JOURNAL OF GEOPHYSICAL RESEARCH, VOL. 106, NO. A8, PAGES 15,635–15,641, AUGUST 1, 2001

The cause-and-effect relationship of solar cycle length and the Northern Hemisphere air surface temperature

Richard Reichel
Department of Economics, Friedrich-Alexander-Universität, Nürnberg, Germany

Peter Thejll and Knud Lassen
Solar-Terrestrial Physics Division, Danish Meteorological Institute, Copenhagen, Denmark

Abstract. It has previously been demonstrated that the mean land air temperature of the Northern Hemisphere could adequately be associated with a long-term variation of solar activity as given by the length of the approximately 11-year solar cycle. In this paper it is shown that the right cause-and-effect ordering, in the sense of Granger causality, is present between the smoothed solar cycle length and the cycle mean of Northern Hemisphere land air temperature for the twentieth century, at the 99% significance level. This indicates the existence of a physical mechanism linking solar activity to climate variations.

1. Introduction

A significant correlation has been reported between the long-term (decadal-scale) variation of the Northern Hemisphere land (NHL) air temperature and the long-term variation of solar activity during the instrumental era from 1861 to 1989 [*Friis-Christensen and Lassen,* 1991] (henceforth FCL91). The close correlation was established using the smoothed solar cycle length (SCL) as an index of long-term variability of the Sun, and it was concluded that this parameter appears to be an indicator of long-term changes in the total energy output of the Sun, which in turn was mainly responsible for the long-term variations of the temperature through the period studied. Subsequently, it has been demonstrated with paleoreconstructions of temperature [*Friis-Christensen and Lassen,* 1992; *Hoyt and Schatten,* 1993; *Hameed and Gong,* 1994; *Lassen and Friis-Christensen,* 1995; *Butler and Johnston,* 1996; *Zhou and Butler,* 1998] that this correlation between SCL and temperature probably has been present for centuries. Before SCL was discussed as a proxy of solar activity, work on the Sun-climate link tended to focus on general aspects (e.g., *Eddy* [1976], who commented on the relationship between the "little ice age" and the Maunder minimum in solar activity), or used the raw sunspot numbers [*Reid,* 1987].

There is interest in the SCL-climate link in the Sun-climate research area; not only have FCL91 and *Lassen and Friis-Christensen* [1995] (henceforth LFC95), and the authors listed above used SCL to compare to climate series, but so have *Mende et al.* [1993] and *Mursula and Ulich* [1998]. However, the interest in SCL in not restricted to the Sun-climate research area; there is interest in SCL in the space physics and astrophysical communities because it may be a broad representation of important aspects of solar activity. *Hoyt and Schatten* [1993] used SCL as part of a solar irradiance reconstruction on the basis of *Granger* [1957], who showed that mean solar activity and sunspot cycle length were inversely proportional, and it has been suggested that SCL and the solar rotation rate are linked [*Mendoza,* 1999] so that records of SCL potentially can be used to determine the solar rotation rate before this was specifically observed. *Soon et al.* [1996] used SCL, among other measures of solar activity, to turn the Sun-climate issue on its head and use terrestrial temperatures to infer past solar irradiance levels.

A significant correlation between two physical parameters is only one of several other necessary requirements for establishing a physical relationship between parameters. An important additional requirement is the identification of a physical mechanism acting between the two parameters.

1.1. Some Possible Physical Mechanisms

At the moment the hypothesis of the Sun being the cause of some variations in the climate is not at the stage where a well established and plausible physical mechanism is agreed upon. While suggestions abound in the literature (see such reviews as *Herman and Goldberg* [1978], *Pittock* [1983], *Pudovkin and Raspopov* [1992] (henceforth PR92), and *Hoyt and Schatten* [1997]), only a few appear testable at the moment. One is the suggestion that the solar signal is imprinted on climate through the variations in the solar luminosity: This

