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**Religiosity and Income: A Panel
Cointegration and Causality Analysis**

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Religiosity and Income: A Panel Cointegration and Causality Analysis

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Abstract. In this paper we examine the long-run relationship between religiosity and income using retrospective data on church attendance rates for a panel of countries from 1925 to 1990. We employ panel cointegration and causality techniques to control for omitted variable and endogeneity bias and test for the direction of causality. We show that there exists a negative long-run relationship between the level of religiosity, measured by church attendance, and the level of income, measured by the log of GDP per capita. The result is robust to alternative estimation methods, potential outliers, sample selection, different measures of church attendance, and alternative specifications of the income variable. Long-run causality runs in both directions, higher income leads to declining religiosity and declining religiosity leads to higher income.

Keywords: religiosity; church attendance, income, panel cointegration, causality.

JEL: N30; O11; C23.

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1. INTRODUCTION

Secularization, broadly understood, describes the phenomenon that later born generations appear to be less religious, revealed, for example, by lower weekly attendance rates at church (see e.g. Wilson, 2003; Voas 2009). The narrow definition of secularization assigns a cause to this time trend, inspired by the observation that secularization occurred in the Western world after the Industrial Revolution and the take off of economic growth (Azzi and Ehrenberg, 1975; Dobbelaere, 1987; Norris and Inglehart, 2004; Bruce, 2011). According to this so called “secularization hypothesis” improving economic conditions have caused the decline in religiosity and the decline of demand for religious services. McCleary and Barro (2006a) provide a concise introduction to the literature on the secularization hypothesis.¹

Religious beliefs and the associated behavior, in turn, may have an impact on economic performance. Recent developments in the theory and empirics of economic growth have increasingly focussed on the “fundamental determinants” of growth, comprising culture, geography, and institutions, which are thought of as causal drivers of the proximate determinants of economic growth, comprising physical and human capital accumulation and technological change (Acemoglu, 2009). Religiosity is certainly an important element of culture but the direction of causality with respect to growth is a priori unclear. On the one hand, increasing religiosity could be conducive to economic growth because it encourages trust and trustworthiness (Guiso et al., 2003) or because it entails an incentive to accumulate human capital (Becker and Woessmann, 2009). On the other hand increasing secularization could be conducive to growth because the turn towards material values and the promised pleasures from consumption induce increasing labor supply and capital accumulation (Lipford and Tollison, 2003; Strulik, 2012). Besides the possibility of forward and backward causation there exists, of course, the possibility that the decline of religiosity is unrelated to economic growth.

There exist already a few quantitative studies on the income – religiosity nexus. McCleary and Barro (2006a) found a causal negative effect of income on religious participation and beliefs across countries, which is also quantitatively important. A one standard deviation increase in

¹In principle, church attendance can be seen as a plausible revealed preference for religion because it entails an opportunity costs. Although declining attendance does not necessarily imply that not attending persons stopped believing, it documents that these persons assign less value to religious services. In this sense their religiosity is declining. Church attendance may actually overestimate religiosity because attending could be driven by secular motives like social capital accumulation. In any case, attendance is the most widely available and most frequently used religiosity variable.

log GDP per capita decreases church attendance by 15 percent. McCleary and Barro (2006b) arrived at similar results. The latter study documents also a significant negative impact of church attendance on economic growth. Paldam and Gundlach (2013) used the World Value survey to compile a religiosity index (14 items from “God is very important in life” to “Churches answer spiritual needs”) and found across countries a causal negative impact of income on religiosity. On average, religiosity falls by 50 percent when countries pass through the transition from being underdeveloped to becoming a developed country. Becker and Woessmann (2013) could not find a causal effect of (teacher-) income on church attendance in Prussia 1886-1911. Lipford and Tollison (2003) documented a bi-causal negative association of income and religious participation across US states. Rupasingha and Chilton (2009) used US county level data on religious adherence and found a causal negative effect of adherence on economic growth. An increase of religious adherence by one standard deviation would reduce growth by 0.4 percent per year.

So far, however, little attention has been paid to the time dimension in the analysis of religiosity and economic growth. On the one hand this is understandable because data on religiosity was not available for sufficiently many time periods to allow for a rigorous time series analysis.² On the other hand this state of the art is deplorable since both economic growth and the decline of religiosity (church attendance) are inherently dynamic phenomena. In particular, secularization is usually not interpreted as differential religiosity across economically diverse countries but as within-country process along the time-dimension. The question is whether an observable decline in religiosity has been caused by a preceding increase of income (or vice versa) and whether countries share a common dynamic relationship between religiosity and income.

In the present paper we try to address these questions. We examine the long-run relationship between religiosity and income using panel cointegration techniques. A feature of cointegration analysis is that the resulting estimates are robust to a variety of estimation problems that often plague empirical work, including omitted variables, endogeneity, and measurement errors (see, e.g., Pedroni, 2007; Herzer, et al., 2012). We identify the direction of causality using Granger-causality tests and impulse response functions, that is by using techniques which are built upon the idea that the cause occurred before the effect.

²McCleary and Barro (2006a) compare estimates from World Value Surveys 1981, 1990, and 1995 and from ISSP data 1991 and 1998 and do not find a significant (autonomous) time trend of church attendance once they control for income and other explanatory variables.

As variable for religiosity we use the estimated national church attendance rates between 1925 and 1990 provided by Iannaccone (2003). The ingenious idea of Iannaccone's study is to estimate historical church attendance rates using contemporary ISSP questionnaires and replies to inquire on church attendance when the respondents were 11 or 12 years old. Naturally, this retrospective method could be harmed by different sorts of biases (e.g. age effects or projection bias). Iannaccone therefore devotes the greater part of his study to carefully demonstrate that there is no reason for concern. The data stands up to numerous test of internal and external consistency. For our cointegration analysis we combine the retrospective attendance data, which is available at five year intervals, with data on GDP per capita from Maddison (2003). Our main finding is a strong negative impact of economic growth on church attendance.

At first sight our results seemingly disagree with those of Franck and Iannaccone (2009) who also used the Iannaccone (2003) data and found little support for a causal impact of income on church attendance. One possible explanation is that the instrument variable technique used by Franck and Iannaccone did not satisfactorily resolve the problems of reverse causality and omitted variables. In contrast, our approach does not require exogeneity assumptions nor does it require the use of instruments. The reason is that, under cointegration, parameter estimates are superconsistent, implying that endogeneity does not affect the results (Engle and Granger, 1987). The superconsistency property holds even in the presence of omitted stationary variables and the presence of non-systematic measurement error.

