Religious Charity, Commitment, and Secularization

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Abstract

This paper studies whether denominational charitable organizations accelerate secularization by raising the cost of religious participation. Exploiting the 1854 Dutch Poor Law (*Armenwet*), which granted religious charities a statutory monopoly on poor relief, I examine how associational density shaped religious affiliation across Dutch municipalities between 1879 and 1930. To address endogeneity, I instrument for post-1854 charitable capacity using the geographic distribution of Catholic monasteries established before the Dutch Reformation of 1578, whose endowments were transferred to Protestant deaconries following the Revolt. Municipalities with greater exogenous charitable association density exhibit persistently higher shares of no religious affiliation and higher religious diversity in the Dutch decennial censuses. These findings are jointly consistent with a commitment-screening mechanism: dense associational networks impose organizational costs that sort out weakly attached affiliates, accelerating secularization, while the locally dominant confession sheds its marginal members proportionally fastest, rebalancing the residual affiliated population toward parity between denominations and raising diversity.

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1 Introduction

Over the past century, Western societies have undergone a dramatic and largely irreversible decline in religious participation. Church attendance has collapsed, denominational membership has fallen, and the share of the population reporting no religious affiliation has risen sharply across Western Europe and, more recently, the United States (Norris and Inglehart, 2004; Voas and Chaves, 2016). Yet beneath this aggregate trend lies a striking puzzle of cross-national heterogeneity. France entered the twentieth century as a largely dechristianized republic, with anti-clerical sentiment institutionalized in the 1905 law on the separation of church and state. The Netherlands, by contrast, maintained near-universal religious affiliation well into the mid-twentieth century, with the share of the population recorded as *geen gezindte* (no religious affiliation) remaining below five percent in 1930. Beginning in the 1960s, Dutch society then underwent one of the most rapid episodes of *ontkerkelijking* (de-churching) in European history. Standard models of religious markets (Iannaccone, 1991, 1998) and cross-country determinants of religiosity (Barro and McCleary, 2003, 2006a) provide a useful framework but struggle to account for such sharp heterogeneity in timing and level. This paper argues that a key piece of the explanation lies in a largely overlooked channel: the institutional bundling of welfare access with religious affiliation.

We study the Netherlands between 1854 and the early twentieth century, a period in which the 1854 *Armenwet* (Poor Law) gave religious charities a statutory monopoly on poor relief. Under this law, municipal authorities were legally permitted to provide secular supplementary assistance only in cases where existing religious charities demonstrably lacked the capacity to meet local need. The consequence was an institutional landscape in which denominational charitable organizations, ranging from poor-relief associations (*armenzorg*) to insurance funds (*verzekeringsfondsen*) and registered welfare societies (*erkende verenigingen*), became the primary infrastructure of local social welfare provision. We study whether the density of these associations, which required active membership, regular financial contributions, and behavioral compliance from their members, affected the subsequent trajectory of religious affiliation and religious diversity. Our main finding runs contrary to the simple welfare-bundling prediction: municipalities with greater exogenous charitable association capacity exhibit persistently *higher* rates of secularization, measured by the share of the population recorded as having no religious affiliation (*geen gezindte*) in the Dutch decennial censuses of 1879–1930, and higher religious diversity. We interpret this as a commitment-screening effect that, in these municipalities, outweighs the countervailing pull of the welfare services bundled with membership: dense associational networks raise the net cost of affiliation and sort out the weakly attached, accelerating secularization. Religious diversity rises at the same time because the locally dominant confession organises most intensively and therefore bears the steepest participation cost; as the more heavily screened denomination it sheds its weakly-attached members proportionally faster than the minority, rebalancing the residual affiliated population toward parity between confessions.

The empirical challenge is that religious charity capacity is endogenous: wealthier or more devout municipalities may have invested more in charitable infrastructure for reasons correlated with the contemporaneous demand for religion. We address this by exploiting the spatial distribution of monasteries and convents established before the Dutch Reformation of 1578 as an instrument for post-1854 charity capacity. Medieval monastic foundations were located according to feudal land grants, ecclesiastical politics, and pilgrimage routes, factors that are, we argue, plausibly orthogonal to the forces driving secularization three centuries later. When the Protestant Republic absorbed the southern and eastern provinces after the revolt against Spain, Catholic monasteries were dissolved; their endowments, comprising land, buildings, and institutional expertise in poor relief, were transferred to Protestant deaconries (*diaconieën*), which subsequently became the primary providers of charitable assistance under the *Armenwet* regime. The geographic distribution of pre-Reformation monasteries thus provides a source of variation in post-1854 charity capacity that is, by construction, predetermined relative to the secularization outcomes we study. This identification strategy is analogous in spirit to Cantoni et al. (2018), who use the distribution of dissolved monasteries in the Protestant Reformation to identify the effect of religious market restructuring on secular outcomes in the German lands.

Our main finding is that greater pre-Reformation monastery density, instrumented for religious charitable association density, is associated with significantly *higher* shares of non-affiliated population and *higher* religious diversity in the census years 1879 through 1930. The secularization effect grows substantially over time and is amplified in IV relative to OLS, consistent with a causal interpretation: by the early twentieth century, a one-standard-deviation increase in association density is associated with a rise in the non-affiliated share on the order of a couple of percentage points, a sizeable shift relative to contemporaneous levels, with the effect growing substantially toward the end of the period. Simultaneously, religious diversity (measured by one minus the Herfindahl-Hirschman Index across denominations) rises, indicating that the remaining affiliated population becomes *less* confessionally concentrated as associational density rises. These two findings are jointly consistent with the commitment-screening mechanism formalized in Appendix B.

This paper makes three contributions. First, we provide new causal evidence that the density of denominational civil society organizations accelerates secularization by screening out low-commitment affiliates. The mechanism operates through participation costs rather than welfare bundling: associations impose organizational requirements that are prohibitive for households with weak religious conviction, sorting the nominally affiliated into the unaffiliated category. This connects the economics of religion literature on club goods and strict churches (Iannaccone, 1994; Berman, 2000) to the sociological literature on nominal versus committed religious identification (Hout and Fischer, 2002, 2014). Second, we offer a unified account of two co-occurring patterns, higher secularization and higher religious diversity in association-dense municipalities, through a single commitment-screening mechanism formalized in Appendix B: the same participation-cost screen that pushes the weakly committed out of religion altogether falls hardest, in relative terms, on the dominant confession, rebalancing the affiliated population toward parity. The long-run prediction is consistent with Gill and Lundsgårde (2004), who document that welfare-state generosity is associated with lower church attendance: once secular alternatives arrive, the pool of weakly committed members retained by high-cost associational networks rapidly exits. Third, we assemble a new municipality-level panel dataset linking pre-Reformation monastery locations, historical charity association density, and census-based religious affiliation for the Netherlands between 1879 and 1930.

Our paper speaks to several strands of literature. The economics of religion literature has, since Azzi and Ehrenberg (1975) and Iannaccone (1991, 1992), modeled religious participation as the outcome of rational household choices over spiritual and material benefits. A central result is that competitive religious markets generate higher aggregate participation than monopolistic ones (Iannaccone, 1991; Stark and Finke, 2000). We introduce a distinct mechanism: organizational density raises the effective cost of affiliation through participation requirements, sorting the population into committed affiliates and unaffiliated exits, and—because the cost falls most heavily on the dominant, most intensively organised confession—rebalancing the remaining affiliated population away from it, raising diversity. Our model formalizes this as an extension of Iannaccone's 1994 strict-church framework, in which the screen is not costly sacrifice per se but the organizational overhead of belonging to a confessional civil society under the *Armenwet* regime. Within the literature on religion and historical development (surveyed by Becker et al., 2021), the closest antecedents are Becker and Woessmann (2009), who show that the Protestant Reformation affected human capital accumulation through religious literacy requirements, and Rubin (2014), who traces the role of printing technology in facilitating the Reformation's diffusion. Finally, on the determinants of cross-country variation in religiosity, Barro and McCleary (2003) and Barro and McCleary (2006b) document that government regulation of religion and the strength of state churches are among the strongest predictors of participation; our paper provides a within-country, causally identified analog of this finding, tracing variation at the municipal level to a specific organizational mechanism.

The paper most closely related to ours is Cantoni et al. (2018). Several aspects separate our contribution. First, our measure of secularization is based on individual-level religious disaffiliation: the share of the population recording no religious affiliation in the Dutch census.[1] Our census-based measure is more precise and behaviorally exact: it registers the actual decision of households to exit organized religion, rather than inferring secularization from aggregate resource flows. Second, Cantoni et al.'s framework operates

[1] Cantoni et al. measure economic secularization as the reallocation of human capital and fixed investment across sectors, with graduates shifting from theology to law and construction shifting from churches to palaces, which captures changes in the structure of economic activity but does not directly measure whether individuals abandon religious identity.

through a political economy channel: religious competition shifts the bargain between secular lords and religious elites, enabling rulers to expropriate monastic wealth and empowering secular relative to ecclesiastical authority, whereas our mechanism operates at the individual level: associational density raises the effective cost of religious membership through participation requirements, screening out weakly committed households regardless of elite politics or state capacity. Third, the long-run channel we identify complements their short-run political economy account.

The remainder of the paper is organized as follows. Section 2 describes the historical context, including the institutional history of the *Armenwet*, the confessional landscape of the nineteenth-century Netherlands, and the transfer of monastic endowments after the Reformation. Section 3 describes the data and the empirical strategy. Section 4 reports the main results and robustness checks. Section 5 concludes. A formal model of the commitment-screening mechanism is developed in Appendix B.

2 Historical Background

2.1 Religion and Society in the Nineteenth-Century Netherlands

At the onset of the nineteenth century, the Netherlands was a deeply religious yet denominationally fragmented society. The dominant confession was the *Nederlands Hervormde Kerk* (Dutch Reformed Church), a legacy of the Calvinist Reformation that had shaped the northern provinces since the late sixteenth century. Catholics constituted roughly 38 percent of the population and were concentrated in the southern provinces of Noord-Brabant and Limburg as well as scattered urban communities in the west. A small but significant Jewish minority was present in the major commercial cities (Knippenberg, 1992; Kossmann, 1978). This confessional geography, a Protestant north and a Catholic south with mixed urban pockets, was not merely a matter of personal belief (Knippenberg, 1998). Denominational identity structured social networks, political alignments, and above all access to charitable relief.

Welfare provision in this period was almost entirely private and religiously organized. The Batavian and French periods (1795–1813) had introduced moments of state intervention, but these remained short-lived. By the early nineteenth century, poor relief was distributed through a patchwork of church deaconries (*diaconieën*), Catholic parish funds, Jewish *parnassim*, and a small number of municipal poorhouses. The state lacked both the fiscal capacity and the ideological mandate to displace this denominational infrastructure.

2.2 Medieval Monasteries and the Reformation Settlement

The religious geography of charitable provision had deep medieval roots. Before the Dutch Revolt, Catholic monasteries and convents were the principal institutions of poor relief across the Low Countries (Lis and Soly, 1979; Spaans, 1997). These institutions held substantial endowments, comprising land, annuities, and movable wealth accumulated over centuries, which financed hospitals, orphanages, and outdoor relief. Their distribution followed the logic of medieval ecclesiastical and feudal politics: proximity to trade routes, episcopal centers, and the patronage networks of noble houses.

The Revolt and subsequent Calvinist ascendancy dismantled this infrastructure but did not destroy it. Following the *Alteratie* of Amsterdam in 1578 and the progressive consolidation of Calvinist control in the northern provinces, monastic properties were confiscated and their endowments transferred to the newly established Protestant deaconries (Gorski, 1993). The 1580s and 1590s saw systematic conversions of monastery assets into the capital base of *diaconieën* across Holland, Zeeland, and Utrecht. The result was a Protestant charitable infrastructure whose geographic extent and resource base derived, paradoxically, from the medieval Catholic settlement. Municipalities that had hosted monasteries before 1578 found themselves, two centuries later, endowed with unusually well-resourced Protestant deaconries, a legacy that would prove consequential when the state reorganized poor relief in 1854.

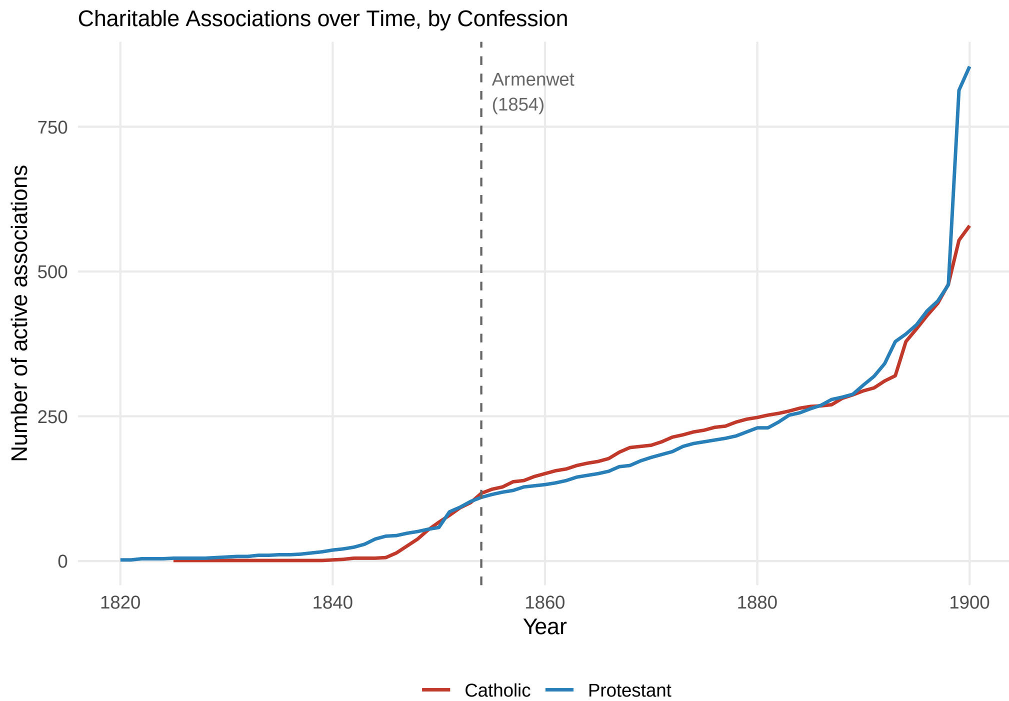
2.3 The *Armenwet* of 1854

The 1854 Poor Law (*Armenwet*) was the decisive institutional intervention of the period. Enacted under the liberal-confessional compromise of Thorbecke's constitutional framework, the law formalized a strict hierarchy of charitable responsibility: the family bore primary obligation; failing that, religious charities (*kerkelijke armenzorg*) held statutory priority; only when both were exhausted could municipal or state relief be invoked (Melief, 1955; van Leeuwen, 2000). This subsidiarity principle effectively conferred a monopoly on poor relief to religious organizations and gave denominations a powerful organizational incentive to expand their associational networks.

