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On the incompleteness of the historical record of North Atlantic tropical cyclones

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[1] There is some question as to whether the historical record of observed North Atlantic tropical cyclones prior to the advent of satellite coverage is complete. This question is central to understanding the historical trend in tropical cyclone activity and the effect of environmental factors on it. To address this question, a statistical model of the relationship between annual cyclone counts between 1870 and 2004 and sea surface temperature and the state of the Southern Oscillation is extended to allow for non-decreasing observation probability prior to 1966. The estimated observation probabilities increase from 0.72 in 1870 to 1 in 1964. Allowing for record incompleteness reduces the estimated effect of sea surface temperature on annual tropical cyclone activity. **Citation:** Solow, A. R., and A. R. Beet (2008), On the incompleteness of the historical record of North Atlantic tropical cyclones, *Geophys. Res. Lett.*, *35*, L11803, doi:10.1029/2008GL033546.

1. Introduction

[2] There is considerable interest in understanding secular variability in the annual frequency of North Atlantic tropical cyclones and the factors that contribute to it. This interest stems in large part from the possibility that the risk posed by such storms to life and property will change in response to large-scale climate change. Although the annual record of North Atlantic tropical cyclone counts extends back to the middle of the 19th century, empirical analysis in this area is complicated by a potential under-count bias in the early part of the record [*Jarvinen et al.*, 1984; *Landsea et al.*, 2004]. Very briefly, prior to the advent of comprehensive satellite coverage in 1966 – and particularly prior to the advent of regular aircraft surveillance in 1944 – it is likely that not all tropical cyclones were recorded.

[3] Incompleteness in the early part of the annual record of tropical cyclone counts poses two related problems to empirical analysis. First, by itself, incompleteness in the early part of the observational record can give the spurious impression of an increasing trend in frequency. Second, because potential explanatory variables like sea surface temperature may themselves exhibit an increasing trend, a spurious trend in observed frequency may suggest a spuriously strong relationship between frequency and such explanatory variables.

[4] The degree to which the record of North Atlantic tropical cyclone counts is incomplete remains an open question. Indeed, this issue is central to a vigorous debate about both historical variations and likely future changes in the frequency of North Atlantic tropical cyclones. In broad terms, the issue of incompleteness has been treated in three ways. First, a number of analyses have attempted to circumvent the problem by focusing on either observation periods or types of storms for which the record may be reasonably assumed to be complete. For example, *Elsner et al.* [2000] focused on the 20th century record of major hurricanes. Although *Elsner et al.* [2000] recognized that the record may not be complete prior to 1943, they argued that “observational bias is smallest for the strong storms considered here”. This argument was challenged by *Landsea* [2001], who pointed out that the period of maximum intensity of a major storm may be too short to be detected without comprehensive coverage.

[5] A second approach has been to incorporate potential incompleteness as part of a statistical analysis of the record. For example, *Solow and Moore* [2000, 2002] developed and applied a test for trend in North Atlantic hurricane frequency that accommodated record incompleteness by assuming that the probability that a hurricane makes landfall in the US (and would therefore be recorded) is constant through time. This formalized an idea of *Fernandez-Partagas and Diaz* [1996]. *Landsea* [2007] made a similar argument in questioning claims by *Holland and Webster* [2007] that the frequency of tropical cyclones has doubled since 1900. As with *Elsner et al.* [2000], this approach is based on the assumption that a subset of the record – in this case, US landfalling storms – is complete. This assumption is open to question [*Holland*, 2007].

[6] In related work, *Solow and Nicholls* [1990] allowed for incompleteness in modeling the influence of the Southern Oscillation on the frequency of tropical cyclones in the Australian region over the period 1910–1965. This involved fitting a Poisson regression model relating tropical cyclone counts to an index of the Southern Oscillation for the period 1910–1988, but allowing for a non-decreasing recording probability for tropical cyclones prior to 1966. *Elsner and Jagger* [2006] adopted a similar approach in developing a seasonal prediction of US landfalling hurricanes. In this case, they adopted a Bayesian approach to fit a Poisson regression model to the historical record of US landfalling hurricanes back to 1851. This analysis, which used indices of the North Atlantic Oscillation (NAO), the Southern Oscillation (SO), and the Atlantic Multi-decadal Oscillation (AMO) as regressors, explicitly accounted for a constant probability that, prior to 1898, US landfalling hurricanes could have gone unrecorded. *Mann et al.* [2007] took an indirect approach to assessing the potential undercount bias in Atlantic tropical cyclones. Briefly, this involved inflating all annual counts prior to 1944 by a fixed

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amount and determining, via a statistical model of tropical cyclone activity, whether the adjusted record was homogeneous over time. The results suggested an undercount bias of around 1.2 cyclones per year.