The finding of cointegration between two nonstationary variables implies that there are no missing relevant nonstationary variables and that no additional nonstationary variables are required to produce unbiased parameter estimates. While a regression consisting of cointegrated nonstationary variables has a stationary error term, any omitted nonstationary variable that is part of the true cointegrating relationship would enter the error term, thereby producing non-stationary residuals and failure to detect cointegration. Since the cointegration property is invariant to extensions of the information set, the finding of a particular cointegration relationship in a small set of variables will also hold in an extended variable set. This invariance of the cointegration property to extensions of the information set implies that the classical omitted variables problem does not exist under cointegration.

These properties are important in light of the secularization debate because another potential driver of church attendance could be the diversity of supply of religious services. If this is the

indeed case – as argued by Finke and Iannaccone (1993) and other proponents of the religious supply hypothesis – then it would not bias our results. The inclusion of additional variables would not destroy the original cointegrating relationship (Lütkepohl, 2007). This is the main justification to consider small “subsystems” (such as the relationship of church attendance and income) if the variables are cointegrated. Of course, this means also vice versa that the strong evidence for a causal impact of income growth on church attendance established in our study does not refute the religious supply hypothesis. If such a supply channel exist, it would be added to the set of explanatory variables without affecting the income channel.

The paper is organized as follows. In Section 2.1 we set up the basic empirical model and discuss some econometric issues. We then describe the data and report summary statistics, including pre-tests for unit-roots and cointegration (Section 2.2). In Section 3 we present the empirical analysis. In Section 3.1 we provide estimates of the cointegrating relationship between religiosity and income, in Section 3.2 we test the robustness of the estimates, and in Section 3.3 we investigate the direction of long-run causality between the two variables. We conclude in Section 4.

2. MODEL AND DATA

2.1. Basic Empirical Model and Econometric Methodology. The basic econometric specification used to examine the long-run relationship between religiosity and income is a conventional bivariate panel cointegration model of the form

$$CHURCH_{it} = a_i + \beta \log(y_{it}) + \epsilon_{it} \quad (1)$$

where $i = 1, 2, \dots, N$ is the country index, $t = 1, 2, \dots, T$ is the time index, and the a_i are country-specific fixed effects. Following previous studies, we use church attendance as our measure of religiosity ($CHURCH_{it}$), while income is measured by real GDP per capita (y_{it}). The income variable is logged, as in most previous studies (see, for example, McCleary and Barro, 2006; Franck and Iannaccone, 2009; Becker and Woessmann, 2013), but in the robustness section, we also estimate the model using the non-log-transformed income variable, as in Lipford and Tollison (2003). Using the log specification allows for an intuitive interpretation of results. To see this, differentiate (1) and obtain the change of church attendance $dCHURCH_{it}$ as a

function of the growth rate of GDP per capita, $dgdip_{it}/gdp_{it}$. Equation (1) thus stipulates that changing religiosity is associated with income growth.

We now turn to econometric issues. The first observation is that the underlying variables are trended; they are nonstationary (as shown in Figures 1 and 2). Given that most economic time series are characterized by a stochastic rather than deterministic nonstationarity, it is plausible to assume that the trends in $CHURCH_{it}$ and $\log(y_{it})$ are also stochastic – through the presence of a unit root – rather than deterministic – through the presence of polynomial time trends. If this assumption is correct, the linear combination of these integrated (or stochastically trending) variables must be stationary, or, in the terminology of Engle and Granger (1987), $CHURCH_{it}$ and $\log(y_{it})$ must be cointegrated. If the two variables are not cointegrated, then there is no long-run relationship between religiosity and income, and Equation (1) would be a spurious regression in the sense of Granger and Newbold (1974). Standard regression output must therefore be treated with extreme caution when variables are nonstationary, since the estimated results are potentially spurious (see also Eberhardt and Teal, 2013). As shown by Entorf (1997) and Kao (1999), the tendency for spuriously indicating a relationship may even be stronger in panel data regressions than in pure time-series regressions. Thus, the necessary condition for the existence of a non-spurious long-run relationship between $CHURCH_{it}$ and $\log(y_{it})$ is that the two integrated variables cointegrate.³

A regression consisting of cointegrated variables has the property of superconsistency such that coefficient estimates converge to the true parameter values at a faster rate than they do in standard regressions with stationary variables, namely at rate T rather than \sqrt{T} (Stock, 1987). The important point in this context is that the estimated cointegration coefficients are superconsistent even in the presence of temporal and/or contemporaneous correlation between the stationary error term, ϵ_{it} , and the regressor(s) (Stock, 1987), implying that cointegration estimates are not biased by omitted *stationary* variables (see, e.g., Bonham and Cohen, 2001).

The fact that a regression consisting of cointegrated variables has a stationary error term also implies that no relevant *nonstationary* variables are omitted. Any omitted nonstationary variable that is part of the cointegrating relationship would become part of the error term,

³The standard time-series approach is to first-difference the variables to remove the nonstationarity in the data and to avoid spurious results. However, this approach precludes the possibility of a long-run or cointegrating relationship in the data and leads to misspecification if a long-run relationship between the levels of the variables exists (see, e.g., Granger, 1988).

thereby producing nonstationary residuals, and thus leading to a failure to detect cointegration (Everaert, 2011).

If there is cointegration between a set of variables, then this stationary relationship also exists in extended variable space. In other words, the cointegration property is invariant to model extensions (see also Lütkepohl, 2007), which is in stark contrast to regression analysis where one new variable can alter the existing estimates dramatically (Juselius, 2006, p. 11). The important implication of finding cointegration is that no additional variables are required to account for the classical omitted variables problem because such a problem does not exist under cointegration; the result for the long-run relationship between religiosity and income would also hold if we included additional independent variables in the model (see also Juselius, 1996).

Of course, there are several other factors such as education, fertility, and government expenditure that may be associated with religiosity and/or income. Therefore, adding further nonstationary variables to the model may, on the one hand, result in further cointegrating relationships. These, however, would have to be identified and estimated (separately). In particular, the difficulty is that if there is more than one stationary linear combination of the variables, identifying restrictions are required to separate the cointegrating vectors. On the other hand, adding further nonstationary variables to the regression model may result in spurious associations. More specifically, if a nonstationary variable that is not cointegrated with the other variables is added to the cointegrating regression, the error term will no longer be stationary. As a result, the coefficient of the added variable will not converge to zero, as one would expect of an irrelevant variable in a standard regression (Davidson, 1998). This justifies a reduced-form model such as Equation (1), given the variables are cointegrated.

