the proportion of households obtaining social assistance was 13.27 percent, 15.14 percent, and 14.28 percent over the three years, with an average of 14.23 percent. Among these, Dibao recipients consistently formed the largest group, accounting for 7.23 percent, 6.96 percent, and 9.12 percent of households annually. This was followed by those receiving educational support (5.24 percent, 7.10 percent, 5.85 percent), while medical assistance programmes served the fewest households (4.86 percent, 5.24 percent, 3.54 percent). Household characteristics revealed stable patterns: Household heads were predominantly male (age 60+ on average). Educational attainment typically ranged between primary and junior high school levels. Average family size grew steadily from 4.80 members (2019) to 5.24 members (2021). In terms of public service provision, the number of residents served per medical staff at village health clinics has gradually increased, while the number of village primary schools has remained stable to ensure children of compulsory education age have access to local schooling.

• Current income status of rural households. According to the *2021 China Rural Statistical Yearbook* (National Bureau of Statistics, 2022) the annual per capita disposable income of rural households reached 11,421.7 yuan (CNY) in 2015, rising to CNY 17,131.50 in 2020 and CNY 18,930.90 in 2021. Over the same period, net transfer income grew from CNY 2,066.30 to CNY 3,937.20, with its share of disposable income increasing from 18.1 percent to 20.8 percent. Overall, both per capita disposable income and net transfer income have shown steady growth. As shown in Table 2, Guangdong presents a paradox: while its 2021 rural disposable income (CNY 22,306) ranked seventh nationally – surpassing the national average – it exhibits severe income inequality. The province's net transfer income (CNY 3,306.8) fell below the national mean (CNY 3,937.20), placing it twentieth among provinces. Strikingly, households receiving social assistance in Guangdong's 119 surveyed villages reported disposable incomes of

Table 1. *Basic characteristics of the variables*

Variables	2019 Mean difference	2019 Standard error	2020 Mean difference	2020 Standard error	2021 Mean difference	2021 Standard error
The Index of Relative Deprivation in income	0.5663	0.2854	0.5776	0.2776	0.5599	0.2747
Sharing Social Assistance	0.1327	0.3393	0.1514	0.3585	0.1428	0.3499
Sharing Dibao	0.0723	0.2590	0.0696	0.2545	0.0912	0.2879
Sharing Medical Assistance	0.0486	0.2150	0.0524	0.2228	0.0354	0.1849
Sharing Educational Assistance	0.0524	0.2229	0.0710	0.2568	0.0585	0.2347
Age of household head	59.1093	12.1054	60.1332	12.0635	60.9064	12.2097
Gender of household head	0.8613	0.3457	0.8584	0.3487	0.8613	0.3457
Educational level of household head	1.6564	0.7498	1.6112	0.7595	1.5911	0.8027
Household head engages in agricultural activities	0.5636	0.4960	0.3639	0.4812	0.1886	0.3913
Family size	4.7960	2.4245	4.9865	2.5171	5.2412	2.6535
Registered population of the village	4,468.73	3,301.98	4,476.10	3,310.08	4,466.64	3,349.74
Distance from villages and towns	5.8540	3.9288	5.9068	4.1463	5.7858	3.9077
Population served by per health worker	2,397.9160	1,640.6250	3,058.6630	2,774.9130	3,097.7180	2,451.1960
Number of village primary schools	1.2413	0.5325	1.2688	0.5640	1.2250	0.5330
Sample size	2863		2959		2906	

Source: Authors' elaboration.

Table 2. *Comparison of disposable income of rural residents in Guangdong and national average in 2021 (unit: Yuan, %)*

Region	Per capita disposable income	Ranking	Per capita net transfer income (Yuan per person)	Ranking	Proportion of per capita net transfer income in disposable income	Ranking
National	18930.9		3937.2		20.8	
Guangdong	22306.0	7	3306.8	20	14.8	27

Source: Authors' elaboration based National Bureau of Statistics (2022).

Table 3. *Per capita disposable income and income sources (unit: Yuan per person)*

Variables	2019 Mean	2019 Standard Deviation	2020 Mean	2020 Standard Deviation	2021 Mean	2021 Standard Deviation
Wage earnings	7,697.12	9,169.00	8,495.04	13,544.66	3,574.89	6,777.54
Net operating income	2,703.95	13,067.68	3,107.62	34,718.19	1,575.88	5,184.23
Net property income	416.19	2,401.15	403.85	2,007.97	568.46	3,277.47
Net transfer income before sharing assistance	1,180.25	5,941.73	948.86	4,014.29	2,766.28	6,173.59
Social assistance funds	203.68	820.74	243.59	836.83	261.27	909.35
Net transfer income after sharing assistance	1,392.09	6,040.65	1,198.87	4,154.21	3,004.27	6,154.52
Total income before sharing assistance	11,038.90	17,247.68	11,436.70	16,990.93	8,574.95	11,405.88
Total income after sharing assistance	11,310.75	17,213.28	11,695.25	16,969.77	9,055.47	13,278.18
The Index of Relative Deprivation in income before sharing assistance	0.5669	0.2852	0.5768	0.2773	0.5489	0.3922
The Index of Relative Deprivation in income after sharing assistance	0.5663	0.2854	0.5776	0.2776	0.5599	0.2747
Sample size	2863		2959		2906	

Notes: (1) The data from the 2019–2021 Guangdong Thousand-Village Survey shows the disposable income and its sources over the past year, with the income calculated based on the prices of the respective year. (2) The calculation formula is as follows: Household disposable income = salary + net operating income + net property income + net transfer income; Per capita household disposable income = household disposable income/ household population. (3) In terms of net transfer income, benefits such as pensions, retirement benefits, social assistance, donations, and support payments from relatives – provided in the form of goods, vouchers, or other non-cash items – have already been valued as cash equivalents and calculated in the survey. (4) The per capita disposable income excludes various transfer expenditures such as social security payments, gift expenditures, support expenditures, donations, and compensation payments.
Source: Authors' elaboration, outlining rural household income structures.

only CNY 9,055.47 (2021) – merely 40.6 percent of the provincial average. Table 3 outlines rural household income structures from 2019 to 2020, highlighting four main components: wage earnings, net operating income,[3] net property income, and net transfer income. Wages dominated household finances throughout this period, averaging CNY 7,697.12 per capita (2019) and CNY 8,495.04 (2020) – consistently representing over two-thirds of post-assistance income. Net operating income followed as the second-largest source at CNY 2,703.95 to CNY 3,107.62 annually, constituting about 25 percent of total income after

3. A measure that calculates total income generated by a business minus all necessary operating expenses.

assistance redistribution. Pre-assistance transfer payments showed notable fluctuation, decreasing from CNY 1,180.25 (2019) to CNY 948.86 (2020), while government social assistance funds gradually increased from CNY 203.68 to CNY 243.59. Property returns remained the smallest contributor at CNY 416.19 and CNY 403.85 respectively. In 2021, due to external economic factors, the total household income declined. Both salary and net operating income decreased, while net property income, net transfer income, and social assistance funds saw increases. Specifically, net property income rose to CNY 568.46 per person, net transfer income before assistance climbed to CNY 2,766.28 per person, and social assistance funds increased to CNY 261.27 per person.

