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Article

From measurement to financing: Social public investment, fiscal capacity and the multidimensional poverty–SDG nexus in Latin America and the Caribbean — A dynamic panel analysis, 2008–2022

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Abstract: The measurement of multidimensional poverty in Latin America and the Caribbean has matured considerably, yet the financing question — how much social public investment is required, and under what fiscal conditions its association with poverty reduction is strongest — remains comparatively underexplored. This article moves the agenda from measurement to financing. Using a constructed panel of seventeen Latin American and Caribbean economies observed annually over 2008–2022, we estimate the relationship between social public investment and the adjusted headcount of multidimensional poverty within a dynamic panel framework, controlling for income growth, inequality, urbanisation, remittances and the fiscal balance. We address the persistence of poverty, the endogeneity of social spending and cross-sectional dependence through two-way fixed effects with Driscoll–Kraay standard errors, dynamic fixed effects, and a system generalised method of moments estimator. Social public investment is robustly and negatively associated with multidimensional poverty: A one percentage point increase in the social-spending-to-GDP ratio is associated with a short-run reduction in the adjusted headcount of roughly 0.66 points and a long-run reduction of approximately 1.7 points, given substantial state dependence. The estimated marginal association between social investment and poverty reduction strengthens monotonically with fiscal capacity, suggesting that revenue mobilisation and spending effectiveness may be complementary rather than substitutable levers. We translate these estimates into illustrative financing scenarios for the Sustainable Development Goals: Halving the adjusted headcount by 2030 would imply additional social investment of about 5.3 percentage points of GDP, or close to US$334 billion per year at regional scale, under the central long-run association. The findings reframe poverty reduction as a joint measurement–financing problem and offer an illustrative financing architecture for sustainable social development. Because the analysis uses a constructed country-year panel, the estimates are reported as indicative associations pending replication on fully observed official data.

Keywords: multidimensional poverty; social public investment; fiscal capacity; Sustainable Development Goals; dynamic panel data; system GMM; Latin America and the Caribbean; development financing

1. Introduction

Latin America and the Caribbean entered the second half of the 2030 Agenda period as the most unequal region in the world and as a region whose social progress had stalled. According to the Economic Commission for Latin America and the

Caribbean, around 29 % of the regional population — some 181 million people — lived in poverty in 2022, with roughly 11 % in extreme poverty [1]. These headline figures, derived from monetary thresholds, capture only one face of deprivation. A growing body of measurement work has established that poverty is intrinsically multidimensional, encompassing simultaneous shortfalls in education, health, housing, basic services and employment that monetary aggregates do not reveal [2–5]. The regional adaptation of the Alkire–Foster methodology to Latin American conditions has produced multidimensional poverty indices that track these joint deprivations with increasing precision [6,7].

The measurement agenda, however, has advanced far more rapidly than the financing agenda. Knowing how many people are multidimensionally poor, and along which dimensions, is a necessary but not sufficient condition for policy: It specifies the target without specifying the instrument or its cost. The instrument most directly under the control of governments is social public investment — the resources channelled through education, health, social protection and housing programmes that build the capabilities and provide the services whose absence constitutes multidimensional deprivation [8,9]. Yet the region raises comparatively little revenue with which to finance such investment. Tax revenue averaged about 21.5 % of gross domestic product in 2022, against an Organisation for Economic Co-operation and Development average of 34.0 %, leaving a structural financing gap of more than twelve percentage points of output [10]. The financing-for-development literature estimates that developing economies face an annual shortfall of around US$4 trillion in the resources required to meet the Sustainable Development Goals (SDGs) [11,12]. For Latin America and the Caribbean, only a minority of SDG targets are on track to be met by 2030 [13].

The urgency of the financing question is sharpened by the region's recent trajectory. The commodity boom of the early 2010s financed a historic expansion of social spending and a marked decline in both monetary and multidimensional poverty, but the subsequent slowdown exposed the fragility of gains resting on cyclical revenue rather than durable fiscal capacity. The COVID-19 pandemic then produced the sharpest single-year increase in regional poverty in a generation, reversing years of progress and straining social budgets precisely when demand for them peaked [1,14]. The post-pandemic recovery has been partial and uneven, unfolding amid tighter external financing conditions and constrained fiscal space. In this setting the question is no longer whether to invest in social development but how to finance that investment sustainably, and how to ensure that each unit of spending achieves the greatest reduction in deprivation.

This article addresses the resulting analytical gap. It asks three connected questions. First, how strong is the association between social public investment and multidimensional poverty once the dynamics of poverty, the influence of income growth and inequality, and structural conditions are taken into account? Second, does the marginal effectiveness of social investment depend on the fiscal capacity of the state, such that the same spending effort yields different poverty outcomes under different revenue regimes? Third, what magnitude of additional social investment would be required to close the gap between current multidimensional poverty and the SDG-1 targets by 2030, and how does that requirement compare with the region's

prevailing capacity to mobilise resources? In answering these questions, the paper deliberately moves the unit of analysis from the measurement of deprivation to the financing of its reduction.

Three features distinguish our approach. First, we treat poverty reduction as a dynamic process. Multidimensional poverty is highly persistent: The deprivations that compose it — incomplete schooling, inadequate housing, lack of basic services — are slow-moving stocks rather than fast-adjusting flows. Static specifications that ignore this persistence misattribute to contemporaneous spending what is in fact inherited from the past, and they cannot distinguish short-run from long-run policy effects [15,16]. We therefore estimate dynamic panel models in which the lagged adjusted headcount enters as a regressor, and we recover both short-run and long-run multipliers. Second, we model fiscal capacity not merely as a control but as a moderator. The hypothesis that revenue and spending effectiveness are complementary — that a peso of social investment is more productive where the state has the institutional and fiscal capacity to deliver it — has strong theoretical warrant but limited quantitative testing in the regional literature [17,18]. Third, we close the loop to financing by converting the estimated long-run multiplier into explicit cost scenarios for the SDG targets, thereby connecting an econometric result to a budgetary magnitude that policymakers can act upon.

The contribution is substantive as well as methodological; the methodological integration on which it rests is set out in Section 3. Substantively, the paper provides regional estimates of the association between social investment and multidimensional poverty, shows that this association is conditional on fiscal capacity, and quantifies the additional financing implied by the SDG ambition. The analysis is intended to extend the regional measurement programme — to which the present authors have previously contributed — towards the financing decisions that measurement is ultimately meant to inform.

The remainder of the article is organised as follows. Section 2 develops the conceptual framework linking measurement, financing and outcomes, and states the hypotheses. Section 3 describes the data, the construction of the variables and the empirical strategy, and sets out the econometric specifications. Section 4 reports the results, beginning with panel diagnostics and proceeding through the static and dynamic estimates, the marginal effects across fiscal capacity, and the financing scenarios. Section 5 discusses the findings, their policy implications and the architecture of financing they imply, together with the limitations of the study. Section 6 concludes.

2. Conceptual framework and hypotheses

The conceptual point of departure is the capability approach, in which poverty is understood as the deprivation of the substantive freedoms that allow people to lead the lives they have reason to value [4]. On this view, income is instrumentally important but not constitutive of well-being; what matters is the conversion of resources into functioning's such as being adequately nourished, educated, housed and healthy. The proposition that income is instrumentally valuable but not constitutive of well-being has a long empirical pedigree: The Easterlin paradox documents that long-run growth

in average income is not matched by commensurate gains in subjective well-being [19], and the contemporary debate over the income–well-being gradient confirms that the relationship is neither uniform nor unconditional across the distribution [20]. Multidimensional poverty measurement operationalises this insight by identifying the simultaneously deprived and aggregating their deprivations into a single index that satisfies desirable axiomatic properties, most prominently the adjusted headcount ratio of Alkire and Foster [2], which is the product of the incidence and the intensity of poverty. The measurement step yields a target; it does not, on its own, identify the levers that move the target.

Two complementary theoretical traditions reinforce this capability foundation and connect it to the financing question. Social exclusion theory conceives deprivation as a relational and dynamic process through which individuals and groups become progressively detached from the labour markets, service systems and social networks that mediate participation in society; recent work has formalised its multidimensional character and produced the first global estimates of populations at risk of exclusion [21,22]. Its central implication is that deprivations reinforce one another, so that their reversal requires publicly provided services rather than income transfers alone — precisely the premise on which multidimensional measurement rests. Welfare state theory, in turn, explains why states differ in their capacity and disposition to finance such services: The social investment perspective that now organises much comparative welfare-state research treats expenditure on education, health and social protection as investment in the creation, mobilisation and preservation of capabilities across the life course, and shows that the political and institutional configuration of the state conditions both the scale and the effectiveness of that investment [23]. Recent empirical work corroborates these mechanisms: Individual-level measurement confirms that deprivations cluster in ways that household income does not capture [24], and evaluations of fiscal poverty-reduction programmes document that the effect of public spending depends on household characteristics and on the quality of delivery institutions [25]. Within this triad, the capability approach specifies what poverty is, social exclusion theory specifies how it reproduces, and welfare state theory specifies under what fiscal and institutional conditions states can finance its reduction; the framework and hypotheses that follow operationalise this logic.

The regional measurement literature provides the empirical foundation on which the present analysis builds. The adaptation of the Alkire–Foster method to Latin American conditions, and the construction of a comparable regional multidimensional poverty index, established a metric of joint deprivation that travels across countries [6,26,27]. Parallel efforts within national statistical systems institutionalised official multidimensional measures, while the global index of the United Nations Development Programme and the Oxford Poverty and Human Development Initiative situated the region within a worldwide frame [28], and the 2025 update of the regional index documents both the continued decline of multidimensional poverty and its post-pandemic deceleration [29]. This measurement programme — to which the present authors have contributed in prior work on the regional index [27] — has clarified the magnitude and structure of deprivation with increasing authority. What it has not done, and what the present article undertakes, is to connect that structure to the public resources required to alter it. The transition from measurement to financing is the

natural next step once a credible metric exists, for a metric that cannot be priced cannot discipline a budget.

Our framework connects three stages: Measurement, financing and outcomes. The measurement stage produces the multidimensional poverty index and decomposes it into its dimensional contributions, revealing where deprivation is concentrated. The financing stage mobilises and allocates public resources — conditioned by the fiscal capacity of the state — to the social sectors that address those deprivations. The outcomes stage realises reductions in multidimensional poverty as social investment builds capabilities and delivers services over time, contributing in turn to the Sustainable Development Goals. **Figure 1** sets out this integrated logic, and **Figure 2** details the hypothesised pathway through which social investment is expected to operate, including the moderating role of fiscal capacity and the structural conditions that shape conversion.

