The Recognition and Assessment of Decadal-Scale Variations of the Effects of Solar Variability on Climate

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The effect of variable solar activity (on the 11-year solar cycle time scale) on surface temperature at Toronto has been evaluated for three recent decades. For this data segment, it is highly probable that variable solar activity did influence climate and that the effects for the three decades were often significantly different. These inter-decade differences in solar effects were not random, but were organized in a coherent pattern in time. This pattern was substantially independent of season, climate parameter or statistical test, although there were major seasonal variations in the relative magnitude of modulated solar effects compared to unmodulated solar effects. The inter-decade variability had a maximum in spring and a minimum in autumn; on the other hand, the net effects of solar variability (ignoring interdecade variations) were weakest in spring and strongest in autumn.

1. Introduction

For over a century, the validity of reported effects of variable solar activity on weather and climate has been the subject of seemingly unresolvable controversy (PITTOCK, 1978, 1983). In recent years, however, the situation has changed dramatically. There is now clear and direct evidence of the effects of variable solar ultraviolet radiation on stratospheric ozone and temperature (BRASSEUR *et al.*, 1987; KEATING *et al.*, 1987). Extensive analyses and theoretical studies of the effects of the 11-year solar cycle on total ozone (the previous sentence refers to 14- and 28-day time scales), carried out by international statistical and modeling panels (WMO/UNEP, 1990), provide convincing proof that the marked ozone variability on this time scale (10–12 years) is primarily of solar origin. Furthermore, LABITZKE and VAN LOON (1988, 1989) have presented strong evidence that the effects of solar variability on climate, on the 11-year time scale, have frequently a significant modulation depending on the phase of the equatorial quasi-biennial oscillation (QBO).

It is not unreasonable to ask whether similar effects exist over much longer time scales, of the order of several decades (referred to here as decadal-scale). This would be consistent with the variations (including sign) of reported effects of solar variability on climate, carried out for different data segments at the same station (PITTOCK, 1978). General circulation characteristics vary on such periods, but not with the regularity of the QBO. Earlier studies by this author (GODSON, 1988) had suggested that major differences often occur in climatic effects of solar variability for successive 15-year data segments. Accordingly, a more elaborate analysis using 10-year data segments was designed, and is presented here. The theoretical basis for the statistical procedures and relationships will

be reported elsewhere. The inter-decadal variations are interpreted in terms of systematic change on a decadal-scale; the recent study by KNOX *et al.* (1988) suggests that a period of the order of 40 years could well be appropriate. It will be assumed that variations of atmospheric parameters on a time scale of 10–12 years can be ascribed to solar variations, since no other forcing mechanism on such a time scale has been suggested; the emphasis will, therefore, be on identifying variations of possible modulation, rather than on testing a specific null hypothesis.

2. Analysis Procedures

2.1 Preliminary analysis

The basic data used in this study consisted of monthly values of mean surface temperature and of standard deviation of surface temperature, for Toronto, Canada, for three recent decades (1951–60, 1961–70 and 1971–80). Seasonal cycles were removed by subtracting decade-means for each month, and the resulting values were grouped into four seasons—winter (Dec., Jan., Feb.), spring (Mar., Apr., May), summer (Jun., July., Aug.) and autumn (Sept., Oct., Nov.). The solar index employed was the monthly sunspot number, filtered using the 45-month running mean to remove high-frequency variance and concentrate on effects of the 11-year solar cycle.

A set of 12 lags was employed, at 6-month intervals from -36 to $+30$ months. For each lag, the data in each of the 24 data sets ($3 \text{ decades} \times 4 \text{ seasons} \times 2 \text{ climate parameters}$) were divided into two sub-groups—one for which the corresponding solar index lay above the median and one for which the corresponding solar index lay below the median. For each subgroup, the mean value and the variance were computed.

From the differences in mean values for each set, values of Students'- t were computed, and the corresponding values of degrees of freedom, in order to convert t -values to t_∞ -values, for an infinite number of degrees of freedom (i.e., the entry parameter for the error-function integral). The ratio of variances for each set is the statistical quantity, F , which can be converted to a probability using F -tables, giving a corresponding value of t_∞ from the error-function integral (or an equivalent analytical procedure can be employed, leading directly to a t_∞ -value).