- The index of relative deprivation in current income. Table 3 also presents the income inequality among rural households. In 2019, social assistance demonstrated measurable success in reducing income disparities, with the relative deprivation index decreasing marginally from 0.5669 to 0.5663 post-intervention. However, this trend reversed in subsequent years: the index increased by 0.0008 (2020) and 0.0110 (2021) after benefit distribution. Notably, this trend diverges significantly from Gini coefficient measurements, highlighting the importance of incorporating subjective perceptions of equity when evaluating social assistance programme effectiveness.

Estimation strategy

The dependent variable in this article is the relative poverty index, strictly bounded between 0 and 1, exhibiting clear left truncation (0) and right truncation (1). Using ordinary least squares (OLS) estimation could result in predicted values exceeding the bounded range and potentially inconsistent coefficient estimates. In contrast, the Tobit model, based on maximum likelihood estimation, is better suited to handle such truncated data. Therefore, this study employs a panel Tobit model to examine the impact of social assistance programmes – such as Dibao, medical assistance, and educational assistance – on reducing income inequality among rural households.

$$RD_{ijt} = \beta_0 + \beta_1 assistance_{\mathrm{it}} + \beta_2 dibao_{\mathrm{it}} + \beta_3 medical_{\mathrm{it}} + \beta_4 education_{\mathrm{it}} + \beta_6 Z_{it} + \varepsilon_{\mathrm{i}} + \gamma_{\mathrm{t}} + \theta_{\mathrm{it}} \quad (2)$$

In Equation (2), RD_{ijt} represents the index of relative deprivation of household i in year t, reflecting the extent of income inequality across households. Wherein, $assistance_{\mathrm{it}}$ is a binary variable indicating whether household i in year t is eligible for sharing social assistance; $dibao_{\mathrm{it}}$. $medical_{\mathrm{it}}$, and $education_{\mathrm{it}}$ represent whether the i is eligible for sharing Dibao, medical assistance, and educational assistance in year t, respectively. Z_{it} represents a set of control variables of the rural household i in year t at the individual, household, and village levels. ε_{i} captures individual fixed effects, while γ_t accounts for time fixed effects. Finally, θ_{it} is the error term that varies across individuals and time.

The targeting efficiency of rural social assistance

Quantitative indicators and eligibility criteria for social assistance

Precise identification of social assistance recipients is a critical step in measuring targeting accuracy. Song, Li and Wang (2020) assessed the targeting efficiency of Dibao using both income-based criteria and multidimensional criteria encompassing household income and assets. When Yan, Xue and Feng (2023) calculated the targeting efficiency of Dibao using the 2021 Guangdong Thousand-Villages Survey, they established measurement standards based on the officially formulated *Guidelines for Verifying the Economic Status and Assessing the Living Conditions of Families Eligible for Dibao in Guangdong Province*. This approach provided a relatively objective representation of the programme's targeting precision. This article references their methodology for identifying assistance recipients. Based on the *Guidelines for Verifying the Economic Status and Assessing the Living Conditions of Families Eligible for Dibao in Guangdong* and the *Measures for Assistance to Families on the Margin of Minimum Living Subsidy and Families Experiencing Expenditure-Based Hardship in Guangdong Province*, we have compiled quantitative indicators and eligibility criteria for Dibao, medical assistance, and educational assistance programmes (as shown in Table 4). Due to the distinct evaluation criteria of various assistance programmes, Dibao targets long-term income poverty, medical assistance addresses expenditure-based poverty caused by major illnesses, and educational assistance focuses on families with school-age children. Using these criteria, the sample is divided into two groups: i) households eligible for assistance and ii) ineligible households.

Exclusion error

As shown in Table 5, the targeted inclusion rate of rural social assistance steadily increased from 20.67 percent in 2019 to 23.40 percent in 2021. Consequently, the under-coverage rate (i.e. those missed or excluded) declined from 79.33 percent to 76.60 percent. Specifically, the targeting efficiency of the Dibao programme followed a similar upward trend. On one hand, the targeted inclusion rate rose from 14.14 percent in 2019 to 17.51 percent in 2021. On the other hand, the under-coverage rate decreased from 85.86 percent to 82.49 percent. For expenditure-based special assistance, targeting efficiency also improved between 2019 and 2021, increasing from 13.65 percent to

Table 4. *Quantified targets and identification criteria for social assistance programmes*

Difficulty type	Indicator	Sub-indicator	Definition and other avoidance rules
Dibao	Family income	Including wage income, net operating income, net property income and net transfer income	
	Family structure	Number of minors and school-age children	(1) Preschool children (2) Students in senior high school, technical secondary school
		Number of elderly and advanced elderly individuals	(1) Elderly aged 70–79 (2) Elderly aged 80+
		Number of persons with disability, severe illness, and loss of capacity	(1) Severely disabled persons, persons with loss of capacity (2) Moderately disabled persons (3) Patients with severe illness, chronic illness, and persons with partial loss of capacity
		Number of persons obligated to provide support, maintenance, or care	(1) Persons aged 18–59 with stable employment income (2) Persons aged 18–59 while engaged in casual work, simple agricultural production, or household labour services (3) the number of elderly, single-parent parents/step-parents, foster carers with maintenance obligations
	Fixed assets and household assets	(1) No self-owned housing (renting or borrowing) (2) Number of fuel-powered motorcycles, electric vehicles (3) Number of air conditioners (4) Number of television sets (5) Number of refrigerators	
	Total	Per capita family income after comprehensive assessment = (Total monthly family income + family structure + relatively fixed assets) /Number of cohabiting family members	
Expenditure-based difficulty	Income	Annual per capita family income is lower than the per capita disposable income of urban (rural) residents in the county (city, district) where the household is registered in the previous year	
	Expenditure	Where such as medical care, education, rehabilitation of disabled family members, or expenses due to disasters or accidents incurred by cohabiting family members reach or exceed 60% of the family's disposable income for the year	
Other conditions to be met for both Assistance programmes	(1)The applicant or a cohabiting family member has local household registration (3)The per capita market value of financial assets is lower than 24 times the local monthly minimum living standard (3) No more than 1 set (building) of commercial housing is owned, and no non-residential real estate for other purposes (including shops, parking spaces, etc.) is owned (4) No motor vehicles (excluding motor vehicles for the disabled, fuel-powered motorcycles, and electric vehicles) (5) No registered business information		

