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"Religious Pluralism and Religious Adherence in U.S. Counties: Assessing the Reassessment"

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Religious Pluralism and Religious Adherence in U.S. Counties: Assessing the Reassessment

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ABSTRACT

We conduct an empirical test of the relationship between religious pluralism and religious participation in U.S. counties using a fixed-effects panel estimation technique. The empirical technique allows us to control for unobserved heterogeneity across counties resulting from various cultural and historical factors. Contrary to prior cross-sectional research on the 1980 and 1990 Glenmary U.S. counties data, we find a significantly positive relationship between pluralism and participation from panel estimation on the same data. However, we also explain how changes between 1980 and 1990 in the composition of denominations in the Glenmary samples can generate a false positive relationship with the panel estimator. The results show the importance for future research on pluralism and participation of data that have a consistent denominational composition across time.

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Religious Pluralism and Religious Adherence in U.S. Counties: Assessing the Reassessment

We conduct an empirical test of the relationship between religious pluralism and religious participation in U.S. counties using a fixed-effects panel estimation technique. The empirical technique allows us to control for unobserved heterogeneity across counties resulting from various cultural and historical factors. Contrary to prior cross-sectional research on the 1980 and 1990 Glenmary U.S. counties data, we find a significantly positive relationship between pluralism and participation from panel estimation on the same data. However, we also explain how changes between 1980 and 1990 in the composition of denominations in the Glenmary samples can generate a false positive relationship with the panel estimator. The results show the importance for future research on pluralism and participation of data that have a consistent denominational composition across time.

INTRODUCTION

For over a decade, debate has continued over the relationship between religious participation and religious pluralism. Finke and Stark (1988) proposed that increased competition in religious markets leads to higher levels of religious participation. Initially adopted as a proxy for the level of competition in religious markets, the "pluralism index"[1] has been a controversial measure fraught with methodological and conceptual pitfalls. The debate has seemingly culminated in Olson's (1999) "reassessment" and Voas, Olson and Crockett's (2002) explanation of "why previous research is wrong," each of which leveled significant methodological criticism at the use of the pluralism index in research on religious participation. This paper addresses the challenges leveled by Olson, Voas, and Crockett and ultimately affirms the need for accurate cross-sectional time series data in order to conduct a proper test of the competition hypothesis using the pluralism index.

[1] As discussed below, the pluralism index is defined as $P = 1\text{-}\Sigma s_i^2$, where s_i represents denomination i's share of all religious adherents.

In this paper, we construct a two-period panel from United States county-level data and employ a widely-accepted panel estimation technique that is methodologically superior to cross-sectional analysis. We then use this methodology to show that a robust positive relationship exists between changes in religious pluralism and changes in religious adherence rates within U.S. counties, results that conflict with prior research using the same data. At least two explanations exist for these results. On one hand, as shown in Voas, Olson, and Crockett (2002), earlier work on pluralism and participation has improperly relied upon cross-sectional analyses. The use of panel data techniques in this study may provide a better statistical assessment than prior cross-sectional work of the role that pluralism plays in these U.S. county religious markets. On the other hand, the data used in this study contain intertemporal variations in sample composition that could induce a mathematically necessary positive relationship between pluralism and participation in our analysis. In the end, our analysis highlights the need for better data in order to conduct reliable tests of the effects of pluralism on participation.

The traditional paradigm in the sociology of religion emphasizes the importance of religious homogeneity. Durkheim ([1897] 1951) argued that multiple competing religious groups discredit each other, with each group's truth claims refuting those of other groups. Berger (1967) extended this theory by asserting that society holds together under a "sacred canopy," which can only exist if all members of a society believe in one faith. To the extent that different groups compete for religious adherents, the "sacred canopy" paradigm asserts that such competition undermines the validity of the varying truth claims, leading to doubts, questions, and ultimately a decline in religious fervor.

Rodney Stark, Roger Finke and others have challenged the traditional paradigm by arguing that religious competition actually increases religious participation. Labeled the "New

Paradigm" by Warner (1993), this religious economies viewpoint conceives of religious denominations as firms that compete with each other to secure market share in the form of adherents. To the degree that a market is unregulated, advocates of the religious economies hypothesis contend that pluralism is the natural state of religion. In this view, pluralism should arise because of diversity in people's preferences for religion, that is, "because of the inability of a single religious firm to be at once worldly and otherworldly, strict and permissive, exclusive and inclusive, expressive and reserved, or . . . austere and loose"[2] (Finke and Stark 2000, 199).

The paper proceeds as follows. The next section summarizes the pluralism/participation debate, with particular emphasis on studies focusing on data from U.S. cities and counties. Next, the data and methodology used in this study are described and the results are presented. After showing a robust positive relationship between pluralism and changes in adherence in U.S. counties, we explain a major flaw in the data and its implications for our results as well as the ongoing study of competition in religious markets.

THE PLURALISM DEBATE

Finke and Stark (1988) sparked the debate over religious pluralism when they used a traditional measure of economic competition, the Herfindahl index, to proxy religious competition in U.S. cities based on Census data from 1906. Used frequently in antitrust enforcement, the Herfindahl index is the sum of the squared market shares of all firms in a market:

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[2] Another key contention of this new paradigm is that monopoly churches tend to be inefficient because they do not comply with market forces (Finke and Stark, 1988, 1998). Verweij *et al.,* (1997, 319) found that "the more a church is supported and controlled by the state, the less people will attend or make use of its services." Iannaccone (1991) showed that, in 12 Protestant nations, pluralism and religious attendance were directly correlated. North and Gwin (2004) provide evidence that lower religious attendance is associated with more regulation of religious markets, including the existence of formally established state churches.

$$H = \sum_{i=1}^{n} s_i^2 ,$$

where s_i is the market share of firm i and n is the number of denominations present in the market. Applied to religious markets, market share is usually defined as the number of adherents in denomination i divided by the total number of religious adherents in the geographical unit. The Herfindahl index ranges between zero and one, and a Herfindahl index near one indicates that religious adherents are highly concentrated in one denomination.[3] The pluralism index is defined as $PI = 1 - H$. Thus, a high pluralism value (i.e., close to one) is the result of a low Herfindahl index, which means that adherents are spread across numerous religious denominations.