Paper number 2001JA900027.
0148-0227/01/2001JA900027$09.00

suggestion is testable in models with fully known methods, namely, by turning the "solar constant" up and down during the model calculation (e.g., *Lean and Rind* [1999]). Model investigations of the effects of luminosity variations tend to show that only a fraction of the observed climate variations could be caused by direct luminosity changes. These results depend on reconstructions of solar luminosity into the past, as only about 20 years of high-quality data on solar luminosity changes are available [*Pap et al.,* 1994]. Because the irradiance reconstructions carry their own arbitrariness, the modeling efforts in this area cannot yet be said to be conclusive. Another testable suggestion is that the forcing through UV and stratospheric ozone changes, induce changes in the stratosphere-troposphere circulation. This idea is currently the subject of development (e.g., *Haigh* [1999]). A complete model of the joint impacts of irradiance changes (visible and UV light), the consequences for ozone of the UV changes, and the dynamical tropospheric response thereto is not yet at hand, but could be of great importance to the Sun-climate field when it is.

Other suggestions of effects on the climate by the Sun (such as modification of cloud cover by particles from the Sun, or galactic cosmic rays [*Svensmark and Friis-Christensen,* 1997] (henceforth SFC97), the flux of which is controlled by the Sun) are not implementable at the moment at some levels of detail, as they either require knowledge of physics not yet developed; such as the microphysics of cloud formation, or require such spatially detailed modeling that present computing power is insufficient. For instance, the detailed three-dimensional modeling of all cloud dynamics and radiative transfer in order to get the clouds' feedback reaction to changes realistically modeled is simply impossible at the moment.

The effects of greenhouse gases on the climate are fairly straightforward to model in comparison. This is mainly because there is less dynamics to model and because the microphysics is easier to describe due to the much more uniform distribution of the well-mixed greenhouse gases. Water vapor is an exception to this rule. The effects of increased amounts of greenhouse gases is also more or less linear, which is not the case for water vapor and clouds.

A common problem for all models of climate is the correct modeling of feedback effects whereby the climate system adjusts to imposed changes. Water vapor is the most important greenhouse gas, and the amount and heterogeneous distribution of water vapor is controlled by the state and dynamics of the climate system. Amount and type of cloud cover are important for the radiative balance, and thus the correct modeling of what the water vapor is doing in a climatic system is a central problem to these issues. At the moment code parameterizations of many such effects take the place of detailed descriptions.

While we wait for breakthroughs in the microphysical understanding of some potentially important processes that influence the radiative forcing of the climate, or wait for faster tools with which to model the existing physical understanding, there is one additional statistical analysis step that can be taken to analyze some of the presently known correlations between climate variations and external factors. Techniques exist for establishing the direction of causality and have previously been used primarily in such fields as econometrics [*Granger,* 1969, 1988; *Granger and Newbold,* 1986], and only lately, with the growing debate over climate change, have such methods been applied in geophysics (e.g., *Andronova and Schlesinger* [1991]).

We will apply a method developed for causality direction testing in econometrics to the problem of solar forcing of temperatures. While the establishment of a possible cause-and-effect ordering of time series does not in itself prove a physical relationship, the opposite is the case: Failure to establish the proper ordering would be damning for the theory that variations in the Sun cause climate variations. Should there be a physical mechanism at work, it is likely (though not certain) that the cause-and-effect ordering will appear in the data.

1.2. Introducing Granger Causality Testing

On the basis of earlier research by *Wiener* [1956], the concept of Granger causality was introduced in 1969 in order to uncover the causality (cause and effect) relationship between two variables (time series) by exploiting lead-lag structures [*Granger,* 1969]. Although Granger's concept is not equivalent to causality in the strict sense of the word, it is a helpful tool to test whether a lead-lag structure conforming to a theory is present in the data. As already stated, Granger causality can be considered a first order (necessary) condition for real (physical) causality [*Granger,* 1980].

Consider two time series X_t and Y_t (t is the time index) and the following bivariate autoregressive model:

$$X_t \quad = \quad c + \sum_{i=1}^{n} \alpha_i X_{t-i} + \sum_{j=1}^{m} \beta_j Y_{t-j} + u_t, \qquad (1)$$

$$Y_t \quad = \quad d + \sum_{k=1}^{p} \gamma_k Y_{t-k} + \sum_{l=1}^{q} \delta_l Y_{t-l} + v_t, \qquad (2)$$

where c and d are constants and i, j, k and l denote the number of lags. The terms u_t and v_t are white noise errors. This unrestricted model regresses the present values of X on the past values of X and Y (equation 1) and the present values of Y on the past values of Y and X (equation 2). On the other hand, a complementary (restricted) version of the model regresses the current values of X and Y on their own past values alone.