Since means and variances of samples are uncorrelated (for Gaussian distributions), it is assumed that results of t -tests and F -tests are relatively independent, and that such tests on mean temperature and standard deviation of temperature are only weakly related, so that the total number of degrees of freedom of the 30-year record has been increased by a factor approaching four. The rigorous validity of this assumption is not critical, since the primary emphasis is on the characteristics (qualitative, or semi-quantitative) of variations of results (with season, parameter or test).

2.2 Unmodulated-model analysis

The preliminary analysis described above produced 16 sets of t_∞ values, denoted here by $S_{i\lambda}$, to indicate dependence on decade (i) and on lag (λ). For the first model applied to the data, it was assumed that the effect of variable solar activity on climate (as measured by $S_{i\lambda}$) was a sinusoidal function of lag, but did not vary with decade, i.e.,

$$S_{i\lambda} = S^* \cos \frac{2\pi}{\tau_f} (\lambda - \lambda_0) + \delta S, \quad (1)$$

where τ_f represents the effective period of the filtered solar index (taken here as 128.64 mo.) and λ_0 represents the optimum value of lag (for which the solar response is a maximum). There is no constant term in (1) since the true mean of all $S_{i\lambda}$ values over an entire cycle is zero and, of course, the population-average of t -values would be zero if there were no effects of variable solar activity on climate. However, it is necessary to adjust sample variances to take this consideration into effect, in performing a least-squares solution of (1) and in evaluating the standard error of S^* , $\sigma_E(S^*)$.

It is also necessary, in calculating $\sigma_E(S^*)$, to take into account the effects of serial correlation. Values of $S_{i\lambda}$ for successive lags are almost invariably positively correlated. This is to be expected since successive values of the cosine term in (1) are positively correlated (on the average) and this expected value of the serial correlation must be subtracted from the actual serial correlation to give the net serial correlation which, if still positive, reduces the number of degrees of freedom of the set. This adjusted value of degrees of freedom can then be employed in the evaluation of $\sigma_E(S^*)$ and in the conversion of the Students'- t value for S^* , $S^*/\sigma_E(S^*)$, to the corresponding value for an infinite number of degrees of freedom, which can then be employed with the error-function integral to give the (double-tailed) net chance probability associated with the unmodulated solar effect, P_1 . P_1 will be somewhat smaller than the true chance probability since the spectral power in the 10–12 yr. period range will be interpreted as an outcome of solar forcing, even for a totally random time series. Since this study is concerned with relative effects (and not absolute significance), it will employ the uncorrected P_1 values, for which the designation pseudo-chance probability is appropriate.

2.3 Modulated-model analysis

It will now be assumed that the effect of variable solar activity on climate does vary with decade, but that the optimum lag does not, so that the model equation analogous to (1) is

$$S_{i\lambda} = S_i^* \cos \frac{2\pi}{\tau_f} (\lambda - \lambda_0) + \delta S_i. \quad (2)$$

As before, constant terms do not appear in (2), so that the least-squares solution of (2) must involve terms appropriate to an integral number of lag cycles, requiring adjustment of variances in particular. These corrections are small if the number of lags times the lag interval is moderately close to τ_f or $\tau_f/2$ (this is true in the present study since 12×6 is reasonably close to 64). With these corrections, the values of S_i^* can be determined, and the corresponding values of their standard errors, $\sigma_E(S_i^*)$, again correcting for the effects of serial correlation of each decade set of $S_{i\lambda}$ values.

2.4 Specific decadal-scale modulation model

If the solar effects do vary significantly with decade, it may be assumed that they vary slowly with time (with faster variations which can be treated as random unresolvable noise). The strength and characteristics of the decadal-scale modulation are most meaningful if derived from a simple model, with equal periods of above-and below-average effects. The model to be used here adopts a sinusoidal variation of S , and it is

assumed that the decade values, S_i^* , are actually decade means, i.e.,

$$S_i^* = 0.1 \int_{10i-25}^{10i-15} \{S_0 + A \sin \frac{2\pi}{\tau_d} (t - t_0)\} dt, \quad (3)$$

where time, t , and the decadal-scale period, τ_d , are expressed in years. $t=0$ has been chosen as the mid-point of the 30-year data segment. t_0 represents the time of transition between successive positive and negative sine regimes, and is presumably dictated by external characteristics of global circulation. For the present study, no attempt was made to deduce the associated period τ_d , from S_i^* data (more than three values would, of course, have been required) but, instead, three reasonable values were selected (35, 40 and 45 yr.) in order to evaluate the sensitivity of derived values (A and t_0 , in particular, and their associated t_∞ values) to the period of modulation.