Finke and Stark (1988) studied cross-sectional data from the U.S. Census Bureau's *Religious Bodies: 1906*, a census of religious denominations taken in 1906. They initially found a negative correlation between pluralism and religious adherence rates in the 150 largest U.S. cities. However, they attributed this effect to the high concentrations of Catholics in some of the cities. After controlling for the share of Catholics in each city's population, Finke and Stark (1988) found a significant positive relationship between pluralism and adherence. They concluded that "the presence of Catholics and the diversity of the religious market both increase the rate of adherents" (1988, 45).

Breault (1989a) challenged Finke and Stark's (1988) findings with results that showed a negative relationship between religious pluralism and religious adherence rates. Breault used church adherence data (gathered by the Glenmary Research Center) from a cross-section of approximately 3100 counties in the U.S. in 1980. Like Finke and Stark (1988), Breault found a

[3] An interesting practical implication of the Herfindahl index is that it represents the unconditional probability that any two religious adherents drawn at random will be members of the same denomination.

negative simple correlation between religious pluralism and religious adherence rates in these counties. In contrast to Finke and Stark, though, a regression of adherence rates on religious pluralism and the Catholic population share continued to yield a highly significant and negative relationship between religious pluralism and adherence. Thus, the control for Catholic population share did not alter the original negative correlation.

Breault (1989a, 1989b) explained that the reason his results differed from Finke and Stark was the strong negative correlation between the pluralism index and the Catholic population share in the U.S. Cities data from 1906. It was this negative correlation that induced the positive relationship between religious pluralism and religious adherence in Finke and Stark's (1988) multivariate regression. Following the first few exchanges between Finke/Stark and Breault, the literature has been inconclusive. Some papers have found a positive relationship between pluralism and religious adherence, some have found a negative relationship, and some have criticized the entire methodology of the model.[4]

Other studies of pluralism using U.S. data have also yielded inconsistent results. Land, Deane and Blau (1991) and Blau *et al.* (1997) examined data on church membership in the U.S. counties compiled by the U.S. Census Bureau in 1906, 1916 and 1926. For each year, they generated cross-sectional results showing a negative relationship between pluralism and church membership. Olson (1999) continued this line of research by analyzing a complete cross-section of U.S. counties compiled in 1990 by the Glenmary Research Center, which updated the data used by Breault (1989a, 1989b). He found that the only positive relationship between pluralism and the adherence rate occurred in counties with a Catholic population share of more than 10

[4] This research includes Finke and Stark (1989), Iannaccone (1991), Finke and Stark (1992), Stark and McCann (1993), Hamberg and Pettersson (1994), Stark and Iannaccone (1994), Stark, Finke and Iannaccone (1995), Finke, Guest and Stark (1996), Pettersson and Hamberg (1997), Finke and Stark (1998), Phillips (1998), Olson and Hadaway (2000), and North and Gwin (2004). An exhaustive review of the literature on religious pluralism is set forth in Chaves and Gorski (2001), who ultimately side with the critics of the New Paradigm.

percent and only when controlling for the Catholic population share. In regressions on all 3,104 counties and on only the counties that had a Catholic population share of less than or equal to 0.10, the estimated relationship between pluralism and adherence was negative. Another study using locality data from the U.S. is Finke, Guest, and Stark (1996), who found a significant positive relationship between religious pluralism and church attendance rates in a cross-section of 858 New York towns in 1865.

A more important component of Olson's (1999) analysis was his expansion of Breault's earlier critique. Olson charged that the disagreement between Breault (1989a, 1989b) and Finke and Stark (1988) arose from Finke and Stark's use of the Catholic population share as a proxy for the substantive effect of Catholicism on adherence. Olson noted that Finke and Stark (1988) found a positive relationship between pluralism and adherence in U.S. cities in 1906 only after controlling for Catholic population share. He further noted that U.S. cities in 1906 had high concentrations of Catholics. Olson asserted that the correlations between Catholic population share and religious pluralism and between Catholic population share and adherence caused a mathematically necessary negative relationship between pluralism (the independent variable) and adherence (the dependent variable). Returning to the 1990 Glenmary data, the only way he was able to generate a positive coefficient estimate on religious pluralism was to control for Catholic population share in a sample of counties with Catholic population shares above 0.10.[5]

Voas, Olson, and Crockett (2002) argued that a positive or negative relationship between pluralism and participation will exist in a cross-section only because of random historical and cultural factors. Their argument can be summarized using the example in Figure 1, which is an

[5] Montgomery (2003) proposes a partial-ordering method of measuring the competitiveness of a religious market, which he tests using the 1990 Glenmary data and the 1865 New York towns data from Finke, Guest, and Stark (1996). While his measure of competition avoids the mathematical problems of the pluralism index that are described by Olson *et al.*, he still obtains the same basic results of a negative cross-sectional relationship in 1990 U.S. counties and a positive relationship in 1865 New York towns.

abbreviated version of an example presented in Voas *et al.* (2002). In Example 1A, only two denominations exist in two towns. The smaller group (Methodists) is the same size in both towns, while the dominant group (Anglicans) is larger in Town B. In Example 1B, both towns have the same percentage of Anglicans, while variation occurs within the minority denomination (Methodists). In both examples, one town has an adherence rate of 0.70, while the other has an adherence of 0.75. However, the cross-sectional variation in adherence in Example 1A produces a negative correlation between pluralism and adherence, while Example 1B produces a positive correlation. Voas *et al.* (2002) contend that such cross-sectional variation is based only on random historical and cultural factors, and that such analyses provide no meaningful information on the causal connection between pluralism and religious participation.

