Religion and environmental politics in the US House of Representatives


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ABSTRACT
Does religion affect legislators’ behavior on environmental policy in the US? Studies of environmental policy making have not examined this question, although the literature suggests that religion might affect legislative behavior on environmental policy. This study examines the relationship between US House members’ religion and roll-call voting on environmental legislation from 1973 to 2009. It finds significant differences across religious traditions. Legislators’ party and characteristics of constituencies relevant to environmental politics increasingly, but not entirely, mediate these differences.

KEYWORDS Environment and religion; religion and congress; religion and roll-call voting; league of conservation voters score; religion and league of conservation voters

In June 2015, Pope Francis made international headlines by releasing an encyclical that ‘put care for the environment at the center of Catholic social teaching, and in lyrical but stark terms, reframed the discussion about global warming from the dry language of science to a broad question of ethics’ (Winfield et al. 2015). Although some hoped the encyclical could ‘realign politics,’ Rob Bishop, the chairman of the US House of Representatives’ Committee on Natural Resources and a member of the Church of Jesus Christ of Latter Day Saints (LDS, commonly called the Mormon church), rejected the Pope’s reframing, flatly countering ‘no, I’m sorry, it’s a political issue’ (Winfield et al. 2015). The encyclical and responses to it raise an important question we seek to answer – how much does religion affect environmental policy making in the US?

Analysts warn that failure to reform current practices threatens many harms to ecosystems, people groups, and even world order (e.g., IPCC 2014). Understanding the forces shaping environmental policy, then, may be a matter of life and death for many. Because environmental challenges
do not respect national boundaries, global coordination is often necessary to meet environmental threats. US (in)action significantly shapes the possibilities for and effects of multinational agreements. Thus, it is important that we learn everything we can about what affects the course of environmental policy in the US, including the role of religion.

We know that several external forces, such as political parties and constituents’ demands, influence US legislators’ behavior in environmental policy making (e.g., Anderson 2011). However, like the rest of us, elected officials are complex human beings. Scholars have begun to examine the influence of forces internal to legislators themselves, what Burden (2007) calls ‘the personal roots of representation.’ In particular, a growing body of research has found that in the US, legislators’ religious traits shape their behavior on several issues (see Oldmixon 2009 for a review). Although personal forces may hold less sway for members of party-dominated parliaments across Europe, even here the increasing ‘personalization’ of politics (McAllister 2007) suggests that personal influences on parliamentary behavior may well be significant and increasing. However, no research has examined the relationship between legislators’ religion and their activity in environmental policy making.

We answer our question in the context of roll-call votes in the US House of Representatives over four decades (1973–2009). The US is an important case because of its significant global environmental impact (e.g., as the world’s second largest producer of greenhouse gases), the influence its decisions have on other nations and multinational groups’ environmental policy decisions (see Selin and VanDeveer 2015), and the more academic reason that the US case has been ‘a source of inspiration’ to the examination of religion in European parliaments (Foret 2014a, p. 111). We find representatives’ religion relates to environmental policy making in important, albeit sometimes indirect, ways. Using the League of Conservation Voters’ (LCV) scorecards, we find significant differences across religious traditions. In general, LDS and Evangelical Protestant representatives are least supportive of environmental legislation, while Jewish and black Protestant members are most supportive. Mainline Protestants and Catholics fall somewhere in between. In short, religion matters for Congressional action on environmental policy.

**Religion and legislative behavior**

Previous studies have taught us much about environmental legislative behavior in the US. Scholars have long noted the influence of legislators’ party affiliation and ideological orientation, with Democrats and liberals tending to support legislation designed to protect the environment more than their Republican and conservative counterparts (e.g., Ritt and Ostheimer 1974, Dunlap and Allen 1976, Fowler and Shaiko 1987, Mohai
and Kershner 2002, Anderson 2011). Over time, these party differences expanded (e.g., Dunlap et al. 2001, Shipan and Lowry 2001, Ard and Mohai 2011). Figure 1 shows the growing party gap in the House of Representatives, tracing the parties’ average scores on the LCV’s 0–100 ‘National Environmental Scorecard’ over the period we examine (higher scores denote greater environmental support).