Figure 2.1 plots the number of active Catholic and Protestant charitable associations by year from 1820 to 1900. Through the 1820s and 1830s both series are flat and low, but they begin to climb in the mid-1840s—already before the *Armenwet*—and continue to rise smoothly and in parallel for the rest of the century, accelerating sharply after about 1885; by 1900 Protestant associations exceed 800 and Catholic associations approach 600, with both series still climbing. The growth of denominational welfare is thus a broad secular trend spanning the whole period rather than a discontinuity at 1854.[2]

[2]This is not a difficulty for what follows. The empirical strategy does not rest on a time-series break at the year of the law: it exploits *cross-sectional* variation in municipalities' inherited capacity to organize religious charity, instrumented by medieval monastic geography (Section 3.2). The role of the *Armenwet* in this design is institutional rather than chronological—by making religious charity the statutory channel for poor relief, it ensured that pre-existing differences in associational capacity across municipalities became economically consequential—and that role does not require the aggregate stock of associations to have jumped in 1854.

Figure 2.1: Charitable associations active per year, by confession, 1820–1900



2.4 Pillarization and the Welfare–Religion Nexus

The post-1854 expansion of denominational charity did not merely extend existing institutions; it embedded poor relief within the emerging system of social segmentation that historians have termed *verzuiling* (pillarization) (Lijphart, 1968; Righart, 1986; Kruijt and Goddijn, 1962). Each confession organized its population into parallel networks of associations covering welfare, education, labor, and eventually politics. Membership in these networks was not purely voluntary: in a world where charitable assistance required denominational affiliation, affiliation carried material stakes. Belonging to a confession meant access to its deaconry, its orphanage, its mutual aid fund.

This institutional bundling had an important implication for the dynamics of religious commitment. A dense associational network raised the effective cost of participation, not just Sunday attendance but involvement in committees, fundraising, and mutual surveillance, while simultaneously providing excludable benefits to those who remained inside. This is the mechanism at the heart of the present paper's argument. Where charitable associations were thicker on the ground, the religious market demanded more from its members, eventually sorting out the weakly attached. The same municipalities that saw rapid associational growth in the decades after 1854 would, by 1909 and 1930, exhibit relatively higher rates of *geen gezindte* (no stated religious affiliation) in the Dutch census (Knippenberg, 1998; van Rooden, 1996).

The variation in associational density that the *Armenwet* activated was itself, however, not random. It was shaped, in part, by the medieval monastic geography described above: municipalities that had inherited well-endowed Protestant deaconries from pre-1578 monasteries had a head start in building the organizational capacity the 1854 law rewarded. This historical accident provides the source of exogenous variation exploited in the empirical analysis below.

The theoretical model developed in Appendix B formalizes this institutional logic and shows that its observable implications are genuine predictions rather than built-in assumptions. Religious associations played two roles at once. As screening devices, they raised the cost of nominal affiliation, attending meetings, contributing financially, and complying with behavioral norms, and pushed weakly committed households toward the exit. But under the *Armenwet* they were also the principal providers of material welfare, so a denser network delivered a larger bundle of tangible services that worked to *retain* marginal members. Affiliation thus reflects a tension between a screening channel and a bundling channel, and the model's first prediction is that associational density raises the overall rate of disaffiliation only where the screening force dominates at the margin, above a threshold density that sufficiently rich networks exceed. A positive empirical relationship between density and *geen gezindte* therefore identifies the municipalities operating in this screening-dominant regime, while in thinly organized municipalities the welfare bundle is net-retentive.

The second prediction concerns the composition of those who remain: in the screening regime, denser networks raise the confessional concentration of the affiliated population. This requires one further, historically grounded ingredient. Because the dominant denomination operated a disproportionate share of local associations, its higher participation costs and thinner committed core would, on their own, let the minority *gain* ground as density rose. Concentration increases only because the minority's base is itself disproportionately weakly attached, sustained by *social-proximity joiners* drawn to the surrounding community rather than to its creed, so that its nominal adherents melt away proportionally faster as costs climb. Together the two propositions predict that municipalities with thicker organizational networks should exhibit both more secularization and a more confessionally homogeneous residual faithful, a joint pattern that is, on its face, paradoxical but follows from a single underlying commitment-screening logic.

3 Data and Empirical Strategy

3.1 Data

The analysis draws on five primary data sources matched at the municipality level using standard Dutch AMCO municipality codes. The unit of observation is the municipality in a given census year. All geographic identifiers are harmonized across sources using the AMCO coding scheme, which provides a consistent municipality classification for the

Netherlands across the nineteenth and early twentieth centuries.

Outcome variables. The main outcome variables are constructed from the *Historische Database Nederlandse Gemeenten* (HDNG), a comprehensive longitudinal dataset covering Dutch municipalities across census years from 1795 to 1971.[3] The HDNG compiles population and social statistics from decennial and quinquennial Dutch censuses, including municipality-level counts of residents by religious denomination: Roman Catholic, various Protestant denominations (Dutch Reformed, Re-Reformed, Lutheran, and others), and, from 1879 onward, those recording no religious affiliation (*geen gezindte*).

The primary outcome variable is the *share of the population reporting no religious affiliation* (*geen gezindte* share), measured in each census year. This variable first entered the Dutch census in 1879, providing a baseline measure of secularization at the onset of the period under study. Subsequent census rounds in 1889, 1899, 1909, 1920, and 1930 allow for a panel analysis of the trajectory of secularization across municipalities. As a secondary outcome, I construct a *religious diversity index* defined as one minus the Herfindahl–Hirschman index of denominational shares, $D_{it} = 1 - \sum_k s_{kit}^2$, where s_{kit} is the share of denomination k in municipality i at census year t. Higher values indicate greater religious pluralism.

Religious charity provision. The central explanatory variable, the intensity of religiously affiliated charitable activity in a municipality, is constructed by aggregating records from four historical databases curated by the Huygens Institute for Dutch History (KNAW). Each database was collected from the respective online repository and matched to AMCO municipality codes, as detailed in Appendix A.

Verzekeringsfondsen (Insurance and Mutual Aid Funds). This database records 1,520 mutual aid societies, occupational guild funds, burial societies, and sick funds active in the Netherlands between approximately 1587 and 1991. Each record contains the fund's name, municipality, province, and active date range. Although this source spans several centuries, its coverage is densest in the eighteenth and nineteenth centuries, capturing the proliferation of corporatist welfare arrangements that predated statutory social insurance. For the present analysis, I use the count of funds per municipality active before 1880 as a measure of pre-existing associational infrastructure.

Erkende Verenigingen (Recognized Associations, 1855–1903). Following the 1855 *Wet op de Verenigingen en Vergaderingen*, all civil-society associations were required to obtain official recognition by Royal Decree before operating. The resulting registry contains 8,984 associations formally recognized between 1855 and 1903. Each record includes the association's name, municipality, date of recognition, Royal Decree number, *Staatscourant* (official gazette) publication date, and founding date. This source provides near-complete

[3] Detailed provenance and construction procedures for all data sources used in this paper are documented in Appendix A; the full replication package is available at https://github.com/basm92/crcs.

coverage of the formal associational sector in the half-century immediately following the 1854 *Armenwet* and is used to identify the number of legally recognized charitable organizations per municipality in the post-*Armenwet* period.

Armenzorg (Poor Relief Associations). This database catalogs 2,584 organizations engaged in poor relief and poverty prevention active primarily between 1838 and 1899. Critically for identification, each record includes the ideological or religious orientation of the organization (neutral, Protestant, Catholic, or Jewish) alongside the founding year, municipality, target beneficiary group (the poor generally, women, youth, the sick, alcoholics), and, in many cases, the gender composition of the membership. This source most directly operationalizes the concept of *religious* charity provision: I construct, for each municipality, the count and share of poor-relief associations with an explicitly religious orientation.

Sociale Zekerheid / Lokale Instellingen (Local Social Welfare Institutions, 1899–1956). This database documents local institutions providing extramural (community-based, non-residential) poor relief and social assistance, as recorded in official charitable guides published across the period 1899–1956. Each institution record contains the institution's name, municipality, years of inclusion in the guide, a categorical affiliation code distinguishing civil/secular (*Burgerlijk*), church-affiliated (*Kerkelijk*), and private (*Particulier*) institutions, and notes on work methods (home visits, casework investigation). This source extends coverage of religious charitable activity into the early twentieth century and is used to construct municipality-level counts of religious versus secular welfare institutions.

Together, the four databases yield measures of the number, density, and religious composition of charitable organizations at the municipality level for the period 1838–1956. The main composite measure of *religious charity intensity* is the share of charitable organizations in a municipality with an explicitly religious affiliation, averaged across sources and weighted by organizational activity years.

Instrumental variables. To address the endogeneity of religious charity provision, namely that religious organizations may be more prevalent in municipalities where religiosity was already high, I exploit a source of exogenous variation in the capacity for religious charity provision.

Pre-Reformation monastery presence. The first instrument is an indicator (and count) of Catholic monasteries present in a municipality before the Dutch Reformation, conventionally dated to the consolidation of the Dutch Republic around 1578–1581. Monastery locations are drawn from the *Kloosterlijst* database.[4] The *Kloosterlijst* covers 762 verified monasteries, convents, and other religious communities in the Netherlands from the medieval period through 1800, including beguinages (*begijnhoven*), houses of the Common Life, mendicant friaries, and enclosed convents of all major orders. Each record contains

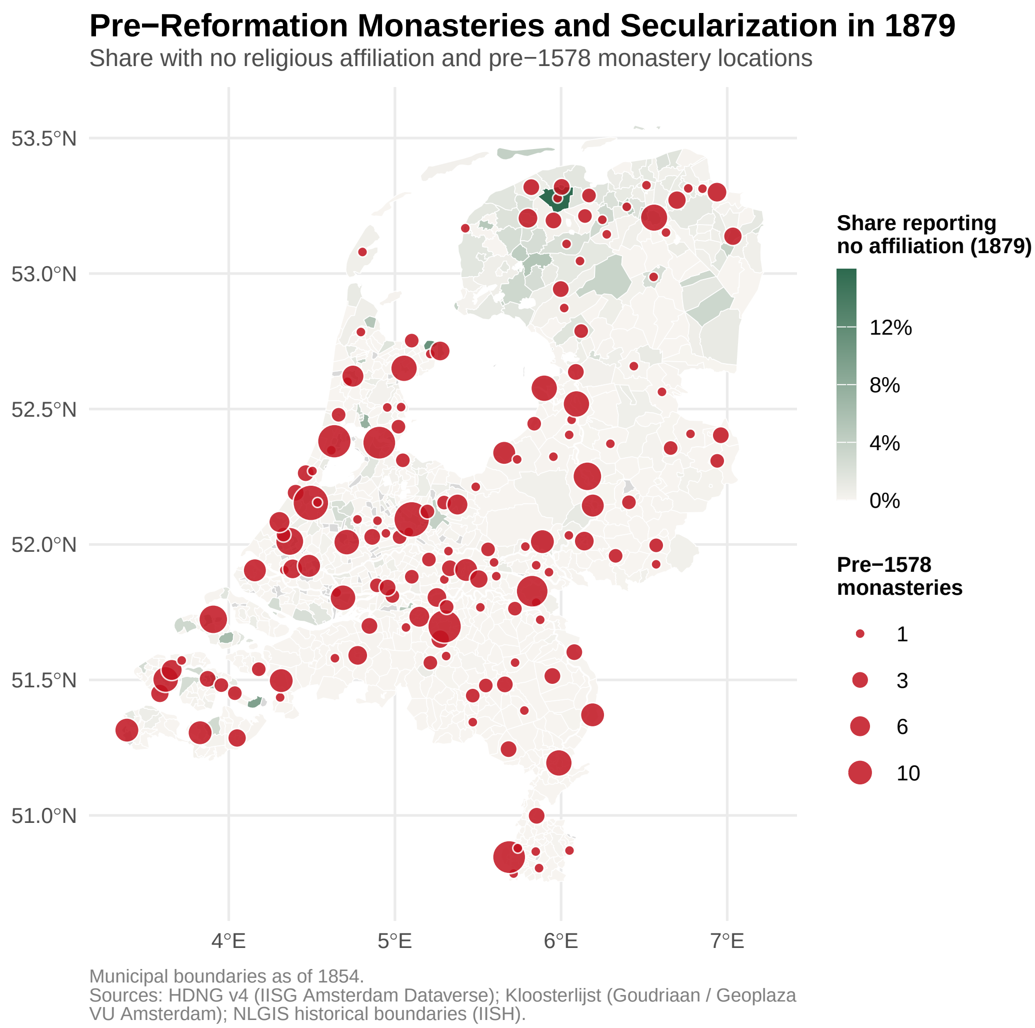
[4]The database was last updated in June 2021. It supersedes the earlier *Monasticon Batavum* of Schoengen (1941/2), correcting numerous errors and adopting the principle that each monastery receives a single record covering all stages of its institutional history. See Appendix A for details.

the monastery's name, historical parish and current municipality, diocese, province or territorial unit, religious order and gender composition, and a dated narrative of institutional stages (*stadia*), together with the year of first documentary mention (*eerste vermelding*) and the year of last mention or dissolution (*laatste vermelding*). The database provides an explicit elimination list (*eliminatielijst*) of institutions mentioned in the secondary literature that could not be confirmed in primary sources; only positively verified records are included in the main file used here. For the purposes of this analysis I restrict attention to institutions whose first documentary mention falls before 1579, constructing municipality-level counts and indicators of pre-Reformation monastic presence. This criterion captures all monasteries that were part of the pre-Reformation monastic landscape whose land, buildings, and institutional expertise in poor relief could plausibly have been transferred to Protestant deaconries following the Revolt—the core mechanism of the instrument. It yields 157 municipalities with at least one pre-1579 monastery, compared to 94 under a more restrictive definition requiring dissolution by 1578. The instrument's validity rests on the following logic: following the Reformation, Catholic monastic properties and their associated endowments were in many localities transferred to or inherited by Protestant deaconries (*diaconieën*), which became the primary institutional vehicle for poor relief under the Dutch Reformed Church. Municipalities with a pre-Reformation monastic presence therefore entered the nineteenth century with a larger endowment base for organized religious charity, independently of contemporaneous religious composition. Monasteries were sited according to medieval settlement patterns and ecclesiastical politics that are plausibly unrelated to the determinants of nineteenth-century secularization conditional on baseline municipality characteristics.

Figure 3.1 illustrates the spatial distribution of these monasteries overlaid by the level of religious non-affiliation in 1879. Pre-Reformation monasteries were concentrated in the urbanized western provinces (Holland, Utrecht, Zeeland) and in the Catholic south (Brabant, Limburg), reflecting patterns of medieval ecclesiastical settlement. The choropleth reveals that secularization by 1879 was most pronounced in the rapidly industrializing cities of the western Randstad and in parts of Drenthe, while the predominantly Catholic provinces of North Brabant and Limburg recorded near-zero rates of religious disaffiliation. The partial overlap between areas of high monastic density and early secularization is consistent with the proposed mechanism: a larger inherited endowment base for organized charity may have sustained religious institutional presence and thereby slowed disaffiliation, but this effect is moderated by the initial confessional composition of the population.