The unrestricted and the restricted model are now compared with respect to their predictive power. If the estimated coefficients on the lagged values of $Y(\beta_j)$ in equation 1 are statistically significant as a group, it is concluded that Y causally influences X in the sense of Granger as changes in Y precede changes in X $(Y \Rightarrow X)$. An analogous comparison can be made between the unrestricted and the restricted versions of equation 2. Should the estimated coefficients on the lagged values of X in equation 2 exhibit statistical significance, X is said to Granger cause Y $(X \Rightarrow Y)$. If in general a causality direction $Y \Rightarrow X$ or $X \Rightarrow Y$ is indicated, "one-way" or "unidirectional" causality exists in Granger's terminology. Bidirectional causality runs between X and Y if both causality directions are not rejected simultaneously (i.e., $Y \Rightarrow X$ and $X \Rightarrow Y$). In this case some sort of feedback mechanism seems to be operative [*Granger,* 1969]. Finally, if neither $X \Rightarrow Y$ nor $Y \Rightarrow X$ is indicated, there is no Granger causality, even in the presence of a significant correlation between X and Y.

In order to test the significance of the estimated coefficients on the lagged Y values (equation 1) respectively the lagged X values (equation 2), the following F test is used

$$F = \frac{(RSS_r - RSS_u)/(df_r - df_u)}{RSS_u/df_u}, \tag{3}$$

where RSS_r and RSS_u are the sum of the squared residuals of restricted and unrestricted models, respectively and df_r, df_u are, respectively, the degrees of freedom in the restricted and unrestricted models. If the F test (given a prespecified significance level) exceeds a critical value, the underlying null hypotheses $(X \not\Rightarrow Y$, $Y \not\Rightarrow X)$ have to be rejected, thus confirming unidirectional causality [*Granger and Newbold,* 1986]. The test equations are estimated using the ordinary least squares method. As the absolute coefficient values $(\alpha_i, \beta_j, \gamma_k$, and $\delta_l)$ are of less importance for the interpretation of causality test results, they are not reported but are available from the authors. Determining the appropriate lag length, two alternative strategies can be applied. The first is an "ad hoc" method where the lags are prespecified. In this case the sequence of test equations should be stopped when (the coefficients of) longer lags become insignificant. However, as causality test results may be theoretically sensitive to the number of lags, a statistical criterion can be employed to determine the optimal lag length. Akaike's final prediction error (FPE) criterion is a useful technique (see *Jin and Yu* [1995] for an explanation and empirical application). The FPE test statistic is given as

$$FPE(i,j) = \frac{T+i+j+1}{T-i-j-1}\frac{RSS(i,j)}{T}, \tag{4}$$

where T is the number of observations. As longer lags (i,j) increase the first term but decrease the second term, the smallest FPE value determines the optimal lag length. Considering the time series used in the following section, the FPE test indicates a lag length of 1 or 2 in the vast majority of cases. For the purpose of comparison as well as to check the stability of the test, we report test results with alternative lags 1, 1-2, and 1-2-3. This strategy is useful to uncover whether the inherent limitations of the Granger test (due to a potentially insufficient number of observations or an inadequate data sampling frequency) which could produce fragile or theoretically unreasonable results are of importance.

2. Is There Granger Causality Between Solar Activity and Terrestrial Temperatures?

In order to test for the underlying cause-and-effect relationship we first formulate the following null hypotheses (here, scl denotes solar cycle length; and temp denotes temperature): Hypothesis H_1 is that solar activity does not Granger-cause terrestrial temperatures (scl$\not\Rightarrow$temp), and hypothesis H_2 is that terrestrial temperatures do not Granger-cause solar activity (temp$\not\Rightarrow$scl). Rejection of H_1 and non-rejection of H_2 would be required to support the interpretation of the observed correlations in the FCL91 and LFC95 studies.

