

Revealing Transient Strain in Geodetic Data with the Radial Basis Function Finite Difference Method

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1 Introduction

Crustal strain rates are fundamentally important quantities for assessing seismic hazard. For one reason, knowing where and how quickly strain is accumulating gives insight into where we can expect stored elastic energy to be released seismically. Secular crustal strain rates, based on Geologic and GNSS data, have been used to constrain seismic hazard models such as UCERF3 (Field et al., 2014). Dense networks of continuous GNSS stations, such as the Plate Boundary Observatory (PBO), have sufficient spatial and temporal coverage for us to feasibly make estimates of transient crustal strain rates, which is considered to be any deviation from the secular strain rate. Transient crustal strain rates can potentially illuminate geophysical signal, which may not be immediately apparent from inspecting the GNSS displacement data. This too is relevant for assessing seismic hazard because there are several instances of major earthquakes being triggered by transient geophysical phenomena, such as slow slip events (Roeloffs, 2006) and postseismic deformation (Freed and Lin, 2001). For these reasons, it is important to develop and improve upon methods for estimating crustal strain rates from GNSS data.

Most methods for estimating crustal strain rates, both secular and time-dependent, from GNSS data assume some parametric form of the deformation signal, and then strain rates are the spatial derivative of the best fitting model. The simplest and traditional method for estimating secular crustal strain rates assumes that GNSS derived velocities can be described with a first order polynomial (i.e., having a constant spatial gradient) over some subnetwork of the GNSS stations (e.g Frank, 1966; Prescott, 1976; Savage et al., 1986; Feigl et al., 1990; Murray and Lisowski, 2000). Linear least squares is then used to find the components of the strain rate tensor for each subnetwork. The uniform spatial gradient assumption is obviously not appropriate when subnetworks span a large area, although it is not clear when a subnetwork is too large. To help overcome this deficiency, Shen et al. (1996) and Shen et al. (2015) used an inverse distance weighting scheme where the estimated strain rate at some point is primarily controlled by observations at nearby stations. However, in formulating the algorithms from Shen et al. (1996) and Shen et al. (2015), it is still assumed that the underlying displacement gradients is uniform over the entire network. The errors in this assumption manifests itself as implausibly low formal uncertainties for the estimated strain rates. Other data-based methods for estimating secular strain rates have parameterized velocities with bicubic splines (Beavan and Haines, 2001), spherical wavelets (Tape et al., 2009), and elastostatic Green’s functions (Sandwell and Wessel, 2016). The choice of parameterization can be subjective. If there are too few degrees of freedom in the parameterization then estimated strain rates will be biased and the uncertainties will be underestimated. If there are too many degrees of freedom then there will not be any coherent features in the estimated strain rates. One could parameterize deformation with a physically motivated model of interseismic deformation (e.g., Meade and Hager, 2005; McCaffrey et al., 2007). In such models the lithospheric rheology and plate interfaces are assumed to be known. Any errors in the assumed physical model could similarly result in biased strain estimates and underestimated formal uncertainties.

The aforementioned studies have all been concerned with estimating secular strain rates. In recent years the Southern California Earthquake Center (SCEC) community has shown interest in developing methods for detecting transient strain. SCEC supported a transient detection exercise, where several research groups tested their methods for detecting transient geophysical signals with a synthetic GNSS dataset. Among the methods tested were the Network Strain Filter (NSF) (Ohtani et al., 2010) and the Network Inversion Filter (NIF) (Segall and Mathews, 1997). the NSF parameterizes the geophysical signal spatially with wavelets and the NIF, which is intended for detecting slow fault slip, uses the elastic dislocation Green’s functions from Okada (1992). For the NSF and NIF, the time dependence of the geophysical signal is modeled as integrated Brownian motion. The method described in Holt and Shcherbenko (2013) was also tested in the transient detection exercise, and they calculate strain rates using a bicubic spatial parameterization of displacements between time epochs. Holt and Shcherbenko (2013) defined a detection threshold based on the strain rate magnitude, and we demonstrate that this is indeed an effective criterion for identifying geophysical signal. One main challenge with using these

methods for detecting transient geophysical signal is that it is difficult to distinguish signal from noise, since the noise on the inferred deformation can be significantly influenced by the choice of parameterization.

Each of the aforementioned methods have assumed a described above

Each of these methods use a spatial parameterization for the geophysical signal. If the and deformation spatial use a parametric model

determine strain rates by spatially parameterizing the and they described a method for detecting transient geophysical signal based on the strain rate magnitude, which were used in this exercise, and their approach identified transient strain events by fitting a set of spatial wavelet basis functions to the deformation field at discrete time epochs, and a Kalman filtering strategy was used to ensure that the coefficients for each basis function varied smoothly in time. Holt and Shcherbenko (2013) calculated time dependent strain by differentiating a bicubic interpolant that was fit to each epoch of a temporally smoothed deformation field.

physical models a physical model that based on an assumption of the earths rheology and the location of plate interfaces, Another approach Model-based estimates of secular strain rate find

Another strategy is to fit a set of interpolating basis functions to the deformation field and then compute the strain from the analytical derivative of the interpolant (e.g. Beavan and Haines, 2001; Tape et al., 2009; Sandwell and Wessel, 2016).