Source: Authors' elaboration, and drawing upon the methodology of Yan, Xue and Feng (2023).

Table 5. *The targeting efficiency of social assistance programmes and the Dibao programme*

	Targeting efficiency of social assistance programmes			Targeting efficiency of Dibao		
	2019	2020	2021	2019	2020	2021
Targeted inclusion rate	20.67%	23.38%	23.40%	14.14%	15.53%	17.51%
Targeted exclusion rate	89.49%	87.69%	88.64%	94.98%	95.64%	93.30%
Under-coverage rate	79.33%	76.62%	76.60%	85.86%	84.47%	82.49%
Leakage rate	10.51%	12.31%	11.36%	5.02%	4.36%	6.70%

Source: Authors' elaboration.

Table 6. *The targeting efficiency of expenditure-based special assistance*

	Targeting efficiency of expenditure-based special assistance			Targeting efficiency of Medical Assistance			Targeting efficiency of Educational Assistance		
	2019	2020	2021	2019	2020	2021	2019	2020	2021
Targeted inclusion rate	13.65%	14.48%	16.08%	9.16%	8.61%	8.82%	10.87%	15.00%	18.75%
Targeted exclusion rate	92.18%	89.48%	92.30%	95.58%	95.06%	96.85%	94.85%	93.12%	94.52%
Under-coverage rate	86.35%	85.52%	83.92%	90.84%	91.39%	91.18%	89.13%	85.00%	81.25%
Leakage rate	7.82%	10.52%	7.70%	4.42%	4.94%	3.15%	5.15%	6.88%	5.48%

Note: If readers are interested in the absolute poverty effects of assistance programmes, they may refer to the analysis on target accuracy rates within the text, which indirectly reflects the coverage of assistance at the absolute poverty level.

Source: Authors' elaboration.

16.08 percent (Table 6). However, its targeting efficiency is weaker, with the under-coverage rate being 0.5 percent to 1.5 percent higher than that of the Dibao programme. Moreover, within expenditure-based special assistance, the under-coverage rate for medical assistance was higher than that of educational assistance, indicating that many rural households are still suffering from high medical expenses without obtaining adequate assistance.

Guangdong has enhanced coordination across social assistance programmes since 2017, yet rural programme targeting remains inefficient, with persistent exclusion errors. Studies attribute these gaps to flawed eligibility criteria, showing an 89.5 percent under-coverage rate (Yan, Xue and Feng, 2023). Notably, the Dibao under-coverage rate for disabled populations reaches 87.8 percent, which is 5–8 percentage points higher than our analysis indicates. Comparative analyses reveal the rural Dibao under-coverage rates to be 89.96 percent (income-based) and 91.51 percent (asset-based), exceeding urban programme rates of 79.79 percent and 86.58 percent respectively (Zhang, Zeng and Yuan, 2017). As Table 5 shows, Guangdong's rural Dibao under-coverage rates improved marginally from 85.86 percent (2019) to 82.49 percent (2021). As China's cornerstone poverty alleviation initiative, the Dibao eligibility conditions directly influence other assistance programmes. This systemic interdependence makes rational criteria formulation critical for enhancing overall social assistance efficacy.

Inclusion error

Programme leakage rates – measuring ineligible recipients' access to benefits – showed minimal fluctuation in Guangdong's rural social assistance system in the period 2019–2021. Exclusion rates remained stable at 89.49 percent, 87.69 percent, and 88.64 percent respectively, while corresponding leakage rates hovered between 10.51 percent and 12.31 percent annually (Table 5 and Table 6). However, trends differed across programmes: while medical assistance saw a decline in leakage, Dibao and Educational assistance increased. By 2021, the Dibao leakage rate had risen 1.68 percentage points from 2019, and educational assistance saw a 0.33 percentage-point uptick over the same period.

The inclusion error in the Dibao has drawn extensive attention. One study indicated that 15.87 percent of Dibao benefits were allocated to ineligible households (Liu and Xu, 2016). Another analysis of rural Dibao distribution revealed a leakage rate of 88.55 percent in 2013 (Han and Gao, 2021). Further research showed urban Dibao leakage rates ranged from 54.59 percent to 69.17 percent (Song, Li and Wang, 2020). A more recent study using data from the 2021 Guangdong Thousand-Village Survey reported rural Dibao leakage rates spanning 29.3 percent to 82.7 percent (Yan, Xue and Feng, 2023). Additionally, research has shown that adjusting for selection bias through propensity score matching (PSM) can enhance targeting accuracy, boosting the effective coverage of rural Dibao from roughly 10 percent to 30 percent (Golan, Sicular and Umapathi, 2017).

Guangdong's rural social assistance programmes demonstrate notably low leakage rates, aligning closely with previous studies. This can likely be attributed to reforms implemented since 2013 (Han and Gao, 2021). Prior to that year, village officials held the authority to determine Dibao recipients. However, since 2013, decision-making has been transferred to the village representative assembly. The two village committees are now responsible for collecting Dibao applications, while social workers from township governments handle the assessment and approval processes. This shift has also set a benchmark for other assistance programmes. As a result, issues such as "relationship-based" or "favour-based" assistance have gradually been reduced, leading to a noticeable decline in the leakage rate.