Summed up, the current state of the pluralism debate is that the continued usefulness of the pluralism index as a proxy for competition is in doubt. Studies of the correlation between religious pluralism and religious participation vary widely in their results, with some data sets showing a positive relationship and others showing a negative relationship. Moreover, from a methodological standpoint, it is unclear that cross-sectional variations in religious pluralism really are able to say anything about competition's effect on religious participation. In a review of the literature, Chaves and Gorski (2001, 274) concluded that "[t]he claim that religious pluralism and religious participation are generally and positively associated with one another . . . is not supported." Our paper steps into the debate at this point and offers a panel data estimation approach that has not previously been used in the pluralism/participation context. In doing so, we demonstrate the possibility of a different and more meaningful interpretation of the relationship between participation and the pluralism index. Unfortunately, compositional variations between the 1980 and 1990 samples of the Glenmary data analyzed in this paper limit the usefulness of applying panel techniques to this data, which we explain below.

DATA AND METHODOLOGY

In this section, we explain the nature of the fixed-effects panel estimator, as well as the benefits that flow from its use in the empirical setting under consideration. Specifically, the use of a fixed-effects panel estimator allows one to interpret the regression coefficient estimate on religious pluralism as a measure of how religious participation changes as pluralism changes *within* a geographical area. Such within-area changes over time capture changing competitive circumstances, and it is reasonable to view any observed statistical relationship with religious participation as causal in nature. Before describing the estimation technique, though, we describe the data used to conduct our empirical tests.

For data on religious adherence rates, we obtained the 1980 Glenmary data previously used by Breault (1989a, 1989b) and the 1990 Glenmary data previously used by Olson (1999) and Montgomery (2003) from The American Religious Data Archive (www.thearda.com). We chose these data because of their apparent fitness for constructing a panel of U.S. counties (especially since both years of data were compiled by the same organization under similar methods), and because prior cross-sectional research on the Glenmary data has consistently yielded a negative relationship between pluralism and adherence. In this way, any results showing a positive relationship would be particularly noteworthy. These data sets contain county-level information on the number of churches, members and adherents in a large array of denominations. From the Glenmary data on adherents and U.S. Census data on total population, we calculated overall adherence rates, the Catholic and LDS adherents' share, and the pluralism index for each available county in 1980 and 1990.

In addition, we gathered various demographic data to assure that any results on the pluralism/participation question were not attributable to observable omitted variables. The demographic variables were drawn from the Census Bureau's USA Counties 1998 CD-ROM, which compiles data on counties from several government agencies including the Census Bureau, the Bureau of Labor Statistics, and the Bureau of Economic Analysis. Summary statistics for all variables are presented in Table 1.

To analyze these data using a fixed-effects panel estimator requires "stacking" the data into a cross section of time-series. Thus, the data consist of two observations (one for 1980, one for 1990) for each county for which data are available.[6] This data structure permits two crucial adjustments to the typical ordinary least squares approach to estimation. First, the regression constant is allowed to vary across each county, so that the regression "lines" for each county are allowed to pass through different intercepts. In this way, the estimator is able to control for otherwise unobserved heterogeneity. Second, the structure of the estimator's error term recognizes the correlation of the error terms within counties. Thus, the equation being estimated is

$$y_{it} = \alpha_i + \boldsymbol{\beta}'\mathbf{x_{it}} + \varepsilon_{it},$$

where y_{it} is religious adherence in county i in year t, α_i is a county-specific constant (which is allowed to be unique for each county), $\boldsymbol{\beta}$ is a vector of coefficients, $\mathbf{x_{it}}$ is a vector of independent variables, and ε_{it} is an error term for each county-year observation. The estimator for $\boldsymbol{\beta}$ is simply a generalized least squares estimator given by $\hat{\boldsymbol{\beta}} = [\mathbf{X}'\mathbf{M_d}\mathbf{X}]^{-1}[\mathbf{X}'\mathbf{M_d}\mathbf{y}]$, where $\mathbf{X}$ is the entire

[6] For some U.S. counties, data on the number of adherents are missing in the Glenmary data.

matrix of independent variables including the county-specific intercepts,[7] $\mathbf{y}$ is the vector of observations on county religious adherence rates, and $\mathbf{M_d}$ is the matrix $\mathbf{M_d} = \mathbf{M^0} \otimes \mathbf{I_n}$. In this formula, $\mathbf{M^0} = \mathbf{I_T} - \mathbf{ii'}/T$, where $\mathbf{I_T}$ is an identity matrix of rank T, $\mathbf{i}$ is a $T\times1$ vector of ones, and T is the number of periods over which the cross-sections are observed (here, $T = 2$). Thus, if there are n counties observed for 2 years each and k explanatory variables including the constant and the fixed effects, then $\mathbf{X}$ is a $2n\times k$ matrix, $\mathbf{y}$ is a $2n\times1$ vector, $\mathbf{M^0}$ is a 2×2 matrix, $\mathbf{I_n}$ is a $n\times n$ identity matrix, and $\mathbf{M_d}$ is a $2n\times2n$ matrix. The matrix $\mathbf{M_d}$ controls for correlation across the error terms within counties. For additional explanation of the fixed effects GLS estimator, see Greene (1997, Ch. 14).

The practical effect of this model is to incorporate time-invariant traits of individual counties into each county's unique intercept term α_i, commonly called a "fixed effect." Thus, if the people in a particular county are particularly prone to being members of a church, the county will have a higher intercept term than other similar counties. Other cultural and social factors should be incorporated into the fixed effect as well. Because the county fixed-effect is constant across time, the unobserved factors being captured by the fixed effect must be constant over time, which is reasonable to assume in this case across a 10 year period. Note that the specifications estimated below also contain a dummy variable for 1990, which serves as a time period fixed effect and allows all counties to have a uniformly different outcome in 1990 than in 1980.