The superconsistency of the cointegration estimation also implies that the potential endogeneity of the regressors does not affect the estimated long-run coefficients; the estimated long-run coefficients from reverse regressions should be approximately the inverse of each other due to the superconsistency (Engle and Granger, 1987). Nevertheless, there are two problems.

First, although the standard least-squares dummy variable estimator is superconsistent under panel cointegration, it suffers from a second-order asymptotic bias arising from serial correlation and endogeneity. As a consequence, its t -ratio is not asymptotically standard normal. To deal with this problem, one has to employ an asymptotically efficient (cointegration) estimator. Examples of such estimators include panel versions of the dynamic OLS (DOLS) and fully

modified ordinary least squares (FMOLS) methods. As shown by Wagner and Hlouskova (2010), the panel DOLS estimator of Mark and Sul (2003) outperforms other asymptotically efficient estimators. Therefore, this estimator is preferred here, but in the robustness section we also present results based on alternative estimation procedures.

Second, although the existence of cointegration implies long-run Granger causality in at least one direction (Granger, 1988), cointegration says nothing about the direction of the causal relationship between the variables. A statistically significant cointegrating relationship between $CHURCH_{it}$ and $\log(y_{it})$ does therefore not necessarily imply that, in the long run, changes in income cause changes in religiosity. The causality may run in the opposite direction, from $CHURCH_{it}$ to $\log(y_{it})$, or in both directions. The empirical implication is that it is important not only to employ an asymptotically efficient cointegration estimator (to account for the potential endogeneity of income), but also to explicitly test the direction of long-run causality. As is common practice in testing long-run Granger causality between cointegrated variables, we use a vector error correction model (VECM) to identify cause and effect in the sense of Granger (1988).

A final econometric issue is the potential cross-sectional dependence in the regression errors due to common shocks or spillovers among countries at the same time. Standard panel (cointegration) techniques assume cross-sectional independence and may be biased if this assumption does not hold. Therefore, we test for cross-sectional dependence in the residuals of the estimated models using the cross-sectional dependence (CD) test developed by Pesaran (2004). In cases where we find evidence of cross-sectional dependence, we employ demeaned variables to control for common effects; that is, in place of $CHURCH_{it}$ and $\log(y_{it})$, we use

$$CHURCH'_{it} = CHURCH_{it} - \overline{CHURCH}_t, \quad \overline{CHURCH}_t \equiv N^{-1} \sum_{i=1}^N CHURCH_{it} \quad (2a)$$

$$\log(y_{it})' = \log(y_{it}) - \overline{\log(y_t)}, \quad \overline{\log(y_t)} \equiv N^{-1} \sum_{i=1}^N \log(y_{it}), \quad (2b)$$

which is equivalent to including time dummies in the model. Moreover, we use a battery of panel unit root and cointegration tests, including so-called second-generation panel unit root and cointegration methods that explicitly allow for cross-sectional dependence.

2.2. Data, Unit root and Cointegration Tests. Data on real per capita GDP are taken from Maddison (2003), available at <http://dx.doi.org/10.1787/456125276116>. Data on church

attendance are from Iannaccone (2003), who uses retrospective survey questions on church attendance rates for respondents and their parents from the International Social Survey Program to construct average weekly church attendance rate of parents and children in 32 countries between 1925 and 1990. Attendance of parents is our preferred measure of religiosity since religious commitments typically develop during adolescent years rather than during early childhood. The attendance rate for children is used in the robustness section. The data are expressed as a percentage of (the parents of) the respondents.

Given that the data set of Iannaccone (2003) spans a long time period, it is inappropriate to use conventional small T panel data models, which ignore the potential nonstationarity of the variables (see, e.g., Phillips and Moon, 2000). The appropriate approach is to use panel time-series techniques to account for the time-series properties of the variables and to avoid spurious results by testing for cointegration. Cointegration estimates are not only robust to omitted variables and endogenous regressors (as discussed above) but also robust to non-systematic measurement errors (Stock, 1987). The latter is an important advantage for applications such as the present one, because it is likely that church attendance rates based on respondents' self-reports are measured with error.

The data on church attendance are available only every five years. That we are forced to use five-year data for our analysis should not be a serious problem because panel cointegration methods exploit both the time-series and cross-sectional dimensions of the data, and can therefore be implemented with a smaller number of time-series observations than their time-series counterparts. The panel cointegration analysis by Madsen et al. (2010), for example, is based on $T = 13$; accurate critical values for panel unit roots and cointegration tests are available even for $T = 10$ (see, e.g., Pesaran, 2007; Banerjee and Carrion-i-Silvestre, 2011). Moreover, it is well known that the total length of the sample period, rather than the frequency of observation, is the important factor when analyzing the integration and cointegration properties of variables (see, e.g., Shiller and Perron 1985; Hakkio and Rush 1991; Lahiri and Mamingi 1995). In addition, several studies show that cointegration estimates are remarkably stable across frequencies (see, e.g., Chambers, 2001; Click and Plummer, 2005; Herzer, 2013).

However, a potential disadvantage is that the estimators we use are designed for balanced panels, while the underlying data sets are unbalanced in the sense that the number of time-series observations per country varies. To construct a balanced panel, which entails a trade-off

between the time span and number of countries in the sample, we select all countries for which complete time-series data are available over the period 1930-1990 (a reasonably long period to conduct cointegration analysis). This sample selection procedure yields a sample of 17 countries and 13 time-series observations per country (221 total observations). In the robustness section, we also estimate the long-run relationship between $CHURCH_{it}$ and $\log(y_{it})$ using different samples over different time intervals ($N = 12, T = 14$ and $N = 20, T = 12$). Table 1 lists the countries in our main sample along with the average values for $CHURCH_{it}$ and $\log(y_{it})$ over the period 1930-1990. The United States had the highest GDP per capita, while Bulgaria had the lowest GDP per capita. Church attendance was highest in Ireland and lowest in Japan. The latter is not surprising, given that Japan is the only country in the sample where Christians are the minority.