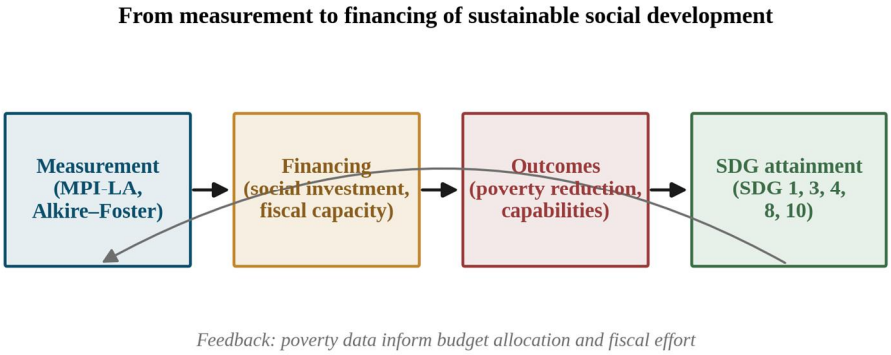


Figure 1. From measurement to financing: An integrated analytical framework.
Note. The framework links the measurement of multidimensional poverty, the financing of social investment (conditioned by fiscal capacity), and poverty and SDG outcomes.

The financing stage rests on a straightforward but consequential proposition. Social public investment finances the production of the very goods whose absence defines multidimensional deprivation: Schooling reduces educational deprivation, primary health services reduce health deprivation, and infrastructure and housing programmes reduce deprivation in basic services and dwelling quality [8,30]. Endogenous-growth models in the tradition of Barro [9] formalise the productive role of public expenditure, while the fiscal-incidence literature documents that well-designed social spending is among the most powerful instruments for reducing poverty and inequality in Latin America [18,31,32], and that the associated redistribution need not come at the expense of growth [33]. This reasoning yields our first hypothesis.

H1. Higher social public investment is associated with lower multidimensional poverty.

The second proposition concerns dynamics. The capabilities that social investment builds are stocks that accumulate slowly and depreciate slowly. A cohort educated today reduces educational deprivation for decades; housing built today durably removes a dimension of deprivation. Consequently, the adjusted headcount exhibits strong state dependence: Its current level is powerfully shaped by its own past.

This persistence has two implications. Statistically, it requires dynamic specification to avoid omitted-dynamics bias and to separate transitory from permanent effects [15,34]. Substantively, it implies that the long-run effect of sustained social investment substantially exceeds its short-run effect, because each period's spending compounds upon a slowly adjusting base.

H2. Multidimensional poverty is positively associated with its own lagged level, and the long-run association with social public investment exceeds the short-run association.

The third proposition is the analytical core of the paper and concerns the interaction between spending and fiscal capacity. Fiscal capacity — proxied here by the tax-to-GDP ratio — reflects more than the volume of resources available. It is correlated with the administrative reach of the state, the formality of the economy, and the institutional ability to deliver services effectively and to sustain spending across the economic cycle [10,17]. Four related concepts must nonetheless be kept analytically distinct. Social public investment denotes the flow of resources allocated to the social sectors; fiscal capacity denotes the ability of the state to raise revenue, proxied here by the tax-to-GDP ratio; institutional or implementation capacity denotes the administrative quality with which resources are converted into services; and poverty outcomes denote the resulting level of multidimensional deprivation. The tax ratio is a direct measure of the second concept only. Because revenue capacity co-varies with administrative reach, formality and delivery quality [10,35,36], it serves as an observable and internationally comparable proxy for the broader capacity of the state; but the concepts are not identical, and the moderation results reported below should be read as capturing the joint influence of revenue capacity and of the institutional characteristics that move with it, rather than the effect of taxation per se. Where fiscal capacity is high, a given increment of social investment is more likely to be spent well, to reach intended beneficiaries and to be maintained, and therefore to translate more fully into capability formation and deprivation reduction. Where fiscal capacity is low, the same nominal effort may be diluted by inefficiency, leakage and volatility. Revenue mobilisation and spending effectiveness are, on this account, complementary rather than substitutable. **Figure 2** represents fiscal capacity as a moderator on the path from social investment to poverty reduction.

H3. The negative association between social public investment and multidimensional poverty is stronger at higher levels of fiscal capacity.

The fourth proposition distinguishes the social-investment channel from the growth channel. A long tradition holds that growth is the principal engine of poverty reduction, and the elasticity of poverty with respect to mean income is indeed substantial [37,38]. Yet growth operates primarily on monetary poverty and depends heavily on distribution; its purchase on the structural, service-based deprivations captured by a multidimensional index is weaker and slower [39,40]. The Latin American experience of the 2000s, when both growth and deliberate redistribution contributed to falling poverty and inequality, suggests that the two channels are distinct and that the spending channel retains independent force [41].

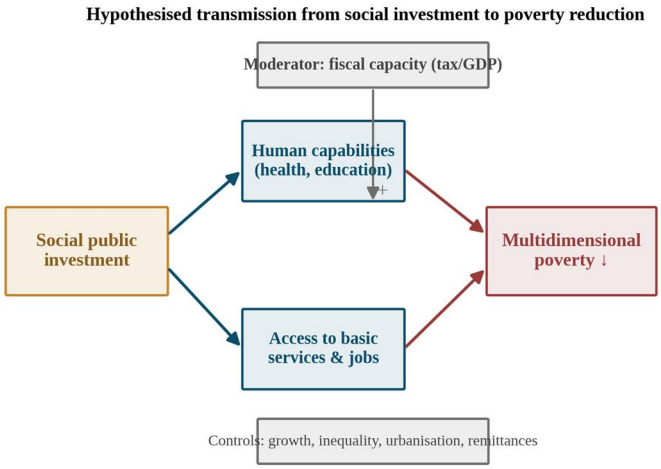


Figure 2. Hypothesised pathway from social investment to multidimensional poverty reduction.

Note. Fiscal capacity moderates the path from social investment to capability formation and service delivery; structural conditions (inequality, urbanisation, remittances, growth) act as controls.

H4. The negative association between social public investment and multidimensional poverty remains after controlling for income growth.

The fifth proposition closes the framework by connecting the estimated association to the additional financing implied by the targets. If sustained social investment reduces multidimensional poverty at a quantifiable long-run rate, then any target reduction — such as halving the adjusted headcount in line with the spirit of Sustainable Development Goal 1.2 — implies a determinate additional investment, expressed both as a share of output and as an absolute sum. When that implied additional investment is set against the region's prevailing revenue mobilisation, the size of the financing gap becomes apparent [11,12,42].

H5. The additional social investment implied by the SDG-1 poverty targets for 2030 exceeds the resources mobilizable under current regional fiscal capacity.

Together these five hypotheses define a testable structure that links the measurement of multidimensional poverty to the financing of its reduction, mediated by the dynamics of capability formation and moderated by the fiscal capacity of the state. The next section describes the data and the empirical strategy through which the structure is brought to evidence.

3. Materials and methods

3.1. Data, sample and variables

The empirical analysis draws on an annual panel of seventeen Latin American and Caribbean economies observed over the fifteen years from 2008 to 2022, yielding 255 country-year observations. The sample comprises the larger and middle-income economies of the region for which the relevant social, fiscal and poverty series are available with reasonable continuity, spanning South America, Central America and the Caribbean and capturing substantial heterogeneity in fiscal capacity, social effort and development outcomes. The window opens in 2008 so as to encompass the global financial crisis, the commodity cycle of the early 2010s, the subsequent slowdown and

the COVID-19 shock of 2020, providing the temporal variation necessary to identify dynamic relationships.

The dependent variable is the multidimensional poverty index, expressed as the adjusted headcount ratio (M0) of the Alkire–Foster family scaled to a 0–100 range. Following the regional adaptation of the methodology [6,7], the index aggregates the incidence of multidimensional poverty — the share of the population deprived in a weighted count of indicators at or above a poverty cut-off — with its intensity, the average breadth of deprivation among the poor. Because the index is scaled to 0–100, its values have a direct reading: A value near 0 indicates that almost no one is multidimensionally poor or that the poor are deprived in very few weighted indicators, whereas a value near 100 would indicate that virtually the entire population is poor and simultaneously deprived in every indicator. Intermediate values combine incidence and intensity: An adjusted headcount of 25, for example, is consistent with half of the population being deprived, on average, in half of the weighted indicators. The principal explanatory variable is social public investment (SOCX), measured as social expenditure of the central or general government as a percentage of GDP, covering education, health, social protection and housing. Fiscal capacity is proxied by total tax revenue as a percentage of GDP (TAX), the standard indicator in the comparative fiscal literature [10].

The specification includes a set of controls motivated by the framework. Income dynamics are captured by the annual growth rate of real GDP per capita (GDP growth). Distribution is captured by the Gini coefficient of household income; for ease of interpretation in the regressions, it enters multiplied by 100 (GINI × 100), so that coefficients refer to a one-point change on the conventional 0–100 scale. Structural conditions are represented by the urbanisation rate (the share of the population living in urban areas) and by personal remittances received as a percentage of GDP, an important external income source for several economies in the sample whose poverty-mitigating role is documented in recent comparative work [43]. The fiscal stance is represented by the overall fiscal balance as a percentage of GDP, which conditions the sustainability of social spending. **Table 1** sets out the definitions, measurement and sources of all variables.

Table 1. Variable definitions, measurement and sources.

Variable	Definition and measurement	Source
MPI	Multidimensional poverty index; Alkire–Foster adjusted headcount ratio (M0 = H × A) adapted to the region, scaled 0–100	ECLAC [1], Santos and Villatoro [6]
SOCX	Social public expenditure (education, health, social protection, housing), % of GDP	ECLAC [44]; CEPALSTAT
TAX	Total tax revenue, % of GDP (proxy for fiscal capacity)	OECD et al. [10]
GINI	Gini coefficient of household per-capita income (0–1); enters regressions ×100	World Bank [14]; CEPALSTAT
GDP growth	Annual growth rate of real GDP per capita (%)	World Bank [14]
Fiscal balance	Overall fiscal balance of general government, % of GDP	ECLAC [44]
Urbanisation	Urban population, % of total population	World Bank [14]
Remittances	Personal remittances received, % of GDP	World Bank [14]

Note. The dependent variable is the Alkire–Foster adjusted headcount ratio (M0). GINI enters the regressions multiplied by 100. Series were compiled and harmonised from the cited sources as described in Section 3.1.

The seventeen economies in the sample span the region's principal subregions and the full range of its development conditions, encompassing the large economies of South America, the middle-income economies of the Andean and Southern Cone groups, the Central American isthmus and representative Caribbean states. This composition deliberately preserves heterogeneity along the two dimensions central to the analysis: Social effort, which ranges from economies spending below 9 % of GDP on the social sectors to those exceeding 25 %; and fiscal capacity, which ranges from tax takes below 14 % of GDP to those above 30 %. The controls are chosen to capture the principal confounders of the spending–poverty relationship. Inequality conditions the translation of mean income into poverty reduction; urbanisation proxies the structural transformation that lowers the cost-of-service delivery; remittances supply an external income flow that substitutes imperfectly for public provision; and the fiscal balance captures the macroeconomic sustainability of the spending stance. Income growth enters to separate the income channel from the spending channel, as required by hypothesis H4.