Control variables. All specifications include a vector of genuinely pre-treatment municipality-level controls $\mathbf{X}_i$ to reduce omitted-variable bias. The control set combines geographic, agricultural, and deep-historical variables that predate the Reformation-era instrument, following the variable-selection approach of [the comparable paper]. *Jewish population*

Figure 3.1


Pre−Reformation Monasteries and Secularization in 1879
Share with no religious affiliation and pre−1578 monastery locations
53.5°N
53.0°N
52.5°N
52.0°N
51.5°N
51.0°N
4°E
5°E
6°E
7°E
Share reporting no affiliation (1879)
12%
8%
4%
0%
Pre−1578 monasteries
1
3
6
10
Municipal boundaries as of 1854.
Sources: HDNG v4 (IISG Amsterdam Dataverse); Kloosterlijst (Goudriaan / Geoplaza VU Amsterdam); NLGIS historical boundaries (IISH).

share in 1809 controls for pre-existing confessional heterogeneity and the cosmopolitan commercial culture associated with tolerant urban trading centers (Johnson and Koyama, 2017; Akcomak et al., 2016; Becker and Woessmann, 2009). *Municipal area* (km^2) accounts for geographic scale effects. *Average elevation* and *terrain ruggedness* capture topographic differences that shaped medieval settlement patterns and agricultural potential. *Crop suitability* (first principal component of barley, wheat, and rye suitability from the GAEZ pre-1500 data) proxies for pre-modern agricultural productivity. *Distance to coast* and *distance to the nearest major river* absorb the commercial and market-access gradients associated with waterborne trade. *Catholic mission presence* is an indicator for municipalities that hosted a *Hollandse Zending* (Dutch Mission) station, capturing pre-Reformation Catholic institutional depth. *Medieval city by 1560* (from Visser 1985) captures pre-Reformation urban status. *Eighty Years' War battle* is an indicator for municipalities that saw at least one battle during the 1568–1648 conflict, absorbing exposure to military disruption during the confessional transition. All controls are measured or fixed before 1579 (the instrument date) or are purely geographic. Post-treatment variables (population, taxes, city rights, migration) are excluded from baseline specifications but reported in robustness checks. All specifications also include *province fixed effects* to absorb the sharp north–south gradient in confessional composition.

Sample construction and summary statistics. The analytical sample comprises all Dutch municipalities for which HDNG religious affiliation data are available in at least one census year between 1879 and 1930 and that can be matched to AMCO municipality codes in the charitable-organization databases. Table 3.1 reports summary statistics for the main variables. The sample covers approximately 1,100 Dutch municipalities observed across census years 1879 to 1930. Panel A reveals that religious disaffiliation was rare in 1879 (mean *geen gezindte* share of 0.4%), but grew substantially over the period: by 1909 the mean had risen to 2%, and by 1930 to 5%, with the distribution becoming markedly right-skewed as secularization concentrated in a subset of municipalities. The religious diversity index similarly increased over time, from a mean of 0.25 in 1879 to 0.38 in 1930, reflecting both secularization and denominational fragmentation. Panel B documents the associational landscape in 1854. Total charitable associations per 1,000 inhabitants averaged 0.15 but with a median of zero, indicating that most municipalities had no formally registered charities at baseline; the long right tail reflects the concentration of associational activity in urban centres. Confesionally disaggregated counts confirm that Catholic associations outnumbered Protestant ones on average, and that the religiously affiliated share of *armenzorg* organizations was 38% (median 31%) among municipalities with any such activity. Panel C shows that pre-1578 monastery presence is rare in the sample: only 8% of municipalities recorded any monastery, and the mean count is 0.28, though the maximum of 26 (for Amsterdam) shows considerable variation in historically urbanized areas. Panels D and E report the vector of genuinely pre-treatment controls. The geo-

Table 3.1: Descriptive Statistics

	Mean	Median	SD	Min	Max	N
Panel A: Outcome Variables						
Share geen gezindte (1879)	0.00	0.00	0.01	0.00	0.16	1119
Share geen gezindte (1909)	0.02	0.00	0.04	0.00	0.35	1118
Share geen gezindte (1930)	0.05	0.01	0.09	0.00	0.51	1077
Religious diversity index (1879)	0.25	0.22	0.20	0.00	0.69	1119
Religious diversity index (1930)	0.38	0.43	0.25	0.00	0.77	1077
Panel B: Independent Variables						
Total assoc. per 1,000 pop. (1849)	0.15	0.00	0.40	0.00	4.24	1117
Catholic assoc. per 1,000 pop.	0.02	0.00	0.09	0.00	1.34	1117
Protestant assoc. per 1,000 pop.	0.01	0.00	0.06	0.00	0.82	1117
Share religious among armenzorg	0.38	0.31	0.40	0.00	1.00	169
Panel C: Instrumental Variable						
N monasteries pre-1578	0.58	0.00	2.46	0.00	27.00	1120
Any monastery pre-1578	0.14	0.00	0.35	0.00	1.00	1120
Panel D: Geographic and Agricultural Controls						
Municipal area (km²)	28.35	16.68	34.74	0.50	337.69	1117
Average elevation (m)	10.63	0.85	26.00	-6.50	201.62	1116
Terrain ruggedness index	2.05	1.08	3.33	0.08	37.51	1107
Crop suitability (PCA)	0.04	-0.47	1.72	-2.55	7.56	1117
Distance to coast (m)	19175.53	16050.41	15450.89	1.98	58464.22	1117
Distance to nearest river (m)	22869.72	16055.81	21369.52	1.75	97651.92	1117
Panel E: Historical and Confessional Controls						
Catholic mission presence	0.21	0.00	0.41	0.00	1.00	1117
Medieval city by 1560	0.11	0.00	0.31	0.00	1.00	1117
Eighty Years' War battle	0.07	0.00	0.25	0.00	1.00	1117
Jewish population share (1809)	0.00	0.00	0.01	0.00	0.11	1120

Descriptive statistics for the main analytical sample of Dutch municipalities. Panel A: Outcome variables — share of population with no religious affiliation (*geen gezindte*) and the religious diversity index (1 minus the Herfindahl–Hirschman concentration index) in selected census years. Panel B: Main explanatory variable — total charitable associations per 1,000 population (active 1854), counts by confession, and the religious share among poor-relief (*armenzorg*) associations. Panel C: Instrument — count and presence of Catholic monasteries documented before 1578. Panels D and E: the vector of genuinely pre-treatment controls included in all specifications, split into geographic/agricultural variables (Panel D) and historical and confessional variables (Panel E). Jewish population share in 1809 is set to zero for municipalities with no recorded Jewish population. Province fixed effects are included in all regressions but not reported here.

graphic and agricultural variables (Panel D) display the expected variation across a small but heterogeneous country: the median municipality covers about 17 km^2, lies close to sea level (median elevation under 1 m), and sits roughly 16 km from both the coast and the nearest major river, with long right tails for the eastern and inland municipalities. The historical and confessional variables (Panel E) are predominantly indicators: 21% of municipalities hosted a *Hollandse Zending* Catholic mission station, 11% held medieval city status by 1560, and 7% saw at least one Eighty Years' War battle, while the 1809 Jewish population share is near zero in most municipalities (mean below 0.01) but reaches 0.11 in the most concentrated trading centres.

3.2 Empirical strategy

The structural equation of interest is

$$Y_{it} \;=\; \alpha + \delta\,\text{AssocPC}_i + \mathbf{X}_i'\boldsymbol{\beta} + \varphi_p + u_{it}, \tag{1}$$

where Y_{it} is either the share of the population reporting no religious affiliation (*geen gezindte*) or the religious Herfindahl–Hirschman Index (HHI) in municipality i at census year $t \in \{1879, 1909, 1930\}$, and $AssocPC_i$ is the number of charitable associations per 1,000 inhabitants (1849 population) active in municipality i in 1854,

The core identification challenge is that municipalities with higher associational density may differ systematically in ways that also affect secularisation. To address this, I instrument the per-capita stock of charitable associations active in 1854 with the count of pre-1578 Catholic monasteries in the municipality. After the Reformation, monastic wealth was inherited by Protestant deaconries (*diaconieën*), generating persistent variation in the institutional capacity for religiously organised poor relief that predates the 1854 *Armenwet* by centuries. This instrument plausibly satisfies the exclusion restriction: pre-Reformation monastery presence affects 19th-century secularisation only through its effect on the supply of religious associations, conditional on province fixed effects. The first-stage equation is

$$\text{AssocPC}_i \;=\; \alpha + \gamma\,\text{Monasteries}_i + \mathbf{X}_i'\boldsymbol{\beta} + \varphi_p + \varepsilon_i, \tag{2}$$

where $AssocPC_i$ is the number of charitable associations per 1,000 inhabitants (1849 population) active in municipality i in 1854, $Monasteries_i$ is the pre-1579 monastery count (first documentary mention before 1579), $\mathbf{X}_i$ is a vector of genuinely pre-treatment geographic and historical controls (Jewish population share 1809, municipal area, elevation, ruggedness, crop suitability PCA, coastal distance, river distance, mission presence, medieval city by 1560, and Eighty Years' War battle indicator), φ_p denotes province fixed effects, and ε_i is an heteroskedasticity-robust error.

Two features of this design warrant comment. First, the endogenous regressor is deliberately measured in 1854, the year of the *Armenwet*. This timing is dictated by the mechanism: the law made the pre-existing stock of religiously organised relief suddenly pivotal by granting it a statutory monopoly, so the quantity that should matter is the charitable capacity in place at the moment the monopoly took effect. Measuring this stock in 1854 also guarantees that the regressor strictly predates the first outcome (the 1879 census), ruling out reverse causation from secularisation to associational supply. Second, because the instrument is a single, time-invariant cross-sectional variable, the identifying content of the design is the *reduced form*, the relationship between pre-1578 monasteries and the outcome; the IV coefficient is simply this reduced form rescaled by the first stage, $\hat{\delta} = \hat{\rho}/\hat{\gamma}$. The choice of measurement year therefore affects the *units* in which the effect is expressed (per association active in a given year), not what is identified. To make this transparent, the baseline tables report the reduced-form and first-stage coefficients on the instrument directly, alongside the IV estimate. A corollary is that the results should be insensitive to the precise pre-treatment year in which the stock is counted, provided the inherited capacity is persistent; Appendix A.6 confirms that re-measuring the regressor in 1850, 1860, or 1870 leaves the estimates essentially unchanged. Replacing the level with the post-1854 *increase* in associations is not an attractive alternative: it would shift the regressor into the partly endogenous, post-treatment period, and Appendix A.7 shows that the instrument predicts this increase nearly as strongly as the level, so differencing isolates no separate source of exogenous variation.[5]

Several robustness checks probe the mechanism and scope of the main results. First, I replace total associations with *Armenzorg*-only organisations to isolate the direct welfare channel mandated by the *Armenwet*. Second, I decompose associations by confession (Catholic, Protestant, neutral) to identify which denominational supply drives secularisation. Third, I use confession-specific associations and population shares as outcomes to test for confessional consolidation alongside disaffiliation. Fourth, I examine heterogeneity by interacting associational density with pre-treatment religious diversity (1809, one minus the HHI), testing whether the effect varies with the municipality's initial confessional composition. Finally, I replace the main IV specification with OLS regressions of the associational HHI (the confessional concentration of local associations) to assess whether the *composition*, rather than the level, of associational supply shapes religious outcomes.

4 Results

4.1 Covariate Balance

Table 4.1 reports standardised differences in pre-treatment covariates between municipalities with and without at least one pre-1578 monastery. The table serves two purposes. First, it documents which observable characteristics distinguish treated from control mu-

[5]The alternative-measurement-year results appear in Tables A.2 and A.3; the diagnostic on the increase specification appears in Table A.4.

nicipalities and thereby clarifies the case for the control vector $\mathbf{X}_i$. Second, it provides a transparent diagnostic for the instrument's plausibility: if pre-1578 monasteries strongly predict pre-*Armenwet* religious composition, the exclusion restriction is suspect.

The standardised differences in Table 4.1 reveal systematic imbalance along historical-urban dimensions. Municipalities with pre-1578 monasteries are substantially more likely to have held medieval city rights (std. diff. 1.48), to have been the site of an Eighty Years' War battle (1.01), and to have had a larger 1849 population (0.83). They also exhibit higher Jewish population shares in 1809 (0.77), consistent with the historical co-location of monasteries with urban Jewish communities in Dutch trading cities, and higher rates of Catholic mission presence (0.58), reflecting shared ecclesiastical geography. These imbalances are precisely what motivate the inclusion of the full control vector in all specifications and the matching exercise reported in Appendix A.8.

Crucially, the confessional composition of 1809 is well balanced across the two groups. Neither the Catholic share (std. diff. -0.12) nor the Protestant share (std. diff. 0.09) differs significantly between monastery and non-monastery municipalities, conditional on province fixed effects. Religious diversity in 1809 is somewhat higher in monastery municipalities (std. diff. 0.38), reflecting greater denominational mixing in the historically urbanised areas where monasteries concentrate. This difference is statistically significant and, unlike the urban-historical imbalances, bears directly on the diversity outcome: because treated municipalities were already more pluralistic before the *Armenwet*, part of the later diversity gap could reflect persistence of pre-existing composition rather than the charity channel. The province fixed effects and control vector absorb the cross-sectional component of this gap, but it tempers a causal reading of the diversity estimates; we return to it in the placebo analysis of Appendix A.5.

4.2 Effect on Secularization

Table 4.2 reports baseline estimates of the effect of charitable association density, measured as the number of associations active in 1854, on the *geen gezindte* share in census years 1879, 1909, and 1930. Columns (1)–(3) present OLS estimates; columns (4)–(6) present IV estimates in which the count of associations is instrumented by the indicator of pre-1578 monastery presence.

Both OLS and IV estimates are positive and statistically significant across all three census years. The OLS coefficients grow over time, from 0.004 in 1879 to 0.029 in 1930, consistent with the secular trend of rising disaffiliation becoming more pronounced in municipalities with denser charitable infrastructure. The IV estimates are uniformly larger in magnitude than their OLS counterparts (e.g., 0.211 versus 0.029 in 1930), pointing to downward attenuation bias in the OLS estimates, likely because the endogenous association count is measured with noise or because municipalites with high unobserved religiosity invested simultaneously in associations and maintained low disaffiliation rates.

The table also reports the two components from which the IV estimate is built: the first-

Table 4.1: Covariate Balance: Municipalities With vs. Without Pre-1578 Monasteries

Variable	No monastery	Monastery	Std. diff.	FE p-value
Municipal area (km²)	26.804	37.843	0.318	5.3e-02*
Average elevation (m)	10.721	10.082	-0.025	0.360
Terrain ruggedness index	2.003	2.312	0.093	1.7e-02**
Distance to coast (m)	18846.874	21200.159	0.152	0.360
Distance to nearest river (m)	22549.109	24844.799	0.107	0.241
Crop suitability (PCA)	-0.004	0.278	0.164	0.111
Catholic mission presence	0.180	0.417	0.578	0.0e+00***
Medieval city by 1560	0.043	0.500	1.482	0.0e+00***
Eighty Years' War battle	0.031	0.282	1.008	0.0e+00***
Catholic share 1809	0.421	0.375	-0.117	0.550
Protestant share 1809	0.575	0.612	0.094	0.893
Jewish share 1809	0.004	0.011	0.768	0.0e+00***
Religious diversity (1–HHI) 1809	0.196	0.270	0.381	0.0e+00***
Population 1849	1764.211	8497.615	0.827	0.0e+00***

Standardised differences are the difference in means divided by the pooled standard deviation. The FE p-value is from a province-FE regression of the covariate on the monastery indicator. N = 1120 municipalities (964 without monasteries, 156 with monasteries). *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

stage coefficient of the instrument on association density and the reduced-form coefficient of the instrument on *geen gezindte*. The first stage is positive and stable across columns—an additional pre-1578 monastery raises charitable density by a similar amount regardless of the outcome year—while the reduced-form effect on disaffiliation is small in 1879 and grows by roughly an order of magnitude by 1930. Because the instrument is a single cross-sectional variable, it is this reduced-form gradient that the design identifies; the IV coefficient simply expresses it per unit of association density. Taken together, the positive first stage and the growing reduced-form effect mean that, once the exogenous component of association density is isolated, the material stake of affiliation embedded in the 1854 welfare monopoly appears to have shaped the long-run trajectory of disaffiliation. Within-province variation accounts for a modest but rising share of the outcome variance (within-R^2 of 0.021 in 1879 rising to 0.133 in 1930 under OLS), indicating that secular time trends are primarily absorbed by province fixed effects while the cross-sectional gradient in charitable density explains a growing share of residual variation.