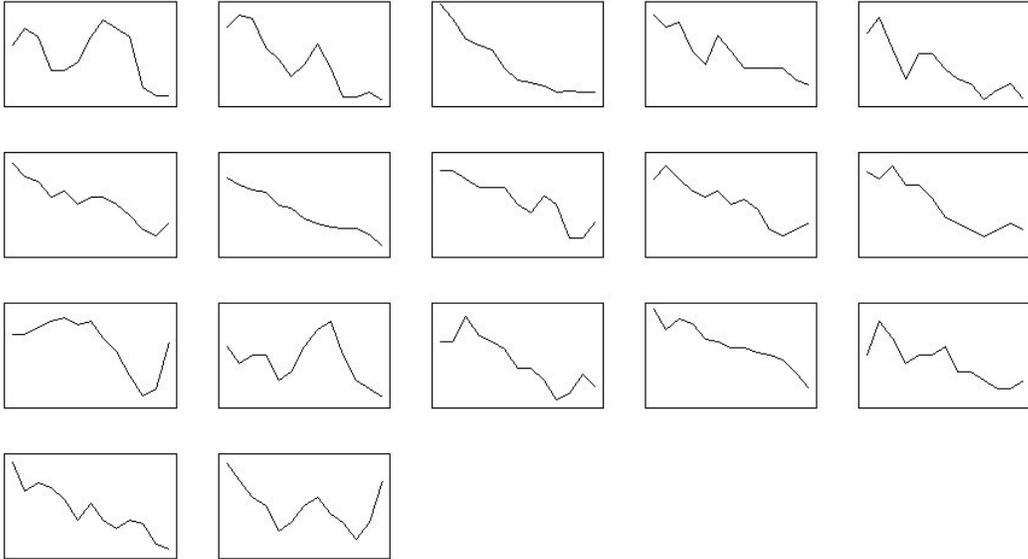
TABLE 1. Countries and Summary Statistics

	Average of $CHURCH_{it}$	Average of $\log(y_{it})$
Australia	31.85	9.13
Austria	50.46	8.75
Bulgaria	19.54	7.93
Chile	42.38	8.33
Denmark	12.08	9.12
France	31.31	8.93
Germany	26.65	8.94
Ireland	94.54	8.50
Italy	60.69	8.71
Japan	7.77	8.49
Netherlands	53.23	9.02
New Zealand	31.31	9.11
Norway	15.00	8.94
Spain	53.85	8.36
Sweden	13.08	9.10
UK	26.38	9.11
United States	55.92	9.38

In Figures 1 and 2, $CHURCH_{it}$ and $\log(y_{it})$ are plotted for each country over the observation period. While GDP per capita increased in all countries, church attendance decreased in all countries between 1930 and 1990. The downward trend in church attendance is especially strong in Austria, Bulgaria, Chile, Denmark, France, Germany, Ireland, Italy, Japan, Norway, Spain, Sweden, and the United Kingdom where church attendance declined relatively steadily, compared to Australia, the Netherlands, New Zealand, and the United States where church

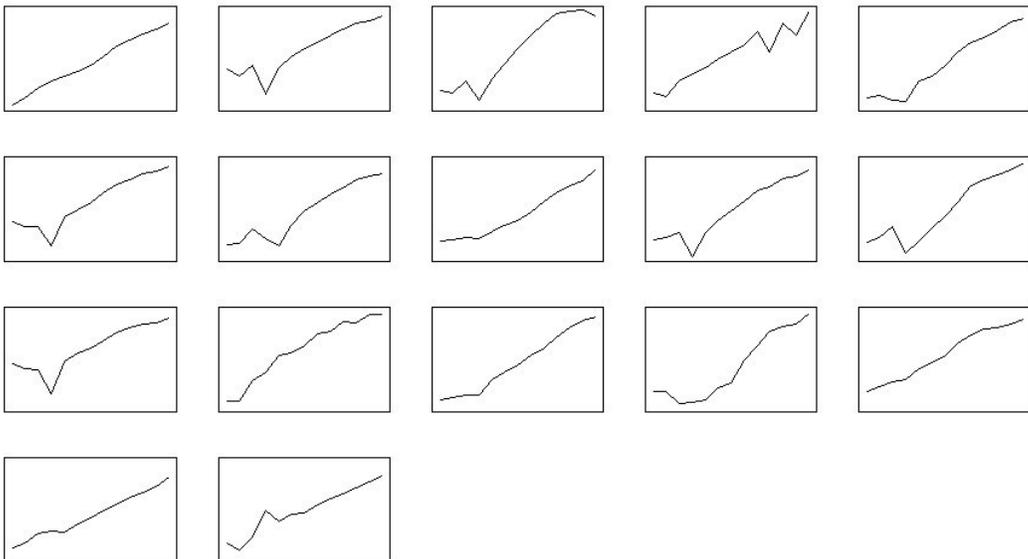
attendance rates exhibit higher volatility. Overall, the time-series evolution is consistent with the possibility that $CHURCH_{it}$ and $\log(y_{it})$ are driven by stochastic trends.

FIGURE 1: CHURCH ATTENDANCE OF PARENTS 1930-1990



The panels in the Figure shows the evolution of church attendance over time for the 17 countries of our preferred sample.

FIGURE 2: LOG OF GDP PER CAPITA:1930-1990 1930-1990



The panels in the Figure shows the evolution of log GDP per capita over time for the 17 countries of our preferred sample.

In order to investigate this issue formally, we conduct panel unit root tests. In recent years, a number of panel unit root tests have been developed. The most commonly used are the

so-called first generation panel unit root tests, such as the ADF-Fisher-type test of Madalla and Wu (1999) (MW), the Breitung (2000) test, the Levin et al. (2002) test, and the Im et al. (2003) test. Hlouskova and Wagner (2006) find that the Breitung panel unit root test generally has the highest power and smallest size distortions of any of the first-generation panel unit root tests. Therefore, we use the Breitung test. Given, however, that the first-generation tests, which assume cross-sectional independence, exhibit severe size distortions in the presence of cross-sectional dependence, we also use second-generation panel unit root tests to allow for cross-sectional dependence. More specifically, we use the panel unit root tests developed by Breitung and Das (2005) and Pesaran (2007). The Breitung and Das test is an extension of the Breitung test and is based on modified standard errors that are robust to cross-sectional dependence. A potential disadvantage of the cross-sectionally robust test is that it requires that $N < T$, which prevents us from using our main sample (with $N = 17$ and $T = 13$). Therefore, we apply the Breitung-Das procedure to an alternative sample of 12 countries over 14 time periods (a sample that also will be used in the robustness section), while the Pesaran panel unit root test is again applied to the main sample. Like the Breitung test and the Breitung-Das test, the Pesaran test is an ADF type test. It is based on an average of the individual country ADF t -statistics and filters out the cross-sectional dependence by augmenting the individual ADF regressions with the cross-sectional averages of lagged levels and first differences of the individual series.