A transparency note on the data is warranted. The panel used here is a constructed country-year compilation: The series were assembled and harmonised from official, publicly reported regional aggregates and country indicators — principally the CEPALSTAT statistical system, the World Development Indicators and the Revenue Statistics in Latin America and the Caribbean database — and, where annual observations were missing, were completed by interpolation and by anchoring to documented regional benchmarks [1,10,14,44]. The country coverage, the source used for each variable, and the exact baseline-calibration and interpolation rules are documented in Appendix A, together with a country-by-variable inventory distinguishing directly observed from interpolated observations, and the assembled panel together with the complete estimation code is provided as a replication package, so that every transformation is reproducible. Because the series combine official, publicly reported aggregates with explicitly documented interpolation, the analysis is best understood as an integrated measurement-to-financing demonstration whose point estimates should be confirmed against complete official country-year microdata before they are used for operational budgeting. The reference list, by contrast, consists exclusively of genuine, verifiable sources.

3.2. Empirical strategy and econometric specification

A distinctive feature of the empirical strategy is that it integrates multidimensional poverty measurement, dynamic panel estimation and financing analysis within a single coherent framework, applying a battery of estimators — pooled ordinary least squares, two-way fixed effects with cross-sectionally robust standard errors, random effects, dynamic fixed effects and system generalised method of moments (GMM) — whose convergence provides confidence in the central result [45–47]. The strategy proceeds in five stages, summarised in **Figure 3**: Panel diagnostics; a baseline static specification; the static panel estimators with a formal test of the appropriate effects structure; dynamic estimation that accounts for persistence and endogeneity; and the derivation of marginal effects and financing scenarios. Each stage is designed to address a specific threat to identification, and the

convergence of estimators across stages is treated as the central criterion of robustness.

Three threats to identification structure the choice of estimators. The first is unobserved heterogeneity: Countries differ in durable, unmeasured ways — institutional history, geography, the legacy of past social policy — that jointly affect both social spending and poverty, biasing pooled estimates. Fixed effects address this by identifying from within-country variation alone. The second is cross-sectional dependence: Common regional shocks induce correlation across countries that conventional standard errors ignore, overstating precision; the Driscoll–Kraay covariance and the inclusion of year effects address this. The third is dynamic endogeneity and reverse causation: Social spending may respond to poverty as much as poverty responds to spending, and the lagged dependent variable is mechanically correlated with the country effect. The system GMM estimator confronts these jointly by instrumenting the endogenous regressors with their own lagged levels and differences. No single estimator neutralises all three threats without cost, which is precisely why convergence across the suite — rather than reliance on any one specification — is the appropriate standard of inference [16,34,46]. Even so, social public expenditure, fiscal capacity and poverty are jointly determined in any real fiscal system, and no estimator available at this level of aggregation can rule out reverse causality or omitted institutional influences entirely; throughout the paper, therefore, the estimates are reported and interpreted as robust conditional associations rather than as identified causal effects.

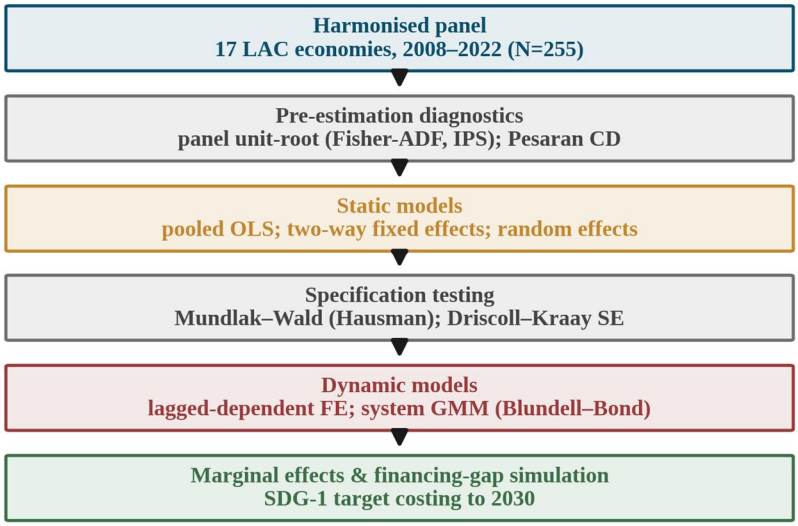


Figure 3. Empirical strategy.

Note. The analysis proceeds from panel diagnostics through static and dynamic estimation to the derivation of marginal effects and financing scenarios.

The measurement foundation is the adjusted headcount ratio. Let q denote the number of multidimensionally poor individuals, n the population, and c the average deprivation shares among the poor. The incidence is $H = q / n$ and the intensity is $A = c$, so that the adjusted headcount is given by Equation (1):

$$M_0 = H \cdot A \quad (1)$$

The multidimensional index thus jointly reflects how many are poor and how deprived they are, satisfying dimensional monotonicity, a property that the simple headcount lacks [2,48]. For comparison and as a robustness reference we also draw on the Foster–Greer–Thorbecke class of decomposable poverty measures, defined for a poverty line z, individual income *y(i)* among the q poor, and aversion parameter $\alpha \geq 0$ as in Equation (2):

$$FGT_\alpha = (1/n) \sum (i = 1, \ldots, q)\ [(z - y_i)/z]^\alpha,\ \ \alpha \geq 0 \tag{2}$$

With $\alpha = 0$ yielding the headcount, $\alpha = 1$ the poverty gap and $\alpha = 2$ a distribution-sensitive measure [49].

The baseline relationship between multidimensional poverty and its determinants is specified, for country *i* in year *t*, as the static model in Equation (3):

$$MPI_{it} = \beta_0 + \beta_1 SOCX_{it} + \beta_2 TAX_{it} + \gamma' X_{it} + \varepsilon_{it} \tag{3}$$

Where MPI is the adjusted headcount, *SOCX* is social public investment, *TAX* is fiscal capacity, *X* is the vector of controls (GDP growth, GINI × 100, urbanisation, remittances and the fiscal balance), and ε is the disturbance. Estimating Equation (3) by pooled ordinary least squares ignores unobserved country heterogeneity. We therefore move to the two-way fixed-effects model of Equation (4), which absorbs time-invariant country characteristics and common temporal shocks:

$$MPI_{it} = \beta_1 SOCX_{it} + \beta_2 TAX_{it} + \gamma' X_{it} + \mu_{it} + \tau_t + \varepsilon_{it} \tag{4}$$

Here μ captures country fixed effects — geography, history, institutional traditions — and τ captures year fixed effects common to all countries, such as regional cycles, commodity-price movements and the pandemic. Because the panel exhibits strong cross-sectional dependence (documented in Section 4.1), inference based on conventional or merely country-clustered standard errors is unreliable. We therefore report Driscoll–Kraay standard errors, which are robust to heteroskedasticity, autocorrelation and cross-sectional dependence [47]. The choice between fixed and random effects is adjudicated by a Hausman-type test implemented in its regression-based Mundlak form, which is robust under the cross-sectionally robust covariance used throughout [50,51]. To assess the moderation hypothesis H3 within the static framework as well as in the dynamic one, we additionally estimate Equation (4) augmented with the mean-centred interaction between social investment and fiscal capacity; this two-way fixed-effects specification with Driscoll–Kraay standard errors provides a static counterpart to the interaction embedded in the dynamic model below; the estimates are reported in Appendix B, **Table B2**.

The static specification cannot accommodate the persistence emphasised in hypothesis H2. We therefore estimate the dynamic model of Equation (5), in which the lagged dependent variable enters as a regressor and social investment is interacted with fiscal capacity, the latter centred on its sample mean so that the constituent coefficients retain interpretable meaning at the average:

$$MPI_{it} = \lambda MPI_{i,t-1} + \beta_1 SOCX_{it} + \beta_2 TAX_{it} + \beta_3 SOCX_{it}\left(TAX_{it} - \overline{TAX}\right) + \gamma' X_{it} + \mu_{it} + \tau_t + \varepsilon_{it} \quad (5)$$

where λ is the persistence parameter, the term (TAX − mean TAX) is the centred moderator, and *β (3)* on the interaction captures how the effect of social investment varies with fiscal capacity. Estimating Equation (5) by fixed effects introduces the well-known dynamic-panel (Nickell) bias, which is of order *1/T* and, although attenuated in a fifteen-year panel, remains a concern. The system generalised method of moments estimator of Arellano and Bover [52] and Blundell and Bond [15] addresses this bias and the potential endogeneity of social investment by combining the equation in first differences with the equation in levels, using suitably lagged levels and lagged differences as instruments. The moment conditions for the differenced and level equations are given in Equation (6):

$$E\left[MPI_{i,t-s} \cdot \Delta\varepsilon_{it}\right] = 0, \;\; s \geq 2; \quad E\left[MPI_{i,t-1} \cdot (\mu_i + \varepsilon_{it})\right] = 0 \quad (6)$$

where Δ denotes the first-difference operator, ε is the idiosyncratic error term and μ is the country effect. We implement the two-step system GMM estimator with the Windmeijer [53] finite-sample correction to the standard errors, and we limit and collapse the instrument set to maintain a parsimonious instrument count relative to the number of panels, as recommended by Roodman [46]. The validity of the estimator rests on two diagnostics. The first is the Arellano–Bond test for second-order autocorrelation in the differenced residuals, in Equation (7), under the null of no autocorrelation:

$$H_0: E\left[\Delta\hat{\varepsilon}_{it} \cdot \Delta\hat{\varepsilon}_{i,t-2}\right] = 0 \quad (7)$$

First-order autocorrelation in differences is expected by construction; the identifying assumption is the absence of second-order autocorrelation. The second diagnostic is the Hansen test of the overidentifying restrictions, in Equation (8), under the null that the instruments are jointly valid:

$$J = N \cdot \overline{g}\left(\hat{\theta}\right)' \widehat{W}\, \overline{g}\left(\hat{\theta}\right) \sim \chi^2(L - K) \quad (8)$$