The results are robust to several supplementary checks reported in the Appendix. Falsification tests confirm that pre-1578 monasteries do not predict Catholic, Protestant, or Jewish population shares in 1809, supporting the exclusion restriction (Appendix A.5).[6]

[6] Religious diversity in 1809 is, however, significantly *higher* in monastery municipalities (Table A.1); since the estimated effect on diversity is also positive, this pre-existing gap is a potential confound rather than a conservative bias, and is the one placebo dimension on which monasteries are not balanced. The province

Table 4.2: Effect of Charitable Associations on Religious Disaffiliation

	OLS			IV (Instrument: Pre-1578 Monasteries)		
Year:	1879	1909	1930	1879	1909	1930
	(1)	(2)	(3)	(4)	(5)	(6)
Charitable associations per 1,000 capita (active 1854)	0.004***	0.015**	0.029**	0.010**	0.065***	0.211***
	(0.001)	(0.006)	(0.013)	(0.004)	(0.023)	(0.069)
Adj. R^2	0.186	0.426	0.487	0.173	0.415	0.486
Within R^2 (adj.)	0.021	0.061	0.133	0.006	0.043	0.132
Observations	1107	1105	1064	1107	1105	1064
Reduced form (Pre-1578 monasteries)	—	—	—	0.000***	0.001***	0.005***
First stage (Pre-1578 monasteries)	—	—	—	0.022***	0.022***	0.022***
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	—	—	—	19.5	19.5	18.6

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Baseline controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns.

Oster (2019) bounds for selection on unobservables yield δ^* estimates ranging from 1.18 to 1.55 for the *geen gezindte* outcomes, indicating that selection on unobservables would need to be 18–55% stronger than selection on observables to eliminate the estimated effect (Appendix A.10). The estimates are also insensitive to the year in which charitable density is measured: re-counting the stock active in 1850, 1860, or 1870 in place of 1854 leaves both the secularization and diversity coefficients essentially unchanged (Appendix A.6), as expected when the instrument fixes the reduced form and the inherited stock is persistent. A related diagnostic asks whether the post-1854 *increase* in associations offers a cleaner regressor than the level; it does not, because the instrument predicts the increase nearly as strongly as the level and the two are collinear, so differencing recovers no separate exogenous variation while sacrificing the pre-treatment timing of the 1854 stock (Appendix A.7). Finally, Mueller–Watson (2024) spatial unit root tests confirm that, although the raw *geen gezindte* series behaves as a spatial unit root, conditioning on province fixed effects and controls renders the OLS residuals spatially stationary—ruling out spurious spatial regression—and the IV estimates are essentially unchanged when a second-degree polynomial in projected centroid coordinates is added to the controls (Appendix A.11).

4.3 Effect on Religious Diversity

Table 4.3 turns to the religious diversity index $D_{it} = 1 - \sum_k s_{kit}^2$, where higher values denote greater pluralism across denominations. The dependent variable is observed in 1879, 1899, and 1930.

The OLS and IV estimates are positive and statistically significant throughout. An additional association per 1,000 inhabitants active in 1854 is associated with an increase in the religious diversity index of about 0.02–0.03 points under OLS and 0.22–0.26 points under IV. The pattern of IV magnitudes exceeding OLS magnitudes mirrors the secularization results: endogeneity leads OLS to understate the effect of religiously embedded

fixed effects and controls partial out its cross-sectional component.

Table 4.3: Effect of Charitable Associations on Religious Diversity (1−HHI)

	OLS			IV (Instrument: Pre-1578 Monasteries)		
Year:	1879	1899	1930	1879	1899	1930
	(1)	(2)	(3)	(4)	(5)	(6)
Charitable associations per 1,000 capita (active 1854)	0.021	0.022*	0.032**	0.257***	0.218**	0.262***
	(0.014)	(0.013)	(0.015)	(0.099)	(0.085)	(0.087)
Adj. R^2	0.470	0.531	0.625	0.472	0.532	0.626
Within R^2 (adj.)	0.122	0.091	0.083	0.125	0.092	0.084
Observations	1107	1107	1064	1107	1107	1064
Reduced form (Pre-1578 monasteries)	—	—	—	0.006***	0.005***	0.006***
First stage (Pre-1578 monasteries)	—	—	—	0.022***	0.022***	0.022***
Province FE	Yes	Yes	Yes	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	—	—	—	19.5	19.5	18.6

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Baseline controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns.

charitable infrastructure on the denominational composition of municipalities. As in the secularization table, the reported reduced-form coefficient of pre-1578 monasteries on the diversity index is positive and stable across census years, and the first stage is positive; the IV estimate is this positive reduced form rescaled by the positive first stage.

The positive sign is what the commitment-screening mechanism of Section 3.2 predicts. Because the locally dominant confession commands the densest charitable network, it imposes the steepest participation costs, and—being already the more heavily screened denomination—it sheds its weakly-attached members proportionally faster than the minority as associational density rises. The affiliated population is thereby pulled toward parity between confessions rather than concentrated within the dominant one, so denominational diversity rises. Greater charitable capacity thus deconcentrates the existing confessional landscape, raising the diversity index, while simultaneously, as Table 4.2 shows, generating elevated disaffiliation: the same screening force that rebalances the affiliated population also drives the weakly committed out of religion altogether. The two findings are therefore complementary rather than contradictory—both are signatures of participation-cost screening falling hardest on the dominant, most heavily organised confession.

Within-province explained variance is again modest (within-R^2 of 0.083–0.125 across years and estimators), and the province fixed effects absorb the bulk of cross-sectional religious geography, leaving the association-density gradient to explain residual variation in pluralism.

The IV estimates for both secularization and religious diversity are qualitatively unchanged when re-estimated on a propensity-score-matched sample that balances the 11 pre-treatment covariates on which monastery and non-monastery municipalities most differ (Appendix A.8). On the matched sample, the IV coefficient on *geen gezindte* in 1930 is 0.318 and the coefficient on the religious diversity index in 1930 is 0.340, both statistically significant and retaining the signs of the full-sample estimates despite the reduction in sample size.[7] Oster δ^* for the diversity outcomes ranges from 0.77 to 1.19 (Appendix A.10),

[7] A purely geographic matching strategy—matching each monastery municipality to its five nearest non-

somewhat below the conventional threshold of unity for the earlier census years, which is consistent with the greater sensitivity of the diversity results to unobserved confessional geography.[8]

4.4 Discussion

The baseline estimates establish two regularities. First, municipalities endowed with greater pre-Reformation monastic infrastructure, and correspondingly denser post-*Armenwet* charitable networks, experienced higher rates of *geen gezindte* by 1930. Second, the same municipalities exhibit higher religious diversity throughout the period. These two findings are complementary rather than contradictory: both follow from participation-cost screening. Denser charitable networks raise the net cost of nominal affiliation; the weakly committed exit into *geen gezindte*, raising secularization, and because the locally dominant confession organises most intensively it bears the steepest cost and sheds its marginal members proportionally fastest, rebalancing the affiliated population toward parity and raising diversity. The growing *geen gezindte* effect over time additionally reflects the long-run erosion of the welfare tie, as secular public provision displaced confessional relief and the material incentive for nominal membership collapsed.

The IV estimates consistently exceed their OLS counterparts, which is consistent with classical measurement error in the association count and with the presence of municipalities where unobserved religiosity simultaneously stimulated associational investment and depressed secularization, biasing OLS toward zero. The growing magnitude of the *geen gezindte* coefficient from 1879 to 1930 reflects the gradual loosening of the welfare tie: as the Dutch welfare state expanded and alternative sources of poor relief became available, the long-run disaffiliation effect of the original charitable endowment became progressively more visible.

5 Conclusion

This paper has studied the long-run relationship between religiously organised poor relief and the trajectory of secularization in the Netherlands between 1879 and 1930. Exploiting

monastery municipalities, so that the comparison holds unobserved regional characteristics approximately fixed—yields the same conclusion, with 1930 IV coefficients of 0.220 (*geen gezindte*) and 0.236 (religious diversity) and a substantially stronger first stage ($F \approx 16$) than the propensity-score match (Appendix A.9).

[8] Mueller–Watson (2024) spatial unit root tests reinforce this caution: unlike the *geen gezindte* residuals, the diversity residuals reject spatial stationarity in 1899 and 1930, indicating that some spatial dependence survives the conditioning set. The hypothesis is nonetheless supported once that dependence is addressed directly. The IV estimates are stable when a second-degree spatial polynomial in centroid coordinates—which absorbs such trends—is added to the controls, and, more demandingly, when the model is re-estimated on Mueller–Watson nearest-neighbour spatially-differenced data, which removes the spatial unit root at its source: the differenced diversity coefficients stay positive and comparable in magnitude throughout and remain significant in 1930, while the secularization estimates are essentially unchanged (Appendix A.11).

the spatial distribution of pre-Reformation Catholic monasteries as an instrument for post-*Armenwet* charitable association density, I find two robust empirical regularities. First, municipalities endowed with greater charitable infrastructure experienced persistently higher rates of religious disaffiliation (*geen gezindte*), with the effect growing substantially from near zero in 1879 to economically meaningful magnitudes by 1930. Second, the same municipalities exhibit higher religious diversity throughout the period, indicating that the affiliated population became *less* confessionally concentrated as associational density rose. The IV estimates are uniformly larger than their OLS counterparts, consistent with attenuation bias arising from measurement error and from unobserved municipal religiosity that simultaneously stimulated associational investment and suppressed disaffiliation.

The first finding is, on its face, paradoxical: if religious charities bundled welfare access with confessional membership, one might expect denser associational networks to sustain affiliation rather than corrode it. The resolution lies in the commitment-screening mechanism formalized in Appendix B. Dense associational networks raised the effective price of nominal membership, through attendance obligations, financial contributions, and behavioral compliance, to a level that households with weak religious conviction found prohibitive. These marginal affiliates exited into *geen gezindte* rather than tolerating the organizational overhead. The same screening force accounts for the second finding: because the dominant denomination commanded a disproportionate share of local associations, it imposed the steepest participation costs and—being already the more heavily screened denomination—shed its weakly-attached members proportionally faster than the minority, rebalancing the affiliated population toward parity and raising diversity. Propositions 1 and 2 of the model show that both effects, higher aggregate disaffiliation and lower confessional concentration among the residual faithful, follow necessarily from a single threshold condition, without requiring auxiliary assumptions about preferences or market structure.

The paper connects to several strands of literature. Iannaccone (1994) demonstrates that high participation requirements produce strong, committed congregations. The present results add a population-level qualification: the intensive-margin gains from screening come at the cost of extensive-margin exits, and in a regime of statutory welfare monopoly the exit option is the unaffiliated category rather than a rival denomination. Barro and McCleary (2003) and Barro and McCleary (2006a) document cross-nationally that government regulation of religion and the presence of state churches are among the strongest predictors of religious participation; the mechanism identified here provides a within-country, causally identified analog, tracing municipal-level variation to a specific organizational channel. The long-run dynamic is consistent with Gill and Lundsgårde (2004), who show that welfare-state generosity is associated with lower church attendance: as Dutch secular public welfare expanded over the twentieth century and the material incentive for nominal affiliation eroded, the reservoir of weakly committed members retained by associational overhead became a source of rapid exits. Finally, the identification strategy follows the tradition of Becker and Woessmann (2009) and Cantoni et al. (2018) in tracing the institutional aftermath of the Reformation through its effects on subsequent social outcomes, but adds a distinct mechanism, not literacy or market restructuring but the transfer of monastic endowments to Protestant deaconries, and a distinct outcome, the long-run trajectory of secularization.

The findings also speak to the present European situation. The Dutch *ontkerkelijking* of the 1960s through the 1980s was among the steepest episodes of mass religious disaffiliation in European history (Knippenberg, 1998; van Rooden, 1996). The evidence presented here suggests that its sharpness was not incidental: municipalities that had built the densest confessional welfare infrastructures under the *Armenwet* regime had, by the mid-twentieth century, accumulated the largest pools of nominally affiliated households held in place by institutional inertia rather than conviction. When the postwar welfare state dissolved the material rationale for maintaining that membership, the exits were correspondingly rapid. The same logic applies, with varying institutional specifics, across much of Northwestern and Central Europe. Countries where religious organizations long served as primary welfare providers, such as Germany, Belgium, Ireland, and the Netherlands, shared the pattern of sustained nominal affiliation followed by sudden collapse once secular alternatives became universally available (Norris and Inglehart, 2004; Voas and Chaves, 2016). The French case is instructive in contrast: the early separation of church and state under the 1905 *loi de séparation* removed the welfare-religion bundle at the outset of the twentieth century, meaning that the weakly committed exited earlier and gradually, leaving a smaller residual pool available for later rapid disaffiliation. Across all these settings, the contemporary rise of religious "Nones" draws disproportionately from those who were affiliated by social habit and institutional access rather than doctrinal conviction (Hout and Fischer, 2002, 2014), precisely the marginal households that the commitment-screening model identifies as the first to exit when participation costs rise or material incentives fall.

The results carry a natural external-validity caveat: the *Armenwet* regime and the specific Reformation-era endowment mechanism are features of the Dutch institutional context, and the quantitative magnitudes should not be extrapolated beyond it. What generalizes is the mechanism: wherever access to social welfare has been historically bundled with confessional membership, the organizational demands of that bundling shape not only the level but the long-run fragility of religious affiliation. Secularization is not simply the product of modernization, rising incomes, or the diffusion of scientific rationalism. It is also shaped by the institutional architecture through which religious organizations once provided the material goods of social life, and by the exit dynamics that architecture sets in motion when secular substitutes eventually arrive.

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A Data Appendix

This appendix documents the construction of the analysis dataset, the provenance of each underlying source, and the procedures used to harmonize them at the municipality level. All raw data, harmonization code, and the assembled panel are available in the replication package.[9] Unless otherwise noted, all sources are matched to municipalities using the standard Dutch AMCO (*Algemene Maatschappelijke Code*) classification, which assigns a stable numeric identifier to each municipality and thereby permits consistent linkage across the frequent boundary changes and amalgamations of the nineteenth and twentieth centuries.

