Precursory remote triggering is absent near the epicenters of impending great earthquakes

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Abstract: Recently, there have been numerous great ($M_w \geq 8$), devastating earthquakes, with a rate in the last 7 years that is 260% of the average rate over the 111-year seismological history. Each great earthquake presents an opportunity to study a major fault at the very end and very beginning of the inferred seismic cycle. In this work, we use these events as both targets and sources to probe susceptibility to dynamic triggering in the epicentral region before and after a large earthquake. This study also carefully addresses the possibility that large earthquakes interact in a cascade of remotely triggered sequences that culminate in further large earthquakes. We seek evidence of triggering associated with the 16 great $M_w \geq 8$ events that occurred between 1998 and 2011, using regional and global earthquake catalogs, to measure changes in inter-event time statistics. Statistical significance is calculated with respect to a non-stationary reference model that includes mainshock-aftershock clustering. In only a few cases do we detect triggering near the epicenters of $M \geq 8$ earthquakes separated by more than 10º. The number of detections is not significant, given the number of detection attempts. Systematic triggered rate changes are less than 15% at 95% confidence, and thus cannot account for the large increase in $M_w \geq 8$ earthquake rate. The catalogs are insufficiently complete to resolve more moderate triggering expected from previous studies. We calculate that an improvement in completeness magnitude from 3.7 to 3.5 could resolve the expected triggering signal in the ISC catalog taken as a whole, but an improvement to M 2.0 would be needed to consistently resolve triggering on a regional basis.

1. Introduction
The last seven years have experienced a surge in great (M ≥ 8) earthquakes relative to the preceding 4 decades [Ammon et al., 2010]. Between the Mw 9 Sumatra earthquake of December 2004 and the Mw 9 Tohoku-Oki earthquake of March 2011, Earth has averaged 1.7 great earthquakes per year, which is 260% of the rate of 0.66 per year over the entire seismological record extending back to 1900 (i.e. a rate increase of 160%). This interval of heightened great earthquake occurrence has prompted many to consider whether the global increase could represent long-range interactions between great earthquakes [Brodsky, 2009; Michael, 2011; Shearer and Stark, 2011]; can the occurrence of one great earthquake increase the likelihood of a subsequent quake in a self-exciting process? This certainly seems to be the case for some of the nearby M ≥ 8 earthquakes. Examples include the 2005 and 2007 Sumatra earthquakes that followed the 2004 Sumatra earthquake, rupturing portions of the plate boundary to the southeast immediately adjacent to and 700 km away from the first event [Nalbant et al., 2005; Wiseman and Burgmann, 2011], as well as the 2006 - 2007 Kuril doublet. However, even after these instances are removed, there remains a surfeit of global great earthquakes – 1.14 per year, almost double the century-long average rate. Could very large magnitude earthquakes have an extended reach, beyond that for conventionally accepted aftershocks, such that they can trigger earthquake cascades at great distance on 1-10 year timescales? The peak dynamic strains associated with the surface waves of great M ≥ 8 earthquakes exceed 10^{-6} at global distances, and such strains are commonly observed to triggered small earthquakes [Gomberg and Johnson, 2005; Hill and Prejean, 2007; Lei et al., 2011; van der Elst and Brodsky, 2010; Velasco et al., 2008]. Indeed, there are reports of triggering of remote activity for many of the recent great events. Examples include triggering in China by the 2003 Tokachi-Oki and 2004 Sumatra earthquakes [Lei et al., 2011; Wu et al., 2011], triggering in Alaska and the continental US by the 2004 Sumatra earthquake [Rubinstein et al., 2011; West et al., 2005], Triggering in California by the 2010 Maule-Chile earthquake [Peng et al., 2010], and triggering throughout the United States by the 2011 Tohoku-Oki earthquake [Rubinstein et al., 2011]). Dynamic triggering of remote events is no longer controversial. However, earthquakes larger than about M 5 have not yet been observed to be dynamically triggered during passage of great event
surface waves [Parsons and Velasco, 2011]. Nevertheless, statistical aftershock models predict that sequences of smaller events may occasionally culminate in delayed events that are larger than the initial events [Felzer et al., 2002; Felzer et al., 2004]. A good example of this process is the M 5.6 Little Skull Mountain earthquake, which occurred 280 km distant and 22.3 hours after the 1992 Landers earthquake, but was preceded by a sequence of small events initiated within the coda of the Landers main shock [Anderson et al., 1994].

This motivates considering the possibility of a global earthquake cascade, which we test in this study. The model is as follows: large earthquakes radiate powerful surface waves that dynamically trigger small events at global distances. These remotely triggered aftershocks trigger their own aftershocks, and so on. Occasionally, secondary aftershocks are larger than their main shock, ramping up the local seismicity rate. This process continues, and may eventually culminate in another great earthquake, at which point the global cycle is renewed. Any particular remotely triggered aftershock sequence is very unlikely to produce an aftershock larger than the initially triggered events, let alone a M ≥ 8 quake, but if enough small remotely triggered aftershock sequences are initiated globally, the cumulative probability of a few sequences producing big events over several years may be large.

A remotely triggered rate increase of 160% may be considered extreme even for the largest of the recent great earthquakes, except if located in highly susceptible regions. At the core of the cascade hypothesis is therefore a fundamental physical question about whether the dynamic triggerability of a fault zone reflects the accumulation of stress on the fault. Since transient stresses generated by seismic waves are much smaller than the total strength of a fault or the stress drop in an earthquake [Gomberg et al., 2004], seismic waves must act as the “straw that breaks the camel’s back,” pushing an already critically stressed fault over the edge to failure. This expectation is guided by laboratory experiments designed to simulate dynamic triggering on stick-slip faults [Beeler and Lockner, 2003; Savage and Marone, 2008] and by numerical studies of theoretical earthquake nucleation models [Gomberg, 2001]. These results imply that when and
where triggered earthquakes are observed, we may be able to infer the presence of critically stressed faults. It is not known, however, whether the rupture zone as a whole is critically stressed before a great earthquake, that is, whether an earthquake knows in advance how big it is going to get. Susceptibility to dynamic triggering may therefore serve as a probe of the state of stress before (and after) a great earthquake.

Here we consider the seismic catalog of the last 13 years, using the cascade model to guide the search for distant triggering. If the recent great earthquakes are part of a global self-exciting cascade, the rate of small earthquakes at the site of some or all impending great earthquakes should have increased at the time of some previous large earthquake(s). We therefore search systematically for triggered rate changes localized near the epicenters of subsequent great earthquakes at the time of earlier great earthquakes in the sequence. We look for triggering by all of the $M \geq 8$ earthquakes identified in the PAGER catalog since 1998, using two global earthquake location catalogs (PDE and ISC), and one regional catalog (JMA (see Data and Resources). We measure rate changes using the inter-event time ratio method [Felzer and Brodsky, 2005; van der Elst and Brodsky, 2010], and evaluate the significance of any detected rate changes with respect to a Poisson process on a region-by-region basis.

1.1 Previous statistical studies on global earthquake rate change

Some studies have looked at whether recent earthquake rates on a global scale are consistent with a stationary, uniform Poisson process [Michael, 2011; Shearer and Stark, 2011]. These two studies examined the statistical distribution of $M \geq 7$ earthquake inter-event times over the past century using a global seismicity catalog, and found that the observed clustering of great earthquakes is not significantly non-Poissonian, aside from regional clustering (as expected for proximate aftershocks).

Here we briefly illustrate this approach and summarize the conclusions. In the PAGER catalog, which is a compilation of large historical earthquakes that emphasizes consistent magnitude determination, 77 $M_W \geq 8$ earthquakes have been recorded in the 111 years since 1900. Twelve of these earthquakes occurred between December 2004 and the
present. Three of these twelve are considered regional aftershocks of other great earthquakes (the 2005 and 2007 Sumatra quakes, and the 2007 Kuril quake), leaving 9 independent events in 7 years since December 2004. The binomial probability of finding 9 of 74 uniformly distributed events in such a 7-year time window is 4.3%. This number would likely decrease further if we also declustered the long-term catalog. However, when we consider the probability of finding some unusual clustering in a long duration catalog, the clustering is not particularly unusual. That is, although the recent cluster of earthquakes has low probability of occurrence for any single 7 year period, finding clustering in one such period out of 111 years is not unexpected [Shearer and Stark, 2011].

On the other hand, inability to falsify the uniform Poisson hypothesis is not proof of constant uniform rate. By convention, scientific hypothesis testing is conservative, and we are usually more concerned with minimizing the probability that the null hypothesis is falsely rejected (Type I error), than falsely accepted (Type II error), and it is preferable to use the simplest model possible as the null hypothesis. However, in considerations that have major societal implications, as in the case of global earthquake clustering, it is important to quantitatively address both types of error, that is, to quantitatively assess our confidence that the rate has not increased.

Addressing the potential for type II error is equivalent to asking: what is the maximum rate change that could still pass the Poissonian hypothesis test at 95% confidence? That is, what are the confidence bounds on the rate change given the limitations of the catalog? Given $N$ observed earthquakes, the 95% confidence bound on the Poissonian rate parameter $\lambda$ is just the highest value of $\lambda$ that would produce $N$ or fewer events at least 5% of the time.