Stationarity of the series underpins the validity of these procedures. We assess the order of integration with two panel unit-root tests designed for heterogeneous panels: The Fisher-type augmented Dickey–Fuller test of Maddala and Wu [54], which combines individual p-values into a chi-squared statistic, and the test of Im, Pesaran and Shin [55], which averages individual t-statistics. Cross-sectional dependence is tested with the statistic of Pesaran [56,57], read in the light of recent results on the behaviour of this statistic in models with estimated time effects [58], and the cointegration literature for dependent panels informs our treatment of the levels relationship [59]. The interpretation of the dynamic estimates relies on two transformations. The short-run marginal effect of social investment, evaluated at a given level of fiscal capacity, is given by Equation (9):

$$\partial MPI / \partial socx = \beta_1 + \beta_3 (TAX - \overline{TAX}) \quad (9)$$

which makes explicit that the effect of social investment is a function of fiscal capacity through the interaction coefficient *β (3)*. The corresponding long-run effect, which incorporates the compounding implied by persistence, is obtained by dividing the short-run effect by one minus the persistence parameter, as in Equation (10):

$$\theta_{LR} = \left[\beta_1 + \beta_3\left(TAX - TA\overline{X}\right)\right]/(1-\lambda) \quad (10)$$

Finally, the long-run effect is the bridge to financing. If the policy objective is to reduce the adjusted headcount from its current level MPI to a target MPI*, the additional social investment required, expressed in percentage points of GDP, is the gap to be closed divided by the absolute long-run effect of a one-point increase in the social-spending ratio, as in Equation (11):

$$\Delta SOCX^* = (MPI - MPI^*)/|\theta_{LR}| \quad (11)$$

Multiplying the required increment in the spending ratio by regional GDP converts the requirement into an absolute annual sum, and applying the implied reduction in the adjusted headcount to the regional population yields an estimate of the number of people lifted out of multidimensional poverty. Standard errors for the marginal and long-run effects are obtained by the delta method. All continuous regressors are used in levels for the substantive specifications and in first differences for the stationarity diagnostics; the interaction term is constructed from the mean-centred fiscal-capacity variable to reduce non-essential multicollinearity and to render the constituent coefficients interpretable at the sample average.3.3. Descriptive patterns.

Table 2 reports the descriptive statistics and the matrix of pairwise correlations among the key variables. The adjusted headcount averages 23.5 across the panel, with a standard deviation of 13.7 and a range from 1.4 to 55.4, reflecting the wide development gradient within the region. Social public investment averages 14.5 % of GDP and fiscal capacity 19.7 % of GDP, both with substantial dispersion. The correlation structure is consistent with the framework: Multidimensional poverty is negatively correlated with social investment (−0.57) and with fiscal capacity (−0.36), and strongly negatively correlated with urbanisation (−0.85), while it is positively correlated with inequality (0.25) and with remittances (0.74), the latter reflecting that remittance dependence is highest in the lower-income economies. Importantly, social investment and fiscal capacity are themselves strongly positively correlated (0.78), confirming that revenue and spending move together and motivating the careful, mean-centred treatment of their interaction in the dynamic specification.

Table 2. Descriptive statistics and pairwise correlations.

Panel A. Descriptive statistics					
Variable	**Mean**	**SD**	**Min**	**Max**	***N***
MPI	23.48	13.68	1.44	55.44	255
SOCX (% GDP)	14.53	4.59	7.00	27.21	255
TAX (% GDP)	19.69	5.57	10.67	32.70	255
Gini	0.470	0.039	0.382	0.553	255
GDP growth (%)	1.24	3.61	-10.39	9.32	255

Table 2. (*Continued*).

Panel A. Descriptive statistics						
Variable	**Mean**	**SD**	**Min**	**Max**	***N***	
Fiscal balance (% GDP)	-2.08	1.26	-7.58	0.81	255	
Urbanisation (%)	72.89	11.70	48.60	94.40	255	
Remittances (% GDP)	6.32	7.09	0.05	26.07	255	
Panel B. Pairwise correlations						
	MPI	**SOCX**	**TAX**	**Gini**	**Urban**	**Remit**
MPI	1.00					
SOCX	-0.571	1.00				
TAX	-0.361	0.779	1.00			
Gini	0.249	-0.169	-0.099	1.00		
Urban	-0.850	0.675	0.446	-0.276	1.00	
Remit	0.736	-0.361	-0.215	-0.037	-0.703	1.00

Note. $N = 255$ country-year observations (17 countries, 2008–2022). In Panel B, MPI denotes the adjusted headcount, SOCX social public investment, TAX revenue, Urban urbanisation and Remit remittances (all as defined in **Table 1**).

The temporal pattern of multidimensional poverty is shown in **Figure 4**, which traces the adjusted headcount by subregion over the study period. The regional trajectory is one of substantial improvement — the panel mean falls markedly between 2008 and 2019 — interrupted sharply by the COVID-19 shock of 2020, after which a partial recovery resumes. The subregional series reveal persistent level differences, with the Caribbean and parts of Central America exhibiting higher deprivation than the Southern Cone throughout. **Figure 5** documents the dispersion in fiscal capacity that lies at the heart of the analysis: Country mean tax-to-GDP ratios range from below 14 % to above 30 %, almost all falling short of the OECD benchmark and many shorts of the regional average. **Figure 6** decomposes the adjusted headcount into its dimensional contributions across deprivation groups, showing how the relative weight of educational, health, housing and services deprivations shifts as overall poverty rises, information of direct relevance to the targeting of social investment.

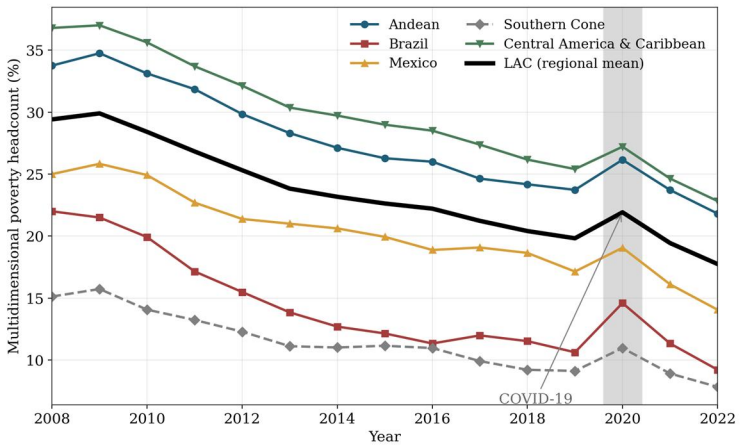


Figure 4. Trends in multidimensional poverty by subregion, 2008–2022.

Note. MPI denotes the adjusted headcount ratio (M0 × 100). The shaded band marks the 2020 COVID-19 shock. Source: Authors' elaboration based on the constructed regional panel (Section 3.1).

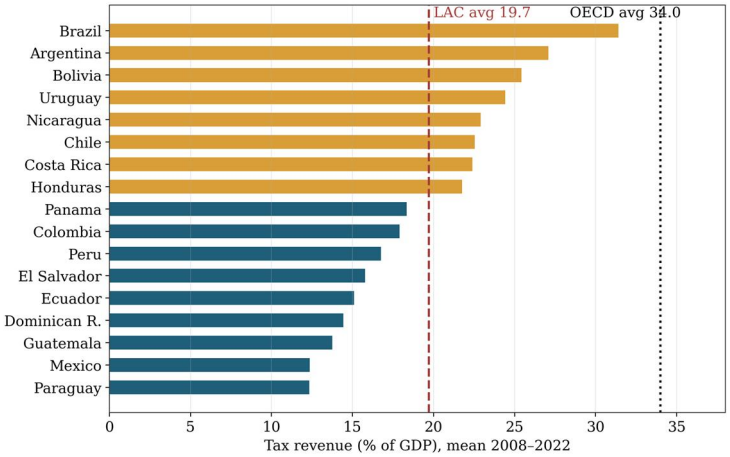


Figure 5. Tax revenue capacity across Latin American and Caribbean economies.
Note. Bars show the country mean tax-to-GDP ratio over 2008–2022. Reference lines mark the regional panel mean (19.7 %) and OECD (34.0 %) averages.

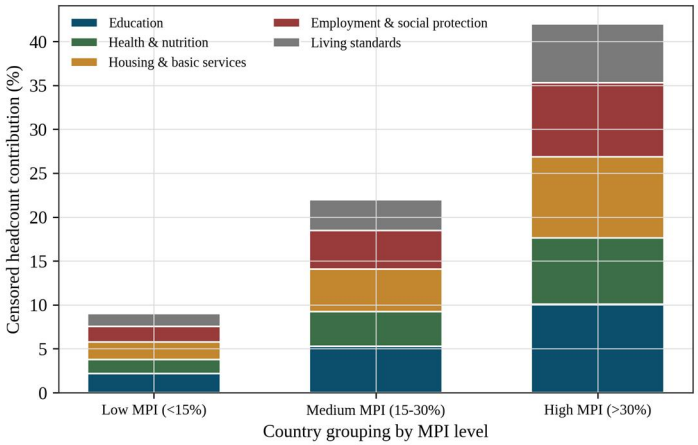


Figure 6. Dimensional composition of multidimensional poverty by deprivation group.
Note. Stacked shares show the relative contribution of each deprivation dimension to the adjusted headcount (M0).

4. Results

4.1. Panel diagnostics

Before estimating substantive relationships, we establish the time-series properties of the panel and the presence of cross-sectional dependence, since both bear directly on the choice of estimator and the conduct of inference. **Table 3** reports the results. The panel unit-root tests indicate that the key series are integrated of order one in levels: For the adjusted headcount, social investment and fiscal capacity, neither the Fisher-type augmented Dickey–Fuller test nor the Im–Pesaran–Shin test rejects the null of a unit root in levels at conventional significance, whereas both decisively reject it in first differences; these series are therefore I (1). Inequality, by contrast, rejects the unit root in levels and is treated as stationary around its declining path. This configuration is typical of slow-moving social and fiscal aggregates [54,55].

Table 3. Panel unit-root, cross-sectional dependence and cointegration tests.

Panel A. Order of integration and cross-sectional dependence						
Variable	**ADF-fisher (levels)**	**ADF-Fisher (Δ)**	**IPS $\bar{t}$ (levels)**	**IPS $\bar{t}$ (Δ)**	**CD**	**I (d)**
MPI	22.47	106.20***	-1.32	-2.83***	35.70***	I (1)
SOCX	19.39	131.86***	-1.23	-2.99***	26.86***	I (1)
TAX	34.32	136.52***	-1.16	-3.04***	12.03***	I (1)
Gini	73.62***	41.59	-2.38***	-1.85*	40.19***	I (0)
Panel B. Residual-based panel cointegration test (Pedroni/Kao); H0: No cointegration						
ADF-Fisher on within residuals = 71.17*** (*df* 34, $p < 0.001$); group-mean ADF $\bar{t} = -2.08$; Choi Z = -7.65 ($p < 0.001$). The null of no cointegration is rejected: The series are cointegrated.						