A.1 Religious affiliation and municipal characteristics: HDNG

Outcome variables and a number of baseline controls are drawn from the *Historische Database Nederlandse Gemeenten* (HDNG, Historical Database of Dutch Municipalities). The HDNG is a longitudinal compilation of population and social statistics for Dutch municipalities covering the period from the early nineteenth century to circa 1970, assembled originally under the direction of Onno Boonstra and subsequently harmonized and distributed through the CLARIAH research infrastructure and the International Institute of Social History (IISH) in Amsterdam. The database aggregates the published returns of the Dutch decennial and quinquennial censuses together with vital-registration and fiscal statistics, reported at the level of the roughly 1,100 municipalities that constituted the Netherlands during the study period.

From the HDNG I extract municipality-level counts of residents by religious denomination. The Dutch census enumerated the population by confession from the early nineteenth century onward, distinguishing Roman Catholics, the principal Protestant denominations (Dutch Reformed, Re-Reformed, Lutheran, Mennonite, and smaller bodies), Jews,

[9]The replication package is available at https://github.com/basm92/crcs.

and, beginning with the census of 1879, persons recording no religious affiliation (*geen gezindte*). The principal outcome variable, the *geen gezindte* share, is the count of residents reporting no affiliation divided by total enumerated population in the census years 1879, 1889, 1899, 1909, 1920, and 1930. The secondary outcome, a religious diversity index, is constructed as one minus the Herfindahl–Hirschman index of denominational population shares within each municipality and census year, so that higher values denote greater confessional pluralism.

The HDNG also supplies several control variables measured prior to the period of interest: the Jewish population share in 1809, municipal population and surface area used to construct population density in 1849, municipal tax revenue per capita in 1859, and the components of cumulative net migration. The indicator for historical city rights (*stadsrechten*) is likewise compiled from the HDNG together with standard reference works on the constitutional history of Dutch towns.

A.2 Charitable associations: the Huygens databases

The intensity of religiously affiliated charitable activity is measured by aggregating four historical registries of welfare organizations curated by the Huygens Institute for Dutch History (KNAW). Each registry digitizes a distinct body of archival and published source material on the associational infrastructure of Dutch poor relief and mutual aid. I describe each in turn and then explain how they are combined.

Armenzorg. The *Verenigingen voor armenzorg en armoedepreventie* (Associations for Poor Relief and Poverty Prevention) database catalogs charitable organizations active principally during the nineteenth century. For identification purposes its most valuable feature is that each organization is coded by ideological or religious orientation (neutral, Protestant, Catholic, or Jewish), alongside its founding year, host municipality, target beneficiary group, and, where known, the gender composition of its membership. This source most directly operationalizes the notion of *religious* charity provision used throughout the paper.

Erkende Verenigingen. The *Erkende Verenigingen* (Recognized Associations) database covers the period 1855–1903. Following the 1855 *Wet op de Verenigingen en Vergaderingen*, associations seeking legal personality were required to obtain recognition by Royal Decree, with their statutes published in the *Staatscourant* (official gazette). The registry provides near-complete coverage of the formally recognized associational sector in the half-century immediately following the 1854 *Armenwet*, recording for each association its name, host municipality, founding date, date of recognition, and Royal Decree number.

Sociale Zekerheid. The *Sociale Zekerheid* (Social Security) database documents local institutions providing extramural, community-based poor relief and social assistance, as recorded in the official charitable guides published over the first half of the twentieth century. Each institution is assigned a categorical affiliation code distinguishing civil or secular (*Burgerlijk*), church-affiliated (*Kerkelijk*), and private (*Particulier*) bodies, which extends the measurement of religious charitable activity into the early twentieth century.

Verzekeringsfondsen. The *Verzekeringsfondsen* (Insurance and Mutual Aid Funds) database records mutual aid societies, occupational funds, burial societies, and sick funds active over several centuries, with coverage densest in the eighteenth and nineteenth centuries. It captures the corporatist welfare arrangements that predated statutory social insurance. For the present analysis I retain funds active before 1880 as a measure of pre-existing associational infrastructure.

Construction of the charity measures. Each database is processed by extracting organization-level records and matching the recorded place name to an AMCO municipality code. The four sources are then pooled into a single organization-level file. The principal explanatory variable is the number of charitable associations per 1,000 inhabitants (1849 population) active in a municipality in 1854. Auxiliary measures used in robustness exercises include counts restricted to *Armenzorg* organizations, counts disaggregated by confession (Catholic, Protestant, neutral), and the Herfindahl–Hirschman index of the confessional composition of a municipality's associations.

A measure of associations "active in 1854" might appear to sit awkwardly with the coverage windows of the underlying registries, two of which begin after that date. The apparent tension is resolved by the construction rule: an organization is counted as active in a given year if its *founding* year is at or before that year and it had not yet been dissolved. Because the rule keys on the founding date rather than on the date of administrative recognition or guide inclusion, organizations founded before 1854 enter the 1854 stock even when they appear in a registry whose nominal window opens later. In practice this means the 1854 stock draws chiefly on the *Verzekeringsfondsen* and *Armenzorg* sources, which document organizations founded throughout the first half of the nineteenth century; the *Erkende Verenigingen* registry (recognition 1855–1903) contributes only those of its associations with a pre-1854 founding date, and *Sociale Zekerheid* (1899–1956), whose institutions are overwhelmingly of later origin, contributes very little at baseline. The 1854 stock is thus a deliberately thin, pre-treatment core; most of the associational activity in the pooled file is founded after 1854, a fact that also bears on the increase specification discussed in Appendix A.7.

A.3 Pre-Reformation monasteries: the Kloosterlijst

The instrument is constructed from the *Kloosterlijst*, a census of monasteries, convents, and other religious communities in the territory of the present-day Netherlands from the medieval period through 1800, compiled at the Vrije Universiteit Amsterdam. The Kloosterlijst supersedes the earlier *Monasticon Batavum* of Schoengen (1941–42), correcting numerous errors and adopting the convention that each institution receives a single record spanning all stages of its history. It documents several hundred verified communities, including beguinages (*begijnhoven*), houses of the Common Life, mendicant friaries, and enclosed convents of all major orders. The database distinguishes positively verified institutions from an explicit elimination list of communities mentioned in the secondary literature but unconfirmed in primary sources; only verified records are used here.

Each record reports the institution's name, its historical parish and current municipality, diocese, religious order and gender composition, and a dated narrative of institutional stages, including the year of first documentary mention and the year of last mention or dissolution. I restrict attention to institutions whose last documentary mention (dissolution) falls before 1578, the conventional date for the consolidation of the Dutch Reformation, and aggregate them to municipality-level counts of pre-Reformation monastic presence after matching parish and place names to AMCO codes.

A.4 Boundary harmonization and matching

Place names recorded in the charity and monastery sources are matched to AMCO municipality codes through a combination of exact string matching against the AMCO reference table and manual resolution of ambiguous or historical spellings. Historical municipal boundaries used for the maps are taken from the NLGIS collection of digitized Dutch municipal boundary files distributed through the IISH. The fully reproducible matching and assembly procedure, together with the resulting municipality-by-census-year panel, is provided in the replication package.

A.5 Placebo tests

Table A.1 reports falsification tests that assess whether pre-1578 monasteries predict religious composition in 1809, three-quarters of a century before the *Armenwet* of 1854. If the exclusion restriction holds, the instrument should not predict pre-treatment outcomes except through channels unrelated to post-1854 charity provision.

Panel A presents reduced-form estimates of the effect of pre-1578 monasteries on Catholic, Protestant, and Jewish population shares and on the religious diversity index in 1809. Neither the Catholic share nor the Protestant share nor the Jewish share is significantly predicted by monastery presence, conditional on province fixed effects and the baseline control vector. Religious diversity in 1809 is, however, significantly *higher* in monastery municipalities, indicating that these already exhibited somewhat greater denominational mixing before the *Armenwet*. Because the estimated effect of association density on diversity is also positive, this pre-existing difference is a potential confound rather than a conservative bias: it is the one placebo dimension on which monastery and

Table A.1: Placebo Tests: Effect of Pre-1578 Monasteries on Pre-Treatment Outcomes

	Placebo (1809):			Reduced form (geen gezindte):			Reduced form (diversity):			
	Catholic 1809	Protestant 1809	Jewish 1809	Diversity 1809	Geen 1879	Geen 1909	Geen 1930	Diversity 1879	Diversity 1899	Diversity 1930
Pre-1578 monasteries	0.001	-0.001	0.000	0.004***	0.000***	0.001***	0.005***	0.006***	0.005***	0.006***
	(0.002)	(0.002)	(0.000)	(0.002)	(0.000)	(0.000)	(0.001)	(0.002)	(0.002)	(0.002)
Adj. R^2	0.699	0.697	0.232	0.476	0.173	0.415	0.486	0.472	0.532	0.626
Within R^2 (adj.)	0.180	0.176	0.156	0.131	0.006	0.043	0.132	0.125	0.092	0.084
Observations	1015	1015	1107	1015	1107	1105	1064	1107	1107	1064

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Placebo tests: Pre-1578 monasteries should not predict pre-Armenwet (1854) religious composition if the exclusion restriction holds. Reduced-form estimates with province fixed effects and baseline controls. Baseline controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle.

non-monastery municipalities are not balanced. The province fixed effects and baseline controls partial out its cross-sectional component, but it counsels caution in reading the diversity estimates as fully causal, and motivates the matching exercise in Appendix A.8.

Panel B reports the reduced-form relationship between pre-1578 monasteries and the main outcome variables for context. The pattern of growing reduced-form effects over census years (1879 to 1930) mirrors the IV results and is consistent with the mechanism of gradually accumulating disaffiliation as the *Armenwet* regime matures.

A.6 Alternative measurement years for charitable density

The baseline measures the endogenous regressor as the stock of charitable associations active in 1854, the year of the *Armenwet*. Section 3.2 argues that, because the instrument is a single time-invariant variable, the design identifies the reduced-form relationship between pre-1578 monasteries and the outcome, and the choice of measurement year only rescales that reduced form by the first stage. It follows that the IV estimates should be insensitive to the exact pre-treatment year in which the stock is counted, provided the inherited capacity is persistent.

Tables A.2 and A.3 confirm this. Each column re-measures the regressor using the stock of associations active in 1850 (before the *Armenwet*), 1854 (the baseline), 1860, and 1870, holding the 1849 population denominator, the instrument, the controls, and the outcomes fixed. The IV coefficients are stable across the 1850, 1854, and 1860 columns for every census-year outcome and for both secularization and diversity. The 1870 column yields uniformly smaller coefficients, but for a revealing reason: by 1870 the first stage is mechanically larger (more associations had accumulated per monastery), so dividing the fixed reduced form by a larger first stage produces a smaller—but identically signed and significant—IV coefficient. The reduced form itself, the object the design actually identifies, is unchanged by the measurement year.

A.7 The increase specification

A natural alternative to the 1854 stock is the *increase* in charitable associations over the following decades. Table A.4 examines whether this increase offers a distinct, cleaner source of exogenous variation than the level. It regresses, in turn, the level of associations per capita active in 1854 (the baseline first stage), the change in that quantity between the

Table A.2: Robustness: Alternative Measurement Years for Charitable Associations (*Geen Gezindte*)

Measurement year:	1850	1854	1860	1870
Geen gezindte 1879	0.011**	0.010**	0.011**	0.007***
Geen gezindte 1909	0.071***	0.065***	0.071**	0.044***
Geen gezindte 1930	0.227***	0.211***	0.231***	0.142***
First stage (monasteries)	0.020***	0.022***	0.020***	0.032***
Kleibergen–Paap F-statistic	18.4	18.6	11.4	19.9
Observations	1,064	1,064	1,064	1,064

IV estimates of the effect of total charitable associations per 1,000 inhabitants on the outcome, instrumented by the count of pre-1578 monasteries. Each column re-measures the endogenous regressor using the stock of associations active in the indicated year; all columns use the 1849 population as denominator. Coefficient rows report the IV estimate for the indicated census-year outcome. All specifications include province fixed effects and the full baseline control vector. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.3: Robustness: Alternative Measurement Years for Charitable Associations (Religious Diversity, 1−HHI)

Measurement year:	1850	1854	1860	1870
Religious diversity 1879	0.278***	0.257***	0.279**	0.172***
Religious diversity 1899	0.236**	0.218**	0.237**	0.146***
Religious diversity 1930	0.282***	0.262***	0.287***	0.176***
First stage (monasteries)	0.020***	0.022***	0.020***	0.032***
Kleibergen–Paap F-statistic	18.4	18.6	11.4	19.9
Observations	1,064	1,064	1,064	1,064

IV estimates of the effect of total charitable associations per 1,000 inhabitants on the outcome, instrumented by the count of pre-1578 monasteries. Each column re-measures the endogenous regressor using the stock of associations active in the indicated year; all columns use the 1849 population as denominator. Coefficient rows report the IV estimate for the indicated census-year outcome. All specifications include province fixed effects and the full baseline control vector. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A.4: The Instrument Loads on Both the Level and the Post-*Armenwet* Growth of Charitable Associations

Dependent variable	Pre-1578 monasteries	Std. error	Model F	Observations
Level: associations p.c. (active 1854)	0.022***	0.007	28.1	1,107
Growth: Δ associations p.c. (1854–1879)	0.018***	0.006	6.7	1,107
Growth | 1854 level	0.015**	0.007	7.3	1,107

Each row is a separate regression of the indicated dependent variable on the count of pre-1578 monasteries, the full baseline control vector, and province fixed effects, with HC1 standard errors. The level (associations per 1,000 inhabitants active in 1854) is the baseline first stage; the growth measure is the change in associations per 1,000 inhabitants between the 1854 and 1879 active stocks; the third row adds the 1854 level as a control. The instrument predicts the growth nearly as strongly as the level, and the growth coefficient survives conditioning on the level, so the increase isolates no exogenous channel distinct from the inherited stock. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

1854 and 1879 active stocks, and the same change conditional on the 1854 level, each on the count of pre-1578 monasteries with the full control vector and province fixed effects.

The instrument predicts the post-1854 increase nearly as strongly as it predicts the level, and its coefficient on the increase survives conditioning on the level. This reflects the construction documented above: because most associations are founded after 1854, the post-1854 growth is dominated by the same persistent, historically urbanised municipalities that hold the larger inherited stock. The two measures are therefore collinear, and an increase specification recovers no exogenous variation separate from the level. Differencing would, however, move the regressor into the post-treatment period, which overlaps the early outcome censuses and is partly endogenous to the secularization process itself. These considerations motivate retaining the pre-treatment 1854 level as the endogenous regressor.

A.8 Propensity score matching

To assess the sensitivity of the IV estimates to functional-form assumptions and common-support concerns, I re-estimate the main specifications on a sample matched on pre-treatment observables. Table A.5 reports covariate balance before and after propensity score matching. Matching is performed via logistic regression of the monastery indicator on the 11 pre-treatment geographic and historical covariates listed in Section 3, using 3:1 nearest-neighbour matching with replacement. Of 1,015 municipalities with complete covariate data, 146 treated municipalities are matched to 198 controls, yielding a matched sample of 344 observations.

Table A.6 reports IV estimates for the *geen gezindte* outcomes on the matched sample. Table A.7 reports the corresponding estimates for the religious diversity index (1−HHI). All specifications include province fixed effects and the full baseline control vector; estimates are weighted by matching weights.