To briefly illustrate this concept, we break down the earthquake record into periods between 1900 and Dec. 2004, and after Dec. 2004 (a prominent change point). This is only an illustration of the concept, and does not constitute a test of the uniform Poissonian hypothesis. We do not consider here the appropriateness of the change point,
nor do we decluster the pre-2004 catalog. The (undeclustered) intensity in the first period is $\lambda_0 = 0.62$ events/yr. With 9 (declustered) events in the subsequent 7 years, we can be 95% confidence that the rate in the second period is greater than 0.67 and less than 2.24 events/yr. The corresponding rate change $\delta \lambda \equiv \Delta \lambda / \lambda_0$ is $8\% - 261\%$, which covers the naïve observation of an apparent rate increase of 160%. These large confidence bounds indicate that the test has very little power to reject the Poisson hypothesis in such a small time window. We cannot statistically rule out a large rate increase, any more than we can rule out uniformity in rate. Thus, purely statistical treatments of the overall earthquake history are not likely to resolve the question of long-range interactions. More direct examination in the context of a physically motivated framework may provide firmer conclusions.

This paper is organized as follows: First, we introduce the inter-event time statistic and define the quantities to be measured. We then describe the spatial and temporal windowing of the catalogs. Before applying the statistic to the real catalog, we calculate the expected rate change at each site, given previously established scaling laws for remote triggering, so that we can interpret the incidence of non-detections. We then proceed to measure actual triggered rate changes near the epicenters of great earthquakes, both before and after they occur, calculating significance with respect to a homogeneous Poisson process. We find no precursory triggering before impending great earthquakes that could explain the global increase in rate. We do find an apparent triggering signal in the wake of great earthquakes, but we show that this apparent post-seismic triggering signal is a consequence of the over-simplified uniform Poisson reference model. We then develop a non-stationary reference model that accounts for mainshock-aftershock clustering, and show that this removes the apparent triggering signal in the post-seismic target regions. Finally, we use the information gathered at each target site to assess how much improvement would need to be made to the global seismic network in order to detect remote triggering at the sites of major subduction zone earthquakes on a consistent basis.

2. Method
2.1 Inter-event time ratios

To capture rate changes in the impending rupture zones of great earthquakes, we take the sample mean of the inter-event time ratio $r$, calculated in regional spatial bins, defined by

$$ r = \frac{t_2}{t_1 + t_2}, \quad (1) $$

where $t_1$ and $t_2$ are the intervals to the first local earthquake before ($t_1$) and after ($t_2$) some reference time – in this case, the time of a distant great earthquake. If the regional earthquake times are uncorrelated with the time of the distant great earthquake, $r$ is uniformly distributed between $[0,1]$. If the timing of regional earthquakes is advanced by the occurrence of the distant great earthquake (i.e., if there is a remotely triggered earthquake or one of its aftershocks in the local bin), $t_2$ will be on average smaller than $t_1$, and the distribution of $r$ will be shifted toward smaller values.

Each great earthquake target zone is divided into a uniform grid, and $r$-values are calculated from pairs of earthquakes within each grid square. The target zones are defined by the first 10 days of aftershocks of each great earthquake, extending no further than 1° from the main shock epicenter. The spatial constraint focuses attention on the plausible nucleation regions for the great events. The grid spacing within this region is optimized to give the maximum number of unique pairs of earthquakes in a region. (Only one value of $r$ can be calculated for each grid square, so the number of unique pairs increases with the grid fineness, up to the point where bins begin to lack earthquake pairs straddling the trigger events.) This optimization is done for each trigger-region pair.

The population of $r$-values for each target zone is averaged, giving a mean $\bar{r}$ (denoted $\bar{r}$), and this mean is compared to the expected mean for a reference model (e.g. uniform Poisson process, Appendix A). The triggering detection confidence is defined as $c = 1 - p_r$, where $p_r$ is the significance, i.e. the probability that a random process would produce a mean as small or smaller than the observed $\bar{r}$, by chance. Note that using this one-sided definition, the triggering confidence equals 50% for zero rate change, and a triggering confidence near zero means that an apparent rate decrease is present. However, this definition keeps us from assigning negative confidence to values of $\bar{r} < 0.5$. 
We also calculate a fractional rate change $\delta \lambda$ consistent with the observed $\bar{r}$, assuming that the earthquake rate is constant and Poissonian, but with a step-change in rate at the time of the trigger (Appendix A). The fractional rate change $\delta \lambda$ is defined as the difference in rates, normalized by the initial rate. We also compute the upper and lower 95% confidence bounds on $\delta \lambda$, which depend on both $\bar{r}$ and the sample size (Appendix A). The upper and lower bounds can be thought of as the highest and lowest rate changes that could still pass a Poissonian hypothesis test. These three values: lower bound, best estimate, and upper bound, are denoted $\delta \lambda_{0.05}$, $\delta \lambda_{0.50}$, and $\delta \lambda_{0.95}$, respectively.

2.2 Triggers and target regions and time windows

The trigger and target regions consist of all $M \geq 8$ earthquakes since the $M_w$ 8.1 Balleny Islands earthquake of 1998 (Figure 1 and Table 1). Earthquake magnitudes are taken from the PAGER catalog, and all correspond to $M_w$ values from centroid moment tensor inversions.

The window for measuring triggering relative to each great event is limited so that triggering windows do not overlap with other trigger times (windows themselves may overlap, but cannot contain more than a single $M_w \geq 8$ trigger event). Allowing more than one trigger in each window could lead to double counting and would allow the impending regional quake itself or its aftershocks to be considered a remotely triggered event. Double counting, or counting the $M_w \geq 8$ earthquakes themselves (which were used to formulate the hypothesis), would invalidate the statistical significance calculations. Each time window is also constrained to be symmetrical around the great event trigger time, as the inter-event time ratio $r$ would be biased if the waiting times $t_1$ and $t_2$ were limited to different ranges.

The choice of target time window and target region do not strongly influence the conclusions of this paper. Using smaller time windows has the effect of diminishing the sample size and reducing statistical robustness, but larger time windows have the effect...
of averaging out any triggered rate increase over the duration of the window. Using a constant, smaller maximum time window of ±59 days, for example, changes the joint statistics by 1-2%, but does not alter any of the conclusions of this study. We choose to use variable-length windows because it maximizes the number of samples overall.

The requirement of non-overlapping windows is a problem for the 2004 Macquarie Islands and 2004 Sumatra-Andaman earthquakes, which are separated by only 2.4 days. This could be taken as a perfect example of one great earthquake being triggered on the heels of a 77º distant earthquake. However, no significant precursory triggering of smaller events due to the Macquarie earthquake is observed in the rupture region of the impending Sumatra quake (there is only one earthquake near Sumatra in the interval between them reported in the ISC catalog). Rather than impose a 2.4-day time limit on both earthquakes, we choose to expand the triggering window to the limits imposed by the previous and next events in the sequence (~95 days rather than 2.4). This means that any triggering by the Sumatra-Andaman quake may be attributed to the Macquarie quake and vice versa. These two triggers are considered as a single event when calculating the joint statistical significance of measured rate changes.

Triggering susceptibility is measured using three target catalogs. We first examine the PDE and ISC global earthquake catalogs. The PDE is a rapidly published compilation of many reporting networks worldwide. It has an overall completeness level of roughly $M_C = 4.1$ and is complete up to the present. (We determine the completeness magnitude as the magnitude above which a linear fit can explain at least 90% of the variance in log-cumulative-number vs. magnitude.) The ISC catalog is the authoritative final catalog combining the best data worldwide. It has a somewhat lower completeness threshold ($M_C = 3.7$), but extends only through 11/2009 at this time. The PDE catalog is therefore the only source of regional target data for the 2010 Chile, and 2011 Tohoku trigger earthquakes. Where the PDE catalog does overlap with the ISC catalog, it provides a qualitative check on the effect of catalog uncertainties on the measured triggering significance, though we consider the ISC catalog to be authoritative. The ISC includes un-reviewed hypocenter data below magnitude 3.5, and the quality of this data is very
non-uniform in time. We therefore restrict the ISC catalog to ‘prime’ quality events above M 3.5. Finally, the regional JMA catalog is used for target regions near Japan. This catalog has a much lower completeness threshold (M_c = 1.3) in the Tokachi-Oki and Tohoku-Oki target regions.

2.3 Binomial and joint significance tests
We compute the significance of \( \bar{r} \) for each of the 256 individual trigger-target pairs (16 triggers at 16 target regions), with respect to the reference model (e.g. uniform Poisson process) and report triggered rate change detection successes as instances with greater than 90% triggering confidence. This represents a relatively low significance threshold, designed to capture relatively low rate changes in small samples, and it requires that we always consider the number of detection thresholds expected by chance when interpreting the results for any subgroup of the data. We therefore also compute the joint significance for several subgroups of the data: 1) All 120 trigger-target pairs that come any time before the regional earthquakes, reflecting long-term precursory triggering; 2) The 15 pairs from Group 1 that come immediately before the regional earthquakes, reflecting short-term precursory triggering; 3) All 120 pairs that come any time after the regional earthquakes; and 4) the 15 pairs from Group 3 that fall immediately after the regional earthquakes. Groups 1 and 2 tell us about any potential earthquake cascade or precursory triggering, and Groups 3 and 4 tell us about changes in triggerability in response to damage induced by a large regional quake.