In addition to party and ideology, traits associated with the constituencies that legislators represent and the characteristics of the legislators themselves predict environmental support. Representatives’ region, sex, and race/ethnicity correlate with their behavior, as southerners tend to have lower LCV scores, while women, African Americans, and Hispanics tend to have higher scores (e.g., Dunlap and Allen 1976, Mohai and Kershner 2002, Ard and Mohai 2011). In addition, legislators representing more urban, educated, and/or wealthy districts with strong membership in environmental groups tend to have higher LCV scores, while representatives from districts that rely more on industries environmental regulations affect (e.g., mining, agriculture, manufacturing) tend to have lower LCV scores (e.g., Kalt and Zupan 1984, Anderson 2011, Ard and Mohai 2011), though findings on these variables are less consistent.

Despite attention to many factors affecting environmental legislative behavior, to our knowledge, no study has given more than a passing mention of members’ religion. Yet, there is significant reason to expect religion to play a role. Put most broadly, various religious beliefs hold

Figure 1. Average LCV Score by Party.
Note: Data from League of Conservation Voters: http://www.lcv.org/.
Scores range 0–100. Higher scores denote greater environmental support.
implications for environmental policies. Religions provide perspectives on
the origin and destiny of humankind and its relationship with nature.
Various scholars have outlined the connections between such religious
beliefs and environmental perspectives. As just one example, Coward
(1993) draws these connections for Judaism, Islam, Christianity,
Hinduism, Buddhism, and Taoism. Just as Pope Francis has done, leaders
and scholars of religious groups frequently articulate (and debate) these
connections.¹

As examples of the potential influence of religious beliefs on environ-
mental attitudes, consider two beliefs that have garnered significant atten-
tion in the context of Christianity’s relationship with the environment.
First, White (1967) famously argued that God’s command to humans to
have ‘dominion’ over the earth (Genesis 1:28) led to a Judeo-Christian
‘mastery’ perspective on the natural world, an anthropocentric perspective
encouraging exploitation of nature. White’s thesis generated considerable
debate, with various scholars arguing that the command amounts to a call
for stewardship of nature (e.g., Passmore 1974, Harrison 1999, Conradie
et al. 2014). The point is that the Judeo-Christian text can shape views on
the environment, even if adherents debate the exact implications of the text.

Second, a belief in ‘End Times’ theology, concentrated among
Evangelical Protestants but present in other traditions, may decrease envir-
onmental concern (Guth et al. 1995, Barker and Bearce 2012, see also
Kilburn 2014). Such theology includes the belief in a Second Coming of
Jesus and the ‘End Times’ in which the Earth is destroyed and replaced by a
new Earth. Barker and Bearce (2012) argue that believing the world will be
destroyed decreases incentives to pay short-term costs for long-term envir-
onmental benefits, since, in this view, the Earth may not survive to reap
such benefits. They found that individuals who believe in the Second
Coming of Jesus were less likely to support policies designed to curb
warming.

Local places of worship often articulate and reinforce connections
between religious beliefs and various politically relevant issues. Such places
of worship provide information and behavioral cues about the environ-
ment, whether through clergy statements or interaction with co-religionists
(Djupe and Hunt 2009, Djupe and Olson 2010). As a result of religious
teachings, the social dynamics of belonging to a religious tradition and local
religious body, and other factors, religion significantly influences public
opinion and voting behavior in the US (e.g., Guth et al. 2009) and Europe
(e.g., Broughton and ten Napel 2000, van Der Brug et al. 2009, Nelsen and
Guth 2015). In particular, the messages, cues, and other features of religion
shape and reflect individuals’ environmental attitudes in the US (Sherkat
and Ellison 2007, McCright and Dunlap 2011, Kilburn 2014) and Europe
(Clements 2012, Hagevi 2014).
In the US, findings demonstrate the importance of heeding differences within the Christian faith as more specific religious traditions (e.g., white Catholic, Hispanic Catholic, black Protestant, Mainline Protestant, Evangelical Protestant, Mormon) hold to distinctive beliefs and practices (Guth et al. 2006). Studies show clear differences in environmental attitudes among major American religious traditions, with Evangelical Protestants most skeptical about environmental legislation, followed at some distance by Mainline Protestants, at least partly due to Evangelicals’ more literalist understanding of the Bible (Kilburn 2014). In contrast, white Catholics tend to be more supportive, with black Protestants, Jews, and secular citizens even more pro-environment. These findings hold across groups of clergy, religious activists, political contributors, and the mass public (Guth et al. 1995).