[7] A third approach has been to assess record incompleteness directly. For example, *Chang and Guo* [2007] attempted such a direct assessment by essentially simulating the ship-borne observation process. Briefly, this involved superimposing recent tropical cyclone tracks on historical ship tracks and determining how often tropical cyclones would not be observed. A similar method was used by *Vecchi and Knutson* [2008]. As with *Solow and Moore* [2000, 2002], this general approach assumes that there has been no systematic change in tropical cyclone tracks, so that the catalogue of recent tracks is representative of the entire period.

[8] In this paper, we apply the general approach of *Solow and Nicholls* [1990] to model the record of North Atlantic tropical cyclone counts over the period 1870–2004. This involves extending the Poisson regression model of *Sabbatelli and Mann* [2007] to allow for record incompleteness. The approach here assumes only that the observation probability for a tropical cyclone is non-decreasing over the period 1870–1965.

2. Method

[9] Let the random variable N_t be the true number of tropical cyclones in year t (t = 1, 2, ..., T). Following *Sabbatelli and Mann* [2007], we assume that N_t has a Poisson distribution with mean:

$$\mu_t = \exp(\beta_o + \beta_1 SST_t + \beta_2 SO_t) \tag{1}$$

where SST_t is the August-October mean sea surface temperature over the main development region (6–18°N, 20–60°W) in year t, SO_t is the December–February Niño3.4 SST ENSO index in year t, and β_o, β_1, and β_2 are unknown regression parameters. Let the random variable Y_t be the observed number of tropical cyclones in year t. In the earlier part of the record, the observed counts are possibly incomplete in the sense that $Y_t < N_t$. Here, we assume that tropical cyclones occurring in year t are observed independently with observation probability p_t. It follows that the conditional distribution of Y_t given $N_t = n_t$ is binomial with n_t trials and success probability p_t. It is straightforward to show that, under this model, the unconditional distribution of Y_t is also Poisson with mean $p_t\ \mu_t$.

[10] To proceed, it is necessary to specify a model for the sequence $p_1, p_2, \ldots, p_T$ of observation probabilities. Here, we make the minimal assumption that this sequence is non-decreasing and that it reaches and therefore remains at 1 at known year T_o. Note that the possibility that observation probability reaches 1 before T_o is not ruled out. The assumption that T_o is known, which is reasonable in the present context, can be relaxed at some computational expense.

[11] Let $\beta = (\beta_o\beta_1\beta_2)$ and $p = (p_1 p_2 \ldots p_{T_o-1})$ denote the unknown regression parameters and observation probabilities, respectively. These parameters can be estimated by the method of maximum likelihood (ML). This involves maximizing the log likelihood:

$$L(\beta, p) = \sum_{t=1}^{T_o-1} (y_t \log(p_t\mu_t) - p_t\mu_t) + \sum_{t=T_o}^{T} (y_t \log \mu_t - \mu_t) \tag{2}$$

subject to the constraint:

$$p_1 \le p_2 \le \cdots \le p_{T_o-1} \tag{3}$$

where y_t is the observed value of Y_t and μ_t is given in (1). The constrained maximization of $L(\beta,\ p)$ must be done numerically. Initial estimates of β and p for use in numerical optimization can be found in the following way. First, fix the value of β. Second, for this fixed value of β, find the least squares estimate $\hat{p}_{LS}\ (\beta)$ of p under the constraint in (3). This estimate is the solution to a standard constrained quadratic programming problem and is therefore quickly found. Third, vary β and repeat the process to find the initial estimate $\tilde{\beta}$ of β for which $L(\beta,\ \hat{p}_{LS}\ (\beta))$ is minimized. The corresponding initial estimate of p is $\tilde{p} = \hat{p}_{LS}\ (\tilde{\beta})$.