We compute the combined significance in these subgroups, using the same method as in the case-by-case basis, i.e. as the probability of obtaining a smaller \( \bar{r} \) by chance for the entire population (the p-value), denoted \( p_r \). However, computing \( p_r \) in this way weights the calculation toward the sites with the most earthquakes, and this may dampen the triggering signal from more sparsely covered regions. For this reason, we also compute the significance of the number of detection success within each group, treating each measurement as a Bernoulli trial with probability of success \( 1 - c_{th} \), where \( c_{th} \) is the detection confidence threshold. The triggering signal could take several forms: occasional high confidence triggering at a few regions, or a systematic small bias toward slightly
positive rate changes. To capture both of these cases, the binomial test is computed for
thresholds of 90% and 50%, with p-values denoted $p_{90}$ and $p_{50}$. The results of these tests
are considered statistically significant if the p-values fall below 0.05.

In this paper, we follow the convention that case-by-case triggering measurements will be
reported in terms of triggering confidence, and all joint statistics derived from multiple
confidence measurements will be reported in terms of significance level $p$.

### 2.4 Expected rate changes from surface wave strain amplitudes

In order to interpret the incidence of non-detections, we also calculate an expected rate
change based on previous regional observations. Van der Elst and Brodsky [2010]
calibrated a relationship between the amplitude of triggering waves and the triggered rate
change using the interevent time ratio $r$ in California. Fitting a power law to the triggered
rate change versus peak surface wave strain $\epsilon$ in that study gives an expected rate change
of

$$\delta \lambda_{cap} = 82 \times \epsilon^{0.43}. \quad (2)$$

An identical analysis of Japan seismicity found that Japan is $\sim$3 times less triggerable
than California. This gives an indication of the variability we should expect in triggering
susceptibility worldwide. The triggering sites in this study are also potentially much
deeper in this study than in the one used to calibrate Eq. 2. We therefore treat Eq. 2 only
as a rough estimate of expected rate change, and likely an upper bound.

Van der Elst and Brodsky [2010] estimated peak surface wave strain using the empirical
surface wave magnitude equation, which relates wave amplitude to earthquake magnitude
and distance. For the large magnitude earthquakes we are considering here, the surface
wave magnitude is saturated and the empirical scaling is inadequate. Instead, we
measure peak velocities directly at nearby seismic stations (see Data and Resources).
Dynamic strain amplitude is estimated as

$$\epsilon = \frac{v_{\text{max}}}{c_s}, \quad (3)$$
where $c_s = 3.5 \text{ km/s}$ is the surface wave phase velocity, and $v_{max}$ is the peak vertical velocity measured on a broadband sensor. Seismic stations are selected within 6 - 10° of the target earthquake epicenter, depending on the density of stations, and the instrument response is removed over a passband of 1 - 50 s. Amplitude at the target earthquake epicenter is extrapolated from the measured values using the distance decay exponent from the surface wave magnitude equation $M_s = \log A + 1.66 \log \Delta + 2$ [Lay and Wallace, 1995]. Where multiple records exist, the amplitudes are distance-corrected and averaged. In the 14 out of 256 cases where no amplitude measurements exist within 10° of the epicenter, the trigger amplitude is estimated as the mean of the distance-corrected amplitudes measured at all the other target sites.

The measured peak ground velocities agree well with the empirical surface wave amplitude equation multiplied by $2\pi/T$ (period $T = 20$ sec) to get velocity, but with somewhat reduced dependence on magnitude, as expected for this magnitude range. There is also antipodal focusing of the surface waves beyond 150° distance that is not captured by the empirical surface wave equation.

Using Eqs. 2 and 3 to transform measured peak velocity into expected rate change, we find a mean expected rate change $\delta \lambda_{exp}$ over all trigger-target pairs on the order of 11% (Figure 2). This expectation is for rate change averaged over a large area, and triggered rate change may be much higher locally. Indeed, it must be higher locally if triggering is to account for the observed increase in the rate of large earthquakes.

Such small rate changes require large sample sizes to establish high significance (Table 2, Appendix A). For example, a sample size of $m = 11$ is required to establish a 100% rate change (doubling) at 90% confidence. A sample size $m = 543$ would be required to establish a rate increase of 10%. To establish a 1% rate change at 90% confidence would require $m = 49,765$. Given the limited completeness of the global seismicity catalog, it may be possible to resolve rate changes of 10-100% on a regional basis, but smaller rate changes (< 10%) will likely go unnoticed.
3. Results

3.1 Example target location

As an example, the target region for the Tohoku-Oki earthquake is shown in Figure 3, using the JMA catalog, for all 16 of the potential trigger earthquakes (Table 1), with triggers 5 and 6 combined. This is the best-instrumented target region out of all 16 targets (M_c = 1.3), and therefore the density of earthquake pairs is very high. Most other sites, using the global earthquake catalogs, are much more poorly sampled.

One trigger beyond 10º distance produces a mildly significant triggering signal (93% confidence) within the target region: the 2003 Peru earthquake (Figure 3, trigger 3). The nearby (3.7º) 2004 Tokachi-Oki earthquake (Figure 3, trigger 4) also clearly triggered earthquakes within the epicentral region of the 2011 Tohoku-Oki quake, with p-value p < 10^{-16}. Examining the effect of the Tohoku-Oki mainshock on its own epicentral region, we see that the M 7 foreshock sequence, initiated two days prior, shows up as a small patch of apparent triggered rate decrease (blue dots, Figure 3, trigger 16), within a robust conventional aftershock sequence that fills in the frame to the northeast and southwest.

3.2 Global ISC catalog – uniform Poisson reference model

We now apply the inter-event time statistic to the global ISC catalog. We measure triggering significance with respect to a stationary, uniform Poisson reference model. As mentioned in the introduction, this is an imperfect model, and it may spuriously map any non-Poissonian clustering into the triggering signal. This is problematic because we know that a Poisson process is a very poor approximation of the aftershock sequences of great earthquakes. We will correct for this in Section 4 by introducing a non-stationary reference model that accounts for aftershocks. However, we first present the data using the over-simplified stationary reference model in order to clearly demonstrate which of our conclusions are model-dependent. The equivalent analyses for the PDE and JMA catalogs, with respect to the uniform Poisson reference model, are included in Appendix B.
Applying the inter-event time test to all sets of triggers and target regions gives a matrix of triggering detections (Figure 4). The triggers are sorted by time of occurrence along the y-axis, and sorted as target regions along the x-axis, such that boxes above the diagonal represent periods after the time of the great earthquake in that region, and boxes below the diagonal represent periods before the occurrence of the regional great earthquake. Trigger-target pairs that are less than 10º distant are marked with red squares, and are not included in the joint remote triggering statistics. We will delay detailed discussion of the triggering matrix for Section 5, after we introduce the improved reference model, and consider only the major features here.

The diagonal of the triggering matrix shows the effect of a trigger earthquake upon its own rupture area (Figure 4). Triggering of conventional aftershocks along this diagonal is detected with very high confidence ($p_r < 10^{-16}$). Beyond 10º distance, however, systematic triggering is not robustly evident, with $p_r = 0.17$ (Table 3).

The lower right triangle of the triggering matrix (Figure 4) is the region to examine for possible triggering precursory to great earthquakes. We do not detect systematic triggering for separations greater than 10º, with a joint triggering significance over the whole precursory population of $p_r = 0.50$ (Table 3). This is the expected value for exactly zero triggered rate change.

The upper left triangle of the matrix (Figure 4) is the region to examine for possible triggering in regions after a great earthquake. This subgroup does appear to show strong triggering, with a joint triggering significance of $p_r = 0.07$ (Table 3). We will now show that this apparent post-seismic triggering is due to the failure of the uniform Poisson reference model, and is not a robust feature of the dataset. The precursory triggering measurements, on the other hand, are not affected by the choice of reference model.

### 4 Methods revisited

#### 4.1 Limitations of the uniform Poisson reference model.
In Section 3.2, the significance $p_r$ was calculated with respect to a uniform Poisson process (Appendix A). However, the regional catalogs are very non-Poissonian due to local mainshock-aftershock clustering. This is a particularly serious problem for this study, because the selection of time windows (Section 2.2) is conditioned on the recent occurrence of large earthquakes – and hence their aftershocks – at the beginning of the time window.

Figure 5 shows the effect of taking a random sample of inter-event time ratios from within an ongoing aftershock sequence, and calculating significance with respect to a uniform Poisson reference model. Even though aftershock rates are decaying – which intuition suggests might produce a spurious rate decrease in the apparent triggering metric – the $r$ statistic actually indicates a spurious rate increase. This is because aftershocks tend to cluster toward the beginning of each target window, preferentially giving large $t_1$’s and small $t_2$’s (Eq. 1).

In order to accurately measure the effect of remote triggering at the time of the trigger, we must compute the significance of the measured $\bar{r}$ with respect to a modeled distribution that accounts for local clustering at the beginning of the time window.

### 4.2 Significance with respect to a non-stationary reference model.

The appropriate distribution of the expected value of $\bar{r}$ is modeled using a simulated catalog that reproduces the heterogeneities in rate due to local mainshock-aftershock sequences. We call this model the non-stationary or “Omori” reference model. The model catalog contains two seismicity components: 1) spontaneous background earthquakes with a statistically constant rate $\mu$, and 2) aftershocks of large ($M \geq 8$) regional and moderate local earthquakes that follow the modified Omori’s law, resulting in a combined model rate $\lambda(t)$ defined by

$$\lambda(t) = \mu + \sum_{i=1}^{n} k_i (t - t_i + c)^{-p} u(t - t_i).$$

Here $k_i$ are the productivity terms for each of the $n$ aftershock sequences, $c$ and $p$ are constants governing the time decay of aftershocks, and the unit step function $u(t - t_i)$ is
zero if $t < t_i$ and one otherwise. The rate $\lambda$ at some time $t$ is hence the sum of the constant background rate and all the aftershock rates from mainshocks with times $t_i < t$.