The value of the fixed effects panel estimator can be seen in the interpretation of the coefficient in the model. Consider a panel regression of religious adherence (y) on the pluralism

[7] In practice, the county-specific fixed effects are estimated by incorporating into the empirical model a constant plus dummy variables for each county save one (to avoid perfect collinearity in the $\mathbf{X}$ matrix).

index (x) along with a vector of fixed effects. A change in the pluralism index within a single county between 1980 and 1990 would have the following effect on the change in adherence:

$$y_{i,1990} - y_{i,1980} = (\alpha_i + \beta x_{i,1990} + \varepsilon_{i,1990}) - (\alpha_i + \beta x_{i,1980} + \varepsilon_{i,1980}),$$

which is equivalent to $\Delta y_i = \beta \Delta x_i + \Delta \varepsilon_i$. For a given change in x_i, the expected change in y_i is determined by the value of the coefficient β, since the expected value of $\Delta\varepsilon_i$ is zero. In the analyses below, the coefficient on religious pluralism (or any other variable) can be interpreted as the partial effect on adherence of a one unit change in pluralism (or any other variable). Because the county fixed effect cancels itself out in this "change" interpretation, the social and cultural factors that can cause cross-county variations in religious attendance are controlled for when interpreting the coefficient β. This provides a substantial methodological improvement over performing ordinary least squares analysis on a cross-section, because the effects of social and cultural factors are estimated through the fixed effects. This allows a cleaner interpretation of the coefficient estimate on religious pluralism, thereby resolving the concerns raised by Olson (1999) and Voas *et al.* (2002).[8]

In terms of the empirical model to be estimated, the religious economies model hypothesizes that $\beta > 0$ for changes in religious pluralism, while the traditional "sacred canopy" view hypothesizes that $\beta < 0$. The examples in Figure 2 make this clear. In Figure 2, we have replicated the examples from Figure 1, except that we view each example as the same town being observed in two different time periods rather than two different towns observed

[8] Indeed, Voas *et al.* acknowledge that panel studies can resolve the concerns they raise. They write that "where available, 'panel' studies of the same set of geographical areas at different periods might enable researchers to investigate whether changes in participation rates over time are related to the pluralism of an area" (Voas *et al.* 2002, 225). However, Voas *et al.* also point out that a panel should ideally have more than two periods. We agree with them on this point. In our case, though, because the compositional bias we discuss below would only be exacerbated by adding additional years of Glenmary data (and/or its predecessors), we have not expanded our current data set beyond the two-period results reported herein.

simultaneously. Under the religious economies model, any gains in adherence should come from successful entry by or gains among the smaller denominations, as in Figure 2's Town B, a statistical effect measured by a positive relationship between change in adherence and change in pluralism. Under the sacred canopy view, adherence should increase when the dominant denominations become more vigorous, as in Town A of Figure 2, where a negative relationship exists between pluralism and adherence. Thus, the methodological advantage provided by a panel estimator allows a clean test of the two hypotheses, which we undertake in the next section.

RESULTS

Our first step in analyzing the Glenmary data for 1980 and 1990 is to generate the cross-sectional results previously found by Breault (1989a) and Olson (1999), which are set forth in Panel A of Table 2. Panel A of Table 2 reports robust OLS estimates from the 1980 and 1990 Glenmary datasets treated as separate cross-sections. The first two columns of Panel A report robust cross-sectional OLS estimates using the 1980 Glenmary data, and they confirm the findings of Breault (1989a) that the estimated pluralism coefficient is negative and significant. While the first column reports the results of a basic regression of the 1980 adherence rate on the pluralism index and a constant, the second column adds Catholic population share and LDS population share to the estimated equation. Both of these shares have positive and significant estimated coefficients, while the sign of the coefficient estimate on the pluralism index remains negative.

The third and fourth columns of Panel A repeat the cross-sectional analysis of the first two columns on the 1990 Glenmary data, and the results agree with Olson (1999). The relationship between pluralism and religious adherence is highly negative for both specifications, showing that this strong negative relationship between pluralism and adherence is robust to inclusion of the Catholic and LDS population shares in the cross-sectional analyses. Thus, as reported in prior research, the cure implemented by Finke and Stark (1988) of controlling for the Catholic population share does not alter the negative cross-sectional coefficient in either of the Glenmary datasets.

In Panel B of Table 2, we report results after pooling the 1980 and 1990 data together. Columns (1) and (2) report the results of a pooled OLS model (that is, without fixed-effects but combining both years of data into the sample), while columns (3) and (4) report the results of a fixed-effects generalized least squares panel estimation. In the pooled OLS estimations reported in the first two columns of Panel B, there continues to be a significant negative relationship between pluralism and religious adherence, which is robust to the inclusion of Catholic and LDS shares of adherents.[9] Both the Catholic and LDS shares of adherents have a significant negative impact on adherence. The results in columns (1) and (2) of Panel B show that the results from the fixed-effects panel estimation are *not* merely a function of pooling the two datasets together.

The results from the generalized least squares panel estimator with county-level fixed-effects are reported in columns (3) and (4) of Panel B. In contrast to the other regressions reported in Table 2, the panel estimator shows a significant positive relationship between pluralism and adherence, both with and without the controls for Catholic and LDS shares of adherents. One possible reason for obtaining different results than the cross-sectional

[9] In order to tell if the estimated coefficient on pluralism is positive due to "including possible substantive causal effects of Catholic presence . . . or whether it [is] due to the mathematically required positive relationship between Catholic population share and adherence rate," Olson (1999, 164) recommended substituting the Catholic share of adherents into the equation in place of the Catholic population share. According to Olson, the result is that there is "no mathematically necessary relationship between Catholic adherence share and the dependent variable, adherence rate" (164).

regressions reported in Table 2 is that the county fixed-effects may provide a simple control for unobserved heterogeneity across counties. As explained in the previous section, the coefficient estimates in columns (3) and (4) of Panel B capture the effect of a change in pluralism within a given county on adherence in that county. Thus, in the 1980 and 1990 Glenmary data, use of a panel data estimation method yields a strong positive relationship between within-county variations in pluralism and within-county variations in religious adherence. As previously explained in connection with Figure 1, this suggests that increases in adherence within counties are fueled by growth among the smaller denominations in those counties. The fixed-effects simply control for each county's unique starting point as of 1980. In clear contrast to past research using the Glenmary data, the results in Table 2 suggest that these data are consistent with the religious economies hypothesis when viewed over time.