[12] Beyond estimating β and p, interest also centers on testing the null hypothesis that the record is, in fact, complete and on constructing a confidence interval for β_1 (which controls the response of the rate of cyclogenesis to sea surface temperature). Let e denote a vector of 1's of length $T_o - 1$. The null hypothesis that the record is complete can be written as H_o : p = e. The likelihood ratio (LR) statistic for testing H_o is:

$$\Lambda_1 = 2\left(L\left(\hat{\beta}, \hat{p}\right) - L\left(\hat{\beta}_o, e\right)\right) \tag{4}$$

where $\hat{\beta}$ and $\hat{p}$ are the ML estimates of β and p allowing for record incompleteness and $\hat{\beta}_o$ is the ML estimate of assuming that the record is complete. That is, $\hat{\beta}_o$ is the ML estimate of β for the model in (1). Due to the length of p and the order restriction on its elements, standard results regarding the distribution of Λ_1 are not valid. Instead, the test can be based on the following parametric bootstrap. Let λ_1 be the observed value of Λ_1. Simulate the time series of tropical cyclone counts from the model fitted under H_o. Find the value of Λ_1 for the simulated series. Repeat the procedure a large number of times and estimate the significance level by the proportion of simulated series for which the value of Λ_1 exceeds λ_1.

[13] A $1 - \alpha$ confidence interval for β_1 is given by the set of values of b for which the null hypothesis $H_o : \beta_1 = b$ cannot be rejected at significance level α. The LR statistic for testing this null hypothesis is:

$$\Lambda_2 = 2\left(L\left(\hat{\beta}, \hat{p}\right) - L\left(\hat{\beta}_o, \hat{p}_o\right)\right) \tag{5}$$

where in this case $\hat{\beta}_o$ is the ML estimate of β with β_1 fixed at b and $\hat{p}_o$ is the corresponding ML estimate of p. The significance of the observed value γ_2 of Λ_2 can be assessed by the same parametric bootstrap described above using time series simulated from the model fitted under H_o. The endpoints of the $1 - \alpha$ confidence interval for β_1 are the

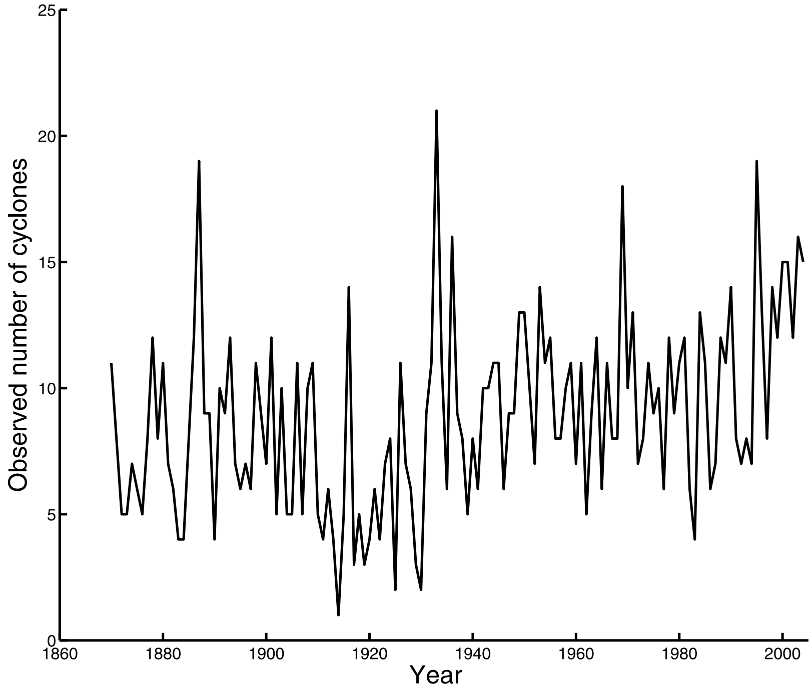

Figure 1. Observed annual North Atlantic tropical cyclone counts, 1870–2004.

smallest and largest values of b for which H_o is not rejected by this test at significance level α.

3. Results

[14] In this section, we present results from applying the methods outlined above to the annual record of observed tropical cyclones for the period 1870–2004. This record, which is shown in Figure 1, was assumed to be complete over the period 1966–2004. We fit the model in (2) to this record allowing for non-decreasing observation probabilities over the period 1870–1965. The ML estimates of β_o, β_1, and β_2 were −11.4, 0.50, and −0.15, respectively, and the estimated observation probabilities are shown in Figure 2. The value of the maximized log likelihood is 1450.5. The estimated probabilities in Figure 2 form a sequence of plateaus increasing from 0.72 in 1870 to 1 in 1964. It is interesting to note that some of the jumps in estimated observation probability correspond to known observation improvements. For example, in 1878, the US Signal Service