Aftershock rates $k_i$ are calculated for local mainshocks with a significant effect on the time series. Significant local mainshocks are identified as the minimum population of large local events sufficient to capture the observed cumulative time series within the range of the modeled time series (e.g., Figure 6). Local mainshocks are modeled only if their inclusion is required to achieve overlap between the modeled and observed cumulative time series, and not if they merely improve the fit by some statistical measure. This captures large non-uniformities in rate without over-fitting the time series, which could mask the signal of remote triggering. The number of significant local mainshocks ranges from $n = 0$ to 5 over the various target regions (Appendix C). The smallest local mainshock magnitude is 5, and the mean magnitude is 6.4. The model rate also includes aftershocks of the local target $M_w \geq 8$ earthquake, as well as any other regional $M_w \geq 8$ earthquakes within $10^\circ$.

To verify that we are not fitting out the signal we are hoping to measure, we also have run tests in which we 1) model only the major $M \geq 8$ quakes with no local mainshocks, 2) model only the subset of local mainshocks that precedes each potential trigger, and 3) model several more or several fewer local mainshocks. None of these alternative tests results in a qualitative change in the joint statistics for any region, trigger, or subset of the triggering matrix, and we do not include these tests in detail here.

The optimal $\mu$, $c$, $p$, and $k_i$ (Eq. 4) are found using the maximum likelihood objective function of [Ogata, 1992] and the Matlab constrained optimization routine $fmincon$, to enforce positivity of the parameters. The productivity terms $k_i$ are not assumed to depend on magnitude through a productivity law as in traditional ETAS models [Ogata, 1992]. Instead, each $k_i$ is treated as a free parameter. This is superior to an $a\ priori$ productivity law because the relatively small epicentral target windows tend to crop aftershock sequences in space, disrupting the usual magnitude-productivity scaling. Our method simply attributes the triggered rate increase within the target region to the first event of
the sequence (mainshock or aftershock) that happens to fall within the target region. This method produces robust inversions as long as the number of \( k_i \) remains small. The optimal parameters for each earthquake catalog, as well as the proportion of aftershocks in each target region, are reported in Appendix C.

As an example, Figure 6 shows the observed and simulated catalogs for the epicentral region of the 2004 Sumatra earthquake, which includes the effects of three regional \( M_w \geq 8 \) trigger earthquakes as well as three moderate local mainshocks.

To compute significance of the observed \( \bar{r} \), we generate 100,000 random catalogs based on the model rate, and then draw simulated samples of inter-event time ratios \( r \) using the same time windows and sample sizes as present in the real catalog. This leads to a model distribution of values for the expected mean, denoted \( \bar{r}_{\text{sim}} \). Significance of the real \( \bar{r} \) is determined as the proportion of modeled \( \bar{r}_{\text{sim}} \) that are smaller than or equal to \( \bar{r} \), i.e.

\[
p_r = F_{\text{sim}}(\bar{r}) \quad \text{(Figure 7).}
\]

When computing the significance of local triggering for the regional \( M_w \geq 8 \) event itself, or for other nearby great earthquakes that contribute to the model rate, we remove these particular \( k_i \) from the modeled seismicity rate.

As discussed in section 2.1, we also estimate a triggered rate change \( \delta \lambda \) based on the measured \( \bar{r} \). However, we want to avoid mapping the contribution from catalog non-uniformity into the rate change. We therefore use a corrected value \( \bar{r}_{\text{corr}} \), defined as the value of \( \bar{r} \) that has the same significance with respect to a uniform Poisson process that the measured \( \bar{r} \) has with respect to the empirical distribution.

\[
F_{\text{Pois}}(\bar{r}_{\text{corr}}) = F_{\text{sim}}(\bar{r}), \tag{5}
\]

where \( F_{\text{Pois}} \) is the cumulative distribution function of \( \bar{r} \) for a uniform Poisson process, and \( F_{\text{sim}} \) is the model cumulative distribution function from the simulations (Figure 7).

Average rate change \( \delta \lambda_{50} \) and confidence bounds \( \delta \lambda_{05} \) and \( \delta \lambda_{95} \) can be calculated from \( \bar{r}_{\text{corr}} \) using the simple Poissonian step-change model (Appendix A). Note that the model distribution in Figure 7 is only applicable for this particular subgroup of the data. The model distribution of \( \bar{r}_{\text{sim}} \) must be recalculated for each particular subgroup.
5 Results: non-stationary Omori reference

5.1 Global ISC catalog

We again apply the inter-event time statistic to the global ISC catalog, but this time compute significance with respect to the non-stationary Omori reference model that accounts for aftershocks of major events. The post-seismic triggering signal reflected by the statistic \( p_r = 0.07 \), Table 3) is essentially eliminated using the more accurate reference model \( p_r = 0.40 \), Table 4). The change in reference model does not significantly alter the other triggering statistics, however, and the triggering matrix appears qualitatively similar (Figure 8).

We now discuss the information presented by the ISC triggering matrix in more detail (Figure 8). The sensitivity of the triggering tests at individual sites is very low, due to the small target regions and poor catalog completeness at most sites. Using Eq. 2 to estimate the expected rate change for each trigger-target pair beyond 10° distance (mean \( \delta \lambda_{\text{exp}} = 9\% \)), we find that in only a single case do we have sufficient data to exclude the small rate changes anticipated by Eq. 2 in the approximately one-degree radius around the regional epicenter. Even a rate increase of 67% (chosen as a representative “large” rate increase) cannot be excluded in any individual case. The sensitivity of the statistic is improved however, by looking at the combined statistics for all triggers at a particular region, or the combined statistics of a particular trigger at all regions. Trigger-target pairs separated by less than 10° are excluded from this combined analysis. We also exclude the 2004 Sumatra earthquake when calculating joint triggering significance, because triggering by the Sumatra earthquake is already reflected in the data for the 2004 Macquarie earthquake, due to the overlapping target time windows (Section 2.2).

The joint triggering confidence for each trigger is depicted in the first column of the triggering matrix (Figure 8). We should be able to resolve joint rate changes of 67% for about 1/2 of the triggers included in the ISC catalog (yellow boxes). The 2006 Kuril Is. quake is identified as a likely trigger with confidence 97.9% at the site of the preceding 2006 Tonga quake, and joint significance \( p_r = 0.016 \) over all regions. However, this
detection success is not significant when we consider the number of detection attempts, with a 21% chance of finding such an extreme significance level by chance, for at least 1 trigger out of 15. We also detect triggering in the converse direction, i.e. by the 2006 Tonga at the site of the 2006 Kuril earthquake, with confidence 99.9%. This is the highest single significance level for any distant trigger-target pair in this study, with only a 12% probability of occurring by chance, given the number of detection attempts.

The joint triggering confidence for each region as a whole is depicted in the bottom row (Figure 8). We do not detect significant triggering for any region and can rule out a 67% rate increase for half of the regions.

We now break the triggering matrix down into precursory and post-regional earthquake subgroup. The bottom right triangle of the triggering matrix is the place to look for precursory triggering. Triggering is detected above 90% confidence at 5 sites beyond 10º distance (Figure 8). (Recall that the Macquarie 2004 and Sumatra 2004 detection successes at the site of the 2010 Chile quake count as a single case (Section 2.2).) These 5 cases of distant triggering cannot be considered significant in the context of multiple detection attempts, however, with $p_{90} = 0.82$ (Table 4).

In the post-seismic, upper left triangle of the triggering matrix, we identify 4 instances of remote triggering. These 4 cases are not significant given the number of detection attempts ($p_{90} = 0.76$).

The estimated rate change (and 95% confidence bounds) for the greater than 10º distant group as a whole (pre and post-seismic) is $\delta \lambda = -1%$ (-14, 15). In the precursory group, the best estimate for triggered rate change is $\delta \lambda = -4%$ (-21, 17). In the post-regional earthquake group, the best estimate for triggered rate change is $\delta \lambda = 4%$ (-17, 29). We therefore conclude that the ISC catalog contains no evidence of systematic earthquake triggering by distant great earthquakes at the site of other great earthquakes.

5.2 Global PDE catalog
The triggering calculations are carried out again for the global PDE catalog (Figure 9). This catalog is less complete than the ISC, with less robust statistics, but it does capture the two most recent triggers. We consider the results of the ISC catalog to be definitive where they replace PDE data. The PDE catalog reveals a pattern similar to the ISC catalog. The statistics are summarized in Figure 9 and Table 5.

The PDE data also do not show significant triggering for quakes > 10° as a whole, with \( p_r = 0.83 \), nor for any of the precursory or post-regional earthquake subgroups. In fact, there are fewer detection successes in the precursory subgroup (lower right triangle, Figure 9) than expected purely by chance (\( p_{90} = 0.9 \)). In the post-seismic subgroup (upper right triangle), the Tohoku-Oki earthquake appears to have triggered in 3 locations at above 90% confidence (Fig, 9), but the joint statistics for this event are again not significant (\( p_{90} = 0.09, p_r = 0.24 \)).

Computing the joint significance region-by-region, we detect apparent triggering in both the 2006 Tonga and 2006 Kuril epicentral regions, but the significance of this number of detections is low when we consider the number of detection attempts (\( p_{90} = 0.45 \)). We can rule out a 67% rate increase for one-third of the target regions.