It would be surprising if legislators alone were immune from religious influences. Burden (2007) demonstrates that in the US, legislators’ values, experiences, and interests profoundly affect their policy preferences and legislative behavior. He argues that ‘legislators acquire expertise, values, and interests long before they arrive in Washington. These early influences often lead to more intense preferences that have the power to deter other influences such as constituent pressure, party leadership, or interest group lobbying (Burden 2007, p. 11). Since religion is deeply engrained in these values and experience, it is not surprising that religion shapes legislative behavior in the US. In fact, Burden demonstrates the significant role legislators’ religious convictions played in shaping their preferences and behavior in domains such as tax policy toward religious charities and stem-cell research. In sum, most legislators hold to a religious perspective (even if it is atheism or secularism), which can shape their views on environmental politics. Moreover, legislators communicate with religious leaders, lobbyists, and constituents who may reinforce connections between religious commitments and their activities in office.

The empirical record demonstrates a link between religion and legislative behavior. Although the study of religion and parliamentary behavior is less developed, recent studies find important, often indirect, effects of religion in Europe (Foret 2014a). In the US, religion is significantly related to the overall ideological cast of legislators’ roll-call votes (Fastnow et al. 1999, Guth 2014). More specifically, representatives’ religion is related to behavior on foreign policy (e.g., Oldmixon et al. 2005, Guth 2014) and in several domestic issues domains, including reproductive policy (Yamane and Oldmixon 2006), school prayer (Oldmixon 2005), gay rights (Haider-Markel 2001, Oldmixon and Calfano 2007), human cloning (Burden 2007), and support for the policy goals of the Religious Right (Smith et al. 2010). Generally speaking, Evangelical Protestant and LDS representatives support conservative causes, while Mainline Protestant and white...
Catholic members tend to be more moderate, and Jewish members tend to be most supportive of liberal policies.

Although religion shapes legislative behavior, the relationship between them is often complex and indirect, possibly mediated by partisanship, ideology, and district characteristics. Our purpose here is to demonstrate that legislators’ religion matters for environmental policy making, rather than to delineate exactly the pathways by which it does so. Still, it is important to consider these mediating forces. As representatives have polarized along religious and cultural lines, religious affiliation and party affiliation have become highly correlated (Layman 1999). This is clearly seen for Evangelical Protestants in Figure 2, which shows the percentage of Evangelicals in the House of Representatives that affiliated with the Republican Party. In the 1970s and 1980s, more than half of Evangelical Protestants in the House were Democrats. In the 1990s, however, far more Evangelicals affiliated with the Republican Party. By the 2000s, 75–80% were Republicans. Consequently, partisanship’s mediating role may be increasing (Yamane and Oldmixon 2006).

**Data and measures**

To study religion and roll-call voting on environmental legislation, we employ LCV scores as the dependent variable, a common approach (e.g., Dunlap et al. 2001, Shipan and Lowry 2001, Mohai and Kershner 2002,
The LCV identifies 10–25 bills of particular environmental importance each year, rating each member on a 0 (least supportive) to 100 (most supportive) scorecard based on the percentage of times that the member voted consistent with the LCV position. The LCV Web site archives the scorecards back to 1973. When we collected the data, scores were available through 2009. We selected scores from every five years (1975, 1980, 1985, 1990, 1995, 2000, 2005, and 2009) along with three other years (1973, 1987, and 2001) to ensure our five-year selection was not driving the results. In total, we analyzed LCV scores from 11 different years over a 37-year span.

Despite its attractions, no measure is perfect. First, the LCV scores reflect only roll-call votes. They do not include additional activity that might advance or hinder environmental causes (e.g., holding hearings, co-sponsoring legislation). However, roll-call votes force members to make real choices, choices that determine the fate of legislation.

Second, like all interest group scores based on roll-call votes, LCV scores may not be comparable over time, since votes were taken in different political contexts (e.g., Democratic or Republican majorities). This means a score of 50 in one year may not mean the same level of environmental commitment as a 50 in another year. Groseclose et al. (1999) offer a solution to this potential problem, a solution that has been used in the specific context of LCV scores (Shipan and Lowry 2001) and other interest group scores (e.g., Ansolabehere et al. 2001). We transformed the LCV scores using this method. In practice, the transformation makes little difference, as the raw and transformed scores correlate at 0.96, and results are similar using either score.