We shall use three different temperature proxies. The first (temp1) is the Northern Hemisphere land temperature series which is available from 1851 to 1998. The second is the combined land and sea series (temp2) available from 1857 to 1998 (both series from P. D. Jones, private communication 2000; previous series discussed by *Jones et al.* [1986], and *Jones* [1988,1994]; see also *Jones et al.* [1999]). (The standard deviation (s.d. = 0.23) of the latter series is significantly lower than the standard deviation of the former (s.d.=0.32) because the variability of marine temperatures is lower than over land areas.) In order to broaden our analysis, we also tested for a potential causal link between the solar cycle length and the Southern Hemisphere temperature (temp3) series (P. D. Jones, private communication, 2000).

The length of the solar cycle was estimated by spectral analysis of sunspot counts of different data frequencies. For a first proxy series (scl1) the wavelength (cycle length) is determined from the power spectrum of daily sunspot counts using a sliding 61-year window. The middle year of the sliding window was considered the center year. This series starts in 1857 and ends in 1967. As a variant, a second series (scl2) was calculated using the same 61-year sliding window but was based on monthly sunspot observations. In order to check whether our results depend on different methods of calculation or a varying length of the time series, scl2 starts in 1851 and ends in 1968, thus exploiting the sunspot counts in 1999. Finally, an annualized ver-

Table 1. Granger Causality Test F-Values Results.

H_0	Lag 1	Lag 1,2	Lag 1,2,3	Conclusion
scl1 $\not\Rightarrow$ temp1	40.43(99%)	14.72(99%)	10.87(99%)	scl1 $\Rightarrow$ temp1
temp1 $\not\Rightarrow$ scl1	0.38	4.89(99%)	1.34	unclear
scl1 $\not\Rightarrow$ temp2	11.82(99%)	6.28(99%)	3.20(95%)	scl1 $\Rightarrow$ temp2
temp2 $\not\Rightarrow$ scl1	0.00	1.02	1.34	no causality
scl1 $\not\Rightarrow$ temp3	8.07(99%)	3.61(95%)	1.97	scl1 $\Rightarrow$ temp3
temp3 $\not\Rightarrow$ scl1	0.07	0.52	0.65	no causality
scl2 $\not\Rightarrow$ temp1	36.68(99%)	11.54(99%)	11.08(99%)	scl2 $\Rightarrow$ temp1
temp1 $\not\Rightarrow$ scl2	0.15	0.10	0.41	no causality
scl2 $\not\Rightarrow$ temp2	8.36(99%)	6.08(99%)	4.11(99%)	scl2 $\Rightarrow$ temp2
temp2 $\not\Rightarrow$ scl2	0.11	0.37	0.34	no causality
scl2 $\not\Rightarrow$ temp3	6.67(95%)	4.93(99%)	3.27(95%)	scl2 $\Rightarrow$ temp3
temp3 $\not\Rightarrow$ scl2	0.11	0.37	0.34	no causality
scl3 $\not\Rightarrow$ temp1	24.24(99%)	7.39(99%)	5.67(99%)	scl3 $\Rightarrow$ temp1
temp1 $\not\Rightarrow$ scl3	0.02	1.02	0.86	no causality
scl3 $\not\Rightarrow$ temp2	8.04(99%)	5.77(99%)	2.98(99%)	scl3 $\Rightarrow$ temp2
temp2 $\not\Rightarrow$ scl3	1.26	1.83	1.06	no causality
scl3 $\not\Rightarrow$ temp3	4.03(95%)	2.66	0.85	weak causality
temp3 $\not\Rightarrow$ scl3	0.00	2.57	1.53	no causality

scl1 (from daily sunspots), scl2 (from monthly mean sunspot numbers), scl3 (from *Lassen and Friis-Christensen,* [1995]), temp1 (Northern Hemisphere land air temperatures), temp2 (land and sea temperatures), temp3 (Southern Hemisphere temperature). H_0 is the statistical hypothesis being tested. The symbols $\Rightarrow$ and $\not\Rightarrow$ stand for the presence and absence of Granger causality in the given direction, respectively. Values in parenthesis indicate the confidence levels.

sion of the 1-2-1 filtered series introduced in LFC95 (their Figure 1) was used (scl3). Since Granger tests require equally spaced values and the original values are unequally spaced, yearly values were calculated by linear interpolation. Although this is a rather crude method, correlation analysis shows a close comovement of the three series. The lowest correlation coefficient is $R_{scl2;scl3} = 0.928$, the highest is $R_{scl1;scl3} = 0.945$.