Methods for estimating strain from geodetic data fall in one of two categories. There are model-based approaches which assume that strain is the result of loading on faults which have a known geometry, and there are data-based approaches which make no assumptions about the source of deformation. We will exclusively consider data-based approaches in this paper. The classic and simplest method for estimating strain is to assume that the strain rate is constant in time and spatially uniform within subnetworks of the geodetic data. Linear least squares is then used to find the components of the strain rate tensor for each subnetwork . Several algorithms have been developed to improve upon this procedure. Shen et al. (1996) and Shen et al. (2015) discuss an algorithm where, instead of using the immediately adjacent stations to calculate strain at a position, the strain is computed with a weighted average over the entire network where the weighting is smaller for more distant stations.

The aforementioned studies have all been concerned with estimating long term strain rates. Time dependent strain would be useful for studying geophysical processes which occur over timescales of days to years such as slow slip events, postseismic relaxation, or volcanic deformation. Ohtani et al. (2010) identified transient strain events by fitting a set of spatial wavelet basis functions to the deformation field at discrete time epochs, and a Kalman filtering strategy was used to ensure that the coefficients for each basis function varied smoothly in time. Holt and Shcherbenko (2013) calculated time dependent strain by differentiating a bicubic interpolant that was fit to each epoch of a temporally smoothed deformation field.

Each of the methods described above are designed to overcome two complications that arise in estimating deformation gradients: (1) geodetic data are noisy and differentiation will only amplify the noise and (2) geodetic data are not observed on a regular grid, which prevents the use of standard finite difference methods for computing derivatives. In this paper we demonstrate that both of these complications can be elegantly handled with the Radial Basis Function-Finite Difference (RBF-FD) method.

The RBF-FD method was developed simultaneously and independently by Tolstykh and Shirobokov (2003), Shu et al. (2003), Cecil et al. (2004), and Wright and Fornberg (2006) as a computationally efficient way to solve large scale partial differential equations over irregular, multi-dimensional domains. The RBF-FD method can be thought of as a generalization of the traditional finite difference method, where the node layout is no longer restricted to regular grids. Indeed, the RBF-FD method can be used to estimate derivatives of discrete data located at arbitrary scattered positions in multi-dimensional space. The RBF-FD method is particularly appealing because it is algorithmically simple, regardless of the domain shape or node layout, and also because the method has performed well in numerous benchmark tests (?, and references therein).

In this paper, we do not use the RBF-FD method to solve a partial differential equation, but rather we use it to spatially smooth geodetic data and to compute deformation gradients. Our smoothing strategy can be viewed as a non-parametric, low-pass filter for scattered data where the degree of smoothness is controlled by a user specifies cutoff frequency. This can be contrasted with interpolation based methods where the resulting interpolant can be largely and perhaps unpredictably controlled by the choice of basis function. This process of smoothing and differentiating can be extended to estimate time dependent strain rates. In that case, we first temporally smooth and differentiate GPS displacement time series to get time dependent velocities. We then spatially smooth and differentiate the resulting velocities for each time epoch to get time dependent strain rates.

The method proposed in this paper has numerous advantages which set it apart from other methods for computing strain rates. The method is computationally efficient and stable (there is no inversion of an ill-conditioned matrix). There are no hyper parameters or penalty parameters that need to be tuned for each application. As opposed to interpolation strategies such as Beavan and Haines (2001), Tape et al. (2009), or Ohtani et al. (2010), our method assumes that velocities are locally rather than globally continuous, which allows us to easily handle discontinuities resulting from, for example, a creeping fault.

We begin this paper by summarizing the RBF-FD scheme and explaining how we construct differentiation matrices for scattered data. We then introduce the RBF-FD filter, which is used to smooth the observed geodetic data prior to differentiation. We provide two real world demonstrations of our method for calculating strain rates. First we calculate the long term strain rates in Southern California from the CMM3 velocity data set (Shen et al., 2011), and we verify that our results are consistent with other studies. We then calculate time dependent strain rates in Cascadia from the GPS data provided by UNAVCO. In Cascadia, we analyze strain resulting from slow slip events and compare it to the long term tectonic strain accumulation. Slow slip events are found to produce compression in the Olympic Peninsula, which is in addition to the compression resulting from tectonic loading. Further south in Oregon, the slow slip events tend to release the compressional strain that is accumulated tectonically. While similar conclusions have been drawn from fault slip inversions for slow slip events, it is important to recognize that slip inversion are the product of inverting an ill-conditioned matrix making it difficult to determine whether slip inferences are real or just an artifact of the inversion. The strain rates presented in this paper are more direct observations and can be interpreted with a higher degree of confidence.

GPR is a bayesian approach which does not rely on dubious parametric assumptions.