As can be seen from Table 2, both the Breitung and the Breitung and Das tests fail to reject the null hypothesis of a unit root at the 10% level for both variables. The Pesaran test does not reject the unit root null hypothesis for $CHURCH_{it}$ at the 10% level, while the unit root null for $\log(y_{it})$ is rejected at the 5% level but not at the 1% level. Given that existing panel studies usually conclude that real GDP per capita contains a unit root (see, e.g., Pedroni, 2007; Herzer, et al., 2012), it is reasonable to assume that both $CHURCH_{it}$ and $\log(y_{it})$ are driven by stochastic trends.

In order to ensure that the relationship between $CHURCH_{it}$ and $\log(y_{it})$ is not spurious, we test for cointegration using the standard panel and group ADF and PP test statistics suggested by Pedroni (1999, 2004). However, these tests do not take account of potential error from cross-sectional dependence, which could bias the results. To test for cointegration in the

TABLE 2. Panel Unit Root Tests

Variable	Breitung (2000)	Breitung and Das (2005)	Pesaran (2007)
$CHURCH_{it}$	1.709	0.758	-3.047**
$\log(y_{it})$	-1.074	-0.744	-2.588

The panel unit-root tests are ADF-type tests. The models allow for individual-specific intercepts and time trends. Given the relatively small number of time series observations, only one lag was used to adjust for autocorrelation. Large negative values lead to rejection of a unit root in favor of (trend) stationarity. The Breitung and the Breitung and Das (2005) statistics are asymptotically distributed as a standard normal. The relevant 10% (5%) [1%] critical value for the Pesaran statistic with an intercept and a linear trend is -2.70 (-2.84) [-3.12] (when $N = 17$ and $T = 12$). The critical values are calculated from the response-surface estimates in Otero and Smith (2013). ** indicates significance at the 5% level.

presence of possible cross-sectional dependence, we also employ the panel cointegration test recently proposed by Banerjee and Carrion-i-Silvestre (2011). This test involves four steps. The first is to estimate the parameters of the cointegrating regression using the pooled common correlated effects (CCE) estimation technique advanced by Pesaran (2006). The pooled CCE estimator accounts for unobserved common factors by augmenting the cointegrating regression with the cross-sectional averages of the dependent and independent variables; these averages are interacted with country-dummies to allow for country-specific parameters. In the second step, the estimated parameters are used to construct the residuals from the long-run relationship, $\hat{\mu} = CHURCH_{it} - \hat{\beta} \log(y_{it})$. Then, these long-run residual series are regressed on country dummies D_i to compute OLS residuals from this regression as $\hat{e}_{it} = \hat{\mu} - \hat{D}_i$. Finally, the Pesaran (2007) unit root test is computed for the estimated OLS residuals. If the presence of a unit root in \hat{e}_{it} can be rejected, it can be concluded that there is a cointegrating relationship between the variables.

The results of these tests are presented in Table 3. The ADF and the PP statistics reject the null hypothesis of no cointegration at the 1% level, and the Banerjee and Carrion-i-Silvestre test rejects the null of no cointegration at least at the 5% level (1% critical values are not available in Banerjee and Carrion-i-Silvestre (2011)). These results indicate that there is a long run relationship between religiosity and income. As discussed above, the finding of cointegration also implies that there are no missing (trending) variables and that therefore no additional variables are required in the model given by equation (1).

TABLE 3. Panel Cointegration Tests

Pedroni (1999, 2004)	
Panel ADF statistic	-5.323***
Group ADF statistic	-7.422***
Panel PP statistic	-4.579***
Group PP statistic	-6.712***
Banerjee and Carrion-i-Silvestre (2011)	-2.971**

Given the relatively small number of time series observations, only one lag was used. Large negative values lead to rejection of the null hypothesis of no cointegration. The ADF statistics and the PP statistics are asymptotically distributed as a standard normal. The 5% critical value for the Banerjee and Carrion-i-Silvestre (2011) statistic (for $T = 15$ and $N = 15$) is -2.56. Banerjee and Carrion-i-Silvestre (2011) do not report 1% critical values. *** (**) indicate a rejection of the null hypothesis of no cointegration at the 1% (5%) level.

3. EMPIRICAL ANALYSIS

3.1. Panel Cointegration Estimates. We use the panel DOLS estimator suggested by Mark and Sul (2003) to estimate the long-run relationship between religiosity and income. The DOLS estimator is superconsistent, asymptotically unbiased, and normally distributed, even in the presence of endogenous regressors. Moreover, recent Monte Carlo evidence by Wagner and Hlouskova (2010) suggests that Mark and Sul’s panel DOLS estimator outperforms other estimators, particularly when the number of time-series observations is small. The idea behind this estimator is to account for possible serial correlation and endogeneity of the regressors by augmenting the cointegrating regression with lead, lag, and current values of the first differences of the I(1) regressors. Accordingly, in our case, the DOLS regression is given by:

$$CHURCH_{it} = a_i + \beta \log(y_{it}) + \sum_{j=-k}^k \theta_{ij} \Delta \log(y_{it-j}) + e_{it}. \quad (3)$$

To ensure that our results are not affected by cross-sectional dependence (due to common shocks or spillovers among countries at the same time), we compute the cross-sectional dependence (CD) test suggested by Pesaran (2004). The CD test statistic is defined as

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\rho}_{it} \right) \quad (4)$$

where

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^T \hat{\mu}_{it} \hat{\mu}_{jt}}{\left(\sum_{t=1}^T \hat{\mu}_{it}^2 \right)^{\frac{1}{2}} \left(\sum_{t=1}^T \hat{\mu}_{jt}^2 \right)^{\frac{1}{2}}} \quad (5)$$

is the sample estimate of the pair-wise correlation of the residuals of the estimated models, $\hat{\mu}_{it}$. The CD test statistic is normally distributed under the null hypothesis of no cross-sectional dependence.

As the CD test in the first row of Table 4 shows, the panel DOLS model does not appear to suffer from cross-sectionally dependent residuals; thus, valid inferences can be drawn from the regression results. The DOLS regression shows a highly significant negative relationship between religiosity and income, and the point estimate implies (if viewed causally) that, in the long run, an increase in per capita GDP by one percent decreases the weekly church attendance rate by 8.9 percentage points on average in our sample (holding all else constant). Although this effect seems large, the magnitude is much smaller than that implied by the McCleary and Barro (2006) estimate.