Table 1 reports the test results (F values) for time lags $i = j = k = l = 1, 2, 3$ and all possible combinations of scl and temp series. (As the critical F values vary with the number of observations and lags, the entries in Table 1 indicate significance at the 99% and 95% level.)

In the large majority of combinations, the null hypothesis of no causality scl $\not\Rightarrow$ temp is clearly rejected, whereas (with an exception reported in Table 1, line 2) the reverse causality direction null cannot be rejected. Thus there is strong evidence for unidirectional Granger causality scl $\Rightarrow$ temp. However, the causality evidence for the Southern Hemisphere series is considerably weaker. This result can be attributed to the poorer data quality due to less coverage as well as to the dampening effects of the sea. On the other hand, with a single exception (Table 1, first two lines)) the empirical evidence indicates no causality running from temperatures to the solar cycle length, which in fact would be theoretically unreasonable. This result supports the different attempts to look for a mechanism relating solar activity and climate. Moreover, it does not depend on the choice of a particular scl series. The lower F statistics which can be generally observed when the temp2 and temp3 series are used indicate that the sensitivity of the combined land and sea temperatures with respect to solar activity changes is considerably weaker. This result could have been expected from a theoretical viewpoint and confirms the reliability of the test. Finally, the test results show that the observed causality pattern is not affected by the different sample sizes and time periods under consideration.

2.1. Confidence Levels

How confident are we that the results of the Granger causality test are unusual? Are we sure that the results could not be the result of a random occurrence?

To answer these questions, we estimate the significance level of the result using a common technique that is based on repeated use of the causality test on a sufficient number of artificially generated time series. If these series are generated in a suitable manner, it is possible to use the occurrence rate of a positive Granger causality result as an estimate of the significance level. If many of the artificial time series give the same result as the real data we cannot be sure that we have an unusual result, while if few of the artificial series give the observed result we can be confident that we have a statistically significant result.

The proper sense of the above words "generated in a suitable manner" must be discussed. At the heart of the problem of significance level estimation lies the correct understanding of the nature of the time series at hand, and in particular the serial correlation in them, that is, the degree to which adjacent points carry the same information. Series in which there is no information available on what the value of the next or previous point is are called "white noise" series. These have, on average, zero autocorrelation at lag 1. Series where there is such information present can be characterized by the structure of their autocorrelation functions. Several strategies exist for the calculation of such simulated "red" series, given information about the autocorrelation. We shall use the assumption that the series can be generated autoregressively on the basis of the autocorrelation, α_1, of the real series at lag 1. In practice the series (x) are generated by iterating the following sequence:

$$x_{i+1} = \alpha_1 x_i + \eta_i \tag{5}$$

where η are random numbers drawn from a white noise random number generator and $x_{i=0}$ is some arbitrary starting value. To avoid spin-up problems, we sample the series only for large values of i; that is, we generate a very long sequence of numbers before we select segments of the required length. This method of estimating levels of significance has been used by, for example, *Mann et al.* [1998] recently. We furthermore require that the lag 1 autocorrelation of the segment chosen is in fact close to the specified value α_1 and discard any segments that do not meet this requirement to a tolerance of less than 0.025.

As the basic question of significance in the present problem is whether the causality test results are significant when the SCL curve is used as the predictor and the temperature as the predictand, we generate random series in the above sense of the SCL curve, and repeatedly apply the test using these simulated SCL series and the real temperature. We perform 100 trials.

The result was that in only one simulated case did the Granger causality method accept that a "Sun to temperature" causality was indicated. This happened when the method's internal rejection level was set at 1%, so the results of the internal significance level test and the exterior Monte Carlo method are consistent. We therefore conclude, from the results of the significance level testing, that the method is robust and has found a significant result linking the SCL curve causally, in the Granger sense, to the temperature.