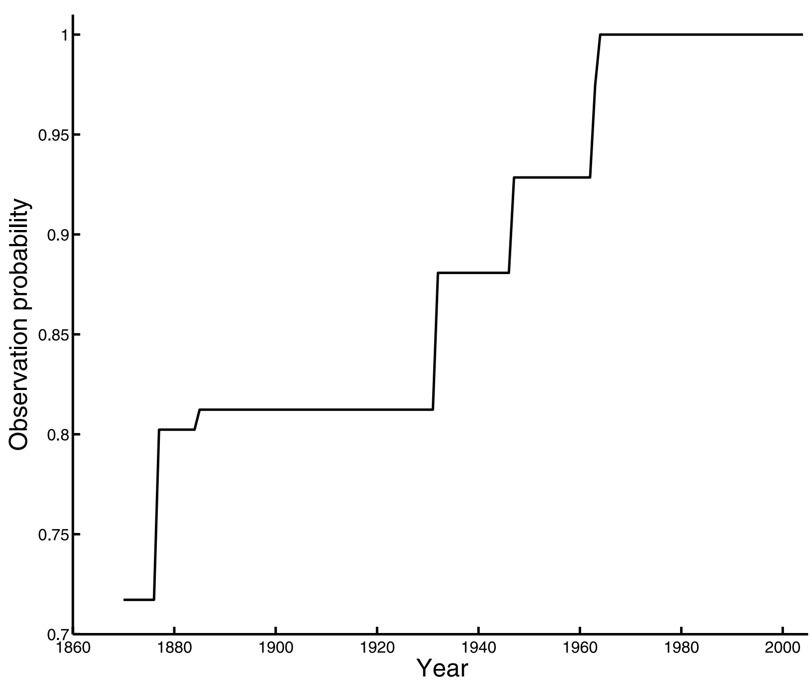

Figure 2. Estimated tropical cyclone observation probabilities, 1870–2004.

began tracking all West Indian hurricanes and systematic aircraft reconnaissance began in 1944. The jump in 1932 does not appear to be connected to such an event. To provide some context for the estimated probabilities in Figure 2, for a year in which there are 10 tropical cyclones and the observation probability of 0.8, the probability that all cyclones are observed is 0.12 with an expected number of missed cyclones of 2. Although not directly comparable, this estimate of the mean number of missed cyclones appears to be slightly larger than that of *Mann et al.* [2007].

[15] The ML estimates of β_o, β_1, and β_2 under the null hypothesis H_o that the record is complete were -15.8, 0.66, and -0.15, respectively, and the maximized value of the log likelihood was 1445.9. The value of the LR statistic Λ_1 in (4) was 9.2. Of 1000 records simulated from the model fitted under H_o, only 48 had values of Λ_1 greater than 9.2, for an estimated significance level of 0.048. Thus, by conventional standards, the null hypothesis that the record is complete can be rejected.

[16] Under the model in (2), the effect of sea surface temperature on cyclogenesis is controlled by the parameter $\beta_{1.}$ Allowing for record incompleteness reduced the estimate of β_1 from 0.66 to 0.50. In rough terms, *SST* increased by 0.8°C between 1870 and 2004. For β_1 = 0.66, this corresponds to a 70% increase in the mean annual number of tropical cyclones, while for β_1 = 0.50, this increase is 50%. This underscores the importance of accounting for record incompleteness in assessing the effect of environmental factors on the frequency of tropical cyclones.

[17] Finally, we used the parametric bootstrap procedure described in the previous section to construct a 0.95 confidence interval for β_1. This interval was (0.26, 0.75). To put this result in context, the corresponding interval for the increase in mean tropical cyclone number due to a 0.8°C increase in *SST* is (23%, 84%).

4. Discussion

[18] The purpose of this paper has been to describe and apply an approach to modeling the dependence of mean annual tropical cyclone number on environmental variables allowing for incompleteness in the observed cyclone numbers. The approach makes the minimal assumption that the probability of observing a cyclone is non-decreasing in time. It has been suggested that observation probability declined during the world wars. However, no significant improvement in model fit was achieved by allowing for this possibility.

[19] Turning to the results of the previous section, this analysis supports the position that the early part of the observed record of tropical cyclone counts is incomplete. Ignoring this incompleteness leads to an over-estimation of the effect of increasing sea surface temperature on mean tropical cyclone frequency. The 0.95 confidence interval for this effect is rather wide. The approach outlined here can be extended in at least two ways. First, it could be applied to the record of US landfalling tropical cyclones for which stronger claims of completeness have been made. Second, while still accounting for record incompleteness, the relationship between true storm count and environmental variables can be refined. For example, *Swanson* [2008] argued that tropical cyclone activity depends not only on SST in the main development region, but also on mean tropical SST.

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