In the second row of Table 4, we also present DOLS results based on demeaned data. While the coefficient on the income variable is still negative and significant, it is smaller (in absolute value) than its counterpart in row 1. However, the CD test indicates the presence of cross sectional error dependence, which could have biased the results. The finding of cross sectional dependence for the demeaned data is consistent with studies showing that the demeaning procedure may introduce cross-sectional correlation among the error terms when it is not already present (see, e.g., Carporale and Cerrato, 2006). In the following, we therefore use demeaned variables only when the CD test indicates the presence of cross sectional dependence.

3.2. Robustness. We perform several sensitivity exercises. First, we examine whether the negative relationship between religiosity and income is robust to alternative estimation techniques. A potential problem with the pooled results (reported in row 1) could be that they are based on the implicit assumption of homogeneity of the long-run parameters. It is well known that, while efficiency gains from the pooling of observations over the cross-sectional units can be achieved when the individual slope coefficients are the same, pooled estimators may yield inconsistent and potentially misleading estimates of the sample mean of the individual coefficients when the true slope coefficients are heterogeneous. Although a comparative study by Baltagi and Griffin (1997) concludes that the efficiency gains from pooling more than offset the biases due to individual country heterogeneity, we nonetheless allow the long-run coefficients to vary across countries by using the group-mean panel DOLS estimator suggested by Pedroni (2001). This estimator involves estimating separate DOLS regressions for each country and averaging the

TABLE 4. Estimates of the Long-Run Relationship Between Religiosity and Income

Estimation method	Coeff. on $\log(y_{it})$	CD stat.	Demeaned	Obs.
(1) Pooled panel DOLS estimator (Mark and Sul, 2003)	-8.944*** (-11.61)	0.47	No	170
(2) Pooled panel DOLS estimator (Mark and Sul, 2003)	-6.545*** (-2.99)	-1.76*	Yes	170
(3) Group-mean panel DOLS estimator (Pedroni, 2001)	-8.295*** (-29.42)	-1.08	No	170
(4) Pooled panel FMOLS estimator (Kao and Chiang, 2000)	-8.821*** (-12.99)	2.55**	No	204
(5) Pooled panel FMOLS estimator (Kao and Chiang, 2000)	-6.401*** (-3.52)	-2.28**	Yes	204

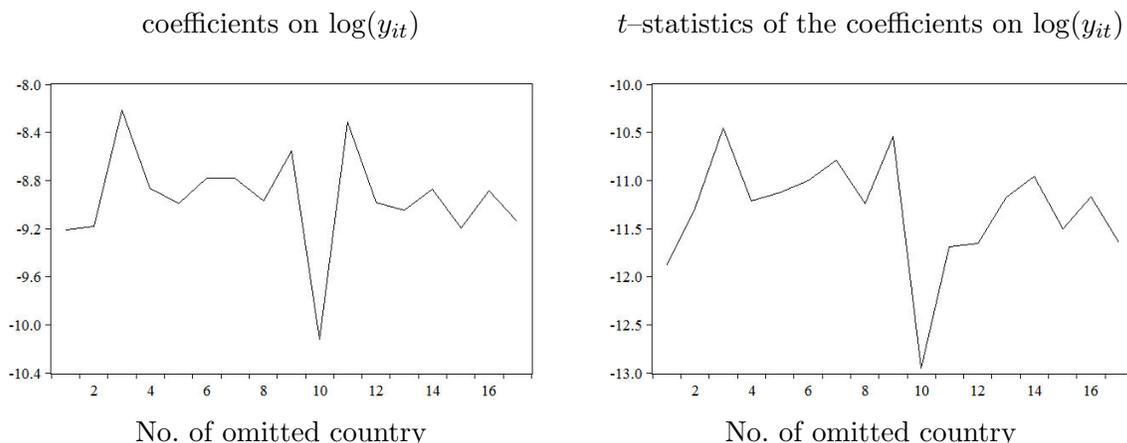
The dependent variable is $CHURCH_{it}$. The DOLS regressions were estimated with one lead and one lag ($k = 1$), given the relatively small number of time series observations. t -statistics are in parenthesis. The CD test statistic is normally distributed under the null hypothesis of no cross-sectional dependence. *** (**) [*] indicate significance at the 1% (5%) [10%] level.

long-run coefficients, $\hat{\beta} = N^{-1} \sum_{i=1}^N$. The corresponding t -statistic is computed as the sum of the individual t -statistics (calculated using heteroskedasticity and autocorrelation-consistent standard errors) divided by the root of the number of cross-sectional units, $t_{\beta} = \sum_{i=1}^N t_{\hat{\beta}_t} / \sqrt{N}$. In addition, we use the panel FMOLS estimator suggested by Kao and Chiang (2000). Like the time series FMOLS estimator, the panel FMOLS estimator incorporates a semi-parametric correction to the OLS estimator, which eliminates the second order bias induced by the endogeneity of the regressors. We report the results of these estimation methods in rows 3-4 of Table 4.

The results show a negative and significant relationship between religiosity and income. Interestingly, the panel and group-mean DOLS estimators produce almost identical coefficients, suggesting that slope heterogeneity is not a serious problem. However, given the relatively small number of time-series observations, and since especially in this case the efficiency gains from pooling are likely to more than offset the potential biases due to individual heterogeneity, we prefer the pooled approach. Specifically, the pooled DOLS estimator is preferred over the pooled FMOLS estimator, since the results of the pooled FMOLS procedure (reported in row 3 and 4) appear to be affected by cross-sectional dependence, even when demeaned data are used, as the CD test statistics show.

Given the relatively small number of countries in our sample, we also need to ensure that the estimated effect is not due to individual outliers. To this end, we re-estimate the DOLS regression, excluding one country at a time from the sample. The sequentially estimated coefficients and their t -statistics are presented in Figure 3. Each number on the horizontal axes represents the country omitted from DOLS regression; on the vertical axes we plot the respective coefficients and t -statistics in the remaining sample. As can be seen, the estimated coefficients are relatively stable and always significant at the 1% level, suggesting that our results are not due to potential outliers.

FIGURE 3: DOLS ESTIMATION WITH SINGLE COUNTRY EXCLUDED FROM THE SAMPLE



Next, we examine whether the negative long-run relationship between religiosity and per capita income is due to sample-selection bias. Sample-selection bias occurs when the selected sample is not random and thus not representative. A potential problem with our sample could be that it consists of relatively rich countries and that, accordingly, the estimated relationship does not apply to poorer countries. Another concern is that the estimated relationship is dominated by highly religious or highly secular countries. We therefore re-estimate the DOLS regression for four subsamples: countries with incomes above the sample average, countries with incomes below the sample average, countries with church attendance rates above the sample average, and countries with church attendance rates below the sample average.