The estimated rate change (and 95% confidence bounds) for the greater than 10° distant group as a whole is \( \delta \lambda = -9\% \) (-23, 7). In the precursory group, the best estimate for triggered rate change is \( \delta \lambda = -35\% \) (-50, -15). The statistics therefore indicate systematic triggered rate decreases in this subgroup. This is probably a catalog artifact, as this observation is not reproduced in the ISC or JMA catalogs. In the post-regional earthquake group, the best estimate for triggered rate change is \( \delta \lambda = 11\% \) (-11, 38). We therefore conclude that the PDE catalog also contains no evidence for systematic earthquake triggering by distant great earthquakes at the site of other impending great earthquakes.

5.3 JMA catalog
Finally, we examine the JMA catalog, which is the most complete in its region of coverage. The JMA catalog covers only the regions of the 2004 Tokachi-Oki and 2011 Tohoku-Oki epicenters, where it is complete to magnitude $M_c = 1.3$.

Triggering at sites greater than $10^\circ$ distant is only observed above 90% confidence for a single trigger: 2001 Peru, but at both target sites in Japan (Figure 10). However, this detection success is not significant given the number of detection attempts ($p_{90} = 0.79$). Neither the precursory nor post-regional earthquake subgroups show significant triggering, with an estimated rate change very near zero for each (Table 6). For the precursory group we get an estimated rate change (with 95% confidence bounds) $\delta \lambda = 0\% (-6, 7)$.

Thanks to the higher completeness of the JMA catalog, we can say with 95% confidence that the triggered rate change in both target regions in Japan is smaller than that expected from Eq. 2, based on triggering in California. This confirms previous studies that have found Japan to be less triggerable than California [Harrington and Brodsky, 2006; van der Elst and Brodsky, 2010].

6. Discussion

6.1 Triggering at sites of impending large earthquakes.

The rate of recent great $M_w \geq 8$ earthquakes has increased with respect to the preceding century by 50 - 160%, depending on the time window and the method used to remove conventional aftershocks. Using the ISC catalog down to $M 3.7$, the best estimate for the systematic rate increase at the site of impending great earthquakes due to preceding great earthquakes is $\delta \lambda = -4\%$ with lower and upper 95% confidence bounds of -21% and 17%.

We can confidently conclude that there is no evidence for large or systematic triggered rate increases at the sites of impending great $M_w \geq 8$ earthquakes. If a long-range triggering process is to explain the occurrence of some great earthquakes, it must do so only rarely, or without systematically affecting the rates of earlier, smaller earthquakes at the time of the trigger.
It is more difficult to rule out triggering on a less systematic, case-by-case basis, as even doublings of rate near the epicenters of most impending earthquakes cannot be excluded with high confidence due to the scarcity of data in most places. Only the JMA catalog has adequate coverage to detect systematic triggering at the level expected from the California productivity scaling for a few cases (Figure 10), and where this coverage exists, triggering is not detected.

6.2 Magnitude of completeness required to detect small rate changes

We are reasonably confident that the earthquake catalogs do not show systematic large triggered rate changes capable of explaining recent large earthquake rates. What about smaller sequences, like those expected from other continental triggering studies, that might tell us something about earthquake processes? We now use the framework developed in this paper to estimate the magnitude of completeness needed to routinely identify small triggered sequences in subduction zones, where most great earthquakes occur. In most cases, the current network coverage is insufficient to detect the small rate changes expected from established scaling laws (Eq. 2). Figure 11 shows the corrected inter-event time statistic $\bar{r}_{corr}$ (Eq. 5), along with the expected value of $\bar{r}$ (Eq. A8) corresponding to the rate change $\delta \lambda$ expected from Eq. 2, as well as the value of $\bar{r}$ required to exceed the 90% confidence threshold given the available number of observations in the ISC catalog. In no case is the number of observations (catalog completeness level) adequate to establish the expected rate change at 90% confidence.

We now estimate the number of observations that would be required to consistently detect rate changes expected from the California scaling (Eq. 2). If we assume that earthquakes follow the Gutenberg-Richter magnitude frequency scaling with $b = 1$, this required sample size can be expressed in terms of the required completeness magnitude. We estimate the required catalog completeness level for 1) the catalog as a whole, 2) each target region as a whole, and 3) for each individual trigger-target pair.

The mean expected rate change, averaged over all distant trigger-target pairs, is $\delta \lambda_{exp} = 9\%$. To have 95% confidence that this rate change would produce a positive signal in the
ISC catalog (i.e. an interevent time statistic $\bar{r} < 0.5$), we would need only a relatively marginal improvement in the magnitude of completeness from 3.7 to 3.5.

On a regional basis, the required magnitude of completeness is dependent on the average seismicity rate, with less active regions requiring a proportionally lower magnitude of completeness to produce the same sample size (Table 7). The required magnitude of completeness for detecting the median expected rate change on a joint regional basis is 2.0. This completeness level is obtained only by the JMA catalog.

On a case-by-case basis (Figure 11), the magnitude of completeness would have to be considerably lower. The median required magnitude of completeness, i.e. the level required to detect distant triggering at half of all sites (median rate change $\Delta \lambda_{exp} = 7\%$) is $M_c = 0.6$. This level of completeness is likely impossible without dense offshore networks of seismic stations.

### 6.3 Limitations

Great earthquake interactions are only considered here in terms of local cascade dynamics. Other possibilities for linking great earthquakes might be: 1) triggering a deep slow slip event [Shelly et al., 2011] that then initiates a great earthquake at the down dip extent of locking; or 2) viscous stress transfer in the lower crust and upper mantle, operating on a much longer timescale than the passage of seismic waves [Freed and Lin, 1998; Pollitz et al., 1998]. A careful search for remotely triggered tectonic tremor down-dip of impending earthquakes could evaluate whether great earthquakes are linked through triggered deep slip.

We have looked only at the largest magnitude potential triggers. The largest triggers are the most likely places to look for a response, but smaller magnitude triggers may be cumulatively just as important in triggering distant rate changes, due to their proportionally greater number [Helmstetter et al., 2005]. Triggering may also be influenced by precise stress orientations of dynamic waves, which vary with fault geometry and relative location of trigger events and target nucleation zones [Hill, 2008].
In some cases, a smaller nearby event may be more efficient for triggering at a particular target.

7. Conclusion

We have systematically measured triggered rate changes at the sites of 16 $M_W \geq 8$ earthquakes, at the times of other $M_W \geq 8$ earthquakes, during the high-great earthquake rate period between 1998 and 2011. We find no evidence that $M_W \geq 8$ earthquakes trigger precursory activity at the site of other impending $M_W \geq 8$ earthquakes, and systematic triggering on a scale capable of explaining a 50-160% increase in the rate of $M_W \geq 8$ earthquakes can be ruled out with 95% confidence.

The data are, however, consistent with expected rate changes given the amplitude of seismic shaking at the target sites, although an improvement in completeness magnitude from 3.7 to 3.5 would be required to definitively establish small rate changes.

There is no apparent increase in regional triggering susceptibility that might serve as a precursory signal of impending earthquakes, at least using current catalogs. However, regional networks with a magnitude of completeness down to 2.0 would be required to adequately resolve expected triggered rate changes on a region-by-region basis.

Data and Resources

Several earthquake catalogs were used in this study. ISC data through 2007 were obtained on CDs from the International Seismological Centre, and supplemented with data through 11/2009 via ftp from www.isc.ac.uk (last accessed December 15, 2011). The PDE catalog was downloaded via ftp from the United States Geological Survey website, http://earthquake.usgs.gov/research/data/pde.php (last accessed November 14, 2011). The JMA catalog was manually copied-and-pasted in 7-day increments from the Japan Meteorological Agency website at www.hinet.bosai.go.jp (last accessed November 22, 2011, account required). The USGS PAGER catalog, used to define the list of $M_W \geq 8$ earthquakes, is available from earthquake.usgs.gov/research/data. Seismograms were
obtained from the IRIS Data Management Center at www.iris.edu and processed in SAC.
Many seismographic networks contributed data through IRIS.

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References


Peng, Z. G., D. P. Hill, D. R. Shelly, and C. Aiken (2010), Remotely triggered microearthquakes and tremor in central California following the 2010 M(w) 8.8 Chile earthquake, *Geophysical Research Letters, 37*


Table 1: List of great earthquakes

<table>
<thead>
<tr>
<th>Date</th>
<th>Name</th>
<th>M_W</th>
<th>T (days)</th>
</tr>
</thead>
<tbody>
<tr>
<td>25-03-1998</td>
<td>Balleny Is.</td>
<td>8.1</td>
<td>83</td>
</tr>
<tr>
<td>16-11-2000</td>
<td>New Ireland</td>
<td>8.0</td>
<td>220</td>
</tr>
<tr>
<td>23-06-2001</td>
<td>Peru</td>
<td>8.4</td>
<td>220</td>
</tr>
<tr>
<td>25-09-2003</td>
<td>Tokachi-Oki</td>
<td>8.3</td>
<td>455</td>
</tr>
<tr>
<td>23-12-2004</td>
<td>Macquarie Is.</td>
<td>8.1</td>
<td>95</td>
</tr>
<tr>
<td>26-12-2004</td>
<td>Sumatra-Andaman</td>
<td>9.0</td>
<td>93</td>
</tr>
<tr>
<td>28-03-2005</td>
<td>Sumatra-Nias</td>
<td>8.6</td>
<td>93</td>
</tr>
<tr>
<td>03-05-2006</td>
<td>Tonga</td>
<td>8.0</td>
<td>196</td>
</tr>
<tr>
<td>15-11-2006</td>
<td>Kuril Is.</td>
<td>8.3</td>
<td>59</td>
</tr>
<tr>
<td>13-01-2007</td>
<td>Kuril Is.</td>
<td>8.1</td>
<td>59</td>
</tr>
<tr>
<td>01-04-2007</td>
<td>Solomon Is.</td>
<td>8.1</td>
<td>78</td>
</tr>
<tr>
<td>15-07-2007</td>
<td>Central Peru</td>
<td>8.0</td>
<td>27</td>
</tr>
<tr>
<td>12-09-2007</td>
<td>So. Sumatra</td>
<td>8.5</td>
<td>27</td>
</tr>
<tr>
<td>29-09-2009</td>
<td>Samoa</td>
<td>8.0</td>
<td>151</td>
</tr>
<tr>
<td>27-02-2010</td>
<td>Chile</td>
<td>8.8</td>
<td>151</td>
</tr>
<tr>
<td>11-03-2011</td>
<td>Tohoku-Oki</td>
<td>9.0</td>
<td>92</td>
</tr>
</tbody>
</table>

* Maximum allowed $t_1$ or $t_2$ in Eq. 1 based on the timing of other $M_w \geq 8$ earthquakes.