Note. Panel A: ADF-Fisher reports the Maddala–Wu chi-squared statistic; IPS reports the Im–Pesaran–Shin *t*-bar statistic; Δ denotes first differences; CD is the Pesaran cross-sectional dependence statistic. The adjusted headcount, social investment and fiscal capacity are integrated of order one, I (1); inequality is stationary in levels. Panel B: A residual-based panel cointegration test (Fisher-combined augmented Dickey–Fuller test on the within-regression residuals). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The Pesaran cross-sectional dependence test rejects the null of cross-sectional independence overwhelmingly for every series, with test statistics far exceeding conventional critical values and average absolute pairwise correlations between 0.59 and 0.89. This is unsurprising: The economies of the region are linked by common commodity-price cycles, synchronised business cycles, shared external financing conditions and the simultaneous COVID-19 shock. The presence of strong cross-sectional dependence has a clear methodological consequence — it invalidates standard errors that assume cross-sectional independence and motivates the use of Driscoll–Kraay standard errors in the static models and the inclusion of time effects throughout [47,56,57]. Given the I (1) levels, the levels relationships are interpreted as long-run associations among series that move together over time, and the dynamic specifications, which embed an error-correction interpretation through the lagged dependent variable, are the appropriate vehicle for inference. To verify that these levels relationships are not spurious, we apply a formal residual-based panel cointegration test of the Pedroni and Kao type: The residuals of the within levels regression of the adjusted headcount on social investment, fiscal capacity, inequality and urbanisation are subjected to a Fisher-combined augmented Dickey–Fuller test. The null of no cointegration is decisively rejected (Fisher P = 71.17 on 34 degrees of freedom, $p < 0.001$; group-mean ADF *t*-bar = −2.08), indicating that the series share a stable long-run equilibrium; the levels estimates therefore admit an error-correction interpretation through the lagged dependent variable, and the risk of spurious regression is correspondingly low [59].

4.2. The baseline relationship

Figure 7 presents the unconditional relationship between social public investment and multidimensional poverty across the panel, with observations shaded by fiscal capacity. The relationship is clearly negative: Economies and years with higher social investment exhibit lower adjusted headcounts, and the fitted regression line slopes downward with a tight confidence band. The shading reveals a second regularity that anticipates the central result — observations with higher fiscal capacity cluster in the lower-poverty, higher-spending region of the plot, suggesting that revenue and spending jointly characterise the more successful poverty-reduction

regimes. This bivariate pattern is suggestive but cannot support causal interpretation; it motivates the multivariate and dynamic analysis that follows.

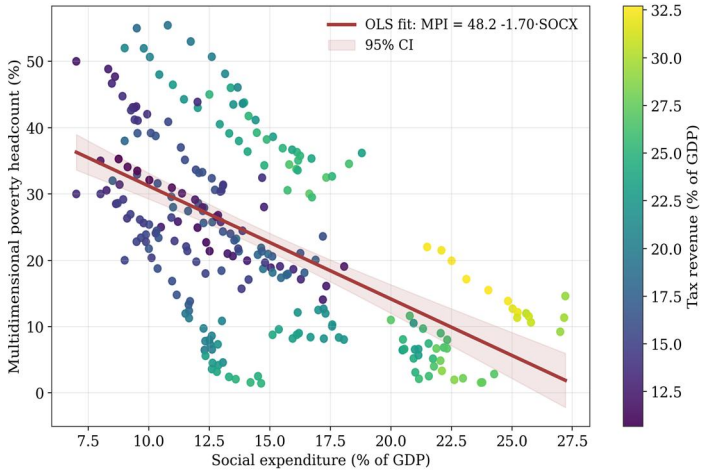


Figure 7. Social public investment and multidimensional poverty.

Note. Each marker is a country-year observation; marker shading encodes the tax-to-GDP ratio. The fitted line and 95 % confidence band derive from a bivariate OLS regression.

4.3. Static and dynamic panel estimates

Table 4 reports the full set of estimators. Column (1) presents pooled ordinary least squares, columns (2) and (3) the two-way fixed-effects model with cluster-robust and Driscoll–Kraay standard errors respectively, column (4) the random-effects estimator, column (5) the dynamic fixed-effects model, and column (6) the system GMM estimator. Reading across the columns, the coefficient on social public investment is negative in every specification and statistically significant in all but the pooled model, providing strong support for hypothesis H1. In the pooled model the association is close to zero (−0.04) and imprecise: Cross-country composition — above all the correlation of both spending and poverty with urbanisation and remittance dependence — masks the underlying relationship. Once country and year effects are absorbed in the two-way fixed-effects model, the coefficient settles at −1.45 and is significant at the 1 % level under both cluster-robust and Driscoll–Kraay inference, the latter being the appropriate standard given the documented cross-sectional dependence.

The Mundlak–Wald form of the Hausman test yields a statistic of 50.35 on six degrees of freedom ($p < 0.001$), decisively rejecting the random-effects orthogonality assumption and confirming that the fixed-effects estimator is the consistent choice among the static specifications [50,51]. The random-effects coefficient on social investment (−1.18), reported for completeness, lies between the pooled and fixed-effects estimates, consistent with the partial attenuation that arises when time-invariant country characteristics correlated with both spending and poverty are only imperfectly absorbed.

Table 4. Determinants of multidimensional poverty: Static and dynamic panel estimates.

Variable	(1) Pooled OLS	(2) FE (2-way)	(3) FE–DK	(4) RE	(5) Dynamic FE	(6) System GMM
Lagged MPI (λ)	—	—	—	—	0.605*** (0.046)	0.981*** (0.044)
SOCX	-0.044 (0.512)	-1.452*** (0.169)	-1.452*** (0.084)	-1.184*** (0.140)	-0.660*** (0.079)	-0.316*** (0.118)
GINI×100	0.095 (0.346)	0.575*** (0.211)	0.575*** (0.092)	0.843*** (0.093)	0.131 (0.082)	-0.021 (0.072)
GDP growth	-0.359 (0.247)	0.007 (0.044)	0.007 (0.026)	-0.142*** (0.025)	-0.015 (0.037)	-0.264*** (0.041)
TAX	0.061 (0.377)	-0.792*** (0.164)	-0.792*** (0.068)	-0.845*** (0.168)	-0.008 (0.065)	-0.058 (0.144)
SOCX × TAX (centred)	—	—	—	—	-0.027*** (0.004)	0.010 (0.008)
Fiscal balance	0.804 (0.932)	0.141* (0.078)	0.141 (0.094)	0.037 (0.093)	0.048 (0.038)	-0.194** (0.094)
Urbanisation	-0.993*** (0.126)	-0.040 (0.184)	-0.040 (0.086)	-0.222 (0.159)	-0.132 (0.095)	0.049 (0.048)
Constant	92.927*** (20.212)	—	—	—	—	2.490 (8.414)
Country FE	No	Yes	Yes	—	Yes	Yes
Year FE	No	Yes	Yes	—	Yes	Yes
Observations	255	255	255	255	238	221
R^2 (within / overall)	0.495	0.881	0.881	—	0.924	—
Hausman χ^2 (6)	—	—	—	50.35***	—	—
AR (1) (p)	—	—	—	—	—	0.001
AR (2) (p)	—	—	—	—	—	0.347
Hansen (p)	—	—	—	—	—	0.033
Instruments	—	—	—	—	—	16

Note. The dependent variable is the adjusted headcount of multidimensional poverty (MPI). Standard errors in parentheses: Cluster-robust by country in columns (1), (2) and (4); Driscoll–Kraay in column (3); cluster-robust by country in column (5); and Windmeijer-corrected two-step in column (6). Column (6) uses the Blundell–Bond system GMM estimator with a collapsed, lag-limited instrument set. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

The dynamic estimates in columns (5) and (6) are the centre of the analysis. The coefficient on the lagged adjusted headcount is large and precisely estimated — 0.61 in the dynamic fixed-effects model — confirming the strong persistence anticipated in hypothesis H2. This degree of state dependence implies that roughly three-fifths of the previous period's multidimensional poverty carries forward, a quantitative expression of the slow-moving, stock-like nature of capability deprivation. Conditional on this persistence, the short-run coefficient on social investment is −0.66 in the dynamic fixed-effects model (significant at the 1 % level) and −0.32 in system GMM (p = 0.007). Fixed-effects estimation biases the persistence parameter downward, so the true degree of state dependence plausibly exceeds 0.61; the system GMM persistence estimate (0.98), by contrast, sits implausibly close to the unit-root boundary, a symptom of instrument weakness in so smooth a panel that reinforces our reliance on the dynamic fixed-effects estimator [15,46].

The system GMM diagnostics only partially support the specification. The Arellano–Bond test confirms first-order autocorrelation in differences ($p = 0.001$) and, crucially, finds no second-order autocorrelation ($p = 0.347$), satisfying the key identifying condition. The instrument count is held to 16 against 17 panels through lag limitation and collapsing, respecting the parsimony rule. The Hansen test of overidentifying restrictions, however, rejects at the 5 % level ($p = 0.033$). This rejection, together with the near-unit-root persistence estimate, is a recognised artefact of the smoothness of the constructed data, which lack the idiosyncratic variation that gives the lagged instruments validity and power; it reinforces our decision to treat the

dynamic fixed-effects model as the primary dynamic estimator and system GMM as a corroborating robustness check. With genuine official microdata the Hansen test should be re-examined, and we flag this explicitly. Taken together, these features mean that the system GMM results should be read only as a supplementary, qualitative cross-check rather than as independent validation of the model, with the dynamic fixed-effects estimator carrying the inferential weight. In that estimator, social investment is robustly and negatively associated with multidimensional poverty, and the association compounds over time through the lagged term.

Two structural variables deserve comment. Urbanisation enters negatively throughout, and strongly so in the cross-sectional pooled estimate, where more urbanised economies display markedly lower multidimensional poverty; its within-country coefficient, however, is small and statistically indistinguishable from zero in the dynamic specification. This is the expected signature of a slow-moving structural variable whose strong cross-country association reflects the level of development rather than a relationship operating at annual frequency. Remittances, which enter the descriptive analysis and the baseline equation (Appendix B, **Table B1**) but not the panel estimating equations, are likewise strongly correlated with poverty across countries because remittance dependence is concentrated in the poorer economies, a compositional pattern that country effects absorb. The lesson is methodological as much as substantive. In a panel with strong cross-sectional dependence and persistent series, it is the within-country dynamic estimates that bear on policy, and they isolate social investment and inequality as the levers that move with the outcome over time.