The most likely source of unobserved heterogeneity in U.S. counties is variation in preferences for denominational affiliation across county populations, which likely results from cultural and historical factors like those described by Voas *et al.* (2002). The bulk of counties where religious adherents are highly concentrated consist of counties that are heavily Catholic, heavily LDS, or heavily Southern Baptist. It is a well-known result in other literature that, in the end, self-proclaimed Catholics and Mormons are more likely to join a church than people who claim affiliation with most Protestant denominations, and it may be true among Southern Baptists as well. However, when viewed in cross-section, this tendency of Catholics and Mormons (and perhaps Southern Baptists) to be more prone to join a church leads to an outcome where the more concentrated counties have higher adherence rates, which generates an observed negative cross-sectional relationship between pluralism and adherence.[10]

It is possible that other variables known to affect religiosity could be responsible for the positive coefficient estimate in columns (3) and (4) of Panel B. For example, various studies of individual religious attendance suggest that several economic and demographic variables can affect religious participation, including age, education, income and ethnicity (e.g., Iannaccone 1998). To account for the impact of such factors on county-level adherence rates, we introduced additional control variables to the specifications reported in columns (3) and (4) of Panel B in Table 2. The results of the fuller estimations are reported in Table 3. Column (1) reports the impact of only the demographic controls without considering pluralism's impact. Columns (2) and (3) then introduce the additional variables of pluralism, Catholic adherents' share, and LDS adherents' share that were examined in Table 2. In Table 3, pluralism remains a positive and significant influence on adherence when controlling for a broad range of demographic data. Similarly, the effects of the demographic controls are not substantively altered by inclusion of the pluralism index or the Catholic or LDS shares. Thus, the fixed-effects GLS estimator is clearly providing robust estimates of the effects of each variable, including the positive sign on the pluralism coefficient estimate.

In general, the coefficient estimates on the demographic variables conform to the expected signs. Education is positive and significant in all three equations. Income has a

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[10] An interesting result in column (4) of Panel B is the strong positive effect of Catholic adherents' share along with the almost significant negative effect of Mormon adherents' share. One possibility is that the Catholic Church in the U.S. has a high level of internal diversity that generates the same effects as external competition among denominations. Such a theory is consistent with the observations of the Catholic Church in Italy made by Diotallevi (2002). In contrast, the LDS church is much more uniform than the Catholics, with much of what transpires in local wards being beamed in via satellite from headquarters in Salt Lake City. This uniformity among Mormons might be an internal competition-reducing factor that leads to the negative coefficient estimate in column (4). This result warrants further study, in light of the insignificance of the LDS share reported in Table 3 and in light of Phillips's (1998) findings that LDS devotion is higher in places where the church is most dominant.

negative effect on adherence (while controlling for education), and unemployment rate has a positive effect, both of which are consistent with the opportunity cost theory of Azzi and Ehrenberg (1975). Percent Black has a positive effect, but percent Hispanic has no significant effect on religious adherence. Surprisingly, the effect of age is opposite its expected sign. Religious participation generally increases with age, but Table 3 shows an insignificant negative effect of Population over 65. Also, the presence of school-age children usually spurs individual attendance, but the coefficient estimate on Population under 18 is negative and significant.

A final factor that has been found to affect religious participation is urbanization. To examine this factor, we controlled for population density and its square, which yielded a positive and concave relationship with adherence. Based on the point estimates, the relationship between adherence and population density is significantly positive up to 34,900 people per square mile. The only county with higher density is New York City, so increased density generally leads to increased adherence in U.S. counties.[11]

At this point, our results appear to strongly confirm the religious economies hypothesis, by demonstrating that an adequate control for Voas *et al.*'s (2002) cultural and historical factors generates a positive relationship between pluralism and participation in the U.S. counties data. These results appear particularly compelling given that these data have, in prior research, consistently generated robustly negative estimates of the relationship. Unfortunately, we cannot make such a bold claim, in light of some seemingly minor variations in the composition of the samples between 1980 and 1990. It is to this topic we now turn.

[11] One note of caution is appropriate here. When population density is included without its square, the coefficient estimate is significant and negative. Thus, the finding on population density is not robust to variations in specifications.

DISCUSSION

Prior cross-sectional research (by Breault, Olson, and others) on the 1980 and 1990 Glenmary religious adherence data for U.S. counties has found a robust negative relationship between the rate of religious adherence and the degree of religious pluralism in these data. Such results are inconsistent with the religious economies model developed by Stark, Bainbridge, Finke, Iannaccone, and others. In this paper, we hoped to significantly move the debate forward by constructing a panel from the 1980 and 1990 data (and other years as well) and then interpreting the results of a fixed effects model. As reported in the previous section, performing these analyses yields a positive relationship between pluralism and the adherence rate.

There is a problem, however, in constructing even a two-period panel from the Glenmary data. The 1980 data contain information on 111 Judeo-Christian church bodies, which constituted an estimated 91 percent of U.S. church membership in 1980. On the other hand, the 1990 data contain information on 133 Judeo-Christian church bodies, and its membership totals exceed the U.S. church membership levels reported in the *Yearbook of American and Canadian Churches: 1990*. At first blush, it seems that compositional variations across years should not be a major problem in our analysis, because the major denominations were included in both years. Thus, compositional changes should have effects on the pluralism index only at the fourth or lower decimal place. Contrary to our *a priori* expectations, though, it turns out that these two data sets are not as similar as we first thought. There are 83 denominations present in both the 1980 and 1990 data sets. Thus, there are 28 denominations present only in the 1980 data, and at the national level these account for 1 percent of all adherents reported for 1980. In addition, there are 50 denominations present only in the 1990 data; at the national level, these accounted for approximately 9 percent of all adherents reported for 1990.