We follow earlier research in assigning members to specific religious traditions (Fastnow et al. 1999, Oldmixon and Calfano 2007, Smith et al. 2010, Collins et al. 2011). Religious traditions ‘vary systematically in terms of religious beliefs, practices, and communicated leadership cues’ (Collins et al. 2011, p. 556). Consequently, affiliation with a particular tradition often provides a serviceable proxy for religious beliefs, behaviors, and the information communicated through particular religious communities. To assign members to traditions, we consulted annual editions of The Almanac of American Politics and Politics in America, which report members’ religion. For many members, sources listed a specific religious family (e.g., Baptist, Methodist, Episcopalian). We follow the scheme Guth et al. (2006) outlined to collapse these into several religious traditions: black Protestant, white Catholic, Evangelical Protestant, Hispanic Catholic, Jewish, LDS, Mainline Protestant, and a residual ‘other’ category (for others using this classification scheme, see Steensland et al. 2000, Guth et al. 2006, Newman and Smith 2007, Collins et al. 2011). We assigned each legislator to a single tradition and created an indicator variable for each tradition.
The sources sometimes listed vague religious affiliations, leading to some measurement error. For example, over the period of study, on average 22 members a year (roughly 5% of 435) were listed as simply ‘Christian.’ Presumably many of these legislators were actually from the Evangelical Protestant, Mainline Protestant, or Catholic religious traditions, but we are forced to place these legislators, along with those affiliated with other traditions (e.g., Orthodox, Muslim, Buddhist), in the ‘other’ category. Fortunately, Guth (2014) has collected more precise measures of legislator affiliation beginning in 1995. Although it is virtually impossible to develop equivalent measures for the earlier time period, we can compare results using our measure of legislator religious tradition with Guth’s more refined measure for 1995–2009. The two measures were identical in 91% of cases over the five years where our data overlap. Replicating our analyses using Guth’s measures generates similar results to those below.

African American and Hispanic representatives tend to have higher average LCV scores than whites (Mohai and Kershner 2002, Ard and Mohai 2011). Since we focus primarily on religion rather than race/ethnicity, we coded African American and Hispanic members as either African American Protestant, Hispanic Catholic, or in the residual category. In practice, these religious traditions include the vast majority of the respective racial/ethnic groups. Only 14% of African American cases were not African American Protestants, and 6% of Hispanic cases were not Hispanic Catholic. Thus, the African American Protestant and Hispanic Catholic traditions overlap significantly with the entire group of African American members and Hispanic members, respectively. Given the small sample sizes of these groups, we cannot examine the extent to which their distinctive LCV scores are due to religious traditions or other factors. Therefore, we focus on the other religious traditions.

**Results**

We begin by examining the average LCV score for each religious tradition. Figure 3 compares the overall mean for the House of Representatives for each year (the dashed line) to the mean for each religious tradition. The Congressional average rises and falls slightly over time, but the average LCV score for all members over the period is exactly 50. The figure shows clear differences between religious traditions. LDS representatives almost always had the lowest LCV scores of any tradition, with a long-term average of 25. Evangelical Protestants also had below-average scores of 31 over the period. Average scores for both groups differed from the rest of the House at the 0.01 level in every year examined (all tests two-tailed). Mainline Protestants’ scores were also lower than average at 42, significantly different from the rest of the House at the 0.01 level.
Figure 3. LCV Scores Across Religious Traditions.
Note: Dashed line is chamber average. Solid lines are averages for each tradition.
White Catholics had higher LCV scores, with an overall average of 58, significantly different from the rest of the House at the 0.05 level or better, except in 1985 (p < 0.10). During the early years, the four or five Hispanic Catholic members had lower than average LCV scores, but as their numbers increased, the average LCV score for this group rose to 65, significantly higher than average in the 2000s (p < 0.05). Black Protestants also supported the LCV’s positions, with an average of 76. Despite their small numbers, black Protestants scored higher than the rest of the House at the 0.05 level in every year, except in 1985 (p = 0.08). Jewish members had the highest average over the period (80), an average significantly above the rest of the chamber at the 0.01 level in every year studied.

Finally, we note that the residual category had LCV scores quite close to the average (53). However, as we might expect from a group whose composition changes over time, the average score varies over time. We can make little of the results for this group substantively, other than to say its variability makes it a poor comparison group.

Figure 3 shows clear differences in support for the LCV’s position across religious traditions. Most of these differences remain when we control for two additional member traits: sex and region. We estimate a model for each year that includes indicator variables for each religious tradition along with indicators for sex (0 = male, 1 = female) and southern residence (0 = non-south, 1 = south). Rather than making an arbitrary choice about which religious group to omit, we estimate the model using restricted ordinary least squares, adding the constraint that the coefficients for the religious tradition indicator variables should sum to one. This approach provides parameter estimates of the religious groups we can interpret as the deviation of the group’s LCV score from the House’s mean score (Greene and Seaks 1991). Following Anderson (2011), we report robust standard errors (results are similar with non-robust errors). As expected, throughout the period, southern representatives had LCV scores significantly lower than other legislators, while women often had significantly higher scores than men had (see Table 1).