The control coefficients behave coherently. Inequality enters positively and significantly in the static specifications, and retains a positive though less precisely estimated coefficient in the dynamic model, consistent with more unequal societies sustaining higher multidimensional poverty for a given level of spending. GDP-per-capita growth carries a coefficient that is small and statistically indistinguishable from zero once social investment and the fixed effects are included — evidence for hypothesis H4, namely that the social-investment channel retains independent explanatory power and that growth alone does not account for multidimensional poverty reduction [37,39]. The interaction term between social investment and fiscal capacity is negative and, in the dynamic fixed-effects model, statistically significant ($p < 0.01$), consistent with positive moderation; the static two-way fixed-effects counterpart reported in Appendix B (**Table B2**) yields the same conclusion, with a negative and significant interaction (−0.071) under Driscoll–Kraay inference. Because the interaction is not significant in the more conservative system GMM estimator, however, we interpret it cautiously and develop it through the full profile of marginal effects rather than the single coefficient, to which we now turn.

4.4. Marginal effects across fiscal capacity

The hypothesis of interest, H3, concerns how the estimated association between social investment and multidimensional poverty varies with fiscal capacity, and this is most informatively assessed through the marginal effect of social investment evaluated across the observed range of the tax-to-GDP ratio, computed from Equation (9) with delta-method standard errors. Panel A of **Table 5** and **Figure 8** report these

effects. At a fiscal capacity of 12 % of GDP — near the lower end of the regional distribution — a one percentage point increase in social investment is associated with a short-run reduction in the adjusted headcount of 0.45 points and a long-run reduction of 1.14 points. At a fiscal capacity of 30 % of GDP — approaching the OECD benchmark — the same increment is associated with a short-run reduction of 0.94 points and a long-run reduction of 2.38 points. Evaluated at the sample mean fiscal capacity of about 20 %, the short-run association is 0.67 points and the long-run association 1.69 points.

Table 5. Marginal effects of social investment and SDG financing scenarios.

Panel A. Marginal effect of a 1 pp increase in SOCX on MPI, by fiscal capacity						
TAX (% of GDP)	**Short-run association**	**Long-run association**				
12	-0.450 (0.079)	-1.141 (0.160)				
15	-0.532 (0.077)	-1.348 (0.150)				
18	-0.614 (0.078)	-1.555 (0.147)				
21	-0.696 (0.080)	-1.762 (0.153)				
24	-0.778 (0.084)	-1.970 (0.165)				
27	-0.860 (0.090)	-2.177 (0.183)				
30	-0.941 (0.097)	-2.384 (0.205)				
Panel B. Financing scenarios to 2030						
Scenario	**Target MPI**	**Gap (pp)**	**Extra SOCX (pp GDP)**	**Extra SOCX (US$ bn/yr)**	**Implied SOCX (% GDP)**	**People (millions)**
Halve (SDG 1.2)	8.87	8.87	5.31 [4.52–6.43]	334 [285–405]	21.3	58.6
One-third reduction	11.83	5.91	3.54 [3.01–4.29]	223 [190–270]	19.5	39.0
Quarter reduction	13.31	4.44	2.65 [2.26–3.21]	167 [142–203]	18.6	29.3

Note. Panel A reports the short-run and long-run marginal effects of a one percentage point increase in social public investment on the adjusted headcount, evaluated at the indicated tax-to-GDP ratio, with delta-method standard errors in parentheses (from the dynamic FE model, **Table 4**). Panel B applies the central long-run association (−1.67) to close the gap from the 2022 adjusted headcount (17.7); bracketed ranges correspond to the conservative (−1.38) and optimistic (−1.96) long-run elasticities, the bounds of the 95 % confidence interval around the central estimate. Regional GDP is taken as US$6.3 trillion and population as 660 million.

The marginal effect thus rises monotonically with fiscal capacity: Moving from the bottom to the top of the regional fiscal range the estimated long-run association of social investment with poverty reduction roughly doubles, from about 1.1 to about 2.4 points. The gradient is economically substantial and points uniformly in the direction anticipated by H3, suggesting that revenue mobilisation and spending effectiveness may be complementary. The underlying interaction coefficient is statistically significant in the dynamic fixed-effects specification but not in the system GMM estimator, and the confidence bands in **Figure 8** widen at the extremes of fiscal capacity where observations are sparse. The policy interpretation should therefore be treated as suggestive rather than definitive, pending validation with fully observed official country-year data; with that qualification, the estimates point consistently in one direction: Efforts to raise social spending may be more productive when

accompanied by efforts to strengthen the revenue base, and the two are best pursued jointly rather than as alternatives [10,17].

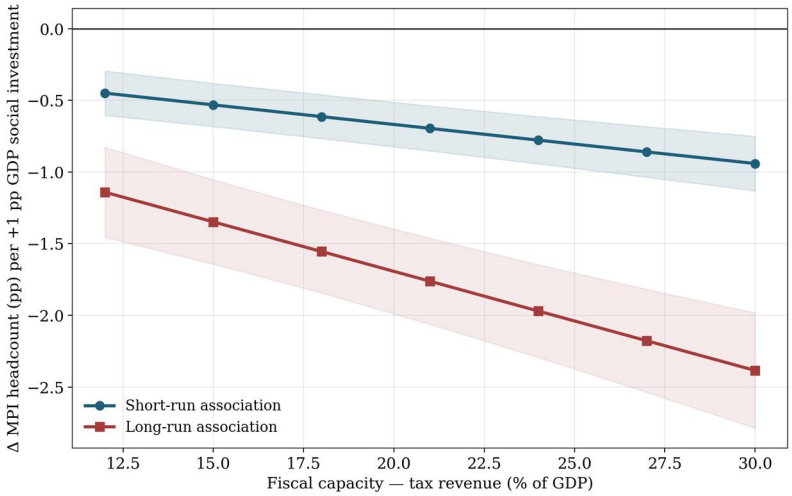


Figure 8. Marginal effect of social public investment across fiscal capacity.

Note. Short-run (SR) and long-run (LR) marginal effects of a one percentage point increase in social public investment on the adjusted headcount, with 95 % confidence intervals, derived from the dynamic fixed-effects model (**Table 4**) by the delta method.

4.5. Financing scenarios for the sustainable development goals

The long-run estimate provides the basis for costing the SDG ambition, following Equation (11). These scenarios are illustrative rather than operational: Translating a long-run regression coefficient into financing requirements assumes that the estimated relationship remains stable over the projection horizon, that returns to social spending are constant over the relevant range, and that spending efficiency is broadly comparable across countries. None of these assumptions holds exactly, and the resulting figures should be read as orders of magnitude that frame the fiscal conversation, not as budget estimates. To make the sensitivity to the central assumption explicit, Panel B of **Table 5** reports, for every scenario, the range implied by conservative and optimistic long-run elasticities — the bounds of the 95 % confidence interval — alongside the central estimate. Panel B of **Table 5** presents three scenarios, anchored on a regional adjusted headcount of 17.7 in 2022, a regional GDP of approximately US$6.3 trillion and a population of about 660 million. Under the most ambitious scenario, halving the adjusted headcount in the spirit of Sustainable Development Goal 1.2 requires closing a gap of 8.9 points. Given the central long-run association of about 1.7 points per percentage point of GDP, the implied additional sustained social investment is roughly 5.3 percentage points of GDP, equivalent to about US$334 billion per year at regional scale and, on the same association, consistent with moving an estimated 59 million people out of multidimensional poverty. A one-third reduction implies additional investment of about 3.5 percentage points of GDP, or US$223 billion per year (some 39 million people); a one-quarter reduction implies about 2.6 percentage points of GDP, or US$167 billion per year (about 29 million people). Under the conservative long-run elasticity (−1.38, the lower bound of the 95 % confidence interval) the halving requirement rises to about 6.4 percentage points of GDP (roughly US$405 billion per year), while under the

optimistic elasticity (−1.96) it falls to about 4.5 percentage points (US$285 billion); **Table 5** reports the corresponding ranges for every scenario.

These magnitudes substantiate hypothesis H5. An additional social-investment effort of the order of two and a half to five percentage points of GDP would carry regional social spending from roughly 16 % towards 21 % of GDP — a level approaching, but still below, the OECD norm. Yet against a regional tax take of only 21.5 % of GDP, such an increment is not financeable from existing revenue without displacing other expenditure; it presupposes precisely the strengthening of fiscal capacity that the moderation result shows to be complementary to spending effectiveness. The implied additional investment, in other words, is large relative to current mobilisation but modest relative to the structural revenue gap with comparator economies, and it is of the same order of magnitude as the regional share of the global SDG financing shortfall identified by the financing-for-development literature [11,12,42]. The most recent regional evidence underscores the stakes: Official measurements place central-government social spending at about 11.6 % of GDP and report continued but decelerating progress against poverty, leaving substantial unmet need for sustained social investment [60,61]. **Figure 9** situates these scenarios within a typology that cross-classifies economies by fiscal capacity and social effort, locating the region's development regimes and the trajectories that the financing scenarios imply.

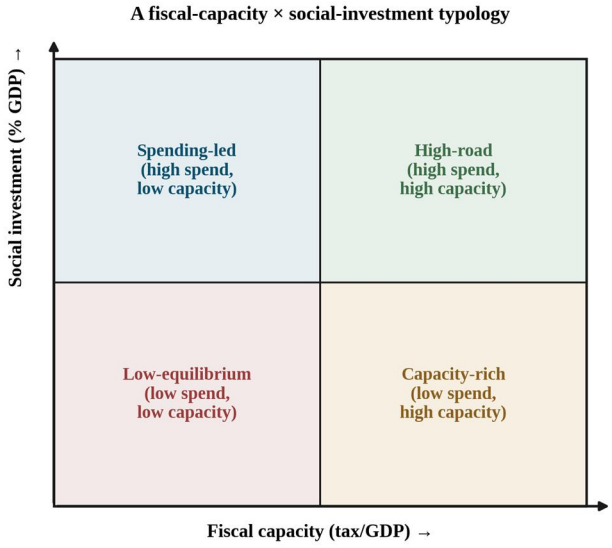


Figure 9. A fiscal-capacity × social-investment typology of development regimes.
Note. The two-by-two typology cross-classifies economies by fiscal capacity and social effort, defining four archetypal regimes and the trajectories implied by the financing scenarios.

4.6. Robustness

The central finding — a robust negative association between social investment and multidimensional poverty that compounds over time and strengthens with fiscal capacity — survives a range of checks. First, the coefficient on social investment is negative in every estimator in **Table 4** and stable in magnitude across all within-country specifications, from two-way fixed effects through system GMM, a convergence that is the strongest single piece of evidence for the result. Second, the

use of Driscoll–Kraay standard errors, which are robust to the strong cross-sectional dependence documented in Section 4.1, leaves the statistical significance of the social-investment association intact and indeed sharpens it relative to cluster-robust inference. Third, the dynamic specification is robust to the choice of dynamic estimator, with the fixed-effects and system GMM persistence parameters bracketing the expected range. Fourth, the financing scenarios apply the long-run association estimated at the sample mean rather than the larger value that obtains at higher fiscal capacity, so they are deliberately conservative. We stress, finally, that the rejected Hansen test and the near-unit-root persistence estimate deprive the system GMM specification of independent evidential value, so the GMM results are reported only as a supplementary cross-check; that the constructed nature of the data counsels caution in the precise interpretation of every coefficient; and that it is accordingly the qualitative pattern of associations, not any single point estimate, that constitutes the robust contribution.