Impacts of social assistance on current income inequality among rural households

Baseline regression analysis

This study employs a panel Tobit model to examine the impact of social assistance on income inequality. Table 7 presents the results. As shown in model 1, social assistance has a positive impact on individual households' relative deprivation, suggesting it may worsen income inequality, a finding significant at the 1 percent level. Models 2 to 4 further examine the effects of Dibao, medical assistance, and educational assistance programmes on relative deprivation respectively. The results indicate that all of the assistance programmes increase individual households' relative deprivation. However, medical assistance shows no significant effect, while the impact of Dibao and educational assistance remains significant at the 1 percent and 5 percent levels, respectively. However, this result differs from that in the descriptive statistics. Based on unconditional mean comparisons across the total sample (8,728 households), descriptive statistics reveal a declining trend in the relative deprivation index from 2019 to 2020, indicating a reduction in income inequality. After controlling for variables such as household size, education, labour force status, and village characteristics in the regression model, the net effect of assistance emerged as positive. This suggests that the negative correlation between assistance and low poverty observed in the descriptive statistics may stem from selection bias, where households with higher poverty levels are more likely to receive assistance. The regression results, by removing this bias, unveil the causal effect of assistance itself, further confirming that the mis-allocation of assistance may worsen relative poverty.

Control variables including householders' age, gender, education, occupation, family size, and village characteristics (population density, urban proximity) reveal additional patterns. Among them, householders' age positively correlates with deprivation, while education and female-led households reduce deprivation. Larger families show lower deprivation. We further controlled for village-level indicators of public service supply in models, such as the number of primary schools and the number of health workers per local residents covered. The results indicate that insufficient public service supply significantly increases

Table 7. *Baseline regression analysis: The impact of social assistance on household current income inequality*

	Model 1	Model 2	Model 3	Model 4
Relative deprivation in income				
Sharing Social Assistance	0.039*** (4.63)			
Sharing Dibao		0.053*** (4.48)		
Sharing Medical Assistance			0.008 (0.58)	
Sharing Educational Assistance				0.028** (2.45)
Age of householder	0.004*** (13.04)	0.004*** (12.91)	0.004*** (12.95)	0.004*** (13.03)
Gender of householder	-0.022* (-1.94)	-0.021* (-1.91)	-0.022* (-1.96)	-0.022* (-1.97)
Educational level of householder	-0.036*** (-7.53)	-0.036*** (-7.47)	-0.037*** (-7.64)	-0.037*** (-7.64)
Householder engages in agricultural activities	0.008 (1.29)	0.008 (1.36)	0.007 (1.22)	0.008 (1.26)
Family size	-0.021*** (-15.86)	-0.021*** (-15.71)	-0.021*** (-15.87)	-0.022*** (-15.95)
Registered population of the village	-0.000** (-2.29)	-0.000** (-2.35)	-0.000** (-2.33)	-0.000** (-2.31)
Distance between village and town	0.002* (1.90)	0.002* (2.00)	0.002* (2.07)	0.002* (2.01)
Population served by per health worker	0.000** (2.06)	0.000** (2.19)	0.000** (2.07)	0.000** (2.08)
Number of village primary schools	0.029*** (4.09)	0.030*** (4.18)	0.030*** (4.19)	0.030*** (4.14)
_cons	0.452*** (15.66)	0.452*** (15.68)	0.458*** (15.84)	0.456*** (15.78)
sigma_u _cons	0.184*** (51.63)	0.184*** (51.75)	0.185*** (51.96)	0.185*** (51.95)
sigma_e _cons	0.183*** (87.73)	0.183*** (87.78)	0.183*** (87.75)	0.183*** (87.77)
N	7014	7014	7014	7014

Notes: t statistics in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Source: Authors' elaboration.

Table 8. *Mechanism analysis: Impacts of coverage errors of social assistance on current income inequality among rural households*

	Model 5	Model 6	Model 7	Model 8
	Eligible households	Ineligible households	Eligible households	Ineligible households
Relative deprivation in income				
Sharing Social Assistance	-0.033*** (-3.60)	0.041*** (3.99)		
Sharing Dibao			-0.046*** (-4.10)	0.053*** (3.42)
Sharing Medical Assistance			-0.035*** (-2.49)	0.003 (0.15)
Sharing Educational Assistance			0.010 (0.73)	0.022 (1.52)
Age of household head	0.001*** (2.95)	0.004*** (9.79)	0.001*** (3.08)	0.004*** (9.73)
Gender of household head	-0.016 (-1.44)	-0.009 (-0.76)	-0.017 (-1.51)	-0.009 (-0.76)
Educational level of household head	-0.004 (-0.66)	-0.032*** (-6.26)	-0.005 (-0.86)	-0.032*** (-6.24)
Household head engages in agricultural activities	0.016*** (2.05)	0.009 (1.34)	0.015*** (2.02)	0.009 (1.39)
Family size	-0.013*** (-8.55)	-0.021*** (-13.53)	-0.013*** (-8.68)	-0.021*** (-13.41)
Registered population of the village	-0.000 (-0.46)	-0.000* (-1.91)	-0.000 (-0.46)	-0.000* (-1.94)
Distance between village and town	-0.001 (-1.42)	0.001 (1.00)	-0.002 (-1.55)	0.001 (1.02)
Population served by per health worker	0.000 (0.27)	0.000* (1.83)	0.000 (0.23)	0.000* (1.94)
Number of village primary schools	0.007 (0.91)	0.031*** (3.92)	0.006 (0.85)	0.032*** (3.99)
_cons	0.830*** (25.22)	0.413*** (13.49)	0.833*** (25.45)	0.412*** (13.45)
sigma_u _cons	0.102*** (19.44)	0.170*** (42.34)	0.102*** (19.19)	0.171*** (42.44)
sigma_e _cons	0.120*** (33.99)	0.178*** (70.55)	0.119*** (33.69)	0.178*** (70.55)
N	1852	5162	1852	5162

Notes: t statistics in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Source: Authors' elaboration.

relative income deprivation but does not alter the influence of key variables on the dependent variable.

Impacts of coverage errors of social assistance on current income inequality

Coverage errors of social assistance. To assess how targeting inaccuracies affect income-related deprivation, we stratified households into two groups: i) those eligible for assistance and ii) ineligible recipients (Table 8). Models 5 and 6 show that receiving assistance reduces relative deprivation for eligible households, a result significant at the 1 percent confidence level. Conversely, when ineligible households receive assistance, their relative deprivation increases – also significant at the 1 percent level. These findings suggest that when ineligible families gain assistance and their income rises, non-recipients perceive a wider income gap, intensifying their sense of deprivation.