2.2. Solar Irradiance and Terrestrial Temperatures

In the analysis conducted above, overall solar activity was approximated by the varying length of the 11-year solar cycle. There are, however, alternative measures of solar activity which could also be used for causality testing, such as the solar-activity-modulated flux of galactic cosmic rays and the 10.7-cm radio flux. Nevertheless, it is hardly possible to use these as they are all too short for the Granger procedure. A similar problem will occur with measured solar irradiance, as only about 20 years of data are at hand. Therefore we decided to base a sequence of additional Granger tests on a reconstruction of solar irradiance by *Hoyt and Schatten* [1993] (hssi; updates for the years after 1992 are available over the Internet at the URL http://users.erols.com/dhoyt1/jgr1993.txt). Hoyt and Schatten have produced a composite solar irradiance model based on three indicators of solar activity; length of the solar cycle, the decay rate of the solar cycle and the average level of solar activity. These indices allow the solar irradiance to be modeled back to the mid-1700s, though for the purpose of comparison we shall use only data from the mid nineteenth century. Solar activity might influence the Earth's climate in many ways, one of which is by variations in incident radiation, so we apply the Granger test to the hemispheric temperatures and the Hoyt and Schatten irradiance model to see if the expected causality direction is present. As discussed in section 2, the test procedure is applied against our three temperature series (time period 1851-1997). Table 2 reports the results.

These results are fully in line with those previously obtained with the solar cycle length as a solar activity proxy. In every case, causality running from solar irradiance to terrestrial temperatures is indicated. The reverse causality direction never occurs. As already shown in Table 1, Granger causality becomes slightly weaker when temperature series which implicitly give greater weight to sea data are used. In order to check whether our results may be artificial due to potential uncertainties of the HS irradiance reconstruction, we additionally tested for causality using the *Lean et al.* [1995] solar irradiance reconstruction. The results were similar, indicating the same causality direction. Therefore we do not report detailed test statistics. All in all, the tests performed on irradiance data fully confirm the results obtained from our three SCL series, thus supporting the reliability and robustness of the test procedure applied.

3. Summary and Conclusion

In this paper we have analyzed causality issues in the question of whether solar forcing, as parameterized by the solar cycle length (or, alternatively, by solar irradiance proxies), causes temperature variations, in the Northern Hemisphere and globally during the twentieth century. The method applied to the test of causality is a statistical method called "Granger causality testing", which is based on statistical analysis of correlation when lags on the predictor are introduced in regressive relationships. Asymmetry of the correlation, under lags of opposite sign, indicates the potential presence of the required physical relationship from the predictor to the predictand, although it does not prove its existence.

The method allows researchers to perform a basic check of those results that are primarily based on cor-

Table 2. Granger Causality Test *F*-values Results - Using Hoyt and Schatten Irradiance Data.

H_0	Lag 1	Lag 1,2	Lag 1,2,3	Conclusion
hssi $\nRightarrow$ temp1	42.83(99%)	11.35(99%)	6.76(99%)	hssi ⇒ temp1
temp1 $\nRightarrow$ hssi	0.13	0.61	1.24	no causality
hssi $\nRightarrow$ temp2	12.37(99%)	4.80(99%)	2.34	hssi ⇒ temp2
temp2 $\nRightarrow$ hssi	0.00	0.66	0.88	no causality
hssi $\nRightarrow$ temp3	9.69(99%)	3.79(95%)	1.61	hssi ⇒ temp3
temp3 $\nRightarrow$ hssi	0.04	1.57	1.82	no causality

hssi (solar irradiance reconstruction from [*Hoyt and Schatten,* 1993]), temp (land air temperatures, see Table 1). Values in parenthesis indicate the confidence levels.

relation for its claim of link between potential factors and their observed consequences, by providing knowledge of the direction of causality. When ample physical understanding of a phenomenon is present the test is not required, but when the suggestion of a link is based mainly on the appearance of a correlation, then the Granger causality test can have an impact on sorting out those correlations that do not even meet the minimum requirements from those that do, which can be a great help in seeking out significant observations for further study. The field of Sun-climate links is an example of this.

Perusal of the literature shows many examples of claimed Sun-climate or Sun-weather links that are not on a firm theoretical basis, but take their support mainly in reported studies of correlations. Some examples include the reported effects of passages of the Earth through the magnetic sector boundaries in the interplanetary magnetic field; there are reports of changes in air pressure during these passages, but no accepted model exists. Another example is the claimed effects of burst of solar particles on atmospheric transparency (PR92), or the effects of ionizing cosmic radiation on the formation of clouds [*Pudovkin and Veretenenko,* 1995; SFC97]. Rainfall in various places has been statistically linked to the solar cycle [*King,* 1974, 1975]. Additional examples can be found in the reviews of the field [*Pittock,* 1978, 1983; and HG78; PR92].