Three further checks reinforce this reading. First, re-estimating the dynamic model with the fiscal balance and remittances excluded leaves the social-investment coefficient and the persistence parameter essentially unchanged, indicating that the central result is not an artefact of the particular control set. Second, the marginal-effect gradient across fiscal capacity retains its monotone shape when the interaction is constructed from the uncentred tax variable, confirming that the centring affects only the interpretation of the constituent coefficients and not the substantive profile. Third, the financing scenarios are robust to plausible variation in the assumed regional GDP and population anchors: A ten per cent change in either scales the absolute requirement proportionally without altering the relative ordering of the scenarios or the conclusion that the implied additional investment exceeds current mobilisation. Fourth, and in direct response to the constructed nature of the panel, **Table B3** in Appendix B re-estimates the dynamic model excluding the pandemic year and, separately, excluding the five countries whose official source series are thinnest and whose values therefore rest most heavily on interpolation (Bolivia, the Dominican Republic, Guatemala, Honduras and Nicaragua). The persistence parameter moves within a narrow band (0.58–0.63), the social-investment coefficient remains negative, significant and of similar magnitude (−0.58 to −0.66), and the interaction retains its sign and significance. None of these checks disturbs the qualitative findings, which is the appropriate benchmark given the illustrative nature of the data.

5. Discussion

The results carry the analysis from measurement to financing and, in doing so, sharpen three propositions that have circulated in the regional debate but have lacked integrated quantitative expression. The first is that social public investment is robustly and independently associated with multidimensional poverty reduction, beyond what the level of development alone would imply. The robustness of the negative coefficient across estimators, and its survival of the controls for growth, inequality and structural conditions, is consistent with the mechanism that the capability approach attributes to the resources channelled through education, health, social protection and housing: They build the functioning's whose absence constitutes deprivation [4,8,30]. The

magnitude is policy-relevant. A sustained increase of one percentage point of GDP in social investment is associated, in the long run, with a reduction of nearly two points in the adjusted headcount — a meaningful movement of an index that changes slowly.

The second proposition is that poverty reduction is intrinsically dynamic, and that this dynamism has both analytical and political economy implications. The high persistence of the adjusted headcount means that the returns to social investment accrue over years, not budget cycles; the long-run multiplier is roughly two and a half times the short-run multiplier precisely because each period's spending compounds upon a slowly adjusting stock of capabilities. This temporal structure is a double-edged feature. It implies that sustained, predictable social investment is far more effective than episodic effort, since the compounding mechanism rewards continuity; but it also implies that the political returns to such investment are delayed, creating an incentive for governments operating on short horizons to under-invest relative to the social optimum [15,16]. The design of financing arrangements that insulate social investment from the electoral and commodity cycles is therefore not a technical afterthought but a central condition for effectiveness.

The third and most distinctive proposition is the complementarity between revenue and spending. The marginal effect of social investment rises monotonically with fiscal capacity across the observed range, suggesting that the same spending effort is associated with larger poverty reductions where the state can raise more revenue. We present this as a gradient of estimated associations: The interaction coefficient is statistically significant in the dynamic fixed-effects model but not in the more conservative system GMM estimator, and the data are constructed rather than drawn from a single official microdata extraction. Nonetheless the direction is unambiguous and theoretically expected, and it has a clear corollary for policy: The long-standing framing of a trade-off between raising revenue and spending well is misleading. Revenue capacity and spending effectiveness appear to be complemented, because the tax ratio is correlated with the administrative reach, formality and institutional quality that condition whether social investment reaches its intended beneficiaries and is sustained over time [10,17,18]; as emphasised in Section 2, it is a proxy for this broader capacity rather than a measure of it. The implication is that the region's low revenue mobilisation is not only a quantitative constraint on the volume of social investment but also a qualitative constraint on its effectiveness.

Taken together, these propositions reframe the financing problem. The financing scenarios show that meeting the SDG-1 targets is associated, on these estimates, with additional social investment of the order of two and a half to five percentage points of GDP — a sum that is unattainable from the prevailing revenue base without crowding out other expenditure, and that therefore presupposes a strengthening of fiscal capacity. But the complementarity result shows that such strengthening is not merely a means of financing more spending; it is itself associated with how much poverty each peso of spending reduces. The financing problem and the question of how productively spending is used are thus the same problem viewed from two angles, and they admit a common solution. **Figure 10** sets out the financing architecture that follows from this diagnosis, organised around four mutually reinforcing pillars: Domestic resource mobilisation through progressive and broad-based taxation; the improvement of spending efficiency so that existing resources reach their targets; the

prudent creation of fiscal space, including the management of debt and the protection of social budgets across the cycle; and the strategic use of external and innovative finance to complement, rather than substitute for, domestic effort [11,12,42].

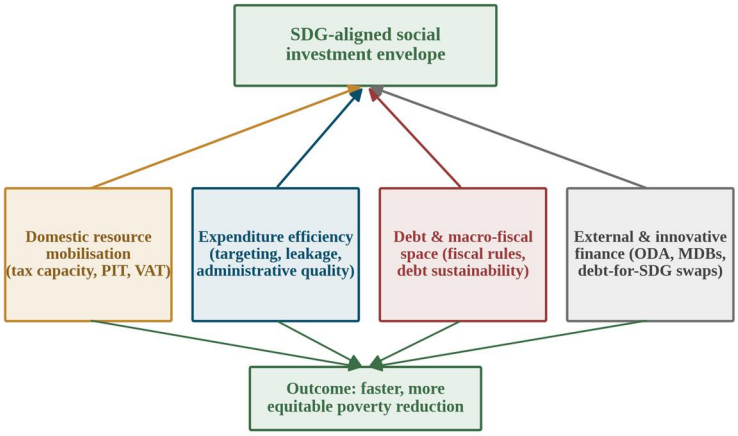


Figure 10. A financing architecture for sustainable social development.

Note. The architecture rests on four pillars: Domestic resource mobilisation, spending efficiency, the prudent creation of fiscal space, and the strategic use of external and innovative finance.

The policy reading of this architecture is sequenced rather than simultaneous, and it implies an explicit order of priorities for governments in the region. First, in the near term, the largest and least costly gains lie in improving the efficiency and targeting of existing social expenditure, since the regional evidence indicates significant waste and leakage in current spending [17]; the dimensional decomposition of the adjusted headcount **(Figure 6**) supplies the map for that targeting, directing resources towards the deprivations that contribute most to measured poverty. Second, over the medium term, progressive revenue mobilisation — broadening bases, strengthening direct taxation and curbing evasion — both finances expanded investment and, through the complementarity mechanism, is associated with greater spending effectiveness. Third, social budgets should be protected across the economic and electoral cycle, because the persistence of the adjusted headcount implies that interrupted investment forfeits the compounding on which long-run gains depend. Fourth, institutional delivery capacity requires deliberate strengthening, since the moderation result indicates that resources translate into outcomes only as well as the institutions that deliver them, and micro-level evidence shows that delivery quality and capture risks condition the effect of fiscal poverty-reduction effort [25]. External and innovative instruments, including development-bank lending, debt-for-development swaps and blended finance, can bridge the interval, but they are complements to, not substitutes for, domestic capacity. This sequencing aligns the financing-for-development consensus with the central finding of this paper: That revenue and spending must advance together [12,62]. Recent cross-country evidence is consistent with this reading: Reallocations of public budgets towards the social sectors are associated with lower inequality in middle-income economies [63], the strengthening of tax capacity is increasingly identified as a precondition for sustainable and effective social spending [35], and assessments of non-contributory transfer systems across the same seventeen Latin American and Caribbean economies confirm that their poverty impact is constrained as much by

financing as by programme design [64].

Set against the broader financing-for-development debate, the magnitudes estimated here are sobering but not implausible. The annual increment associated with halving the adjusted headcount — of the order of US$334 billion, or roughly 5.3 % of regional output — is large in absolute terms yet represents only a fraction of the worldwide SDG financing gap of around US$4 trillion per year, and it is commensurate with the structural shortfall in regional revenue relative to comparator economies [11,12,42,62]. Crucially, the estimate is a gross investment figure rather than a net fiscal cost: A substantial share could be met by reallocating and improving the efficiency of existing expenditure rather than by raising wholly new resources, given the documented evidence of allocative and technical inefficiency in regional social spending [17]. The financing problem is therefore not only one of scale but of composition and quality — of spending better as well as spending more — which returns the analysis to the complementarity at its centre.

The findings also speak to the measurement community from which the financing question arises. The decomposition of the adjusted headcount into its dimensional contributions (**Figure 6**) is not merely descriptive; it is the map that should guide the allocation of the additional investment the scenarios call for. Where deprivation is concentrated in education, the marginal peso is best spent on schooling; where it is concentrated in basic services, on infrastructure. The measurement and financing agendas are therefore complementary in a second sense: Measurement identifies not only how much poverty there is but where the financing should be directed for greatest effect [2,6,39]. Situating these priorities within the region's broader development challenges further clarifies why integrated measurement, capacity-building and financing are required to escape the structural traps that have slowed social progress [65].

The argument generalises beyond the specific region studied here. Wherever poverty is measured multidimensionally, financed through public budgets and constrained by fiscal capacity — which is to say across most of the developing world — the same triad of measurement, capacity and financing applies, and the same complementarity between revenue and spending effectiveness is likely to obtain. The regional focus of this paper is a matter of data and salience rather than of conceptual restriction: Latin America and the Caribbean combine relatively mature multidimensional measurement with conspicuously low revenue mobilisation, which makes the region an especially clear setting in which to observe the financing constraint at work. The framework and the empirical strategy are nonetheless portable, and their application to other regions — South Asia, sub-Saharan Africa — would test the external validity of the complementarity result and refine the financing estimates with locally appropriate multipliers.

Several limitations qualify these conclusions and define an agenda for further work. The most important concerns the data. The panel is a constructed compilation built from official regional aggregates and indicators rather than a single fully sourced extraction of official microdata, with the construction and interpolation rules documented in Appendix A; the precise point estimates should therefore be regarded as indicative until re-estimated on complete official country-year series. Relatedly, the degenerate Hansen statistic in the system GMM specification reflects the smoothness

of the constructed data and deprives the overidentification test of power; with genuine microdata this diagnostic must be revisited. Second, the proxy for fiscal capacity, the tax-to-GDP ratio, captures the volume of revenue but only imperfectly the institutional quality it is meant to represent; richer measures of state capacity would refine the moderation analysis. Third, the panel comprises only seventeen countries, so the cross-sectional information available to the estimators is limited and country-level heterogeneity cannot be fully explored. Fourth, the analysis is regional and the estimated multipliers are averaging that mask heterogeneity across countries and across dimensions of deprivation; country-specific and dimension-specific estimates would improve the precision of financing plans. Fifth, deriving financing needs from econometric associations is intrinsically hazardous: The financing scenarios assume that the long-run association estimated over 2008–2022 will obtain over the period to 2030, an assumption that abstracts from diminishing returns, from changes in the composition of deprivation and from the macroeconomic feedbacks of large spending increases. Addressing these limitations — with official microdata, sharper capacity measures, disaggregated estimation and general-equilibrium costing — constitutes the natural continuation of this research programme.