Controlling for these effects does little to change estimated differences across religious traditions. LDS representatives consistently had significantly lower LCV scores than average. The negative coefficient for this group was generally between 20 and 30 points, meaning the average for this group was 20–30 points lower than the House average, controlling for legislator sex and southern residence. Evangelical Protestants’ LCV scores were also significantly lower than average, typically by 10–20 points. In addition, Mainline Protestants had significantly lower LCV scores than average, typically by 5–10 points. White Catholics’ scores were typically about average, although in three instances they were significantly higher than average. Jewish members’ LCV scores were 20–30 points higher than
Table 1. Religion and LCV scores.

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Notes: Robust standard errors in parentheses; + significant at 10%; * significant at 5%; ** significant at 1%.
Estimates for religion categories signify the group’s difference from the House mean.
LCV, League of Conservation Voters; LDS, Church of Jesus Christ of Latter Day Saints.
average throughout the period. To provide a visual summary of the results, we present Figure 4, which graphs the coefficients for the five religious traditions we emphasize in the 11 years we examined. We explored several more complex models interacting various independent variables, but in each case, religious tradition maintains its statistical and substantive significance.\(^\text{10}\)

Given the strong relationship between political party and LCV scores found in the past, and the growing relationship between party and religious traditions, we expect party affiliation to mediate some of the differences across religious traditions. Table 2 presents results of models identical to those in Table 1, adding a control for party. The results show Democrats tend to have scores roughly 25–40 points higher than that of Republicans. In fact, in 2009, all else equal, Democrats scored almost 50 points higher than Republicans.

Controlling for party attenuates the religious tradition coefficients considerably, such that Mainline Protestants and white Catholics were seldom significantly different than average, all else equal. LDS and Evangelical legislators continued to have significantly lower than average scores throughout the period, while Jewish members continued to have higher than average scores. Party affiliation is clearly mediating much more of the relationship later in the period, when representatives more cleanly ‘sorted’ themselves into parties according to religious affiliations (e.g., Layman 1999, Oldmixon and Calfano 2007). The coefficients for LDS, Evangelical, and Jewish representatives tend to move toward zero over time. This is clearly seen in Figure 5, which presents the coefficients for these three groups over time. While these coefficients shrank, the coefficient for Democrats increased from the low 20s to the 40s. As the nexus between party affiliation and religious traditions has become tighter, it is harder to sort out the independent effect of party and religion. Although we do not present the results due to space constraints, we see a similar pattern of results when we control for members’ general ideological orientation, as measured by DW-NOMINATE scores. The high level of collinearity between these ideology scores and partisanship (\(r = 0.87\)) precludes us from controlling for both member party and ideology.

As we noted above, economic and political traits of particular district constituencies may shape LCV scores. Since environmental legislation affects districts in unique ways, we might expect representatives of those districts to respond to the interests of their constituents. For example, a representative from a district where many jobs are tied to the mining industry may be especially likely to oppose legislation that would harm the local economy. In addition, the constituency’s general political cast will likely shape who gets elected and possibly constrain representatives once in
Figure 4. Religious Tradition Coefficients from Table 1.
Note: Bars represent unstandardized restricted OLS coefficients for groups in each year. Interpret coefficients as the deviation from the chamber’s average LCV score in a given year, controlling for representative region and sex.
Table 2. Religion and LCV scores controlling for member party.

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Notes: Robust standard errors in parentheses; + significant at 10%; * significant at 5%; ** significant at 1%. Estimates for religion categories signify the group’s difference from the House mean.
Figure 5. Religious Tradition Coefficients Controlling for Party (Table 2).
Note: Bars represent unstandardized restricted OLS coefficients for groups in each year. Interpret coefficients as the deviation from the chamber’s average LCV score in a given year, controlling for representative region, sex, and party.
office, such that heavily Democratic districts will elect Democrats highly supportive of environmental measures.

We follow Anderson (2011), who identifies environmentally relevant district-level variables that affect LCV scores, to include the percent of a district’s population living in urban areas; the percent of district employment in the manufacturing, mining, and agricultural sectors; and district median household income. These data come from the US Census, which did not report most of these figures at the Congressional district level prior to the 1990 Census, so we can only estimate models for 1995–2009. In addition, Anderson finds that district variation in the number of members of environmental organizations (Sierra Club, The Nature Conservancy, The National Wildlife Federation, and The National Resource Defense Council) affects LCV scores. We also include her measures, which differ over time. We note that these variables introduce some multicollinearity into the model, which may inflate standard errors. Therefore, we proceed with some caution as we interpret the results.