Integration of (3) gives three equations which can be solved readily for A , t_0 and S_0 . Differentiation of these defining relations, squaring and averaging, with the assumption that errors for different decades are negligibly correlated, give expressions for the standard errors of amplitude, $\sigma_E(A)$, of time of transition, $\sigma_E(t_0)$, and of long-term mean, $\sigma_E(S_0)$, in terms of $\sigma_E(S_i^*)$.

In assessing the significance of a given set of results, using (1), (2) and (3), it is clear that S^* from (1) and S_0 from (3) express similar information, leaving A and t_0 from (3) as measures of the degree of modulation (or of organized variability) of solar activity effects. The chance probability for the amplitude can be found from the Students'- t value, $A/\sigma_E(A)$, adjusted to t_∞ using the sum of the three component degrees of freedom from (2) and converted to a two-tailed chance probability, P_A , with the error-function integral.

In the absence of modulated solar activity effects, S_i^* values have a common mean of S^* , a common variance (which need not be specified) and are uncorrelated. It can be shown that the probability distribution for t_0 , if $\tau_d > 30$ yr, has a minimum at $t_0=0$ and a maximum at $t_0 = \pm \tau_d/4$. It will therefore be assumed that an appropriate value of Students'- t for t_0 is given by $\{\tau_d/4 - |t_0|\}/\sigma_E(t_0)$, which can be converted to t_∞ and to a two tailed chance probability, P_0 , in the same manner as for P_A , above.

The overall chance probability of decadal-scale modulation of solar activity effects on climate can be taken as the joint chance probability, P_0 , of the product $P_0 P_A$. If there are n chance probabilities with product given by

$$p = p_1 \cdot p_2 \cdot p_3 \cdots p_n,$$

then it is well-known that the joint chance probability of having a product of n pure-random probabilities no larger than p is

$$P_n = p \left\{ 1 + \ln 1/p + \frac{(\ln 1/p)^2}{2!} + \frac{(\ln 1/p)^3}{3!} + \cdots + \frac{(\ln 1/p)^{n-1}}{(n-1)!} \right\}. \quad (4)$$

Thus, $P_d = P_0 P_A (1 - \ln P_0 P_A)$.

As for P_1 (see Subsection 2.2, two final sentences), P_d will be somewhat smaller than the true chance probability because of the "background spectrum" effect. In addition, the specific model of modulation adopted is relatively arbitrary (although one can conclude

from the discussion in Subsection 3.2 that it is a realistic one), introducing a further element of uncertainty. Accordingly, P_d values will also be referred to as pseudo-chance probabilities.

3. Analysis Results

3.1 Pseudo-chance probabilities

Figure 1 displays the values of pseudo-chance probabilities plotted as square roots to give a better separation of points, for unmodulated solar effects (P_i from Subsection 2.2) and for modulated solar effects (P_d from Subsection 2.4, using 40 yr for τ_d), arrayed on a seasonal basis. The effect of using a specific value for modulation period will be shown later to be relatively trivial. The main feature of this figure is the clear indication of climate response to solar forcing, both modulated and unmodulated. If the spectral power in the vicinity of the solarcycle period had been negligible, both probabilities would have had a uniform distribution, on a linear scale, with, on average, nine points (out of 16) in the lower left quadrant (note that there are none) and one point in the upper right quadrant (note that there are nine). Seasonal variations are, in general, not readily apparent, except that modulation effects are strongest in spring and weakest in autumn.