Suggestions with various degrees of detail exist for how some of these effects could work; common to them all is that an application of the Granger causality test could help sort out the best candidates for further study. The results of this paper show that there does exist the necessary causality direction for solar forcing (represented by solar cycle length variations) to cause climate variations. This should, however, not be misinterpreted in the sense that effects of other climate determinants (e.g., greenhouse gases) are downplayed. It is outside the range of statistical causality tests to disentangle the effects of different climate-forcing variables.

However, this is the first time this causality has been demonstrated in this problem, and underlines that it is potentially fruitful and important to continue seeking the physical mechanism for solar control of climate change in order that separation of the natural and anthropogenic climate changes can be made. If a detailed physical model could be built of the long-term solar influence on climate, it would be possible to better constrain the other parameterizations used in climate models. At the moment it is possible to "fit" model parameters to observations in many ways because the influences of several of the physical factors are similar. Reduction of the uncertainty in one arena will narrow the possible range for "free" model parameters, such as those used in parameterizations from cloud feedback and in the models used for aerosol cooling.

Acknowledgments. This work was supported by the Danish Climate Centre. We thank Phil Jones of the Climatic Research Unit at the University of East Anglia for providing land-only temperature series.

Michel Blanc thanks Marie-Lise Chanin and the other referees for their assistance in evaluating this paper.

References

Andronova, N. G., and M. E. Schlesinger, The application of cause-and-effect analysis to mathematical models of geophysical phenomena. 1. Formulation and sensitivity analysis, *J. Geophys. Res., 96,* 941-946, 1991.

Butler, C. J., and D. J. Johnston, A provisional long mean air temperature series for Armagh Observatory, *J. Atmos. Solar Terr. Phys., 58,* 1657-1672, 1996.

Eddy, J. A., The Maunder minimum, *Science, 192,* 1189–1202, 1976

Friis-Christensen, E., and K. Lassen, Length of the solar cycle: An indicator of solar activity closely associated with climate, *Science, 254,* 698-700, 1991.

Friis-Christensen, E., and K. Lassen, Global temperature variations and a possible association with solar activity variations, *Sci. Rep. 92-3,* Dan. Meteorol. Inst. Copenhagen, 1992.

Granger, C. W. J., A statistical model for sunspot activity, *Astrophys. J., 126,* 152-158, 1957.

Granger, C. W. J., Investigating causal relations by economic models and cross-spectral methods, *Econometrica, 37,* 424-438, 1969.

Granger, C. W. J., Testing for causality - a personal viewpoint, *J. Econ. Dyn. Control, 2,* 329-352, 1980.

Granger, C. W. J., Some recent developments in a concept of causality, *J. Econometr., 39,* 199-211, 1988.

Granger, C. W. J., and P. Newbold, *Forecasting Economic Time Series,* Academic, San Diego, Calif., 1986.

Haigh, J. D., Modeling the impact of solar variability on climate, *J. Atmos. Solar Terr. Phys., 61,* 63-72, 1999.

Hameed, S., and G. Gong, Variation of spring climate in lower-middle Yangtse river valley and its relation with solar-cycle length. *Geophys. Res. Lett., 21,* 2693-2696, 1994.

Herman, J. R., and R. A. Goldberg, Sun, weather and climate, *NASA Spec. Publ. SP-426*, 1978.

Hoyt, D. V., and K. H. Schatten, A discussion of plausible solar irradiance variations 1700-1992, *J. Geophys. Res., 98,* 18,895-18,906, 1993.

Hoyt, D. V., and K. H. Schatten, *The Role of the Sun in Climate Change*, Oxford Univ. Press, New York, 1997.

Jin, J. C., and E. S. H. Yu, The causal relationship between exports and income, *J. Econ. Dev., 20,* 131-140, 1995.

Jones, P. D., Hemispheric surface air temperature variations: Recent trends and an update to 1987, *J. Clim., 1,* 654-660, 1988

Jones, P. D., Hemispheric surface air temperature variations: A reanalysis and an update, *J. Clim., 7,* 1794-1802, 1994.