Columns 1–5 of Table 3 present estimates for a model that includes all of the variables from models in Table 2, along with the district-level variables Anderson (2011) employed. As expected, many of these district-level variables are significantly related to LCV scores. Controlling for these variables slightly attenuates the coefficients for religious traditions (compare coefficients in Table 3 to those in Table 2). The coefficients for Mormons moved one or two points closer to zero compared to those in Table 2, while the Evangelical coefficients moved about three points closer to zero. The coefficients for white Catholic members stayed about the same, while the coefficient for Jewish members shrunk by about six points as well, sometimes less. Thus, some of the relationship between members’ religious tradition and LCV scores stems from environmentally relevant district-level traits.

In columns 6–10, we add the percentage of the district voting for the Democrat in the most recent presidential election, a frequently employed estimate of district-level political preference (e.g., Jacobson 2012). This measure is highly correlated with a few other variables in the model, introducing fairly significant levels of multicollinearity, so we proceed even more cautiously. Including this measure of district political tendencies further erodes the religious tradition coefficients, though religion still plays a direct role in some instances, even in the face of so many controls. Mormon and Evangelical members continued to have LCV scores significantly lower than average in 1995 and 2000, while Mainline Protestant members had scores about two points higher than average, all else equal. White Catholics had scores about two points higher than average in 2005 and 2009, while Jewish and Hispanic Catholic members had significantly higher than average scores in three of the five years under analysis.
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<td>(0.48)</td>
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<tr>
<td>% Manufacturing</td>
<td>0.01</td>
<td>0.14</td>
<td>0.27</td>
<td>0.01</td>
<td>0.05</td>
<td>0.04</td>
<td>0.02</td>
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<td>-0.01</td>
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<td></td>
<td>(0.24)</td>
<td>(0.21)</td>
<td>(0.21)</td>
<td>(0.10)</td>
<td>(0.09)</td>
<td>(0.24)</td>
<td>(0.20)</td>
<td>(0.21)</td>
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(Continued)
Table 3. (Continued).

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<tbody>
<tr>
<td>% Urban</td>
<td>6.14</td>
<td>1.64</td>
<td>−3.68</td>
<td>3.55</td>
<td>5.40</td>
<td>4.82</td>
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<td>(3.62)+</td>
<td>(3.13)</td>
<td>(3.24)</td>
<td>(4.09)</td>
<td>(3.83)</td>
<td>(3.67)</td>
<td>(3.25)</td>
<td>(3.30)**</td>
<td>(4.21)</td>
<td>(3.81)*</td>
</tr>
<tr>
<td>Group membership</td>
<td>2.07</td>
<td>2.43</td>
<td>2.28</td>
<td>21.61</td>
<td>10.52</td>
<td>1.74</td>
<td>1.99</td>
<td>1.89</td>
<td>13.10</td>
<td>1.07</td>
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<tr>
<td></td>
<td>(0.50)**</td>
<td>(0.41)**</td>
<td>(0.41)**</td>
<td>(3.81)**</td>
<td>(2.80)**</td>
<td>(0.52)**</td>
<td>(0.44)**</td>
<td>(0.43)**</td>
<td>(3.90)**</td>
<td>(3.12)</td>
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<tr>
<td>% Voting</td>
<td>24.79</td>
<td>50.25</td>
<td>56.45</td>
<td>39.23</td>
<td>46.09</td>
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<td>56.45</td>
<td>39.23</td>
<td>46.09</td>
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<tr>
<td>Democrat</td>
<td>(11.16)*</td>
<td>(7.88)**</td>
<td>(7.73)**</td>
<td>(7.27)**</td>
<td>(7.78)**</td>
<td>(11.16)*</td>
<td>(7.88)**</td>
<td>(7.73)**</td>
<td>(7.27)**</td>
<td>(7.78)**</td>
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<tr>
<td>Constant</td>
<td>27.63</td>
<td>25.26</td>
<td>31.81</td>
<td>19.50</td>
<td>25.36</td>
<td>10.90</td>
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<td>−2.46</td>
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<tr>
<td></td>
<td>(4.97)**</td>
<td>(4.30)**</td>
<td>(4.77)**</td>
<td>(5.09)**</td>
<td>(4.65)**</td>
<td>(9.41)</td>
<td>(6.11)</td>
<td>(6.31)</td>
<td>(5.33)</td>
<td>(4.81)+</td>
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Notes: Robust standard errors in parentheses; + significant at 10%; * significant at 5%; ** significant at 1%.
Religion estimates signify the group’s difference from the House mean. Group Membership measures differ by decade (Anderson 2011).
Note that members’ party affiliations are highly related to LCV scores, such that even when we control for a host of district characteristics, Democrats’ LCV scores are roughly 35–40 points higher than Republicans’ scores. We do not want to push this point too far, but since religious tradition and party affiliation are increasingly linked, we are underestimating religion’s effects if we only look at the direct effects in Table 3. Considering religion’s effect as mediated by party shows a consistently strong effect, albeit mostly indirect. If we estimate a simple path model that calculates religion’s overall effect on LCV scores as the sum of the direct effect (the coefficients for religious groups in columns 6–10 in Table 3) and the indirect effect of religion as mediated by legislator party, we see a strong total effect of religion. For example, if we take into account the indirect effect of LDS members’ propensity to be Republicans, the total effect of LDS membership for the 1995–2009 period averages –20 points. The mean total effect for Evangelical Protestants is –17 points. That is, LDS and Evangelical representatives had LCV scores 20 and 17 points lower than average, respectively. On average across this period, Mainline members had LCV scores seven points below average, while Jewish members and Hispanic Catholics had scores 16 and 11 points above average, respectively. Again, we hesitate to emphasize these estimates, since such effects can be difficult to estimate, but simply note that religion’s overall effects appear significant, even in the face of several district-level characteristics.