The matched-sample estimates retain the sign and significance of the full-sample results. For *geen gezindte* in 1930, the IV coefficient is 0.318 (compared to 0.211 in the

Table A.5: Covariate Balance Before and After Propensity Score Matching

Variable	Mean (control), pre	Mean (treated), pre	Std. diff., pre	Mean (control), post	Mean (treated), post	Std. diff., post
Municipal area (km^2)	27.770	36.429	0.245	43.461	36.429	-0.199
Average elevation (m)	9.752	9.738	-0.001	12.913	9.738	-0.128
Terrain ruggedness index	1.914	2.308	0.124	2.659	2.308	-0.111
Distance to coast (m)	19668.658	22167.617	0.163	21172.874	22167.617	0.065
Distance to nearest river (m)	23123.220	25072.842	0.090	26402.683	25072.842	-0.061
Crop suitability (PCA)	0.044	0.285	0.140	0.380	0.285	-0.055
Catholic mission presence	0.190	0.432	0.579	0.482	0.432	-0.120
Medieval city by 1560	0.039	0.507	1.516	0.473	0.507	0.111
Eighty Years' War battle	0.029	0.295	1.062	0.251	0.295	0.173
Catholic share 1809	0.421	0.376	-0.112	0.384	0.376	-0.020
Jewish share 1809	0.004	0.012	0.778	0.012	0.012	-0.006

Nearest-neighbour matching with propensity score estimated via logistic regression, 3:1 ratio with replacement. Pre-matching standardised differences use raw means; post-matching standardised differences use matching weights.

Table A.6: IV Estimates on Matched Sample: Religious Disaffiliation

Share geen gezindte:	1879	1909	1930
	(1)	(2)	(3)
Charitable associations per 1,000 capita (active 1854)	0.017*	0.087*	0.318**
	(0.009)	(0.044)	(0.156)
Adj. R^2	0.225	0.577	0.483
Within R^2 (adj.)	-0.014	0.172	0.234
Observations	344	344	337
Province FE	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	6.0	6.0	6.0

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$

Baseline controls: Jewish population share 1809; Municipal area (km^2); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns. Matched sample via propensity score matching (3:1 nearest-neighbour with replacement). Estimates weighted by matching weights.

Table A.7: IV Estimates on Matched Sample: Religious Diversity (1−HHI)

Religious diversity (1−HHI):	1879	1899	1930
	(1)	(2)	(3)
Charitable associations per 1,000 capita (active 1854)	0.224*	0.247**	0.340**
	(0.126)	(0.125)	(0.158)
Adj. R^2	0.470	0.548	0.615
Within R^2 (adj.)	0.084	-0.017	0.049
Observations	344	344	337
Province FE	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	6.0	6.0	6.0

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Baseline controls: Jewish population share 1809; Municipal area (km^2); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns. Matched sample via propensity score matching (3:1 nearest-neighbour with replacement). Estimates weighted by matching weights.

full-sample IV specification) and for the religious diversity index in 1930 it is 0.340 (compared to 0.262 in the full sample); both are statistically significant. The larger magnitudes on the matched sample reflect the up-weighting of comparable control municipalities; the first-stage F-statistic falls to approximately 6.0, at the boundary of the Stock–Yogo 10% critical value for a single instrument, reflecting the reduced sample size and the elimination of the extreme urban contrasts that drive much of the full-sample first-stage variation. The 1879 results are the weakest and only marginally significant, consistent with the full-sample pattern in which effects grow over time as the *Armenwet* mechanism compounds.

A.9 Geographic matching

The propensity score match of the previous section balances monastery and non-monastery municipalities on a vector of *measured* pre-treatment characteristics. A complementary concern is balance on *unmeasured* but spatially structured confounders—regional soil and market access, the administrative legacy of the historical provinces and quarters, and the diffusion of confessional culture across contiguous parishes—which a covariate-based score cannot directly target. To address this I implement a purely geographic matching strategy: each municipality with at least one pre-1578 monastery is matched to its five nearest municipalities without a monastery, where distance is the planar distance between municipal centroids projected onto the Dutch national grid (EPSG:28992). Matching is performed with replacement, so a control municipality bordering several monastery municipalities may serve as a neighbour for more than one of them; control weights are accumulated accordingly and normalised so that the total treated and control weights are equal. The main IV specifications are then re-estimated on this spatially matched sample.

Table A.8: IV Estimates on Geographically Matched Sample: Religious Disaffiliation

Share geen gezindte:	1879	1909	1930
	(1)	(2)	(3)
Charitable associations per 1,000 capita (active 1854)	0.006	0.061**	0.220***
	(0.005)	(0.025)	(0.084)
Adj. R^2	-0.307	0.243	0.142
Within R^2 (adj.)	-0.697	-0.613	-0.428
Observations	663	663	634
Province FE	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	16.6	16.6	15.6

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Baseline controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns. Geographically matched sample: each municipality with a pre-1578 monastery is matched to its 5 nearest municipalities without one (centroid distance, with replacement). Estimates weighted by matching weights.

The logic is that of a local comparison. By construction, a monastery municipality and its five nearest non-monastery neighbours sit within a small geographic radius—the median matched neighbour lies roughly six kilometres away—and therefore share the slowly varying regional environment almost by definition. Differences in post-1854 secularization between a monastery municipality and its immediate neighbours are thus far less plausibly driven by broad regional confounders, since those are held approximately fixed within each matched neighbourhood. The design trades the covariate balance of the propensity score for balance on *location*, and the two exercises are therefore informative precisely because they hold different things fixed.

Of 1,117 municipalities with a defined treatment indicator and usable centroid coordinates, 156 monastery municipalities are matched to 512 distinct control municipalities. Table A.8 reports the IV estimates for the *geen gezindte* outcomes on this sample, and Table A.9 the corresponding estimates for the religious diversity index (1−HHI); both retain province fixed effects, the full baseline control vector, and matching weights.

The geographically matched estimates reproduce the full-sample pattern closely. For *geen gezindte* in 1930 the IV coefficient is 0.220, almost identical to the full-sample value of 0.211 and considerably below the up-weighted propensity score estimate of 0.318; the 1909 coefficient (0.061) is likewise significant, and the 1879 coefficient is small and insignificant, again consistent with the gradual compounding of the *Armenwet* mechanism. For the religious diversity index the 1930 coefficient is 0.236 (full sample 0.262), with the 1879 and 1899 estimates positive and significant or marginally so. Two features distinguish this exercise from the propensity score match. First, because geographic matching does not discard the within-region contrast between historically urbanised monastery

Table A.9: IV Estimates on Geographically Matched Sample: Religious Diversity (1–HHI)

Religious diversity (1–HHI):	1879	1899	1930
	(1)	(2)	(3)
Charitable associations per 1,000 capita (active 1854)	0.233**	0.188*	0.236**
	(0.117)	(0.100)	(0.101)
Adj. R^2	-0.132	0.014	0.202
Within R^2 (adj.)	-0.941	-1.104	-1.005
Observations	663	663	634
Province FE	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	16.6	16.6	15.6

* p < 0.1, ** p < 0.05, *** p < 0.01
Baseline controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns. Geographically matched sample: each municipality with a pre-1578 monastery is matched to its 5 nearest municipalities without one (centroid distance, with replacement). Estimates weighted by matching weights.

towns and their rural neighbours, it preserves much more of the variation that identifies the first stage: the Kleibergen–Paap F-statistic is roughly 16, comfortably above the Stock–Yogo threshold, rather than the value of about 6 obtained under propensity score matching. Second, the matched-sample magnitudes are very close to the full-sample magnitudes rather than inflated, indicating that the full-sample estimates are not an artefact of comparing distant, structurally dissimilar municipalities. That the result survives a comparison restricted to immediate geographic neighbours is reassuring evidence against confounding by unobserved regional characteristics.

A.10 Oster (2019) bounds for selection on unobservables

Tables A.10 and A.11 report Oster (2019) bounds for selection on unobservables, computed using the `robomit` package for R.[10] δ^* is the degree of selection on unobservables, relative to selection on observables, that would be required to drive the estimated treatment effect to zero. $\delta^* > 1$ is the conventional threshold for robustness: it indicates that selection on unobservables would need to be *stronger* than selection on observables to eliminate the result. β^* is the bias-adjusted treatment effect under the assumption of equal selection on observables and unobservables ($\delta = 1$). $R_{\max}$ is set to $\min(1, 1.3 \times \tilde{R}^2)$ following Oster's recommendation.

For the *geen gezindte* outcomes (Table A.10), δ^* ranges from 1.18 (1879) to 1.55

[10]Because `robomit` uses `lm` internally, the province fixed effects are partialled out by pre-residualisation before estimation; the resulting R^2 values correspond to the within-R^2 from the `fixest` specifications.

Table A.10: Oster (2019) Bounds for Selection on Unobservables: Religious Disaffiliation. δ^* is the degree of selection on unobservables (relative to observables) needed to drive the treatment effect to zero. $\delta^* > 1$ indicates the result is robust to unobserved confounding of similar or smaller magnitude than observed confounding. β^* is the bias-adjusted treatment effect under $\delta = 1$ (equal selection on observables and unobservables). $R_{\max} = \min(1, 1.3 \times \tilde{R}^2)$ following Oster (2019).

Year	$\dot{\beta}$	$\tilde{\beta}$	$\dot{R}^2$	$\tilde{R}^2$	$R_{\max}$	δ^*	β^*	N
1879	0.0038	0.0038	0.024	0.031	0.040	1.18	0.0043	1107
1909	0.0156	0.0146	0.046	0.070	0.091	1.55	0.0124	1105
1930	0.0448	0.0290	0.073	0.142	0.185	1.51	0.0142	1064

Province fixed effects partialled out before estimation. Controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Estimates via robomit R package (Oster 2019, JBES).

(1909), exceeding unity for all three census years. The bias-adjusted treatment effects β^* remain positive throughout, although they are attenuated relative to the controlled estimates: for 1930, $\beta^* = 0.014$ compared to the controlled coefficient of 0.029. For the religious diversity outcomes (Table A.11), δ^* ranges from 0.77 (1879) to 1.19 (1930). The diversity results are thus somewhat more sensitive to unobserved selection, particularly in the earlier census years where the overall variation in religious diversity explained by the model is modest.

A.11 Spatial unit root diagnostics and polynomial controls

A potential concern with cross-sectional regressions on geographically distributed units is spatial non-stationarity. If both an outcome and its predictor follow a spatial random walk—a spatial $I(1)$ process in the sense of Mueller and Watson (2024)—ordinary least squares may detect a significant association purely as an artefact of correlated spatial trends, even when no causal relationship exists (Mueller and Watson, 2024). This section addresses that concern in two ways.

Table A.12 reports Mueller–Watson (2024) spatial unit root diagnostics for each of the six main outcome variables, computed using the `spatialunitroot` package for R. Centroid coordinates are projected onto the Dutch RD New grid (EPSG:28992, metres) so that Euclidean distances match actual ground distances. For each outcome the table reports the

Table A.11: Oster (2019) Bounds for Selection on Unobservables: Religious Diversity (1–HHI). δ^* is the degree of selection on unobservables (relative to observables) needed to drive the treatment effect to zero. $\delta^* > 1$ indicates the result is robust to unobserved confounding. β^* is the bias-adjusted treatment effect under $\delta = 1$. $R_{\max} = \min(1, 1.3 \times \tilde{R}^2)$.

Year	$\dot{\beta}$	$\tilde{\beta}$	$\dot{R}^2$	$\tilde{R}^2$	$R_{\max}$	δ^*	β^*	N
1879	0.0713	0.0206	0.032	0.130	0.169	0.77	-0.0070	1107
1899	0.0642	0.0222	0.025	0.100	0.130	0.94	-0.0015	1107
1930	0.0717	0.0320	0.033	0.093	0.120	1.19	0.0064	1064

Province fixed effects partialled out before estimation. Controls: Jewish population share 1809; Municipal area (km^2); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Estimates via robomit R package (Oster 2019, JBES).

LR test statistic and Monte Carlo p-value (10,000 replications) for four test variants:

- *I(1) raw*: H_0 = spatial unit root present; rejection indicates spatial stationarity of the raw series.
- *I(0) raw*: H_0 = spatial stationarity; rejection indicates a spatial unit root.
- *I(1) resid.*: same H_0 applied to OLS residuals after conditioning on the baseline controls and province fixed effects.
- *I(0) resid.*: H_0 = residuals are spatially stationary; non-rejection is the "good" outcome consistent with no spurious regression.

The raw-series tests in the first two columns deliver an unambiguous verdict: for all six outcomes the $I(1)$ null cannot be rejected (Monte Carlo $p \geq 0.42$) while the $I(0)$ null is decisively rejected ($p \leq 0.005$). The raw *geen gezindte* and diversity series therefore behave as spatial unit roots, reflecting the strong, slowly varying north–south and urban–rural gradients that characterise Dutch religious geography. Taken at face value, this is precisely the configuration under which a naive cross-sectional regression of one such series on another would be at risk of spurious correlation.

The residual tests in the last two columns ask whether the baseline conditioning set removes that risk, and the answer differs across the two outcomes. For *geen gezindte*, conditioning on province fixed effects and the baseline controls renders the residuals spatially stationary: the $I(0)$ null is never rejected ($p = 0.26, 0.99, 0.96$), and the $I(1)$ null is rejected in the two later census years that carry most of the identifying signal (1909, $p = 0.001$; 1930, $p = 0.03$). The secularization regressions are thus not spurious—the spatial trend common to outcome and regressors is absorbed by the controls and fixed effects, leaving stationary residuals. For the religious diversity index the evidence is weaker: the residual $I(1)$ null is not rejected in any year, and the residual $I(0)$ null is itself rejected in 1899 ($p = 0.03$) and 1930 ($p = 0.01$), indicating that some spatial dependence survives the conditioning set. This residual spatial structure in the diversity outcome is of a piece with the other diagnostics that mark the diversity results as the more fragile of the two—the below-unity Oster δ^* in the earlier years (Appendix A.10) and the pre-existing 1809 diversity imbalance documented in the placebo analysis (Appendix A.5)—and counsels treating the diversity estimates with corresponding caution.

Tables A.13 and A.14 provide the constructive counterpart to these diagnostics. Each re-estimates the baseline IV specification after adding a second-degree polynomial in the projected centroid coordinates—the terms x, y, x^2, y^2, and xy, where x and y are standardised RD New easting and northing—to the control vector. These five terms span the space of smooth quadratic surfaces over the Netherlands and absorb arbitrary north–south and east–west gradients at a finer scale than the province indicators, directly targeting the kind of smooth spatial trend that the diversity residual tests detect. The IV estimates are nonetheless stable: the 1930 coefficient is 0.216 for *geen gezindte* (against 0.211 in the baseline) and 0.233 for religious diversity (against 0.262), both significant at the 1% level, and the first-stage Kleibergen–Paap F-statistics remain strong at roughly 18. The instrument's identifying variation thus operates through the within-province, within-smooth-surface contrast between monastery and non-monastery municipalities rather than through the broad regional gradients the polynomial absorbs, and the diversity result in particular survives explicit control for the spatial trend that its residual $I(0)$ test flags.