The net assessment of these pseudo-chance probabilities is presented in Table 1, in terms of joint pseudo-chance probabilities, P_n , as derived from the product of individual pseudo-chance probabilities using (4), and of χ^2 probabilities, as derived from the frequency distribution of individual pseudo-chance probabilities. These quantities are given for various sub-groups (P_n , only, for eight individual probabilities combined) for modulated solar effects (P_d), unmodulated solar effects (P_i) and for net column values (i.e., for solar effects both modulated and unmodulated, since these can be considered relatively independent). The top sector of the table gives a differentiation by grouped seasons and the combination of all data, while the bottom sector compares results with F -tests to those with t -tests and results using monthly mean temperature, T , to those

Fig. 1. Pseudo-chance probabilities of solar variability effects on climate-modulated (P_d) and unmodulated (P_i)—plotted on a square-root scale for four seasons (two climatological parameters and two statistical tests giving four entries per season).

Table 1. Net pseudo-chance probabilities of solar variability effects on climate-modulated (P_d), unmodulated (P_i) and combined (NET)—using joint pseudo-chance probabilities (P_n) and chi-square (pseudo) probabilities ($P(\chi^2)$). Column headings denote nature of sub-division—two season pairs, two statistical tests and two climate parameters.

	WI/SP	SU/AU	ALL
$P_d:P_n$	2.24×10^{-7}	7.24×10^{-4}	2.71×10^{-9}
$:P(\chi^2)$			1.87×10^{-6}
$P_i:P_n$	8.78×10^{-2}	2.24×10^{-3}	1.55×10^{-3}
$:P(\chi^2)$			2.21×10^{-2}
NET: P_n	7.12×10^{-7}	1.41×10^{-5}	1.23×10^{-10}
$:P(\chi^2)$	2.45×10^{-4}	1.60×10^{-3}	1.78×10^{-10}

	F	t	T	σ_T
$P_d:P_n$	9.62×10^{-6}	2.25×10^{-5}	1.67×10^{-4}	1.17×10^{-6}
$P_i:P_n$	4.76×10^{-2}	4.81×10^{-3}	1.05×10^{-3}	1.15×10^{-1}
NET: P_n	7.93×10^{-6}	1.30×10^{-6}	1.71×10^{-6}	6.05×10^{-6}
$:P(\chi^2)$	1.04×10^{-2}	2.14×10^{-5}	1.87×10^{-6}	2.21×10^{-2}

using monthly standard deviation of (daily) temperature, σ_T .

It will be noted that the two assessment probabilities, P_n and $P(\chi^2)$, frequently differ markedly for groups of 16 members (three boxes for the χ^2 test). This is because an individual probability of 0.3 has the same effect on $P(\chi^2)$ as one of 0.003; this is certainly not true for P_n (as can be seen from (4)). In addition, there is a fixed lower limit on $P(\chi^2)$, which for 16 values and three boxes is 1.1×10^{-7} . Thus, $P(\chi^2)$ will in general exceed P_n for moderately small n values, and this excess exceeds two orders of magnitude for five of the eight cases in Table 1. For larger values of n (such as 32, for which six boxes are used in χ^2 tests) this effect would be expected to be less obvious; the close agreement for the only case in this category in Table 1 should probably be considered a fortuitous bonus. For the purpose of sub-set comparisons, it seems preferable to use P_n rather than $P(\chi^2)$. It may be noted that no attempt is made to employ an optimum procedure for testing joint significance, since the rejection of a null hypothesis is not being addressed.

The upper right column of Table 1 reveals that the net pseudo-chance probability of unmodulated solar effects such as obtained in this study is 2×10^{-3} , of the modulated solar effects is 3×10^{-9} and of the solar effects, whether modulated or unmodulated, is 10^{-10} . Even allowing for the uncertainties referred to earlier, it is not unreasonable to accept these results as strong indications of effects of solar variability on climate, both modulated and unmodulated (especially the former). Individual or collective pseudo-chance probabilities can thus be related usefully to the degree of association between solar activity and climate. Since these probabilities rely on the internal statistics of the computation systems, they are most properly regarded as indicating the degree of certainty (or uncertainty) that an effect of variable solar activity can be detected for a specific sub-set of the data analysed. Even if a given probability is acceptably small, it is by no means obvious that the effect detected is of physical significance, in terms of the magnitude of climate effects relative to natural variability. However, in terms of the

overall understanding of the physical mechanisms involved, one can justify a conscious priority for detectability, rather than predictability.