The resulting coefficients are listed in Table 5. Regardless of which subsample is used, the null hypothesis of no cross sectional dependence cannot be rejected (as the CD statistics show), and the long-run relationship between religiosity and income is negative and significant at the

1% level. From this, it can be concluded that the negative coefficient on $\log(y_{it})$ is not due to sample-selection bias. Clearly, it would be desirable to also assess whether there are significant differences in the effects of income on religiosity between low-income and high-income countries or between countries with high levels of church attendance and those with low levels of church attendance. In particular, it appears that the magnitude of the negative income effect on religiosity is somewhat larger when the sample is restricted to countries with high church attendance rates. However, the small sample sizes do not allow statistically meaningful comparisons in this regard.

TABLE 5. DOLS Estimates for Subsamples

Estimation method	Coeff. on $\log(y_{it})$	CD stat.	No. countries	Obs.
(1) Countries with incomes above the sample average	-9.108*** (-7.90)	0.19	10	100
(2) Countries with incomes below the sample average	-8.819*** (-8.45)	0.32	7	70
(3) Countries with church attendance rates above the sample average	-10.588*** (-8.66)	1.49	7	70
(4) Countries with church attendance rates below the sample average	-7.881*** (-8.45)	-1.55	10	100

The dependent variable is $CHURCH_{it}$. The DOLS regressions were estimated with one lead and one lag ($k = 1$), given the relatively small number of time series observations. t -statistics are in parenthesis. The CD test statistic is normally distributed under the null hypothesis of no cross-sectional dependence. *** (**) [*] indicate significance at the 1% (5%) [10%] level.

Finally, we examine whether the results are robust to alternative measures of religiosity and alternative specifications of the income variable. Franck and Iannaccone (2009) measure religiosity not only by the parental rate but also by the church attendance rate of children. In the study by Lipford and Tollison (2003), income enters in non-logarithmic form. Table 6 presents the results of the DOLS regressions using these two different variables, labeled $CHURCH_{children_{it}}$ and y_{it} , both separately and jointly. As can be seen, all coefficients are negative and statistically significant. Summarizing, the negative effect of income on religiosity is robust to different estimation techniques, potential outliers, sample selection, different measures of church attendance, and alternative specifications of the income variable.

3.3. Long-run Causality. The above interpretation of the estimation results is based on the assumption that long-run causality runs from income to religiosity. However, while cointegration

TABLE 6. DOLS Estimates Using Different Measures of Religiosity and Alternative Specifications of the Income Variable

	Coefficient of the income variable	CD stat.	Obs.
(1) Regressand: $CHURCHchildren_{it}$ Regressor: $\log(y_{it})$	-13.063*** (-9.67)	0.34	170
(2) Regressand: $CHURCH_{it}$ Regressor: y_{it}	-0.001*** (-7.60)	0.72	170
(3) Regressand: $CHURCHchildren_{it}$ Regressor: y_{it}	-0.002*** (-10.59)	-1.15	170

The dependent variable is $CHURCH_{it}$. The DOLS regressions were estimated with one lead and one lag ($k = 1$), given the relatively small number of time series observations. t -statistics are in parenthesis. The CD test statistic is normally distributed under the null hypothesis of no cross-sectional dependence. *** (**) [*] indicate significance at the 1% (5%) [10%] level.

implies causality in at least one direction, it says nothing about the direction of the causal relationship between the variables, as discussed above. Causality may run in either direction, from income to religiosity or from religiosity to income, or in both directions. To test the direction of long-run causality, we follow common practice in the applied panel cointegration literature and employ a two-step procedure. In the first step, we use the (DOLS) estimate of the long-run relationship (from the first row of Table 4) to construct the disequilibrium term

$$ec_{it} = CHURCH_{it} - [\hat{a}_i - 8.944 \log(y_{it})]. \quad (6)$$

In the second step, we estimate the error correction model

$$\Delta CHURCH_{it} = c_{1i} + a_1 ec_{it-1} + \sum_{j=1}^k \varphi_{11,j} \Delta CHURCH_{it-j} + \sum_{j=1}^k \varphi_{12,j} \Delta \log(y_{it-j}) + e_{it}^{CHURCH} \quad (7a)$$

$$\Delta \log(y_{it}) = c_{2i} + a_2 ec_{it-1} + \sum_{j=1}^k \varphi_{11,j} \Delta CHURCH_{it-j} + \sum_{j=1}^k \varphi_{12,j} \Delta \log(y_{it-j}) + e_{it}^{\log(y)} \quad (7b)$$

The error-correction term, ec_{it-1} , represents the error in, or deviation from, the equilibrium, and the adjustment coefficients a_1 and a_2 capture how $CHURCH_{it}$ and $\log(y_{it})$ respond to deviations from the equilibrium. From the Granger representation theorem (Engle and Granger, 1987) it follows that at least one of the adjustment coefficients must be nonzero if a long-run relationship between the variables is to hold. A statistically significant error correction

term also implies long-run Granger causality from the explanatory variables to the dependent variables (Granger, 1988), and thus that the dependent variables are endogenous in the long run. An insignificant error correction term implies long-run Granger non-causality, and thus that explanatory variables are weakly exogenous (Hall and Milne, 1994). Given that all variables in the model, including ec_{it-1} , are stationary (because the level variables are cointegrated), a conventional likelihood ratio chi-square test can be used to test the null hypothesis of weak exogeneity, $H_0 : a_{1,2} = 0$.

TABLE 7. Tests for Long-run Causality / Weak Exogeneity

Weak exogeneity of	$\chi^2(1)$ (p values)	CD stat.	Demeaned	Obs.
(1) $CHURCH_{it}$ Dependent variable: $\Delta CHURCH_{it}$	49.080 0.000	0.71	No	187
(2) $\log(y_{it})$ Dependent variable: $\Delta \log(y_{it})$	14.213 0.000	7.51***	No	187
(3) $\log(y_{it})$ Dependent variable: $\Delta \log(y_{it})$	7.823 0.006	-1.41	Yes	187

The number of degrees of freedom ν in the standard $\chi^2(\nu)$ tests correspond to the number of zero restrictions. Given the relatively small number of time series observations, the lag length was set to $k = 1$. The CD test statistic is normally distributed under the null hypothesis of no cross-sectional dependence. *** indicate a rejection of the null of no cross-sectional dependence at the 1% level.