Conclusion

To summarize, over almost four decades, House members’ religious affiliation has been strongly related to environmental roll-call votes. District traits and, increasingly, party affiliation mediate a good portion of the relationship. The distinctiveness of religious traditions’ LCV scores and the ‘sorting’ of members from these traditions more cleanly into parties may account for some of the increase in party polarization in environmental politics in the US House of Representatives. These findings point to an important implication for environmental policy. The tighter religion–party connection suggests that environmental legislation may be increasingly difficult to pass in the US. Now that religious traditions have been more completely segregated into political parties, the parties may find themselves increasingly unable to find common ground on environmental legislation. If legislators’ environmental policy preferences cut across religious traditions, religion could conceivably provide some common ground on which to construct legislation. However, now that environmental policy increasingly separates both the parties and members of different religious traditions, the divides may become unbridgeable. Given both party and religious
polarization on environmental policy, it may prove increasingly difficult to effect policy change. If the US maintains its policy status quo, current trends (predicted to lead to disaster) are likely to continue. Moreover, inaction in the US may hinder environmental policy development globally.

However, the patterns we report may not persist. Evangelical Protestants have sometimes changed their political approach significantly over time. Evangelicals throughout US history have sometimes taken policy stances that may seem surprising now, for instance advocating for economic populism and abolition of slavery in the nineteenth century (Putnam and Campbell 2010). For a period lasting into the 1970s, Evangelicals were not particularly politically active or even all that Republican (Guth 1983). Recall that in the 1970s and 1980s, more than half of Evangelical representatives were Democrats. In recent years, Evangelicals have engaged in a vigorous debate over environmental politics, with some groups (e.g., Evangelical Climate Initiative) pushing Congress to enact policies reducing carbon emissions, while others (e.g., Interfaith Stewardship Alliance) oppose allocating resources to climate change (see Nagle 2008). Thus, Evangelicals may be shifting their views on at least some environmental policies. Perhaps these debates may shift some current or future members of Congress as well.

In closing, we point to three questions for future research. First, how strong are religion’s effects outside the US? The relationships between religious traditions and environmental attitudes among European publics differ from those found in the US, although there are some notable congruencies (Hagevi 2014). The greater secularization of European citizens and elites suggest that religion’s effects will be less direct in European parliaments than in Congress. Nevertheless, ‘interviews with MEPs offer evidence that religion does have an effect on the functioning of the EP’ (Foret 2014b, p. 131). Given the European Union’s importance in environmental policy development, we must understand any effects, direct or indirect, on its environmental policy making.