Endogeneity test. Table 9 reports both the Tobit coefficients and average marginal effects, which are closely aligned in their values. This indicates a low proportion of observations near model boundaries, suggesting robust and stable model estimation. The results of model 9 and model 11 reveal that the sign of the effects for both groups aligns closely with the main regression conclusions in Table 8.

Table 9. *Tobit regression results and marginal effects based on Kernel-matched samples*

	Model 9	Model 10	Model 11	Model 12
Variable	**Eligible households**	**ME**	**Ineligible households**	**ME**
Sharing Social Assistance	-0.045*** (0.009)	-0.045*** (0.009)	0.051*** (0.010)	0.051*** (0.010)
Sharing Dibao	-0.062*** (0.010)	-0.062*** (0.010)	0.061*** (0.015)	0.061*** (0.015)
Sharing Medical Assistance	-0.058*** (0.015)	-0.058*** (0.015)	0.037* (0.019)	0.037* (0.019)
Sharing Educational Assistance	-0.030** (0.013)	-0.030** (0.013)	0.033** (0.013)	0.033** (0.013)
N	1852	1852	5162	5162

Notes; (1) The estimates are based on kernel-matched weighted samples. The Tobit model is specified with a lower limit of 0 and an upper limit of 1. (2) Standard errors are reported in parentheses. $^{***}p < 0.001$, $^{**}p < 0.01$, $^{*}p < 0.05$. (3) Eligible households refer to those meeting the policy criteria for social assistance application, while ineligible households are coded as 1. (4) The sample size varies slightly across assistance types; the maximum sample size is reported here.

Source: Authors' elaboration.

For households eligible for assistance, the marginal effects of all assistance types are negative, indicating that assistance effectively reduces relative poverty among these households. Specifically, the poverty-reduction effect of Dibao is the most significant, reflecting its institutional role as a direct income supplement. The effects of medical and educational assistance, while slightly smaller compared to Dibao, were both significantly negative. These findings confirm the effectiveness of

Table 10. *Robustness test: Year-by-sample regression*

	Eligible households			Ineligible households		
	Mode13	Model 14	Model 15	Model 16	Model 17	Model 18
	2019	2020	2021	2019	2020	2021
Relative deprivation in current income						
Sharing Social Assistance	-0.060*** (-4.02)	-0.034*** (-2.52)	-0.027* (-1.72)	0.040** (2.09)	0.052*** (2.96)	0.061*** (3.22)
Age of household head	0.001* (1.82)	0.001 (1.29)	0.001 (1.38)	0.002*** (4.43)	0.002*** (4.65)	0.004*** (8.01)
Gender of household head	0.006 (0.35)	-0.002 (-0.12)	-0.034* (-1.84)	-0.011 (-0.58)	-0.007 (-0.38)	-0.004 (-0.24)
Educational level of household head	-0.010 (-0.98)	-0.011 (-1.15)	0.002 (0.22)	-0.034*** (-4.21)	-0.035*** (-4.35)	-0.038*** (-4.90)
Household head engages in agricultural activities	0.019 (1.54)	-0.009 (-0.74)	-0.045** (-2.34)	0.073*** (6.05)	0.033*** (2.72)	-0.008 (-0.51)
Family size	-0.015*** (-5.85)	-0.016*** (-6.72)	-0.009*** (-3.77)	-0.030*** (-11.78)	-0.023*** (-9.63)	-0.013*** (-5.45)
Registered population of the village	-0.000 (-0.11)	0.000 (1.53)	-0.000** (-1.97)	0.000 (0.54)	-0.000 (-0.19)	-0.000*** (-3.56)
Distance between village and town	-0.002 (-1.51)	-0.002 (-1.52)	0.000 (0.19)	0.001 (-0.69)	-0.001 (-1.01)	0.006*** (3.38)
Population served by per health worker	0.000 (0.83)	-0.000 (-1.63)	0.000* (1.68)	0.000 (0.11)	0.000 (0.12)	0.000*** (3.06)
Number of village primary schools	0.015 (1.39)	-0.022* (-1.70)	0.007 (0.58)	0.035*** (2.64)	0.032** (2.56)	0.037*** (3.13)
_cons	0.840*** (17.44)	0.924*** (19.17)	0.805*** (14.02)	0.456*** (9.89)	0.484*** (10.77)	0.308*** (6.72)
sigma _cons	0.147*** (34.53)	0.145*** (34.92)	0.166*** (34.58)	0.238*** (56.78)	0.245*** (59.26)	0.250*** (59.42)
N	620	634	598	1623	1772	1767

Notes: t statistics in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Source: Authors' elaboration.

Table 11. *Robustness test: Lagged dependent variables as control variables and Switching Tobit models*

	Model 19	Model 20
	Eligible households	Ineligible households
Relative deprivation in current income		
Sharing Social Assistance	-0.034*** (-4.13)	0.021** (2.44)
Age of household head	0.000 (0.88)	0.001*** (4.06)
Gender of household head	-0.004 (-0.41)	0.004 (0.42)
Educational level of household head	0.004 (0.73)	-0.005 (-1.36)
Household head engages in agricultural activities	-0.003 (-0.42)	0.009 (1.43)
Family size	-0.005*** (-3.51)	-0.006*** (-5.70)
Registered population of the village	0.000 (0.20)	-0.000** (-2.40)
Distance between village and town	-0.001 (-1.28)	0.000 (0.02)
Population served by per health worker	0.000 (1.08)	0.000 (0.90)
Number of village primary schools	-0.000 (-0.06)	0.014** (2.40)
Relative deprivation of income in 2019	0.360*** (20.28)	0.653*** (60.86)
2019		
2020	0.024*** (2.83)	-0.009 (-1.38)
2021	-0.020** (-2.07)	-0.013* (-1.73)
_cons	0.549*** (16.90)	0.129*** (5.85)
sigma _cons	0.138*** (55.98)	0.182*** (94.11)
N	1609	4449

Notes: *t* statistics in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Source: Authors' elaboration.