Table 7 reports the results. As can be seen from row 1, the null hypothesis of weak exogeneity is rejected for $CHURCH_{it}$ at the 1% level, and the CD test suggests that this inference is not affected by cross-sectional dependence. Since the null hypothesis of no cross-sectional dependence is decisively rejected in row 2 for the residuals from the $\Delta \log(y_{it})$ equation, row 3 uses the demeaned data to account for the cross-sectional dependence through common time effects. Row 3 indicates that we can also reject the null hypothesis of weak exogeneity of $\log(y_{it})$. From this it can be concluded that the statistical long-run causality is bidirectional, implying that declining religiosity is both a consequence and cause of economic growth.

To check the robustness of this conclusion, we perform a standard panel Granger causality test based on a levels VAR regression with fixed effects. While it is well known from the time series literature that in general the asymptotic distributions of the Wald (or likelihood ratio) test for Granger causality in levels VARs with integrated variables are nonstandard (see, e.g., Toda and Phillips, 1993), Lütkepohl and Reimers (1992) show that the conventional Wald test

for a bivariate cointegrated VAR model is asymptotically distributed as chi-square and therefore valid as a test for Granger causality.

We report the p -values of the Granger causality chi-square statistics using one and two lags, respectively, of each variable in Table 8. The number of lags was determined by the Schwarz criterion (with a maximum of two lags), as is common practice in testing for Granger causality in a VAR model. As can be seen from the first row, the null hypothesis of no Granger causality from $\log(y_{it})$ to $CHURCH_{it}$ is decisively rejected, and the sum of the coefficients on lagged income is negative in the church attendance equation. This confirms our result that increasing income leads to declining religiosity.

Rows 2 and 3 of Table 8 show that the Granger causality test rejects the null hypothesis that one lag of $CHURCH_{it}$ does not help predict $\log(y_{it})$ at the 1% level and that the coefficient on lagged church attendance is also negative. While the CD test rejects the null hypothesis of no cross sectional dependence in row 2, the results in row 3, using the demeaned data, do not appear to suffer from cross sectional dependence. These results support the conclusion that increasing religiosity leads to declining income.

TABLE 8. Panel VAR Causality Tests

Null hypothesis	Lags	p -value	Sum of coeff. of causal variable	CD stat.	Demeaned	Obs.
(1) $\log(y_{it})$ does not cause $CHURCH_{it}$ Dependent variable: $CHURCH_{it}$	2	0.000	-2.722	1.12	No	187
(2) $CHURCH_{it}$ does not cause $\log(y_{it})$ Dependent variable: $\log(y_{it})$	1	0.000	-0.015	8.81***	No	204
(3) $CHURCH_{it}$ does not cause $\log(y_{it})$ Dependent variable: $\log(y_{it})$	1	0.006	-0.008	-1.54	Yes	204

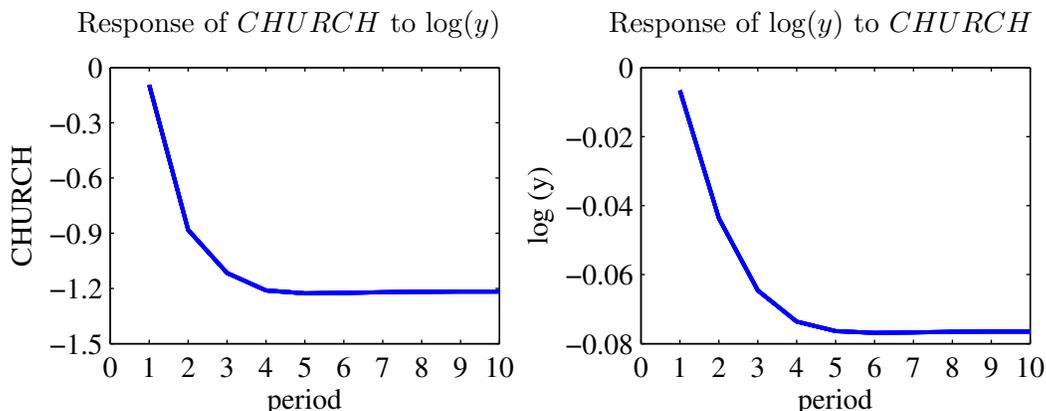
This table reports the p -values of Granger-causality VAR tests. The null hypothesis is that one or two lags of the (demeaned) series of $\log(y_{it})$ ($CHURCH_{it}$) do not help predict the series of $CHURCH_{it}$ ($\log(y_{it})$). The number of lags was determined by the Schwarz criterion with a maximum of two lags. The CD test statistic is normally distributed under the null hypothesis of no cross-sectional dependence. *** indicate a rejection of the null of no cross-sectional dependence at the 1% level.

Finally, Figure 4 presents generalized impulse response functions based on a one-lag panel vector error correction model using the raw data over a 50-year horizon (10 five-year periods).⁴ As can be seen in the left panel of Figure 4, a one-standard-deviation shock in income results in a gradual and permanent decrease in church attendance and reaches its full impact after 20

⁴The demeaned data produce qualitatively similar results.

years (4 periods). The right panel shows that income gradually and permanently decreases in response to a one-standard-deviation innovation in church attendance and that the full impact is reached after 25 years (5 periods). The impulse response functions are thus consistent with the Granger causality tests reported above.

FIGURE 4: IMPULSE-RESPONSES



4. CONCLUSION

In this paper we took into account that secularization is a usually understood as a dynamic process and investigated the time series properties of church attendance rates and income per capita. We showed that there exists a cointegration relationship between both variables and estimated, in our preferred specification, that a one percent increase in the log of income per capita is associated with a decline in church attendance by about 9 percentage points. A further advantage of cointegration analysis is that results are not biased by omitted variables. This means in our context that results are not biased by the missing consideration of measures of religious supply or other drivers of the demand for religion. This an important feature with respect to the demand–supply debate in the theory of religion.

We have documented that our result on the income–religiosity nexus is robust to alternative estimation methods, potential outliers, sample selection, different measures of church attendance, and alternative specifications of the income variable. We found that long-run causality runs in both directions, higher income leads to declining religiosity and declining religiosity leads to higher income. Secularization appears to be both cause and consequence of economic development.

One limitation of our analysis is the relatively small sample of countries, consisting mostly of developed Western countries primarily inhabited by (former) Christians. Whether and how the results are transferable to developing countries and countries inhabited predominantly by non-Christians is an intriguing question, which we are looking forward to address when time series data on religiosity will be available for these countries as well.

Eventually, with respect to future developments, the strong linear relationship between church attendance and the log of income has to become non-linear if income continues to grow because church attendance is bounded from below. Our analysis, comprising most of the 20th century, however, has shown that a fading trend of secularization is not yet visible in the data.

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