The total trigger time window is $2T$, centered on each great event.
Table 2: Sample size required to detect a triggered rate change

<table>
<thead>
<tr>
<th>Rate change $\delta \lambda$ [%]</th>
<th>Expected $\bar{r}$</th>
<th>Required sample size $m$ for 90% (95%) confidence</th>
</tr>
</thead>
<tbody>
<tr>
<td>1</td>
<td>0.498</td>
<td>49.765 (81.980)</td>
</tr>
<tr>
<td>5</td>
<td>0.492</td>
<td>2.071 (3.411)</td>
</tr>
<tr>
<td>10</td>
<td>0.484</td>
<td>543 (895)</td>
</tr>
<tr>
<td>67</td>
<td>0.415</td>
<td>20 (32)</td>
</tr>
<tr>
<td>100</td>
<td>0.386</td>
<td>11 (18)</td>
</tr>
<tr>
<td>160</td>
<td>0.345</td>
<td>6 (10)</td>
</tr>
</tbody>
</table>

Table 3. Statistical significance: ISC catalog – Poisson reference

<table>
<thead>
<tr>
<th>$r$</th>
<th>$m^a$</th>
<th>$p_r$</th>
<th>$p_{50}$</th>
<th>$p_{90}$</th>
<th>$\delta \lambda_{50}(\delta \lambda_{95,95})$ [%]</th>
</tr>
</thead>
<tbody>
<tr>
<td>All</td>
<td>0.379</td>
<td>621</td>
<td>$&lt;10^{-16}$</td>
<td>0.10</td>
<td>0.04 110 (87, 136)</td>
</tr>
<tr>
<td>All &gt;10</td>
<td>0.486</td>
<td>371</td>
<td>0.17</td>
<td>0.23</td>
<td>0.85 9 (-6, 26)</td>
</tr>
<tr>
<td>Pre</td>
<td>0.500</td>
<td>206</td>
<td>0.50</td>
<td>0.55</td>
<td>0.92 0 (-18, 22)</td>
</tr>
<tr>
<td>Post</td>
<td>0.467</td>
<td>165</td>
<td>0.07</td>
<td>0.13</td>
<td>0.59 22 (-3.52)</td>
</tr>
</tbody>
</table>

*Sample size $m$ is the combined total number of local earthquake pairs ($t_1$, $t_2$) in each subgroup, which can be greater than the number of great earthquake trigger-target pairs in the subgroup.*

Table 4. Statistical significance: ISC catalog – Omori reference

<table>
<thead>
<tr>
<th>$\bar{r}$</th>
<th>$\bar{r}_{corr}$</th>
<th>$m^a$</th>
<th>$p_r$</th>
<th>$p_{50}$</th>
<th>$p_{90}$</th>
<th>$\delta \lambda_{50}(\delta \lambda_{95,95})$ [%]</th>
</tr>
</thead>
<tbody>
<tr>
<td>All</td>
<td>0.379</td>
<td>0.423</td>
<td>621</td>
<td>$&lt;10^{-16}$</td>
<td>0.13</td>
<td>0.02 59 (42, 79)</td>
</tr>
<tr>
<td>All &gt;10</td>
<td>0.486</td>
<td>0.501</td>
<td>371</td>
<td>0.53</td>
<td>0.36</td>
<td>0.85 -1 (-14, 15)</td>
</tr>
<tr>
<td>Pre</td>
<td>0.500</td>
<td>0.506</td>
<td>206</td>
<td>0.62</td>
<td>0.64</td>
<td>0.82 -4 (-21, 17)</td>
</tr>
<tr>
<td>Post</td>
<td>0.467</td>
<td>0.494</td>
<td>165</td>
<td>0.40</td>
<td>0.20</td>
<td>0.76 4 (-17, 29)</td>
</tr>
<tr>
<td>1st pre</td>
<td>0.483</td>
<td>0.417</td>
<td>16</td>
<td>0.12</td>
<td>0.77</td>
<td>0.52 66 (-19, 233)</td>
</tr>
<tr>
<td>1st post</td>
<td>0.425</td>
<td>0.464</td>
<td>33</td>
<td>0.24</td>
<td>0.34</td>
<td>0.47 24 (-24, 103)</td>
</tr>
</tbody>
</table>

*Sample size $m$ is the combined total number of local earthquake pairs ($t_1$, $t_2$) in each subset, which can be greater than the number of great earthquake trigger-target pairs in the subset.*
### Table 5. Statistical significance: PDE catalog – Omori reference

<table>
<thead>
<tr>
<th></th>
<th>$\bar{r}$</th>
<th>$\bar{r}_{corr}$</th>
<th>$m$</th>
<th>$p_r$</th>
<th>$p_{50}$</th>
<th>$p_{90}$</th>
<th>$\delta \lambda_{50} (\delta \lambda_{95})$ [%]</th>
</tr>
</thead>
<tbody>
<tr>
<td>All</td>
<td>0.399</td>
<td>0.432</td>
<td>467</td>
<td>$&lt;10^{-16}$</td>
<td>0.71</td>
<td>0.01</td>
<td>50 (32, 72)</td>
</tr>
<tr>
<td>All &gt;10</td>
<td>0.502</td>
<td>0.517</td>
<td>286</td>
<td>0.83</td>
<td>0.97</td>
<td>0.40</td>
<td>-9 (-23, 7)</td>
</tr>
<tr>
<td>Pre</td>
<td>0.564</td>
<td>0.572</td>
<td>113</td>
<td>1.00</td>
<td>1.00</td>
<td>0.90</td>
<td>-35 (-50, -15)</td>
</tr>
<tr>
<td>Post</td>
<td>0.461</td>
<td>0.483</td>
<td>173</td>
<td>0.21</td>
<td>0.50</td>
<td>0.13</td>
<td>11 (-11, 38)</td>
</tr>
<tr>
<td>1st pre</td>
<td>0.518</td>
<td>0.478</td>
<td>15</td>
<td>0.38</td>
<td>0.86</td>
<td>0.57</td>
<td>14 (-45, 135)</td>
</tr>
<tr>
<td>1st post</td>
<td>0.474</td>
<td>0.506</td>
<td>31</td>
<td>0.54</td>
<td>0.86</td>
<td>0.57</td>
<td>-3 (-42, 61)</td>
</tr>
</tbody>
</table>

### Table 6. Statistical significance: JMA catalog – Omori reference

<table>
<thead>
<tr>
<th></th>
<th>$\bar{r}$</th>
<th>$\bar{r}_{corr}$</th>
<th>$m$</th>
<th>$p_r$</th>
<th>$p_{50}$</th>
<th>$p_{90}$</th>
<th>$\delta \lambda_{50} (\delta \lambda_{95})$ [%]</th>
</tr>
</thead>
<tbody>
<tr>
<td>All</td>
<td>0.473</td>
<td>0.466</td>
<td>3958</td>
<td>$&lt;10^{-16}$</td>
<td>0.70</td>
<td>0.21</td>
<td>23 (18, 29)</td>
</tr>
<tr>
<td>All &gt;10</td>
<td>0.496</td>
<td>0.499</td>
<td>2534</td>
<td>0.42</td>
<td>0.80</td>
<td>0.41</td>
<td>1 (-5, 7)</td>
</tr>
<tr>
<td>Pre</td>
<td>0.497</td>
<td>0.499</td>
<td>2090</td>
<td>0.46</td>
<td>0.25</td>
<td>0.23</td>
<td>0 (-6, 7)</td>
</tr>
<tr>
<td>Post</td>
<td>0.494</td>
<td>0.494</td>
<td>444</td>
<td>0.33</td>
<td>0.97</td>
<td>0.77</td>
<td>4 (-9, 19)</td>
</tr>
<tr>
<td>Region</td>
<td>Median number of pairs in catalog</td>
<td>Number of pairs required to resolve rate change $\delta \lambda_{\text{exp}}$</td>
<td>ISC completeness magnitude $M_C$</td>
<td>$M_C$ required to resolve rate change $\delta \lambda_{\text{exp}}$</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>---------------</td>
<td>---------------------------------</td>
<td>-------------------------------------------------</td>
<td>-----------------</td>
<td>-----------------</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>Balleny Is.</td>
<td>6</td>
<td>1596</td>
<td>3.7</td>
<td>0.5</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>New Ireland</td>
<td>7</td>
<td>959</td>
<td>3.7</td>
<td>2.2</td>
<td></td>
<td></td>
<td></td>
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<tr>
<td>Peru</td>
<td>8</td>
<td>842</td>
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<td><strong>3.5</strong></td>
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</table>
Figure Captions

Figure 1. Locations of $M_w \geq 8$ great earthquakes used in this study. Crosses mark locations; numbers give the order of occurrence (Table 1).