social assistance within its target groups. When assistance resources accurately cover eligible households, they can indeed reduce relative deprivation and bridge income gaps. For households that are ineligible for assistance due to policy criteria, the marginal effects of assistance are all positive, indicating that the mis-allocation of assistance (i.e. "incorrect coverage") worsens these households' relative poverty. Specifically, Dibao shows the strongest positive effect, suggesting that allocating limited assistance resources to ineligible households not only wastes public

Table 12. *Regional heterogeneity analysis: Sample of eligible rural households*

	Model 21	Model 22	Model 23	Model 24
	Pearl River Delta	Eastern Guangdong	Western Guangdong	Northern Guangdong
Relative deprivation in current income				
Sharing Social Assistance	0.001 (0.05)	-0.030* (-1.73)	-0.055*** (-3.77)	-0.025 (-1.36)
Age of household head	0.001 (1.39)	0.001* (1.69)	0.001*** (2.03)	-0.000 (-0.12)
Gender of household head	0.034 (1.26)	-0.033 (-1.49)	-0.046** (-2.57)	0.015 (0.64)
Educational level of household head	-0.021 (-1.46)	-0.018 (-1.35)	0.018* (1.75)	-0.010 (-0.85)
Household heads engage in agricultural activities	0.016 (0.86)	0.034** (1.97)	-0.005 (-0.38)	0.036** (2.41)
Family size	-0.009** (-2.07)	-0.011*** (-3.54)	-0.014*** (-5.56)	-0.015*** (-5.17)
Registered population of the village	0.000** (2.28)	-0.000 (-1.22)	-0.000** (-2.33)	0.000* (1.78)
Distance between village and town	-0.000 (-0.13)	-0.005 (-1.49)	-0.001 (-0.84)	-0.000 (-0.07)
Population served by per health worker	-0.000*(-1.75)	0.000(1.44)	0.000(-0.84)	-0.000***(-0.07)
Number of village primary schools	-0.030(-0.99)	-0.023(-1.24)	0.015(1.29)	0.039(1.57)
_cons	0.756*** (8.20)	0.885*** (13.33)	0.878*** (16.44)	0.869*** (11.45)
sigma_u _cons	0.092*** (7.29)	0.105*** (10.49)	0.090*** (9.27)	0.104*** (9.99)
sigma_e _cons	0.117*** (14.60)	0.109*** (15.35)	0.126*** (20.63)	0.117*** (16.43)
N	305	423	666	458

Notes: t statistics in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.0$.

Source: Authors' elaboration.

resources but may also amplify relative deprivation by altering income distribution and triggering social comparison psychology. The findings reveal that “mis-allocated subsidies” not only fail to alleviate poverty but may increase the relative poverty of households outside target groups.

Interestingly, the absolute values of the marginal effects for the eligible and ineligible groups are roughly equivalent. This symmetry further strengthens the

Table 13. *Regional heterogeneity analysis: Sample of ineligible rural households*

	Model 25	Model 26	Model 27	Model 28
	Pearl River Delta	Eastern Guangdong	Western Guangdong	Northern Guangdong
Relative deprivation in current income				
Sharing Social Assistance	0.024 (1.00)	0.056*** (2.82)	0.042** (2.56)	0.015 (0.65)
Age of household head	0.004*** (4.56)	0.004*** (5.89)	0.003*** (4.48)	0.003*** (3.23)
Gender of household head	0.019 (0.75)	-0.057** (-2.54)	-0.011 (-0.49)	0.023 (0.79)
Educational level of household head	-0.033*** (-2.94)	-0.027*** (-3.12)	-0.032*** (-3.47)	-0.021 (-1.64)
Household head engages in agricultural activities	0.043*** (2.91)	0.012 (0.90)	0.002 (0.16)	-0.036** (-2.24)
Family size	-0.023*** (-5.61)	-0.024*** (-8.76)	-0.024*** (-9.11)	-0.017*** (-4.80)
Registered population of the village	0.000 (0.82)	-0.000** (-2.05)	-0.000 (-0.52)	0.000 (1.29)
Distance between village and town	-0.004 (0.67)	0.000 (-1.05)	-0.001 (-1.91)	0.003 (0.58)
Population served by per health worker	0.000 (1.22)	0.000 (1.41)	0.000 (0.14)	0.000 (0.46)
Number of village primary schools	0.021 (0.81)	0.053*** (3.45)	0.008 (0.64)	0.007 (0.24)
_cons	0.329*** (4.43)	0.415*** (8.00)	0.550*** (9.54)	0.423*** (5.45)
sigma_u _cons	0.171*** (19.42)	0.166*** (23.03)	0.159*** (22.29)	0.168*** (16.64)
sigma_e _cons	0.170*** (32.03)	0.168*** (37.28)	0.177*** (39.75)	0.200*** (30.79)
N	1047	1505	1577	1033

Notes: t statistics in parentheses. $^{*}p < 0.1$, $^{**}p < 0.05$, $^{***}p < 0.01$.

Source: Authors’ elaboration.

credibility of causal inference, as the effects of assistance exhibit a “mirror-image” relationship across the two groups. It precisely confirms the policy designers’ expectation that assistance should benefit the target group, while mis-allocation would lead to the contrary result.

Robustness check. To ensure robustness, we conducted regression analyses using year-by-sample regression and further tested the results by adding lagged dependent variables as control variables, as shown in Table 10 and Table 11. Models 13 to 15 show, all else being equal, that receiving social assistance reduces relative deprivation for eligible households – effects significant at the 1 percent and 10 percent confidence levels, consistent with the findings of model 5. Conversely, models 16 to 18 reveal that when ineligible households receive support (after controlling for other factors), their relative deprivation rises, worsening income inequality – significant at the 1 percent to 5 percent levels. While the years 2020 and 2021 coincide with the COVID-19 pandemic, as shown in Table 10, the results indicate that the mis-allocation in social assistance and income inequality from 2019 to 2021 remained consistent with the main regression findings. Table 11 shows Switching Tobit Models and using the relative deprivation in income (Kakwani index) from the base period (2019) as a control variable to test the impact of historical poverty status on current assistance receipt. Results indicate that the coefficient signs and significance levels of the core explanatory variable (Sharing Social Assistance) were not altered (see models 19–20), indicating that reverse causality did not significantly distort the estimation results.