Jones, P. D., S. C. B. Raper, R. S. Bradley, H. F. Diaz, P. M. Kelly and T. M. L. Wigley, Northern hemisphere surface air temperature variations: 1851–1984, *J. Clim. Appl. Meteorol., 25,* 161-179, 1986.

Jones, P. D., M. New, D. E. Parker, S. Martin, and I. G. Rigor, Surface air temperature and its changes over the past 150 years, *Rev. Geophys., 37,* 173-199, 1999.

King, J. W., Weather and the earth's magnetic field, *Nature, 247,* 131–134, 1974.

King, J. W., Sun-weather relationships, *Astronaut. Aeronaut., 13,* 10–19, 1975.

Lassen, K., and E. Friis-Christensen, Variability of the solar cycle length during the past five centuries and the apparent association with terrestrial climate, *J. Atmos. Solar Terr. Phys., 57,* 835-845, 1995.

Lean, J., and D. Rind, Evaluating sun-climate relationships since the little ice age, *J. Atmos. Solar Terr. Phys., 61,* 25-36, 1999.

Lean, J., J. Beer, and R. S. Bradley, Reconstruction of solar irradiance since 1610: Implications for climate change, *Geophys. Res. Lett., 22,* 3195-3198, 1995.

Mann, M. E., R. S. Bradley, and M. K. Hughes, Global-scale temperature patterns and climate forcing over the past six centuries, *Nature, 392,* 779-787, 1998.

Mende, W. R., R. Stellmacher and M. Oldberg, Relationships between relative sunspot numbers and air temperature, *Meteorol. Z., 2,* 121–126, 1993.

Mendoza, B., Solar rotation and cycle length, *Solar Phys., 188,* 237-243, 1999.

Mursula, K. and T. Ulich, A new method to determine the solar cycle length, *Geophys. Res. Lett., 25,* 1837-1840, 1998.

Pap, J. M., R. C. Willson, C. Frölich, R. F. Donnelly and L. Puga, Long-term variations in solar irradiance, *Solar Phys., 152,* 13-21, 1994.

Pittock, A. B., A critical look at long-term Sun-weather relationships, *Rev. Geophys., 16*, 400-420, 1978.

Pittock, A. B., Solar variability, weather, and climate: an update, *Q. J. R. Meteorol. Soc., 109,* 23-55, 1983.

Pudovkin, M. I., and O. M. Raspopov, The mechanism of action of solar activity on the state of the lower atmosphere and meteorological parameters (a review), *Geomagn. Aeron., 32,* 593-608, 1992.

Pudovkin, M. S. and S. Veretenenko, Cloudiness decreases associated with Forbush-decreases of galactic cosmic rays, *J. Atmos. Solar Terr. Phys., 57,* 1349-1355, 1995.

Reid, G. C., Influence of solar variability on global sea surface temperatures, *Nature, 329,* 142–143, 1987.

Soon, W. H., E. S. Posmentier, and S. Baliunas, Inference of solar irradiance variability from terrestrial temperature changes, 1880-1993: An astrophysical application of the sun-climate connection, *Astrophys. J., 472,* 891-902, 1996.

Svensmark, H., and E. Friis-Christensen, Variation of cosmic ray flux and global cloud coverage-a missing link in solar-climate relationships, *J. Atmos. Solar Terr. Phys., 59,* 1225-1232, 1997.

Wiener, N., The theory of prediction, in *Modern Mathematics*, edited by E. F. Beckenback, 165-190, McGraw-Hill, New York, 1956.

Zhou, K., and C. J. Butler, A statistical study of the relationship between the solar cycle length and tree-ring index values, *J. Atmos. Solar Terr. Phys., 60,* 1711-1718, 1998.

K. Lassen and P. Thejll, Danish Meteorological Institute, Solar-Terrestrial Physics Division, Lyngbyvej 100, DK-2100 Copenhagen Ø, Denmark. (lassen@dmi.dk and pth@dmi.dk)

R. Reichel, Friedrich-Alexander-Universität, Lange Gasse 20, D-90403 Nürnberg, Germany. (richard.reichel@wiso.uni-erlangen.de)

(Received July 4, 2000; revised January 29, 2001; accepted February 4, 2001.)