A second and more demanding remedy for the residual spatial dependence is to attack the spatial unit root at its source rather than to control for it. Table A.15 re-estimates the main IV specifications on data transformed by the Mueller–Watson (2024) nearest-neighbour spatial differencing operator. This transformation replaces every variable—the outcome, the endogenous regressor, the instrument, and each control—by its deviation from a local average of its nearest spatial neighbours, the spatial analogue of first differencing a time series, thereby removing the spatial unit root before any regression is run. Because the differencing itself absorbs the low-frequency spatial trend, the province fixed effects are dropped; identification now rests entirely on highly local, neighbour-to-neighbour contrasts. The results are supportive of the hypothesis on both outcomes. The differenced IV coefficients are positive in every census year and remain comparable in magnitude to the baseline estimates, and the first stage stays informative (Kleibergen–Paap $F \approx 11$–13). For secularization the estimates retain conventional significance in the two later years (*geen gezindte* coefficient 0.197, $p < 0.01$ in 1930; 0.056, $p = 0.04$ in 1909). For religious diversity the 1930 coefficient is 0.189 and significant at the 10% level ($p = 0.06$); the 1879 and 1899 coefficients (0.146 and 0.139) keep the same sign and order of magni-

Table A.12: Spatial Unit Root Diagnostics

Outcome	Raw series		OLS residuals	
Outcome	I(1)	I(0)	I(1), resid.	I(0), resid.
No religion share, 1879	5.716 [0.423]	2.751*** [0.004]	852.884 [0.747]	5.691 [0.261]
No religion share, 1909	4.440 [0.668]	3.699*** [0.000]	2075.779*** [0.001]	3.623 [0.990]
No religion share, 1930	2.383 [0.971]	5.094*** [0.000]	1713.597** [0.029]	4.539 [0.957]
Relig. diversity, 1879	5.712 [0.424]	3.257*** [0.001]	1343.129 [0.210]	5.965 [0.185]
Relig. diversity, 1899	5.124 [0.534]	3.439*** [0.000]	1130.002 [0.424]	7.284** [0.031]
Relig. diversity, 1930	4.286 [0.699]	3.689*** [0.000]	951.428 [0.632]	9.568** [0.011]

Mueller–Watson (2024) spatial unit root tests (`spatialunitroot` package). Centroid coordinates are in the Dutch RD New projection (EPSG:28992, metres). Columns I(1) and I(0) test the raw outcome series under H_0: spatial unit root present [I(1)] and H_0: spatial stationarity [I(0)], respectively. Columns I(1) resid. and I(0) resid. apply the same tests to residuals from the baseline OLS specification (baseline controls and province fixed effects partialled out). Each cell reports the LR statistic and, in brackets, the Monte Carlo p-value (50 000 replications). *, **, *** denote significance at 10%, 5%, 1%.

tude but are no longer significant at conventional levels once the spatial trend is purged. That the secularization results are essentially untouched while the diversity results survive in attenuated but still positive form is exactly the pattern the unit root diagnostics anticipate: the secularization residuals were already spatially stationary, whereas part—though by no means all—of the raw diversity gradient reflected the common spatial trend that the differencing removes. Even under this stringent transformation, the surviving estimates continue to point in the direction predicted by the participation-cost mechanism.

B A Model of Religious Associations, Commitment Screening, and Secularization

B.1 Motivation

A large literature in the economics of religion treats religious organizations as providers of club goods, and religious markets as subject to the same supply-and-demand logic as secular markets (Azzi and Ehrenberg, 1975; Iannaccone, 1998). The canonical result—that "strict churches are strong"—rests on a screening logic: costly participation requirements filter out free-riders, leaving behind a core of genuine believers (Iannaccone, 1994; Berman, 2000).

The Dutch evidence confronts us with the population-level consequences of this mechanism. Municipalities with denser networks of charitable associations exhibit *higher* rates of secularization and *greater* religious diversity—lower confessional concentration—over 1879–1930. These two findings are jointly consistent with commitment screening operating through a single, observable dimension: the *relative* organizational density of the dominant denomination. A household with weak religious conviction optimally exits the denominational system when the cost of nominal affiliation exceeds its returns. And be-

Table A.13: IV Estimates with Spatial Polynomial Controls: Religious Disaffiliation

Share geen gezindte:	1879	1909	1930
	(1)	(2)	(3)
Charitable associations per 1,000 capita (active 1854)	0.010**	0.064***	0.216***
	(0.004)	(0.023)	(0.072)
Adj. R^2	0.180	0.421	0.499
Within R^2 (adj.)	0.014	0.053	0.154
Observations	1107	1105	1064
Reduced form (Pre-1578 monasteries)	0.000**	0.001***	0.005***
First stage (Pre-1578 monasteries)	0.021***	0.021***	0.021***
Province FE	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	18.5	18.5	17.7

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Baseline controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns. A second-degree spatial polynomial in projected centroid coordinates (standardised Dutch RD New easting x and northing y: terms x, y, x^2, y^2, xy) is included as an additional control to absorb smooth spatial trends in religious composition.

Table A.14: IV Estimates with Spatial Polynomial Controls: Religious Diversity (1−HHI)

Religious diversity (1−HHI):	1879	1899	1930
	(1)	(2)	(3)
Charitable associations per 1,000 capita (active 1854)	0.222**	0.179**	0.233***
	(0.100)	(0.087)	(0.088)
Adj. R^2	0.498	0.562	0.657
Within R^2 (adj.)	0.168	0.151	0.160
Observations	1107	1107	1064
Reduced form (Pre-1578 monasteries)	0.005***	0.004**	0.005***
First stage (Pre-1578 monasteries)	0.021***	0.021***	0.021***
Province FE	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	18.5	18.5	17.7

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$
Baseline controls: Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Province fixed effects included in all columns. A second-degree spatial polynomial in projected centroid coordinates (standardised Dutch RD New easting x and northing y: terms x, y, x^2, y^2, xy) is included as an additional control to absorb smooth spatial trends in religious composition.

Table A.15: Spatially-Differenced IV Estimates (Nearest-Neighbour Transformation)

	Share geen gezindte			Religious diversity (1–HHI)		
Year:	1879	1909	1930	1879	1899	1930
	(1)	(2)	(3)	(4)	(5)	(6)
Charitable associations per 1,000 capita (active 1854)	-0.001	0.056**	0.197***	0.146	0.139	0.189*
	(0.007)	(0.027)	(0.068)	(0.104)	(0.107)	(0.100)
Adj. R^2	0.018	0.086	0.171	0.142	0.103	0.092
Observations	1107	1105	1064	1107	1107	1064
Spatial differencing (nearest-neighbour)	Yes	Yes	Yes	Yes	Yes	Yes
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes
Kleibergen–Paap F-statistic (first stage)	11.2	11.2	12.6	11.2	11.2	12.6

* p < 0.1, ** p < 0.05, *** p < 0.01
Baseline controls (spatially differenced): Jewish population share 1809; Municipal area (km²); Average elevation (m); Terrain ruggedness index; Crop suitability (PCA); Distance to coast (m); Distance to nearest river (m); Catholic mission presence; Medieval city by 1560; Eighty Years' War battle. Estimated on Mueller–Watson (2024) nearest-neighbour spatially-differenced data: the outcome, the endogenous regressor (charitable associations per capita), the instrument (pre-1578 monasteries), and all controls are spatially differenced to remove the spatial unit root before estimation. Province fixed effects are dropped because the differencing absorbs the low-frequency spatial trend. Standard errors are HC1.

cause the dominant denomination commands a disproportionate share of local associations, it imposes the steepest participation costs and sheds its weakly-attached members *proportionally faster* than the minority does; the affiliated population is therefore pulled toward parity, and diversity rises even as aggregate affiliation falls. Both signs flow from the same force—no auxiliary assumption about the composition of either denomination's base is required.

Crucially, religious associations were not only screening devices: under the 1854 *Armenwet* they were the principal providers of *material* welfare—poor relief, schooling, sick funds, burial insurance. A denser associational network therefore delivers a larger bundle of tangible services, which *retains* marginal members rather than screening them out. The model below incorporates this welfare benefit explicitly, so that affiliation reflects a genuine tension between a *screening* channel (participation costs rise with density) and a *bundling* channel (welfare benefits rise with density). The empirical sign of the net effect is then a substantive prediction rather than a built-in assumption: Proposition 1 holds precisely when, at the margin, the cost of participation rises faster than the value of the services bundled with it, a condition guaranteed for sufficiently dense networks. The companion prediction that diversity *rises* (Proposition 2) follows from the same screening force—the dominant denomination's higher relative participation cost makes it shed members proportionally faster—and likewise holds in the screening regime, so that both signs are predictions the data adjudicate rather than assumptions.

B.2 Setup

Consider a municipality with a unit mass of households $i \in [0,1]$ and two denominations $k \in \{D, M\}$ (dominant and minority). Let $A > 0$ denote total associational intensity (associations per 1,000 inhabitants), and let $\lambda \in (1/2, 1)$ denote the dominant denomination's share of local associations, so that denomination D operates λA associations and denomination M operates $(1-\lambda)A$.

Households. Each household i is characterized by a pair (θ_i, δ_i):

- **Spiritual taste** $\theta_i \sim U[0,1]$, drawn independently across households, captures intrinsic valuation of denominational membership (beliefs, family tradition, social networks) net of material considerations.
- **Denominational affinity** $\delta_i \in \{D, M\}$ indicates the household's natural denomination, i.e., the one it would join absent cost differences. A fraction $\mu > 1/2$ of households has affinity D (the dominant denomination is also the majority confession).

We assume θ_i and δ_i are independent, so the distribution of spiritual taste is identical across denomination-affinity groups.

Participation Costs. Membership in denomination k requires active participation—attendance, financial contributions, behavioral compliance—proportional to the denomination's associational density. Following Iannaccone (1994), we model the participation cost for a member of denomination k as:

$$c_k = c(\lambda_k A), \qquad c' > 0, \quad c(0) = 0, \tag{3}$$

where $\lambda_D = \lambda$ and $\lambda_M = 1-\lambda$. Because $\lambda > 1/2$, the dominant denomination is *costlier* to join: $c_D \equiv c(\lambda A) > c((1-\lambda)A) \equiv c_M$ whenever $A > 0$.

Welfare Benefit. Affiliation also confers access to the material services the denomination provides through its associations—poor relief, schooling, sick and burial funds. Because under the *Armenwet* relief was administered confessionally (Catholic charities served the Catholic poor, Protestant deaconries the Protestant poor), the benefit a member of denomination k obtains scales with that denomination's *own* associational density:

$$b_k = b(\lambda_k A), \qquad b' > 0, \quad b'' \leq 0, \quad b(0) = 0, \tag{4}$$

so the welfare bundle expands with density but at a diminishing rate (congestion of relief rolls, saturation of need). It is convenient to collect the cost and benefit into a *net participation cost*

$$g(x) \;\equiv\; c(x) - b(x), \qquad g'' = c'' - b'' \geq 0, \quad g(0) = 0, \tag{5}$$

which is convex because c is convex and b is concave. Its derivative $g'(x) = c'(x) - b'(x)$ measures whether, at the margin, the participation cost of additional density outweighs the extra services it buys. Since c' is non-decreasing and b' is non-increasing, g' is non-decreasing. We assume an Inada-type condition—$\lim_{x\to\infty} c'(x) > \lim_{x\to\infty} b'(x)$ (it suffices that $c'(x) \to \infty$)—so that g' is eventually positive and therefore crosses zero at a finite *regime threshold* $x^\dagger \geq 0$ with $g'(x^\dagger) = 0$; the crossing is unique when g is strictly convex. For $x > x^\dagger$ screening dominates ($g' > 0$), for $x < x^\dagger$ bundling dominates ($g' < 0$).

Payoffs. Remaining unaffiliated (*geen gezindte*) yields a reservation utility normalized to zero:

$$U_i^0 = 0. \tag{6}$$

Affiliating with denomination k yields:

$$U_{ik} = \theta_i - c_k + b_k + \mathbf{1}[\delta_i = k] \cdot \phi = \theta_i - g(\lambda_k A) + \mathbf{1}[\delta_i = k] \cdot \phi, \tag{7}$$

where $\phi > 0$ is the *affinity premium*—the additional utility a household derives from joining its natural denomination (shared liturgical tradition, family ties, social network overlap). This premium introduces a switching cost that disciplines denomination choice: households prefer their natural denomination unless the net-cost differential is large enough to overcome ϕ. The welfare benefit b_k, by contrast, accrues to every member regardless of spiritual taste θ_i; it therefore shifts the affiliation margin uniformly and works *against* the screening force embodied in c_k.

B.3 Equilibrium

Denomination Choice Conditional on Affiliation. A household with affinity $\delta_i = D$ prefers denomination D over M if and only if:

$$\theta_i - g(\lambda A) + \phi \;\geq\; \theta_i - g((1-\lambda)A) \quad \Longleftrightarrow \quad \phi \;\geq\; g(\lambda A) - g((1-\lambda)A) \equiv \Delta g(A, \lambda).$$

In the screening-dominant region ($\lambda_k A > x^\dagger$) the net cost g is increasing, so $\Delta g > 0$ because $\lambda > 1/2$. For households with affinity D, the condition holds whenever $\phi \geq \Delta g$, which we assume throughout (the affinity premium is large relative to the cross-denomination net-cost gap). Symmetrically, a household with affinity M prefers denomination M whenever $\phi \geq \Delta g$, which holds by the same assumption.

Implication: In equilibrium, every affiliated household joins its natural denomination. This clean separation allows us to analyze affiliation and denomination choice independently.

The Affiliation Threshold. Household i with affinity $\delta_i = k$ affiliates with denomination k if and only if:

$$\theta_i - g(\lambda_k A) + \phi \;\geq\; 0 \quad \Longleftrightarrow \quad \theta_i \;\geq\; g(\lambda_k A) - \phi.$$

Define the affiliation thresholds:

$$\theta_k^*(A) \equiv \max\{g(\lambda_k A) - \phi,\ 0\} = \max\{c(\lambda_k A) - b(\lambda_k A) - \phi,\ 0\}. \tag{8}$$

A household with affinity k and spiritual taste $\theta_i < \theta_k^*(A)$ exits into *geen gezindte*; one with $\theta_i \geq \theta_k^*(A)$ affiliates with denomination k. The welfare benefit *lowers* the affiliation threshold: the larger is b, the more weakly-attached households are retained by the material services their denomination bundles with membership.

In the screening-dominant region ($\lambda_k A > x^\dagger$, so g is increasing) and since $\lambda > 1/2$ implies $g(\lambda A) > g((1-\lambda)A)$, we have:

$$\theta_D^*(A) \geq \theta_M^*(A) \geq 0,$$

with strict inequality whenever A is large enough that $g((1-\lambda)A) > \phi$ holds—that is, both denominations impose strictly positive affiliation thresholds, and the dominant denomination's is higher.[11]

B.4 Main Results

Proposition 1 (Associations Increase Secularization in the Screening Regime)**.**

The equilibrium share of unaffiliated households is strictly increasing in total associational intensity A if and only if the screening channel dominates the bundling channel at the margin, i.e.

$$\mu\,\lambda\,g'(\lambda A) + (1-\mu)\,(1-\lambda)\,g'((1-\lambda)A) > 0, \qquad g' = c' - b'. \tag{9}$$

A sufficient condition is $A > x^\dagger/(1-\lambda)$, so that $g'(\lambda_k A) > 0$ for both denominations; and under the maintained Inada condition the threshold density $\bar{A} \equiv x^\dagger/(1-\lambda)$ is finite.