Heterogeneity analysis. The impact of social assistance on rural households’ income-related relative deprivation varies across socioeconomic groups, leading to heterogeneous effects. To explore these regional differences in rural income inequality, we classify households into four groups: Pearl River Delta, Eastern Guangdong, Western Guangdong, and Northern Guangdong, with results in Table 12 and Table 13. For eligible households, receiving assistance reduces relative deprivation – but this effect is significant at the 1 percent level only in Western Guangdong. By contrast, no significant impact emerges in the Pearl River Delta, Eastern Guangdong, or Northern Guangdong.

Ineligible households also show varied regional responses: in Eastern and Western Guangdong, receiving support increases their relative deprivation, worsening income inequality – findings significant at the 1 percent and 5 percent levels, respectively. However, in the Pearl River Delta and Northern Guangdong, social assistance has no statistically significant effect on either reducing or exacerbating households’ relative deprivation.

Discussion and conclusion

This study evaluates the targeting efficiency of social assistance programmes and examines their impact on rural household income inequality. Focusing on policy implementation outcomes, we assess programme effectiveness across four key dimensions: how accurately benefits reach eligible recipients (targeted inclusion), miss vulnerable groups (under-coverage), exclude ineligible individuals (targeted exclusion), and inadvertently include non-target populations (leakage).

Findings show an overall improvement in targeting efficiency. First, the correct inclusion rate for rural social assistance has risen steadily, accompanied by lower under-coverage – trends notably evident in Dibao and educational assistance programmes. However, medical assistance saw a slight uptick in under-coverage, signalling a need to revisit health care eligibility criteria. Since the 2013 policy reforms, leakage rates have declined, reflecting stronger programme integrity. Unlike earlier studies, our analysis finds relatively low leakage in Guangdong's rural assistance system, contradicting prior high-leakage assumptions.

After receiving social assistance, rural households saw a drop in relative deprivation indices, indicating tangible improvements in income inequality metrics. However, regression results show no significant statistical link between social assistance and reduced within-group income disparities. This paradox stems primarily from coverage errors in the rural social assistance system. On the one hand, eligible households that access support experience lower relative deprivation, directly narrowing income gaps. Yet high under-coverage rates mean many eligible families still lack access, limiting the overall equalizing effect. On the other hand, misallocations – where ineligible households receive benefits – boost their incomes, worsening perceived income inequality among non-recipients. Overall, enhancing the targeting accuracy of social assistance programmes is critical to addressing rural income disparities. To strengthen the effectiveness of social assistance, three key strategies are recommended.

First, improve targeting by refining eligibility criteria to better identify the "invisible poor" – households facing unrecognized hardships – and granting social workers limited discretion to swiftly enrol vulnerable families. This should be supported by a shift toward service-oriented assistance, strengthened third-party oversight, and differentiated support for Dibao recipients, near-poor households, and those impoverished by sudden expenditures. To be specific, for low-income households with working capacity, provide developmental support such as employment training and public welfare jobs to help them gradually reduce dependence on assistance. For those without working capacity (such as the elderly and the severely disabled), increase assistance standards and connect them with long-term care services. While for households with short-term unexpected expenses but stable long-term income (such as migrant worker families incurring temporary debt for their children's schooling), introduce a "borrow-and-defer" mechanism. This provides interest-free loans or allows deferred payment of certain fees, rather than direct cash assistance. Second, optimize programme coordination by consolidating overlapping initiatives, expanding coverage where needed, and integrating assistance with social insurance schemes – such as linking medical assistance to critical illness insurance and nursing support to long-term care insurance. Timely emergency assistance should also be prioritized to respond to evolving needs. Third, mobilize broader social participation by encouraging talents and resources to flow into rural areas, strengthening charitable organizations through simplified registration and third-party evaluations, incentivizing social enterprises with tax benefits, and enhancing the skills and motivation of volunteers and social workers through targeted training and support. Together, these efforts aim to create a more inclusive, flexible social assistance system that combines strong government leadership with active grassroots engagement.

Although this study has employed multiple robustness tests wherever possible, causal inferences still require careful consideration. Due to limited data availability, this study can only examine the short-to-medium-term effects of social assistance on income relative deprivation; it fails to assess long-term effects as would be possible with a comprehensive, mature, nationwide survey such as those currently available in China. In terms of reverse causality, households with relatively higher levels of poverty may be more likely to actively apply for assistance, which potentially leads to reverse causality bias. However, we have not identified suitable instrumental variables to more effectively address endogeneity issues. Instead, further validation can only be conducted via Propensity Score Matching (PSM) and time-lagged variable models. Furthermore, this article aims to examine the collective sense of deprivation that may arise from institutionalized distributional inequity. It does not address the deprivation that may stem from individual psychological differences and cultural norms in China. Consequently, research on income inequality in the "Post-Poverty Era" should include these discussions. For instance, research should introduce localized measurement tools to construct models capturing the interaction between culture and institutions. We may incorporate cultural psychological variables, such as "reputation sensitivity" and "social comparison tendencies", into questionnaire designs. A second approach involves supplementing qualitative research, by using semi-structured interviews to explore the sources of rural residents' perceptions of "fairness" in welfare distribution, and identifying distinctive dimensions of relative deprivation in China – such as intergenerational support and social exclusion. Additionally, this article does not directly estimate the impact of assistance on extreme poverty. Perhaps future research could further examine the differences in the effects of assistance on extreme poverty versus relative poverty or discuss whether moderation effects may exist.

Data availability statement

The Guangdong Thousand-Village Survey are open and available dataset from Jinan University. The dataset can be requested from the website of Survey Data Center, Jinan University.

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BOOK REVIEW

Cabib, Ignacio. **Avoiding retirement in Chile: Extending working lives in an uncertain and precarious context.** London, UK, Routledge. 2025. 192 pp. ISBN 9781032777542.

Population ageing and concerns about the sustainability of pension systems have brought renewed attention to the issue of extended working lives. Opinions vary on this issue between those heralding it as a pathway to the better use of untapped labour, which will support healthy labour markets and pension system sustainability, to those voicing concerns about decent work and social protection for older workers.[1] Typically, the existing academic literature examines policy reforms in advanced welfare states, where governments have introduced incentives to delay retirement and encourage labour market participation in later life. It is a welcome addition, then, that in *Avoiding retirement in Chile*, Ignacio Cabib offers a valuable contribution by shifting the focus to a different context: a Latin American country where older people increasingly remain economically active despite the absence of explicit policies promoting longer working lives.