Figure 2. Peak surface wave velocity and expected rate change $\delta \lambda_{\exp}$ (Eq. 2) for each $M \geq 8$ earthquake, at each target site. The legend gives the earthquake number from Table 1 (e.g. 1: 1998 Balleny Islands, 6: 2004 Sumatra-Andaman, 16: 2011 Tohoku-Oki). The black curve shows the empirical surface wave magnitude equation, multiplied by $2\pi/T$ (period $T = 20$ sec) to obtain velocity, for a $M_S$ 8.2 earthquake (the median trigger magnitude in this study). The dashed line shows the 10º limit used to define the farfield. The increase in amplitude beyond 150º is due to antipodal focusing.

Figure 3. The Tohoku-Oki target region for all 16 triggers (numbered in the bottom right of each panel). The top left panel is the context map. The crosshairs mark the epicenter of the future Tohoku-Oki quake, and the polygon outlines the first 10 days of Tohoku-Oki aftershock activity within ~1º of the epicenter. The region is gridded on a fine scale (optimized for each trigger), and each colored dot represents one pair of earthquakes within that grid – one before and one after the trigger time. The color corresponds to the $r$ value (Eq. 1); warmer colors indicate larger triggered rate change.

Figure 4. Triggering matrix for the ISC catalog, with significance measured with respect to a stationary, uniform Poisson process. The triggering matrix shows regions with rate increases above 90% confidence (grayscale). Triggers are sorted by date along the y-axis, and sorted as target regions on the x-axis. Yellow squares show where a rate increase of 67% could theoretically be detected/excluded. Red squares mark events within 10º of each other. The left-most column gives the combined triggering confidence for each trigger over all target regions (excluding those in red squares), and the bottom row shows the combined confidence for each region over all triggers (excluding red squares). The bottom left square gives the triggering confidence for the entire catalog. Crosses indicate no data for that trigger-target pair.
**Figure 5.** Illustration of the effect of early aftershocks on apparent triggering as measured by the inter-event time statistic $r$. The grey line is the cumulative number of earthquakes in a simulated catalog containing 50% background events and 50% aftershocks decaying according to Omori’s law. The squares show measurements of $\overline{r}$ for hypothetical distant earthquakes occurring at 100 day intervals. The simulated target window is $2T = 200$ days long. The simulation contains no long-range triggering effects, therefore the skew downward of $\overline{r}$ near the beginning of the aftershock sequence, and associated skew upwards of the apparent triggering confidence (circles), are effects of the aftershock sequence itself on the statistic.

**Figure 6.** Cumulative observed seismicity (ISC catalog) in the 2004 Sumatra target region (black line) along with 1000 simulated time series (gray lines) generated from the non-stationary Omori model (Eq. 4). The times of all the $M_w \geq 8$ potential triggers are shown as large crosses. The ISC catalog extends only to 11/2009. The local 2004 $M_w 9$ Sumatra earthquake and the nearby 2005 Sumatra-Nias and 2007 Southern Sumatra earthquakes are labeled by trigger number (Table 1). Aftershocks from these three great triggers, as well as three significant local earthquakes, evident as small jumps in seismicity rate, are included in the modeled local seismicity rate.

**Figure 7.** Modeled cumulative density function (cdf) of $\overline{r}_{lim}$ based on a non-stationary Omori model rate (Eq. 4) that includes aftershocks of major regional and local events (solid black line). The dashed line shows the theoretical distribution for a stationary Poisson model (Appendix A). Inset: probability density functions. This particular example is for all trigger-target pairs greater than $10^\circ$ distant (ISC catalog). The arrows show the transformation from the observed statistic $\overline{r}$ to the corrected statistic $\overline{r}_{corr}$ (Eq. 5).

**Figure 8.** Triggering matrix for the ISC catalog, using the non-stationary Omori reference model. (Compare with Figure 4, using the uniform Poisson reference model.) The triggering matrix shows regions with rate increases above 90% confidence (grayscale). Triggers are sorted by date along the y-axis, and sorted as target regions on
the x-axis. Yellow squares show where a rate increase of 67% could theoretically be
detected/excluded. Red squares mark events within 10 degrees of each other. The left-
most column gives the combined triggering confidence for each trigger over all target
regions (excluding those in red squares), and the bottom row shows the combined
confidence for each region over all triggers (excluding red squares). The bottom left
square gives the triggering confidence for the entire catalog. Crosses indicate no data for
that trigger-target pair.

Figure 9. Triggering matrix for the PDE catalog, using a non-stationary Omori reference
model. See Figure caption 8 for details.

Figure 10. Triggering matrix for JMA catalog, using a non-stationary Omori reference
model, showing the Tokachi-Oki and Tohoku-Oki regions only. Green squares show
where the expected rate change (Eq. 2) should be resolvable, and yellow squares show
where a 67% rate change should be resolvable. See caption of Figure 9 for other features.

Figure 11. Observed $\bar{r}_{corr}$ (filled circles) compared with the $\bar{r}$ needed to exceed the 90%
confidence threshold given the number of samples (open circles), for trigger-target pairs
greater than 10º distance, as a function of trigger strain. Filled squares are cases where the
observed rate change exceeds 90% confidence. The dashed line shows the expected $\bar{r}$
and associated rate change from Eq. 2. In no case is the expected rate change (dashed
line) above the threshold for significance (open circles) for isolated trigger-target pairs.
Figures.

Figure 1.

Figure 2.
Figure 3
Figure 4.

Figure 5.
Figure 6.

Figure 7.
Figure 8.

Figure 9.
Figure 10.

Figure 11.
Appendices

Appendix A: Distribution of the sample mean of $r$

We use the sampling distribution of the mean inter-event time ratio $\overline{r}$ to calculate the significance of $\overline{r}$ with respect to a Poisson process, as well as to find the minimum rate increase at which the Poisson process could be rejected. The distribution of $\overline{r}$ in the case of zero rate change is known, as $r$ is uniform on $[0,1]$. The distribution of the sum $s = m\overline{r}$ of a sample of uniform distributed random variables is [Uspensky, 1937]

$$f_s(s|m) = \sum_{k=0}^{m} (-1)^k \frac{m}{k!(m-k)!} (s-k)^{m-1} u(s-k).$$ \hspace{1cm} (A1)

Here $u$ is the unit step function. In this Appendix, we use the convention that distributions are subscripted by their applicable variable in uppercase. The distribution of the mean is the distribution (A1) rescaled over the range $[0,1]$, i.e.

$$f_R(\overline{r}) = mf_s(m\overline{r}).$$ \hspace{1cm} (A2)

$$f_R(\overline{r}|m) = \sum_{k=0}^{m} (-1)^k \frac{m^2}{k!(m-k)!} (m\overline{r}-k)^{m-1} u(m\overline{r}-k).$$ \hspace{1cm} (A3)

This distribution very rapidly approaches the normal distribution as $m$ increases (Figure A1).

The distribution of $\overline{r}$ for an arbitrary rate change $\delta \lambda$ is required to compute confidence bounds, i.e. the minimum and maximum rate changes that could pass the Poisson hypothesis test. Van der Elst and Brodsky [2010, Appendix A] derived the expectation of $r$ for a step change in a Poisson process. Here we give the full distribution function.

The distribution of $r$ (defined in Eq. 1, main text) with $t_1$ and $t_2$ drawn from distinct Poisson processes with intensities $\lambda_1$ and $\lambda_2$, respectively, is

$$f_R(r|\lambda_1, \lambda_2) = \frac{\lambda_1 \lambda_2}{[(1-r)\lambda_1 + r\lambda_2]^2}.$$ \hspace{1cm} (A4)

This equation follows from transforming the joint probability distribution of two independent, exponentially distributed times $t_1$ and $t_2$, with means $\lambda_1^{-1}$ and $\lambda_2^{-1}$, into a function of the random variable $r$ [van der Elst and Brodsky, 2010].
Defining fractional rate change as
\[ \delta \lambda \equiv \frac{\lambda_2 - \lambda_1}{\lambda_1}, \]  
(A5)
and substituting (A5) into (A4) gives
\[ f_R(r|\delta \lambda) = \frac{\delta \lambda + 1}{(1 + \delta \lambda \cdot r)^2}. \]  
(A6)

The distribution of the sample mean \( \bar{r} \) (i.e., the first moment) as a function of fractional rate change \( \delta \lambda \) and sample size \( m \) cannot be obtained analytically, as the moment generating function for (A6) is expressible only with special functions. For small samples \( (m < 10) \), we use numerical convolution to obtain the distribution and cumulative distribution of the sum \( s \) [Grinstead and Snell, 1997], and then rescale this distribution (Eq. A2) to obtain the distribution of the mean \( \bar{r} \).