Proof. The overall unaffiliated share is:

$$s(A) = \mu\,\theta_D^*(A) + (1-\mu)\,\theta_M^*(A),$$

since the uniform distribution on $[0,1]$ makes the exiting fraction of each affinity group equal to its threshold. With interior thresholds, $\partial\theta_k^*/\partial A = \lambda_k\, g'(\lambda_k A)$, so

$$\frac{\partial s}{\partial A} = \mu \cdot \lambda\, g'(\lambda A) + (1-\mu) \cdot (1-\lambda)\, g'((1-\lambda)A),$$

which is the left-hand side of (9). Each term carries the sign of $g'(\lambda_k A)$. Since g' is increasing with unique zero $x^\dagger$, both terms are positive when $\lambda_k A > x^\dagger$ for $k \in \{D, M\}$; the binding case is the minority denomination ($\lambda_M = 1-\lambda$ smallest), giving the sufficient

[11] This is consistent with the maintained assumption $\phi > \Delta g = g(\lambda A) - g((1-\lambda)A)$: combining $g((1-\lambda)A) > \phi$ with $\phi > \Delta g$ yields $g(\lambda A) < 2\,g((1-\lambda)A)$, a mild restriction satisfied for interior values of A.

condition $A > x^\dagger/(1-\lambda)$. The Inada condition guarantees g' crosses zero at a finite $x^\dagger$, hence $\bar{A} < \infty$. ■

Two opposing forces. The condition makes the economics transparent. Additional associations exert two opposing forces on affiliation. The *screening* force raises the participation cost $c(\lambda_k A)$ and pushes weakly-attached households out; the *bundling* force raises the welfare benefit $b(\lambda_k A)$ and pulls them back in. The net effect on secularization is positive exactly when marginal cost outruns marginal benefit, condition (9). Because c is convex (congestion, organizational complexity), b concave (diminishing returns to relief), and marginal cost eventually outstrips marginal benefit (the Inada condition), the cost curve must overtake the benefit curve: there is a finite density $\bar{A}$ beyond which screening necessarily dominates, and below which marginal associations are net-*retentive*. The sign of the relationship between associational density and secularization is thus a substantive prediction of the model, not a maintained assumption: it turns on whether a municipality sits above or below $\bar{A}$. A positive empirical association between density and the *geen gezindte* share identifies the relevant municipalities as operating in the screening-dominant region $A > \bar{A}$, while the net-retentive prediction for $A < \bar{A}$ matches the welfare role of confessional charity in thinly organized municipalities.

Interpretation. In the screening regime, rising associational density increases *net* participation costs for both denominations faster than it expands the welfare bundle. Households with weak spiritual conviction—those with low θ_i—find that the cost of nominal membership now exceeds its combined material and social return, and optimally exit into *geen gezindte*. This formalizes the "nominal exit" mechanism documented by Hout and Fischer (2002) and Hout and Fischer (2014) for the United States: the growth of religious "Nones" draws disproportionately on those who were affiliated by social habit rather than conviction.

Proposition 2 (Associations Raise Religious Diversity)**.**

In the screening regime ($A > \bar{A}$), an increase in total associational intensity A lowers *the dominant denomination's conditional share among affiliated households—and hence* lowers *the Herfindahl–Hirschman Index (HHI) of denominational concentration, raising the diversity index $D = 1 - \mathrm{HHI}$—because the dominant denomination sheds its members proportionally faster than the minority,*

$$\underbrace{\frac{\lambda\, g'(\lambda A)}{n_D(A)}}_{\rho_D} > \underbrace{\frac{(1-\lambda)\, g'((1-\lambda)A)}{n_M(A)}}_{\rho_M}, \tag{10}$$

where $n_D(A) = 1 - \theta_D^*$ and $n_M(A) = 1 - \theta_M^*$ are the denominations' surviving bases.

Proof. The conditional share of denomination D among affiliated households is

$$\sigma_D(A) \;=\; \frac{\mu\, n_D(A)}{\mu\, n_D(A) + (1-\mu)\, n_M(A)},$$

with $n_k(A) = 1 - \theta_k^*(A)$ the survival rate of denomination k and, for interior thresholds, $\partial n_k/\partial A = -\lambda_k\, g'(\lambda_k A) < 0$. Write the *proportional* exit rates as $\rho_k \equiv -(\partial n_k/\partial A)/n_k$, so that $\rho_D = \lambda\, g'(\lambda A)/n_D$ and $\rho_M = (1-\lambda)\, g'((1-\lambda)A)/n_M$.

Lemma. *If c is weakly convex ($c'' \geq 0$), b weakly concave ($b'' \leq 0$), and $\lambda > 1/2$, then in the screening regime $\rho_D(A) > \rho_M(A)$ whenever $\theta_D^*(A) > \theta_M^*(A) > 0$.* *Proof of Lemma.* The net cost $g = c - b$ has $g'' \geq 0$, so g' is non-decreasing; with $\lambda > 1-\lambda$ this gives $\lambda\, g'(\lambda A) \geq (1-\lambda)\, g'((1-\lambda)A)$—the dominant denomination's absolute marginal loss is the larger. Simultaneously $n_D = 1 - \theta_D^* < 1 - \theta_M^* = n_M$ because $\theta_D^* > \theta_M^*$: the dominant denomination starts from a smaller, more heavily pre-screened base. A weakly larger numerator over a strictly smaller denominator yields $\rho_D > \rho_M$. □

Differentiating the conditional share,

$$\frac{\partial \sigma_D}{\partial A} = \frac{\mu(1-\mu)\big(n_M\, \partial n_D/\partial A - n_D\, \partial n_M/\partial A\big)}{\big(\mu n_D + (1-\mu) n_M\big)^2},$$

and the numerator equals $-n_D n_M\,(\rho_D - \rho_M)$. By the Lemma $\rho_D > \rho_M$, so the numerator is *negative* and $\partial\sigma_D/\partial A < 0$: the dominant share falls toward parity as density rises.

The HHI among affiliated households is $H(A) = \sigma_D(A)^2 + (1-\sigma_D(A))^2$, which is strictly increasing in σ_D for $\sigma_D > 1/2$. Because the dominant denomination is the majority confession ($\mu > 1/2$), it remains the plurality among the affiliated throughout the empirically relevant region, so $\sigma_D > 1/2$; as σ_D falls toward $1/2$ the HHI falls and the diversity index $D = 1 - H$ rises with A. ■

Interpretation. The dominant denomination is both costlier at the margin—its associations are denser, so $\lambda\, g'(\lambda A) > (1-\lambda)\, g'((1-\lambda)A)$ whenever $\lambda > 1/2$—and more heavily pre-screened, since its higher threshold θ_D^* leaves it a thinner committed core. Both forces make it shed members *proportionally* faster than the minority as participation costs climb. The affiliated population is therefore drawn toward parity between the confessions: rising associational density does not consolidate the dominant denomination but erodes its relative predominance, raising measured diversity. This is the mirror image of the secularization result—the same screening force that pushes the weakly-attached out of religion altogether falls hardest, in relative terms, on the confession that organizes most intensively. No assumption about an asymmetrically weak minority base is needed; the result follows from $\lambda > 1/2$ and the convexity of the net cost alone.

Corollary (Joint Predictions)**.**

In the screening regime ($A > \bar{A}$), an increase in total associational intensity A (i) raises the unaffiliated share $s(A)$ and (ii) lowers the HHI of confessional concentration among the affiliated population—raising religious diversity. The secularization effect (i) is larger in municipalities with higher dominant-denomination concentration λ.

Proof. Part (i) follows from Proposition 1. Part (ii) follows from Proposition 2. For the comparative static in λ: $\partial^2 s/\partial A \partial\lambda = \mu\, g'(\lambda A) + \mu\lambda\, g''(\lambda A) A - (1-\mu)\, g'((1-\lambda)A) - (1-\mu)(1-$

$\lambda)\, g''((1-\lambda)A)A$, which is positive under $g'' \geq 0$, $g' > 0$, and $\mu > 1/2$, so the secularization effect steepens with λ. A higher λ also widens the marginal-cost gap $\lambda\, g'(\lambda A) - (1-\lambda)\, g'((1-\lambda)A)$, deepening the gap $\rho_D - \rho_M$ in (10); the diversity effect therefore also strengthens with λ. ■

B.5 Discussion

Strict Churches and the Population-Level Dark Side of Screening. Iannaccone (1994) shows that high participation requirements produce committed memberships and strong organizational performance. Proposition 1 does not contradict this result: within the affiliated population, municipalities with denser associations retain more committed members by construction ($\theta_i \geq \theta_k^*$). The point is that the *marginal* effect on population-level secularization is positive, because the intensive margin of commitment is purchased at the cost of extensive margin exits.

Competition, Monopoly, and the *Armenwet* Context. Iannaccone (1991) and Stark and Finke (2000) document that competitive religious markets exhibit higher aggregate participation than monopolistic ones. Our model is consistent with this finding but adds a within-market dimension: even under the statutory welfare monopoly of the 1854 *Armenwet*, higher organizational intensity sorts out the weakly attached and, because it taxes the dominant denomination most heavily, erodes that denomination's relative predominance among the affiliated. The monopoly context removes the safety valve of denominational switching (which our model captures via the large ϕ assumption), channeling exits directly into the unaffiliated category rather than redistributing them across denominations.

Bundling, Screening, and the Welfare Role of Charity. The material benefit $b(\cdot)$ makes the two channels of the model explicit. Under the *Armenwet*, confessional charity was the chief route to poor relief, so b was large and the bundling channel partly offset screening: dense networks retained marginal members who valued the services more than they resented the participation cost. Proposition 1 holds only where the screening channel nonetheless dominates—empirically, the positive reduced-form sign tells us it did.

Long-Run Secularization. Propositions 1 and 2 characterize static equilibrium sorting within the *Armenwet* regime, where b is high. As the Dutch welfare state expanded and secular alternatives to religious poor relief became universally available over the twentieth century, the material benefit of confessional membership eroded—formally, $b(\lambda_k A) \to 0$, so the net cost $g \to c$ and every affiliation threshold θ_k^* jumps upward. Municipalities with higher associational density had retained a larger pool of weakly committed members precisely because their bundle of services was larger; once that bundle was replaced by the secular state, this pool became a reservoir of rapid exiters. The reduction of b

both raises the exit threshold and removes the force that had been retaining the marginal members in the first place, consistent with the sharp and geographically uneven *ontkerkelijking* (de-churching) documented for the Netherlands from the 1960s onward (Becker and Woessmann, 2009; Gill and Lundsgårde, 2004). This comparative static is formalized in the dynamic sketch of Section B.6.1 below.

B.6 Extensions (Sketches)

The two extensions below are deliberately informal. They show that the static screening model nests naturally into a dynamic "reservoir" account and into a regional account that maps onto the Protestant north / Catholic south divide. Full treatments are left for future work; the aim here is to verify that the central comparative statics survive and to derive their additional testable implications.

B.6.1 A Two-Period Dynamic Extension

Index time by $t \in \{1, 2\}$. Period 1 is the *Armenwet* regime; period 2 is the era of the secular welfare state. The only primitive that changes is the welfare benefit:

$$b_1(\lambda_k A) = b(\lambda_k A) > 0 \qquad \Longrightarrow \qquad b_2(\lambda_k A) = (1-\kappa)\, b(\lambda_k A), \quad \kappa \in (0, 1],$$

where κ measures the degree to which the secular state crowds out confessional welfare ($\kappa = 1$ is full replacement). Spiritual taste θ_i and affinity δ_i are fixed; households re-optimize each period. The net cost rises from $g_1 = c - b_1$ to $g_2 = c - (1-\kappa)b > g_1$, so every threshold climbs:

$$\Delta\theta_k^* \;=\; \theta_k^*(A)\big|_{t=2} - \theta_k^*(A)\big|_{t=1} \;=\; \kappa\, b(\lambda_k A) \;>\; 0.$$

The mass of households that exits between periods in denomination k is exactly $\mu_k\, \kappa\, b(\lambda_k A)$ (with $\mu_D = \mu$, $\mu_M = 1-\mu$). Two implications follow. First, the *flow* of new *geen gezindte* between $t = 1$ and $t = 2$ is increasing in A: high-density municipalities had banked a larger retained pool—those with $\theta_i \in [\theta_k^*|_1, \theta_k^*|_2)$—and disgorge it when b collapses. This is the "reservoir" mechanism stated verbally in the conclusion, now with a sign: the disgorged mass in denomination k is $\mu_k\, \kappa\, b(\lambda_k A)$, so $\partial(\Delta s)/\partial A = \kappa\big[\mu\lambda\, b'(\lambda A) + (1-\mu)(1-\lambda)\, b'((1-\lambda)A)\big] > 0$ wherever the thresholds are interior—the driver here is the benefit gradient $b' > 0$, not the net-cost gradient. Second, the period-2 jump in diversity rises with A by the same *proportional* logic as Proposition 2: as b collapses, the dominant denomination—costlier at the margin and more heavily pre-screened—disgorges its remaining members proportionally faster than the minority, so the affiliated population deconcentrates further and the diversity index rises. The model thus predicts that the cross-sectional association documented for 1879–1930 should *steepen* after mid-century— a prediction testable with post-war census waves and consistent with an event-study reading of the coefficient path.

B.6.2 Regional Heterogeneity: Protestant North versus Catholic South

Let the dominant-denomination share vary by region, $\lambda \in \{\lambda_N, \lambda_S\}$, with the historically dominant confession differing in identity (Protestant in the north, Catholic in the south) but both regions satisfying $\lambda_r > 1/2$. Three comparative statics carry over from the Corollary and sharpen the empirics.

1. **Concentration amplifies secularization.** Since $\partial^2 s/\partial A \partial \lambda > 0$, the secularizing effect of associational density is stronger in regions with more lopsided confessional dominance: where one confession is overwhelmingly dominant (λ_r near 1), the screening tax falls almost entirely on that confession's own members. The diversity effect is likewise stronger in such regions: a higher λ_r deepens the dominant denomination's relative marginal losses, widening the gap $\rho_D - \rho_M$ in (10) that drives deconcentration of the affiliated population.

2. **The instrument operates through the dominant confession.** The pre-1578 monastery instrument raised the endowment base of the *Protestant* deaconries that inherited Catholic wealth at the Reformation. In northern municipalities the dominant confession (D) is itself Protestant, so the instrument loads directly onto λA; in southern municipalities the Reformation transfer endowed what became the local *minority* institution, loading onto $(1 - \lambda)A$. The model predicts a stronger first stage—and a stronger reduced form—in the north, because there the instrument moves the costlier, higher-threshold denomination.

3. **Sign of the diversity effect is robust to which confession dominates.** Proposition 2 depends on $\lambda > 1/2$ and the convexity of g—but not on the *identity* of the dominant denomination. The HHI therefore falls with A (diversity rises) in both regions, even though the household exiting at the margin is Protestant in the north and Catholic in the south.

A full regional model would let ϕ, μ, and the benefit function b differ across r; the sketch above isolates the role of λ alone, which is the margin the data speak to most directly through the province fixed effects and the north/south robustness splits.