The book begins by presenting a clear empirical puzzle. Over the past two decades, Chile has experienced one of the largest increases in labour force participation among people aged 65+ in Latin America. Using ILOSTAT data for context, most countries in the region have shown either a decline or only a slight increase in participation rates among those aged 65+, whereas in Chile the rate rose from 15.9 percent in 2002 to 25.4 percent in 2019. This rise likely underestimates the true extent of older-age employment in Chile, as it excludes individuals who continue to work while receiving a pension as well as those engaged in non-standard employment. Yet this development cannot easily be explained by the institutional mechanisms commonly highlighted in studies of Member states of the Organisation for Economic Co-operation and Development (OECD), such as rising statutory retirement ages, financial incentives to postpone retirement or active ageing policies. Instead, Cabib argues that extended working lives in Chile are shaped primarily by long-term labour market conditions and life-course experiences. In particular, the accumulation of advantages and disadvantages across key social domains, such as health, family formation and education, play important roles in shaping older workers' attachment to the labour market.

The volume focuses on the current generation of older workers (the baby boomer generation). These individuals entered the labour market in the 1970s and 1980s, a period during which Chile underwent profound institutional transformations. These included the military dictatorship, the restructuring of the pension system to a defined contribution model based on individual accounts administered by private pension funds, as well as subsequent phases of rapid economic growth and democratic transition in the late 1980s and 1990s. Although this cohort was expected and legally

1. See OECD. 2025. *Employment outlook*. Paris, Organisation for Economic Co-operation and Development; ILO. 2024. *Asia Pacific employment and social outlook: Promoting decent work and social justice to manage ageing societies*. Bangkok, International Labour Office; Cooke, F. L.; Rogovsky, N. 2025. Extending the productive working life of older workers: The role of human resource management. Geneva, International Labour Organization.

entitled to retire during the 2010s, a large proportion have continued working well beyond retirement age.

The book's eight chapters guide the reader from the broader research field through to empirical findings and their policy implications. It begins by situating extended working lives within the existing literature, before outlining the study's objectives, data and methods. The core chapters examine life-course trajectories, perceptions of retirement, and the role of key domains such as caregiving, health and education, all within the context of Chile's liberal institutional framework. The final chapters bring these strands together by highlighting how precarity and uncertainty reshape the retirement transition, and by discussing the implications for policy design in similar contexts. All in all, the book offers a synthesis of a prolific research agenda disseminated across several academic publications.

The analysis is based on an ambitious mixed-methods research design. The book combines quantitative evidence from life-course calendar surveys undertaken in Santiago, Chile's capital, with longitudinal qualitative interviews conducted with older workers across the country. This methodological approach allows the author to reconstruct individuals' employment, family and health trajectories across the life course while also exploring how older people perceive the transition to retirement. The integration of these methods provides a rich empirical basis for understanding both the structural and subjective dimensions of work in later life.

While each chapter can be read separately as a stand-alone contribution, a central argument presented in the book is that retirement is viewed increasingly by citizens in Chile as financially unattainable rather than it being postponed. The Chilean pension system relies heavily on individual retirement accounts funded by workers' contributions, which presupposes relatively stable and formal employment trajectories. However, many workers have experienced fragmented careers, periods of informality and unstable labour market participation. As a result, pension incomes often remain insufficient to ensure financial security in old age. In this context, remaining economically active becomes a strategy to supplement income and maintain living standards.

However, this is just half of the story. Cabib develops a framework centred on two analytical concepts: uncertainty and precarity. Despite the Chilean labour market being among the most formal and productive in Latin America, the notion of uncertainty highlights the instability and unpredictability that have characterized the country's labour markets over several decades. It refers both to external conditions and the ways individuals process them. Hence, exogenous uncertainty captures the instability and limited institutional protection that current older workers have faced over the life course, including informality, economic crises and weak policy support. However, endogenous uncertainty refers to individuals' capacity to tolerate, interpret and actively manage these conditions through different strategies. As argued, uncertainty is understood not simply as a constraint, but as a condition shaping agency and decision-making over the life course. Current older workers entered the labour market during periods of economic crisis in the 1970s and 1980s which saw structural adjustment and deregulation, and they often faced unstable employment opportunities and limited institutional protection. These experiences shaped how individuals have approached work and retirement throughout their lives.

The concept of precarity complements this perspective by capturing both the material and subjective dimensions of insecurity. On the one hand, many older individuals face objective risks such as inadequate pensions, rising living costs and ongoing financial responsibilities toward family members. On the other hand, subjective experiences of marginalization, declining job quality, and social isolation also influence decisions to remain economically active. While the objective dimensions of uncertainty and precarity are staple elements of most accounts of Latin American labour markets and pension systems, by focusing on individual perceptions and strategies, the book argues that these factors contribute to a situation in which retirement is not simply delayed – but increasingly avoided.

This conceptual innovation is combined with another strength of the study: the adoption of a life-course perspective, which highlights the dynamic relationship between labour market trajectories and experiences in family life, health and education. The analysis shows how employment in later life is closely linked to earlier experiences across these domains. Gender differences are particularly salient. Women's labour market participation is often shaped by unpaid caregiving responsibilities and intermittent employment histories, while men more frequently follow formal full-time trajectories. These cumulative experiences contribute to the diverse employment patterns observed among older workers in Chile. In this vein, interview evidence suggests that some older women avoid retirement not only for financial reasons but also to resist the expectation of taking on additional domestic responsibilities, such as caring for grandchildren or elderly spouses.

The book also offers insights into the role of welfare regimes in shaping later-life employment. Chile's highly market-oriented social protection system places considerable responsibility on individuals and families to manage risks related to work, health and old age. In such a context, extended working lives emerge primarily not as a policy objective but as a response to the limitations of existing social security arrangements.

Overall, *Avoiding retirement in Chile* provides an important contribution to international debates on ageing, labour markets and pension systems. By highlighting the life-course and institutional conditions that shape extended working lives in a middle-income country, the book broadens the geographical scope of existing research and raises important questions for policy makers concerned with retirement security, decent work, and labour market inclusion in later life.

A Spanish-language edition of the book (*Evitando la jubilación en Chile: trayectorias laborales extendidas en contextos inciertos y precarios*. Santiago, Ediciones UC. 2025) has also been published, which will extend the international reach of Cabib's important findings.

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