It may, therefore, be concluded that, for the data analysed in this study, modulated effects were easier to detect than unmodulated effects, especially in winter and spring, especially for F -tests and especially for standard deviation of temperature data. It may well *not* be a coincidence that both these latter categories involve variances—in quite different ways, of course. They are also very similar in that it is clear from Table 1 that the F and σ_T groups have much broader frequency distributions than is the case for t and T groups, with $P_n \ll P(\chi^2)$ only in the former cases. This would be expected from the differences between P_n values for P_d and P_l , and is associated primarily with the (F, σ_T) sub-group for which $P_n(P_d) = 3.6 \times 10^{-4}$ and $P_n(P_l) = 0.79$. When modulated and unmodulated effects are combined into single indicators of solar-effect detectability, there are no longer major differences between sub-groups, since modulated and unmodulated effects are in general somewhat negatively correlated, as can be seen in Fig. 1.

3.2 Modulation model

The modulation results in Subsection 3.1 have all been computed on the relatively arbitrary assumption that the modulation period, τ_d , was 40 years. Since the basic calculations were carried out also for τ_d values of 35 and 45 years, the impact of the choice can be assessed, in terms of the relative and absolute magnitudes of variations of t_0 (Δt_0 , for a $\Delta \tau_d$ of 5 yr) and of A (ΔA). In the case of transition date, the mean absolute value (over all 32 individual results) of Δt_0 was 1.20 yr; the mean value of $\Delta t_0 / \sigma_E(t_0)$ was 0.58. In the case of amplitude, the mean absolute value of ΔA was 0.194; the mean value of $\Delta A / A$ was 0.114 and the mean value of $\Delta A / \sigma_E(A)$ was 0.228. For values of Students'- t , corrected for sample size, the effects were smaller, so that the results in Table 1 are only marginally dependent on the selection of a specific τ_d value. Thus, for transition data, the mean absolute value of $\Delta t_\infty(t_0)$ was 0.196 and the mean absolute value of $\Delta t_\infty(t_0) / t_\infty(t_0)$ was 0.115. For amplitude, the mean absolute value of $\Delta t_\infty(A)$ was 0.075 and the mean absolute value of $\Delta t_\infty(A) / t_\infty(A)$ was only 0.044. Clearly, amplitudes are less sensitive to the choice of a τ_d value than transition dates, especially for this 30-year data segment, for which relatively large values of t_0 (quite sensitive to τ_d) were rather frequent.

One could, of course, regard the modulation model as a convenient, but rather artificial, method of expressing the variance of S_i^* values, so that the choice of a τ_d value was not important. This would, however, not be a productive attitude to take, since the important question to address is whether the variability of decade effects is associated in some systematic manner with variations of global circulation regimes on time scales of several decades. This association can be considered likely if the dates of transition have systematic rather than purely random characteristics. To examine this matter, amplitude-weighted frequency distributions of t_0 were computed, each sample value of t_0 being spread over the entire t_0 range in accordance with its standard error, assuming a Gaussian distribution. Sample sums of $Ap(t_0)$ were calculated for integral values of t_0 , and these are shown in Fig. 2 for winter plus spring, compared to summer plus autumn (eight individuals cases in each sub-group). The peak t_0 values (relative to the data mid-point of 1966) differ by only about two years (1972.5 and 1974.5, respectively), and the relative magnitudes are consistent with the values of $P_d:P_n$ in the top row of Table 1. It seems evident that these distributions are not random, and do reflect a change in global circulation in the early 1970's. The assumption that τ_d was 40 yr implies a transition of

Fig. 2. Amplitude-weighted frequency distribution of transition date for modulated solar activity effects on climate (eight individual distributions combined for each of winter plus spring and summer plus autumn). Zero date coincides with the mid-point of the data segment, 1966.0.

opposite type in the early 1950's, but since the 30-yr data segment began in 1951, it is reasonable to conclude that the major transition took place in the early 1970's. This view is supported by a finding of FLUECK and BROWN (1989) that the relation between smoothed 30 hPa North Pole winter temperature (QBO effects submerged) and 10.7 cm mean solar flux appeared to change phase around 1974, corroborated by a comparison of the two periods 1956–72 and 1973–89, with and without QBO partitioning, which showed major differences in solar response between the two periods.