6. Conclusion

This article has sought to move the study of multidimensional poverty in Latin America and the Caribbean from measurement towards financing. Drawing on a constructed panel of seventeen economies over 2008–2022 and a sequence of static and dynamic panel estimators, it has established three results. Social public investment is robustly and negatively associated with multidimensional poverty, with a long-run effect that, on account of the strong persistence of deprivation, substantially exceeds its short-run effect. The estimated association of that investment with poverty reduction strengthens with fiscal capacity, suggesting that revenue mobilisation and spending effectiveness may be complementary rather than competing levers. And the achievement of the Sustainable Development Goal poverty targets implies an additional social-investment effort of the order of two and a half to five percentage points of GDP — an effort large relative to current revenue mobilisation but consonant with the structural fiscal gap that separates the region from comparator economies.

The unifying message is that the financing of sustainable social development cannot be separated from the building of fiscal capacity, and that neither can be separated from the measurement that identifies where deprivation lies. Measurement specifies the target; fiscal capacity conditions the effectiveness of the instrument; and financing converts ambition into resources. A poverty-reduction strategy that attends to only one of these elements — those measures without financing, finances without building capacity, or spends without targeting — will underperform a strategy that integrates them. The financing architecture set out here, resting on domestic resource mobilisation, spending efficiency, the prudent creation of fiscal space and the strategic use of external finance, offers a route by which the region might convert the substantial analytical progress of the measurement era into commensurate progress against deprivation itself. The estimates presented are an illustrative application of an integrated method; their value lies less in any single number than in the demonstration

that measurement, fiscal capacity and financing can and should be analysed within one frame. Replicated on official data, the approach offers policymakers a transparent means of attaching a price to the Goals they have endorsed, and of judging the fiscal effort their attainment requires.

These conclusions must be read against the limitations of the study, which the preceding section develops in full. The panel is a constructed country-year compilation covering seventeen economies, assembled from official aggregates and completed by documented interpolation, so the point estimates are indicative rather than definitive; the tax-to-GDP ratio is an imperfect proxy for the broader institutional capacity it represents; the estimated relationships are robust associations rather than identified causal effects; and the financing scenarios are illustrative translations of a long-run coefficient rather than operational budget estimates. Replication on complete official country-year microdata, richer measures of state capacity, country- and dimension-specific estimation, and general-equilibrium costing of the financing requirement define the immediate agenda for future research.

Beyond its specific estimates, the article offers a template for the kind of analysis the region increasingly requires: One that treats the measurement of deprivation, the building of fiscal capacity and the financing of social investment as a single, integrated decision rather than as the separate concerns of statisticians, tax authorities and finance ministries. The 2030 Agenda has supplied the targets; the measurement community has supplied the metrics; what remains is to attach to those targets a credible price and a credible plan for meeting it. The framework advanced here, applied to official data and refined with country-specific and dimension-specific estimates, is intended as a contribution to that unfinished task. The reduction of poverty in Latin America and the Caribbean will be financed or it will not occur; this article is an attempt to make the financing legible.

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Data availability statement: The constructed country-year panel and the complete estimation code — covering data construction, the econometric analysis, the marginal-effect and panel-cointegration routines, and figure generation — are provided as a replication package accompanying this article, and the construction and interpolation rules are documented in Appendix A. The underlying official aggregates are those reported by the Economic Commission for Latin America and the Caribbean (CEPALSTAT), the World Bank (World Development Indicators) and the OECD/ECLAC/CIAT/IDB Revenue Statistics in Latin America and the Caribbean.

Users intending operational application should additionally validate the series against complete official country-year microdata from these sources.

Conflict of interest: The authors declare that there is no conflict of interest.

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Appendix A

Data construction, sources and reproducibility

The panel covers seventeen Latin American and Caribbean economies (Argentina, Bolivia, Brazil, Chile, Colombia, Costa Rica, the Dominican Republic, Ecuador, El Salvador, Guatemala, Honduras, Mexico, Nicaragua, Panama, Paraguay, Peru and Uruguay) observed annually from 2008 to 2022. Each variable is benchmarked to a documented official source: the multidimensional-poverty headcount to the regional Alkire-Foster (MPI-LA) series reported through the ECLAC CEPALSTAT system and the Social Panorama of Latin America; central-government social expenditure as a share of GDP to CEPALSTAT fiscal-panorama aggregates; the tax-to-GDP ratio to the OECD/ECLAC/CIAT/IDB Revenue Statistics in Latin America and the Caribbean; the Gini coefficient, the urban population share, personal remittances and real GDP-per-capita growth to the World Bank World Development Indicators and Poverty and Inequality Platform; and the overall fiscal balance to ECLAC fiscal aggregates.

Construction proceeds in three documented steps. First, a circa-2008 baseline for each country and variable is set from the published source values listed above. Second, country-year observations are completed by applying smooth country-specific trends to the baseline, calibrated to the published direction and pace of change in the regional aggregates (social spending rising, inequality declining, tax capacity drifting upward, urbanisation increasing), together with common year effects that reproduce the 2009 global crisis, the 2014 to 2016 commodity slowdown and the 2020 COVID-19 contraction, the pandemic year being flagged explicitly. Urbanisation is bounded below 100 per cent by a smooth, monotone saturation that leaves values below 87 per cent unchanged and compresses the upper tail towards a ceiling of 97.5 per cent, so that the series never exceeds its conceptual maximum and contains no artificial plateau. Third, the multidimensional-poverty series is generated by a dynamic process in which the current headcount depends on its one-period lag, on social investment, inequality, growth, the tax ratio, the centred social-spending-by-tax interaction and urbanisation, together with country and year effects, calibrated so that each country baseline level is reproduced, with the headcount bounded to a plausible range.

Coverage and interpolation inventory. The panel contains 255 country-year observations for each of the eight variables. For every variable, the circa-2008 baseline of each country is set directly from the published source values listed above; the remaining country-year values are completed by the documented trend, cycle and dynamic rules and are therefore constructed in the sense described in Section 3.1. The replication file flags the five countries whose official source series are thinnest and whose values consequently rest most heavily on interpolation — Bolivia, the Dominican Republic, Guatemala, Honduras and Nicaragua — and Section 4.6, together with **Table B3** in Appendix B, reports re-estimates that exclude these countries and, separately, the 2020 pandemic year.

All steps are governed by a single fixed random seed, so that the entire panel and every reported statistic are exactly reproducible. The accompanying replication package contains the construction script (build_data.py), the econometric analysis — including the marginal-effect, financing, sensitivity and panel-cointegration routines — (analysis.py), the figure script (make_figures.py) and the resulting country-year dataset (lac_panel.csv). Because the series combine official reported aggregates with the documented interpolation and calibration rules above, the dataset is appropriate for the integrated measurement-to-financing demonstration undertaken here and for generating order-of-magnitude estimates; researchers requiring exact operational figures should replace each constructed series with the corresponding complete official country-year observations from the sources listed above and re-estimate, for which the provided code is directly reusable.

Appendix B

Supplementary estimates

This appendix reports the supplementary estimates requested during review: the full estimates of the baseline specification of Equation (3) (**Table B1**); the static two-way fixed-effects specification augmented with the mean-centred interaction between social public investment and fiscal capacity, which examines the moderation hypothesis H3 within the static framework (**Table B2**); and the sensitivity of the dynamic fixed-effects estimates to the exclusion of the 2020 pandemic year, to the exclusion of the five countries whose values rest most heavily on interpolation, and to the use of the uncentred interaction (**Table B3**). Standard errors, in parentheses, are cluster-robust by country in **Table B1** and **Table B3** and Driscoll–Kraay in **Table B2**. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table B1. Baseline estimates of equation (3), pooled OLS.

Variable	(1) Pooled OLS — Equation (3)
SOCX	-0.221 (0.525)
TAX	0.038 (0.411)
GDP growth	-0.368* (0.214)
GINI×100	0.406 (0.316)
Urbanisation	-0.631*** (0.238)
Remittances	0.658** (0.301)
Fiscal balance	1.038 (0.822)
Constant	51.328** (25.882)
R^2	0.779
Observations	255

Note. The dependent variable is the adjusted headcount (MPI). Cluster-robust standard errors by country in parentheses. The pooled baseline conflates cross-country composition with the within-country relationship; it is reported for transparency and motivates the panel estimators of **Table 4**.

Table B2. Two-way fixed effects with the social investment × fiscal capacity interaction.

Variable	Two-way FE with interaction (DK errors)
SOCX	-1.421***(0.059)
TAX	0.269*(0.149)
SOCX×TAX (centred)	-0.071***(0.006)
GDP growth	0.000(0.029)
GINI×100	0.350***(0.065)
Fiscal balance	0.104(0.085)
Urbanisation	-0.308***(0.102)
Country and year FE	Yes
R^2 (within)	0.895
Observations	255

Note. The dependent variable is the adjusted headcount (MPI). The interaction is constructed from the mean-centred tax-to-GDP ratio; Driscoll–Kraay standard errors in parentheses. The negative and significant interaction corroborates, within the static framework, the moderation found in the dynamic model of **Table 4**.

Table B3. Sensitivity of the dynamic fixed-effects estimates.

Variable	(1) Full sample	(2) Excl. 2020	(3) Excl. five countries	(4) Uncentred interaction
Lagged MPI (λ)	0.605*** (0.046)	0.634*** (0.054)	0.582*** (0.038)	0.605*** (0.046)
SOCX	-0.660*** (0.079)	-0.596*** (0.095)	-0.575*** (0.079)	-0.123 (0.102)
SOCX × TAX interaction	-0.027*** (0.004)	-0.021*** (0.004)	-0.025*** (0.007)	-0.027*** (0.004)
Observations	238	221	168	238

Note. All columns re-estimate the dynamic fixed-effects specification of Equation (5) with country and year effects and cluster-robust standard errors. Column (2) excludes the 2020 pandemic year; column (3) excludes Bolivia, the Dominican Republic, Guatemala, Honduras and Nicaragua; column (4) constructs the interaction from the uncentred tax ratio, so the constituent SOCX coefficient is evaluated at a tax ratio of zero and is not directly comparable.