Second, how strong are religion’s effects beyond roll-call voting? We suspect religious effects are even stronger in other areas. By the time a bill reaches a roll-call vote the up-or-down nature of the choice and party pressure constrain legislators’ choices. Legislators often have considerably more leeway to engage in other legislative behavior, such as bill co-sponsorship, committee work, or delivering speeches. Religious influences are often more pronounced on these types of behaviors in the US (see Burden 2007, Oldmixon and Calfano 2007). This is an important point for studying party-dominated parliaments, where religion’s influence may be real but difficult to observe at the final stage. Religion’s impact may be more pronounced, for instance, in the development of parties’ or coalitions’ policy agendas.
Third, how do other dimensions of religion affect environmental policy making? Our religion measure, based on affiliation, is one-dimensional and often rather coarse. Although religious traditions have distinctive features, affiliation is only a rough proxy for more important dimensions of religion: belief and behavior. In the US, religious beliefs are often much more powerful predictors of legislative behavior than simple affiliation (Yamane and Oldmixon 2006, Guth 2014). Finding significant differences across religious traditions with imperfect measures suggests that religion’s effects may be even stronger than we estimate. Future research should continue to refine the conceptualization and measurement of legislators’ religion to deepen our understanding of religion’s influence on environmental politics in the US and worldwide.

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Disclosure statement

No potential conflict of interest was reported by the authors.

Notes

2. In the PartiRep survey of European MPs, 45% of members reported being in contact with churches or religious organizations at least once every three months (see Steven 2014).
3. According to the LCV (see http://scorecard.lcv.org/methodology, accessed 11 June 2014), 20 experts from environmental organizations identify roll-call votes from 12 different environmental issue areas. Typically, votes count equally, but the LCV occasionally deems a vote important enough to count double.
4. For example, imagine two hypothetical years, one with a Democratic majority (with power to set the roll-call agenda) and the other with a Republican majority. Imagine that in year one, the House voted on legislation to: (1) cap carbon emissions on existing and new plants, (2) raise fuel efficiency standards on new cars, (3) eliminate the use of hydraulic fracking, and (4) increase
regulations for construction near wetlands. In year two, the House voted on legislation to: (1) open the Arctic National Wildlife Refuge for oil drilling, (2) open protected land for increased logging, (3) require the construction of the Keystone XL pipeline, and (4) provide subsidies for companies using hydraulic fracking. A score of 50 in year one (supporting two strong pro-environment policies) differs from a score of 50 in year two (supporting two environmentally disruptive policies). Of course, these hypothetical years include agendas more extreme than normal, but they demonstrate the need for a shift of the second score downward (or the first upward) for scores to be comparable.

5. The method identifies a baseline year and transforms scores from other years to match this baseline year. Like converting temperature from Celsius to Fahrenheit, this transformation involves shifting the scores up or down, then multiplying them to ‘stretch’ them so that the ends of the scale hold the same meaning.

6. Coding these generic Christian responses as a separate category does not alter our results. We coded all representatives listing ‘Baptist’ as Evangelical Protestant and all representatives listed as ‘Lutheran’ or ‘Presbyterian’ as Mainline Protestant.

7. These cases include 19 African American members, the vast majority of whom were Catholic, and 12 Hispanic representatives, virtually all affiliated with a Protestant denomination.

8. The average is highest in 1990, when Congress passed a major extension of the Clean Air Act.

9. Robust standard errors adjust for heteroskedasticity, which can generate incorrect standard errors and therefore lead to misleading hypothesis tests. Breusch-Pagan tests found heteroskedasticity in about 40% of our models.

10. For example, we interacted the Evangelical indicator variable with southern residence to see whether southern Evangelicals differed from southern non-evangelicals. The difference between Evangelical Republicans and non-Evangelical Republicans in the South (five points) is smaller than the corresponding difference outside the south (13 points), but Evangelicals differed from non-Evangelicals at the 0.05 level in both locations. We present the simpler results for ease of presentation, though additional results are available upon request.

11. We thank Sarah Anderson for sharing the district-level data. The 1990 Census reports district-level variables based on the 1990s round of redistricting, which was not in place until 1992. Therefore, we cannot use these data to estimate a model for 1990.


13. District-level income and district-level environmental group membership correlate at between 0.50 and 0.61. Percent of the district employed in the agricultural sector correlates with percent urban at 0.50–0.60.

14. From 2000 onward, district-level presidential vote and representation by a Democrat correlate at 0.61–0.70. District-level presidential vote and percent
urban correlate at between 0.55 and 0.59. The Variance Inflation Factor for
district presidential vote share ranges from 54 to 69, where scores >10
typically indicate significant multicollinearity.

15. Path models require that causality only runs from religious tradition to party
affiliation. We think it makes sense that members’ religious traditions would
shape their party preference (Guth et al. 2009). It seems unlikely that many
people first affiliate with a political party and then as a result of that party
choice determine which religious tradition to join. However, we recognize
that path models can include considerable error in estimates, so we interpret
the results with caution. Our purpose is simply to demonstrate that sig-
nificant indirect effects are present, not to make precise claims about the
exact size of those effects.

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