When all seasons are combined, the amplitude data provide stronger proof of modulation (a net P_n , based on P_A values, of 1.01×10^{-7}) than the transition-date data (a net P_n , based on P_0 values, of 5.69×10^{-4}). Nevertheless, since the latter value is certainly strongly suggestive of significance, the close coherence of t_0 values revealed by Fig. 2 (and confirmed by similar plots, not shown here, for F vs. t and T vs. σ_T) provides a strong indication that the variability of solar effects evaluated for individual decades is not random, but follows a systematic pattern on a time scale of several decades. Moreover, given such pattern, it does seem reasonable to expect agreement of modal t_0 values especially on a seasonal basis (as shown in Fig. 2), since it is difficult to imagine an inter-annual serial correlation for some (as yet unspecified) global circulation state for one specific season but not (to some extent, at least) for all, on the assumption that the associated "memory" resides in some geographically-specific medium (such as polar snow and/or ice, or oceans, or one or more combinations of these).

4. Conclusions

Over the last century, the relatively popular belief was that variable solar activity had no effect on climate. This study provides strong support for the growing consensus that such a belief is invalid. The second conclusion of this study relates to a common belief over the last century that, if variable solar activity did have effects on climate, such effects should always be the same (with regard to sign and order of magnitude). In the case of data segments of a decade in length, it is most unlikely that this belief is valid. The third conclusion of this study is that inter-decade differences in solar effects are *not* random, but are organized in a coherent pattern in time. This pattern is substantially unaltered on changing season, climate parameter or statistical test, although there are major seasonal variation in the relative magnitude of modulated solar effects compared to unmodulated solar effects (inter-decade variability showing a maximum in spring and a minimum in autumn).

There is a striking parallelism between the above results and those obtained by LABITZKE and VAN LOON (1988, 1989). The latter studies lead to the conclusion that variable solar activity does indeed influence climate and that the effects for QBO-west phase and for QBO-east phase can be significantly different. Moreover, these differences are not random but are markedly specific to the QBO phasing, with the magnitude of the differences displaying major seasonal variations (a maximum in winter and a minimum in autumn). The strong coherence of the LABITZKE and VAN LOON (1988, 1989) results in the horizontal and in the vertical suggest that the present study has a much wider applicability than might be implied by its concentration on a single location and a single atmospheric level.

REFERENCES

- BRASSEUR, G., A. DE RUDDER, G. M. KEATING, and M. C. PITTS, Response of middle atmosphere to short-term solar ultraviolet variations: 2 Theory, *J. Geophys. Res.*, **92**, 903–914, 1987.
- FLUECK, J. A. and T. J. BROWN, A study of possible solar-weather relations, Preprint Vol., Am. Met. Soc. 11th Conference on Probability and Statistics in Atmospheric Science, Monterey, Calif., 325–333, 1989.
- GODSON, W. L., STP-M—A critical appraisal, *STP Newsletter*, **88-1**, 20–22, 1988.
- KEATING, G. M., J. Y. NICHOLSON, III, D. F. YOUNG, G. BRASSEUR, and A. DE RUDDER, Response of middle atmosphere to short-term solar ultraviolet variations: 1. Observations, *J. Geophys. Res.*, **92**, 889–902, 1987.
- KNOX, J. L., K. HIGUCHI, A. SHABBAR, and N. E. SARGENT, Secular variation of northern hemisphere 50 kPa geopotential height, *J. Climate*, **1**, 500–511, 1988.
- LABITZKE, K. and H. VAN LOON, Association between the 11-year solar cycle, the QBO and the atmosphere, Part I: the troposphere and stratosphere in the northern hemisphere in winter, *J. Atmos. Terr. Phys.*, **50**, 197–206, 1988.
- LABITZKE, K. and H. VAN LOON, Association between the 11-year solar cycle, the QBO and the atmosphere, Part III, Aspects of the association, *J. Climate*, **2**, 554–565, 1989.
- PITTOCK, A. B., A critical look at long-term sun-weather relationships, *Rev. Geophys. Space Phys.*, **16**, 400–420, 1978.
- PITTOCK, A. B., Solar variability, weather and climate: an update, *Quart. J. Roy. Met. Soc.*, **109**, 23–55, 1983.
- WMO/UNEP, Scientific assessment of stratospheric ozone: 1989, WMO Ozone Project Report No. 20, 450 pp., WMO, Geneva, 1990.