Unfortunately for the vitality of our results, these variations in the composition of the sample introduce a compositional bias that generates a positive within-county relationship between pluralism and adherence. We demonstrate the problem through an example set forth in Figures 3A and 3B, which expands upon the examples in Figure 1. In Figure 3A, we imagine a hypothetical U.S. county that has a constant population of 10,000 in both 1980 and 1990, and whose church adherents are spread among the four denominations listed in Figure 3A: the Evangelical Covenant Church of America, the Anglican church, the Methodist church, and the Church of God Prophecy.[12] However, we assume that data on the Evangelical Covenant denomination are present only in the 1980 Glenmary data, and that data on the Church of God Prophecy denomination are present only in the 1990 Glenmary data (and indeed these assumptions are true for these two denominations). The first column of Figure 3A shows the "true" status of pluralism and adherence in our hypothetical county for 1980 and 1990. By assumption, there is no change in the number of adherents nor in the pluralism index. The second column shows that, if the Church of God Prophecy were omitted from the data in 1980, the observed adherence rate would be 49.66 percent and the observed pluralism index would be 0.420. The third column shows that, if the Evangelical Covenant Church were omitted from the data in 1990, the observed adherence rate would be 54.03 percent and the observed pluralism index would be 0.502. Using these two points to calculate a simple slope estimate that mimics the point estimate in column (3) of Table 2, Panel B, we find that our example yields a positive slope estimate of 53.3 despite the county having experienced absolutely no change across the

[12] In fact, we derived the numbers in Figure 2A in order to conform to several conditions derived from either the data or the examples from Figure 1. Specifically, the hypothesized numbers assure that (1) the Evangelical Covenant Church is 1% of the adherents measured in 1980, which corresponds to the 1980-only denominations' national share in the 1980 data; (2) the Church of God Prophecy is 9% of the adherents measured in 1990, which corresponds to the 1990-only denominations' national share in the 1990 data; and (3) the ratio of Anglicans to Methodists is 5:2, as in the examples in Figure 1.

hypothesized ten-year period.

Having discovered the existence of this compositional bias in our constructed panel, we took a variety of steps in an effort to resolve the problems. First, we limited the analysis to only those 83 denominations that are present in both the 1980 and 1990 samples. The fixed effects model on these 83 denominations led to a negative and significant relationship between pluralism and adherence, which seemingly refutes the religious markets hypothesis and confirms the sacred canopy hypothesis. However, the sample thus constructed omits many smaller denominations while being more heavily weighted toward long-standing mature denominations. Under the religious markets hypothesis, as explained in Stark and Finke (2000, Ch. 8), the engine of growth in religious markets is competition from smaller denominations seeking to move into larger, more mainstream niches. As a result, the negative coefficient estimate obtained from the sample of denominations present in both years cannot adequately refute the religious markets hypothesis, because it ignores the very types of religious denominations that the theory posits are the source of vitality. Our search for a solution to the composition bias had to continue.

Next, we imputed county-level values for the denominations missing from either the 1980 or 1990 samples. To do so, we assumed that such denominations had the same population share in both years, using the population share in the available year to compute the population share for the missing year. The population shares of the missing denominations allowed us to calculate new values for the total number of adherents in each county, leading to recalculated values in each county for the adherence share and the pluralism index. Applying the fixed effects model on the recalculated data yielded the same negative and significant relationship between pluralism and adherence that was generated by the regression on the sample of 83 denominations present in both 1980 and 1990. This is not surprising because we implicitly assumed no major change in the status of the missing denominations, so that any changes in pluralism and adherence were driven primarily by the 83 denominations present in both years. For the reasons stated in the preceding paragraph, then, the recalculated sample of all missing denominations did not allow us to make any reasonable inference about the religious markets or the sacred canopy hypotheses.

Our final attempt to resolve the problem was to consider the possibility of calculating an upper bound on the size of any composition bias. That is, we considered calculating the largest possible positive coefficient that could be generated from the Glenmary data as a sole result of the composition bias. If a reasonably calculated upper bound on the coefficient were smaller than the coefficient estimate actually generated by the fixed effects model, then we could argue that the coefficient estimates reported in Tables 2 and 3 were too large to be attributed solely to composition bias. The example presented in Figure 3A made clear that such a strategy was not feasible. After all, the example we constructed in Figure 3A as a simplified situation consistent with the data yielded a composition-biased slope estimate of 53.3, which is larger than any of the positive fixed effects coefficient estimates in Tables 2 and 3. Moreover, we constructed another "typical" county with as many denominations in each category (present only in 1980, present only in 1990, and present in both years) as the average for all counties in the Glenmary data.[13] We then assumed that all denominations within each category were of equal size, thereby generating a hypothetical county with the highest possible level of pluralism that is otherwise consistent with our data. Performing an analysis similar to that in Figure 2, we calculated a

[13] Specifically, the county-level average number of denominations present only in 1980 was 0.8; the average number present only in 1990 was 2.4; the average number present in 1980 for denominations present in both years was 14.6; and the average number present in 1990 for denominations present in both years was 15.2.

composition-biased slope estimate of 471.2. The slopes found in the examples we constructed are sufficiently large that there is little chance that an upper-bound strategy will have any success in resolving the composition bias problem.

Thus, we are left somewhat farther along in the pluralism/participation debate than where we began. Importantly, we have demonstrated an effective empirical methodology – the fixed-effects panel estimator – that has the ability to differentiate between the religious markets hypothesis and the sacred canopy hypothesis in a way that does not involve the critiques of Olson (1999) and Voas, Olson, and Crockett (2002). Moreover, we believe that the results obtained from using this technique to examine the effects of various demographic variables (such as race, age, income, education, and so forth) on aggregate measures of religious participation are reliable. After all, the compositional changes across sample years in the Glenmary data only introduce a small measurement error into the estimate of the overall adherence rate, and this measurement error is not of necessity correlated with any of the demographic variables in the same way that the pluralism index is. Finally, we have also demonstrated that, unfortunately, the Glenmary county adherence data are not well-suited to being made into a panel for purposes of testing the pluralism/participation question. As a result, we sound a call to identify existing data that avoid the composition bias problem, or to generate new longitudinal data that can eliminate the compositional problems inherent in the Glenmary data. Moreover, our results strongly suggest that constructing an even longer panel from the various adherence data on U.S. counties available from 1952 to 2000 would not be a productive enterprise for purposes of analyzing the pluralism/participation question.[14]

[14] The American Religious Data Archive includes denominational data for U.S. counties from 1952, 1971, 1980, and 1990, the two latter data sets being the Glenmary data discussed in this paper. There is also a Glenmary 2000 data set currently available. In light of the compositional problems we found in connection with the 1980 and 1990

In closing, we note two important points. First, the panel data approach can solve the Voas/Olson/Crockett critique, given the right data set. However, any such data must contain no variation (or at most, very little variation) in denominational composition across time, other than the variation that truly occurs as a result of the actual patterns of changing denominational adherence within a geographical area. Second, the Voas/Olson/Crockett critique shows the weakness of a testing method, not of a theory. The data flaws we uncover in this paper similarly point only to the weakness of a testing methodology, not of the theory being tested. To show that empirical tests using the pluralism index are not reliable in certain specific ways is neither to undermine nor to confirm the validity of the religious markets hypothesis. Rather, our findings and those of Voas/Olson/Crockett are calls to find better empirical tests of the theory, including better ways to model and measure the competitiveness of religious markets.