For larger samples \( (m > 10) \), we take advantage of the central limit theorem to obtain an analytical approximation of the distribution of the mean. The central limit theorem guarantees that the distribution of the mean of any random variable with mean \( \mu \) and finite variance \( \sigma^2 \) approaches a normal distribution with mean \( \mu \) and variance \( \sigma^2 / m \) as sample size \( m \) increases [Casella and Berger, 2002]. The mean of \( r \) as a function of \( \delta \lambda \) can be found by computing the expectation from (A6):
\[ \mu_r(\delta \lambda) = \int_0^1 r f_R(r|\delta \lambda) dr. \]  
(A7)
\[ \mu_r(\delta \lambda) = \frac{1}{\delta \lambda^2} \left[ (\delta \lambda + 1) \ln(\delta \lambda + 1) - \delta \lambda \right], \]  
(A8)
and the variance (expressed as a function of the mean for simplicity) is
\[ \sigma_r^2(\delta \lambda) = \frac{1}{\delta \lambda} \left[ 1 - 2 \mu_r \right] - \mu_r^2. \]  
(A9)
In the limit of \( \delta \lambda = 0 \), the mean \( \mu_r \) and variance \( \sigma_r^2 \) can be shown to be 1/2 and 1/12, respectively, consistent with a uniform distribution on [0,1]. We find that the normal approximation using (A8) and (A9) is very good for \( m > 10 \) and \( \delta \lambda < 1 \).
The significance of \( \bar{r} \) is defined as the probability \( p_r \) that a uniform random process would produce a smaller mean than that observed, by chance:

\[
p_r \equiv F_{\bar{r}}(\bar{r}),
\]

(A10)

where \( F_{\bar{r}} \) is the cumulative distribution function. Using the normal approximation with (A8) and (A9),

\[
F_{\bar{r}}(\bar{r}) \equiv \Phi\left[ \frac{\bar{r} - \mu_r(0)\sigma_r^2(0)}{\sigma_r^2(0)/m} \right],
\]

(A11)

Where \( \Phi \) is the cumulative normal distribution.

Inverting the statistic (A10) gives the confidence bounds on the rate change parameter \( \delta \lambda \) [Casella and Berger, 2002]. The maximum bound on the rate change \( \delta \lambda \), at significance level \( \alpha \) (e.g. 0.05), is the value of \( \delta \lambda \) for which

\[
\alpha = F_{\bar{r}}(\delta \lambda | \bar{r}) \equiv \Phi\left[ \frac{\bar{r} - \mu_r(\delta \lambda)\sigma_r^2(\delta \lambda)}{\sigma_r^2(\delta \lambda)/m} \right].
\]

(A12)

Given \( \bar{r}, m, \) and \( \alpha \), we solve iteratively for \( \delta \lambda \). To obtain the minimum bound on \( \delta \lambda \), we solve (A12) for \( \alpha' = 1 - \alpha \).

**Figure A1** Distribution of the sample mean \( \bar{r} \) for a stationary Poisson process (Eq. A3), for various sample sizes \( m \). Solid lines are the exact solution given by (Eq. A3), and the dashed lines show the approximation using a normal distribution.
Appendix B: Analysis of PDE and JMA catalogs using uniform Poisson reference

Figure B1. Triggering matrix for the PDE catalog, with significance measured with respect to a stationary, uniform Poisson process. The triggering matrix shows regions with rate increases above 90% confidence (grayscale). Triggers are sorted by date along the y-axis, and sorted as target regions on the x-axis. Yellow squares show where a rate increase of 67% could theoretically be detected/excluded. Red squares mark events within 10° of each other. The left-most column gives the combined triggering confidence for each trigger over all target regions (excluding those in red squares), and the bottom row shows the combined confidence for each region over all triggers (excluding red squares). The bottom left square gives the triggering confidence for the entire catalog. Crosses indicate no data for that trigger-target pair.

Table B1. Statistical significance: PDE catalog – Poisson reference

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<tr>
<th></th>
<th>r</th>
<th>m</th>
<th>p_r</th>
<th>p_{50}</th>
<th>p_{90}</th>
<th>δλ_{50}(δλ_{45.95}) [%]</th>
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<td>2x10^{-14}</td>
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<td><strong>26</strong> (2, 57)</td>
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<td>0.31</td>
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<td>17 (-30, 94)</td>
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Figure B2. Triggering matrix for JMA catalog, using a stationary, uniform Poisson reference model, showing the Tokachi-Oki and Tohoku-Oki regions only. Green squares show where the expected rate change (Eq. 2) should be resolvable, and yellow squares show where a 67% rate change should be resolvable. See caption of Figure B1 for other features.

Table B2. Statistical significance: JMA catalog – Poisson reference

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Appendix C. Omori model parameters

This appendix contains the parameters used in the non-stationary Omori reference models (Eq. 4). It also lists the total fraction of the simulated catalog made up of aftershocks (local, regional, and target events), the fraction that is aftershocks of the target $M \geq 8$ earthquake itself, and the number of local significant mainshocks used to fit the time series.
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<th>$p$</th>
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<th>Frac. direct aftershocks</th>
<th>Num. local mainshocks$^a$</th>
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<td>0.69</td>
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<td>0.87</td>
<td>0.83</td>
<td>1</td>
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<td>0</td>
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<tr>
<td>So. Sumatra</td>
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<td>0.93</td>
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<td>Tohoku-Oki$^b$</td>
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<td>0.01</td>
<td>0.86</td>
<td>0.49</td>
<td>0.00</td>
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</tr>
</tbody>
</table>

$^a$ local mainshocks with $M < 8$, only

$^b$ Does not include the local $M \geq 8$ earthquake itself due to catalog time limit
### Table C2. PDE Catalog

<table>
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<tr>
<th>Region</th>
<th>$\mu$ [yr$^{-1}$]</th>
<th>$c$ [days]</th>
<th>$p$</th>
<th>Fraction aftershocks</th>
<th>Frac. direct aftershocks</th>
<th>Num. local mainshocks$^a$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Balleny Is.</td>
<td>0.7</td>
<td>0.19</td>
<td>1.38</td>
<td>0.49</td>
<td>0.49</td>
<td>0</td>
</tr>
<tr>
<td>New Ireland</td>
<td>8.4</td>
<td>0.04</td>
<td>0.95</td>
<td>0.42</td>
<td>0.35</td>
<td>1</td>
</tr>
<tr>
<td>Peru</td>
<td>3.8</td>
<td>0.44</td>
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<td>0.69</td>
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</tr>
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<td>Tokachi-Oki</td>
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<td>0.04</td>
<td>0.95</td>
<td>0.70</td>
<td>0.55</td>
<td>4</td>
</tr>
<tr>
<td>Macquarie Is.</td>
<td>0.1</td>
<td>0.06</td>
<td>1.21</td>
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<td>0.82</td>
<td>0.92</td>
<td>0.26</td>
<td>4</td>
</tr>
<tr>
<td>Sumatra-Nias</td>
<td>3.3</td>
<td>0.09</td>
<td>0.89</td>
<td>0.87</td>
<td>0.73</td>
<td>2</td>
</tr>
<tr>
<td>Tonga</td>
<td>12.4</td>
<td>0.92</td>
<td>1.26</td>
<td>0.57</td>
<td>0.47</td>
<td>5</td>
</tr>
<tr>
<td>Kuril Is.</td>
<td>8.1</td>
<td>0.10</td>
<td>1.01</td>
<td>0.77</td>
<td>0.42</td>
<td>3</td>
</tr>
<tr>
<td>Kuril Is.</td>
<td>1.0</td>
<td>0.58</td>
<td>1.56</td>
<td>0.94</td>
<td>0.20</td>
<td>0</td>
</tr>
<tr>
<td>Solomon Is.</td>
<td>4.4</td>
<td>0.46</td>
<td>1.40</td>
<td>0.78</td>
<td>0.36</td>
<td>5</td>
</tr>
<tr>
<td>Central Peru</td>
<td>4.8</td>
<td>0.13</td>
<td>1.18</td>
<td>0.71</td>
<td>0.69</td>
<td>1</td>
</tr>
<tr>
<td>So. Sumatra</td>
<td>1.5</td>
<td>0.04</td>
<td>0.85</td>
<td>0.90</td>
<td>0.64</td>
<td>2</td>
</tr>
<tr>
<td>Samoa</td>
<td>2.4</td>
<td>0.08</td>
<td>1.17</td>
<td>0.51</td>
<td>0.51</td>
<td>0</td>
</tr>
<tr>
<td>Chile</td>
<td>11.8</td>
<td>0.56</td>
<td>1.21</td>
<td>0.75</td>
<td>0.68</td>
<td>1</td>
</tr>
<tr>
<td>Tohoku-Oki</td>
<td>12.8</td>
<td>0.62</td>
<td>1.02</td>
<td>0.77</td>
<td>0.65</td>
<td>4</td>
</tr>
</tbody>
</table>

$^a$local mainshocks with M < 8, only

### Table C3. JMA Catalog

<table>
<thead>
<tr>
<th>Region</th>
<th>$\mu$ [yr$^{-1}$]</th>
<th>$c$ [days]</th>
<th>$p$</th>
<th>Fraction aftershocks</th>
<th>Frac. direct aftershocks</th>
<th>Num. local mainshocks$^a$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Tokachi-Oki</td>
<td>175</td>
<td>0.07</td>
<td>0.86</td>
<td>0.59</td>
<td>0.45</td>
<td>4</td>
</tr>
<tr>
<td>Tohoku-Oki$^b$</td>
<td>2037</td>
<td>0.05</td>
<td>0.94</td>
<td>0.20</td>
<td>0.00</td>
<td>2</td>
</tr>
</tbody>
</table>

$^a$local mainshocks with M < 8, only

$^b$Does not include the local M $\geq$ 8 earthquake itself because there are no subsequent triggers.