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data, we believe it would be pointless to construct a longer panel from these data sources. The compositional bias would only be exaggerated by adding in multiple years.

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Figure 1. Examples Derived from Voas *et al.* (2002)

Example 1A				
Town A	Anglicans	50%	Participation Rate	70%
	Methodists	20%	Pluralism Index	0.41
Town B	Anglicans	55%	Participation Rate	75%
	Methodists	20%	Pluralism Index	0.39
Example 1B				
Town A	Anglicans	50%	Participation Rate	70%
	Methodists	20%	Pluralism Index	0.41
Town B	Anglicans	50%	Participation Rate	75%
	Methodists	25%	Pluralism Index	0.44

In this figure, based on an example in Voas *et al.* (2002), there are two towns composed of only two religious groups. In Example 1A, variation occurs in the dominant group, which leads to a negative relationship between participation and pluralism. However, in Example 1B, variation occurs in the small group, which leads to a positive relationship between participation and pluralism.

Figure 2. Examples of Within-County Variations in Pluralism and Participation

Town A				
1980	Anglicans	50%	Participation Rate	70%
	Methodists	20%	Pluralism Index	0.41
1990	Anglicans	55%	Participation Rate	75%
	Methodists	20%	Pluralism Index	0.39

Town B				
1980	Anglicans	50%	Participation Rate	70%
	Methodists	20%	Pluralism Index	0.41
1990	Anglicans	50%	Participation Rate	75%
	Methodists	25%	Pluralism Index	0.44

This figure slightly alters Figure 1. In Town A, growth in participation over time occurs in the dominant group, which leads to a negative relationship between the change in participation and the change in pluralism. This pattern reflects the operation of the sacred canopy hypothesis. In contrast, in Town B, variation occurs in the small group, which leads to a positive relationship between the change in participation and the change in pluralism. Such a pattern reflects the effects posited by the religious markets hypothesis. More over, unlike cross-sectional analyses, it is far more plausible to view the within-county relationship between pluralism and participation as causal in nature.

Figure 3A. Example Demonstrating Compositional Bias in Glenmary-Based Panel

	True Number	Measured in 1980	Measured in 1990
Evangelical Covenant	50	50	0
Anglican	3511	3511	3511
Methodist	1405	1405	1405
Church of God Prophecy	487	0	487
Total Adherents	5453	4966	5403
Adherence Rate	54.53	49.66	54.03
Pluralism Index	.511	.420	.502
Slope of Regression Line			53.3

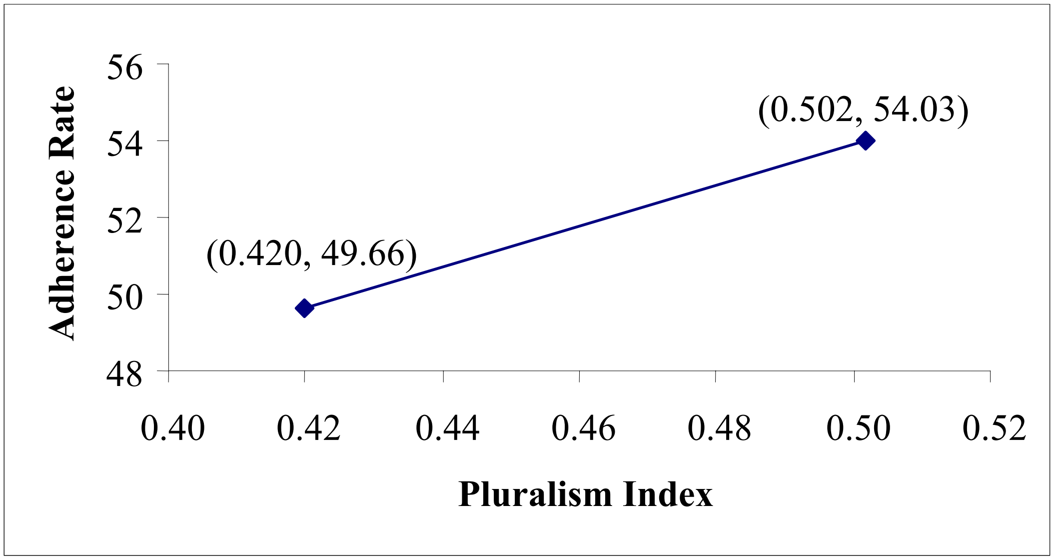
In this example, we assume a county with a constant population of 10,000 people in both 1980 and 1990. Moreover, we assume that there are only four denominations in the county, and that the number of adherents in each denomination is constant between 1980 and 1990. If measured accurately in both years, the county would show no change in either adherence rate or pluralism.

Assume though that in 1980, the Church of God Prophecy did not respond to the Glenmary questionnaire, and that in 1990 the Evangelical Covenant Church of America did not respond. Thus, we would have no data on the number of adherents in the county's Church of God Prophecy in 1980 or on the county's Evangelical Covenant Church in 1990.

The numbers in the example are derived so that (1) the Evangelical Covenant Church is 1% of the adherents measured in 1980, which corresponds to the 1980-only denominations' national share in the 1980 data; (2) the Church of God Prophecy is 9% of the adherents measured in 1990, which corresponds to the 1990-only denominations' national share in the 1990 data; and (3) the ratio of Anglicans to Methodists is 5:2, as in the examples in Figure 1. (All are rounded to the nearest whole number.)

The Evangelical Covenant Church of America was present in only the 1980 data and is still active today. From 1980 to 2004, the Evangelical Covenant Church went from 92,765 members to 148,296 members. The Church of God Prophecy, which was present in only the 1990 data, has been a denomination since the resolution of a legal battle in 1952.

Figure 3B. Graphical Representation of Figure 3A.



This graph plots the within-county pluralism-adherence combination "observed" in Figure 2A for 1980 and 1990. The slope of the "regression line" is:

$$\text{Slope} = \frac{54.03 - 49.66}{0.502 - 0.420} = 53.3\,.$$

Table 1. Summary Statistics

Variable	Entire Sample	1980 Only	1990 Only
Religious Adherence Rate	57.44	55.37	59.49
	(19.52)	(18.79)	(20.01)
Pluralism Index	0.6912	0.6809	0.7016
	(0.1567)	(0.1615)	(0.1511)
Catholic Population Share	12.89	12.71	13.07
	(15.17)	(15.10)	(15.25)
LDS Population Share	1.84	1.73	1.96
	(8.90)	(8.69)	(9.10)
Percent High School Graduates	64.50	59.37	69.62
	(12.54)	(12.41)	(10.39)
Median Household Income	17,866.45	17,297.97	18,433.82
	(4644.66)	(4113.44)	(5057.54)
Percent Black	8.47	8.49	8.45
	(14.29)	(14.34)	(14.24)
Percent Hispanic	4.15	3.80	4.50
	(10.66)	(10.19)	(11.10)
Percent under 18	28.22	29.53	26.91
	(3.77)	(3.53)	(3.55)
Percent over 65	14.03	13.19	14.86
	(4.38)	(4.16)	(4.43)
Unemployment Rate	6.80	7.41	6.19
	(3.21)	(3.33)	(2.96)
Population Density	196.88	194.50	199.26
	(1491.30)	(1560.41)	(1419.21)
Number of Observations	6202	3098	3104

Numbers in parentheses are standard deviations.
Median Household Income is reported in 1982-1984 dollars.

Table 2. Adherence and Pluralism

Panel A. Cross Section Results.

	1980 Data		1990 Data	
	(1)	(2)	(3)	(4)
Pluralism Index	-27.22	-13.15	-43.54	-36.26
	(.000)	(.000)	(.000)	(.000)
Catholic Population Share		0.4450		0.2736
		(.000)		(.000)
LDS Population Share		0.2068		-0.0259
		(.000)		(.446)
Constant	73.92	5.83	90.04	81.41
	(.000)	(.000)	(.000)	(.000)
R-squared	0.0546	0.1721	0.1080	0.1490
F-statistic	142.81	251.77	291.34	203.35
	(.000)	(.000)	(.000)	(.000)

Panel B. Pooled Sample.

	Pooled OLS		Fixed Effects GLS	
	(1)	(2)	(3)	(4)
Pluralism Index	-34.84	-40.19	16.40	31.33
	(.000)	(.000)	(.000)	(.000)
Catholic Share of Adherents		-0.0405		0.3642
		(.000)		(.000)
LDS Share of Adherents		-0.1838		-0.1251
		(.000)		(.101)
Year = 1990	4.826	4.993	3.787	3.607
	(.000)	(.000)	(.000)	(.000)
Constant	79.11	84.22	44.21	26.26
	(.000)	(.000)	(.000)	(.000)
Within R^2			0.1275	0.1691
Between R^2			0.0945	0.0229
Overall R^2	0.0888	0.1013	0.0229	0.0111
F-Statistic	254.38	128.98	226.21	157.44
	(.000)	(.000)	(.000)	(.000)

p-values are in parentheses. Dependent variable is the percentage of the county's population who are religious adherents. In Panel B, columns 1 and 2 reflect OLS estimation on the 1980 and 1990 data pooled together but without fixed effects; columns 3 and 4 reflect a generalized least squares panel estimator with county-level fixed effects. For all OLS estimations, *p*-values are based on robust standard error estimates.

Table 3. Determinants of County-Level Adherence Rates

	(1)	(2)	(3)
Pluralism Index		11.2225	26.85914
		(0.000)	(0.000)
Catholic Share of Adherents			0.3767
			(0.000)
LDS Share of Adherents			-0.0524
			(0.482)
Percent High School Graduates	0.4252	0.4005	0.3640
	(0.000)	(0.000)	(0.000)
Median Household Income	-0.2865	-0.3053	-0.3401
	(0.003)	(0.001)	(0.000)
Percent Black	0.3707	0.3562	0.4834
	(0.006)	(0.008)	(0.000)
Percent Hispanic	0.0346	0.0566	0.0930
	(0.775)	(0.639)	(0.430)
Percent under 18	-0.4833	-0.4098	-0.4037
	(0.000)	(0.002)	(0.002)
Percent over 65	-0.1404	-0.1210	-0.1481
	(0.367)	(0.436)	(0.328)
Unemployment Rate	0.7357	0.7156	0.6865
	(0.000)	(0.000)	(0.000)
Population Density	0.0060	0.0056	0.0049
	(0.114)	(0.142)	(0.188)
Population Density Squared	$-7.88\text{x}10^{-8}$	$-7.51\text{x}10^{-8}$	$-7.02\text{x}10^{-8}$
	(0.020)	(0.026)	(0.033)
Year = 1990	-0.1003	0.0584	0.2687
	(0.900)	(0.942)	(0.731)
Constant	41.4956	33.4872	16.62411
	(0.000)	(0.000)	(0.036)
Within R^2	0.1674	0.1717	0.2130
Between R^2	0.0320	0.0535	0.0320
Overall R^2	0.0159	0.0293	0.0181
F-Statistic	62.09	58.17	64.22
	(0.000)	(0.000)	(0.000)
Number of Observations	6202	6202	6202

p-values are in parentheses. Dependent variable is the percentage of the county's population who are religious adherents. All equations are estimated using a fixed effects panel estimator.