2 Method

In this section, we describe how GNSS displacement observations are used to identify transient crustal strain rates, which we denote as $\dot{\epsilon}(z)$, where z is the ordered pair (x, t) , x are spatial coordinates in \mathbb{R}^2 , and t is time. We consider $\dot{\epsilon}$ to be spatially and temporally coherent deviations from the steady rate of strain accumulation from plate tectonics. We determine $\dot{\epsilon}$ by first identifying transient displacements, $\mathbf{u}(z)$, which we then spatially and temporally differentiate. As we will show in Section X, estimates of $\dot{\epsilon}$ turn out to be more effective at illuminating geophysical signal than estimates of \mathbf{u} or $\dot{\mathbf{u}}$. We make a prior assumption that each component of \mathbf{u} is a Gaussian process,

$$u_i(z) \sim \mathcal{N}(0, C_{u_i}), \quad (1)$$

where $C_{u_i}(z, z')$ is a covariance function indicating how we expect $u_i(z)$ to covary with $u_i(z')$. For simplicity, we treat each component of displacement independently and ignore any potential covariance. We then drop the component subscripts with the understanding that the same analysis is being repeated to estimate the east, north, and vertical components of \mathbf{u} . We further assume that C_u can be separated into positive definite spatial and temporal functions as

$$C_u((x, t), (x', t')) = X(x, x')T(t, t'), \quad (2)$$

The appropriate choice for X and T may vary depending on the geophysical signal we are trying to describe (e.g. postseismic deformation or deformation from slow slip events), and we discuss this matter in the next section.

We constrain u with GNSS data, which records u as well as other physical and non-physical processes which we are not interested in. We describe GNSS observations at position x_i and time t_j as a realization of the random variable

$$\begin{aligned} d_{ij} = & u(x_i, t_j) + v(x_i, t_j) + w_{ij} + a_i^{(1)} + a_i^{(2)}t_j + \\ & a_i^{(3)} \sin(2\pi t_j) + a_i^{(4)} \cos(2\pi t_j) + a_i^{(5)} \sin(4\pi t_j) + a_i^{(6)} \cos(4\pi t_j), \end{aligned} \quad (3)$$

where $a_i^{(1)}$ is an offset that is unique for each GNSS monument, $a_i^{(2)}$ is the steady rate of tectonic deformation at x_i , and the sinusoids describe seasonal deformation (using units of years for t). We use w_{ij} to denote normally distributed, uncorrelated noise. Correlated noise which does not have a parametric representation is denoted by v . For example, v can consist of temporally correlated noise describing benchmark wobble (e.g., Wyatt, 1982, 1989), and/or spatially correlated noise describing common mode error (e.g., Wdowinski et al., 1997). For now, we will only assume that $v \sim \mathcal{N}(0, C_v)$. We consider the six coefficients in eq. (3) to be uncorrelated random variables distributed as $\mathcal{N}(0, \kappa^2)$ in the limit as $\kappa \rightarrow \infty$ (i.e., the coefficients have diffuse priors). Of course, the tectonic deformation, $a_i^{(2)}$, is spatially correlated and we could invoke a tectonic model to form a prior on $a_i^{(2)}$. However, in our application to Cascadia, we will be using displacement time series which are long enough to sufficiently constrain $a_i^{(2)}$ for each station, and there is no need to incorporate a prior. Likewise, the seasonal coefficients may be spatially correlated (Langbein, 2008), and it may be worth exploring and exploiting such a correlation in a future study.

We now consider the column vector of n GNSS observations, \mathbf{d}_* . Let \mathbf{y} be the set of (x_i, t_j) pairs describing where and when each of the GNSS observations have been made. Let \mathbf{a} be the vector of coefficients from eq. (3) for each of the m GNSS stations. We use \mathbf{P} to represent the $n \times 6m$ matrix of corresponding basis functions evaluated at each point in \mathbf{y} . We also denote the vector of uncorrelated noise for each observation as \mathbf{w} , whose

standard deviations are given by the formal data uncertainty σ . The observations \mathbf{d}_* can then be viewed as a realization of the random vector

$$\mathbf{d} = u(\mathbf{y}) + v(\mathbf{y}) + \mathbf{w} + \mathbf{P}\mathbf{a}, \quad (4)$$

which is distributed as $\mathcal{N}(\mathbf{0}, \Sigma + \kappa^2 \mathbf{P}\mathbf{P}^T)$, where

$$\Sigma = C_u(\mathbf{y}, \mathbf{y}) + C_v(\mathbf{y}, \mathbf{y}) + \text{diag}(\sigma^2). \quad (5)$$

It should be understood that the notation $f(\mathbf{x})$ and $f(\mathbf{x}, \mathbf{y})$ is shorthand for $[f(x_i)]_{x_i \in \mathbf{x}}$ and $[f(x_i, y_j)]_{(x_i, y_j) \in \mathbf{x} \times \mathbf{y}}$, respectively.

We now describe how to condition the prior for each component of transient displacements, u , with \mathbf{d}_* to form a posterior estimate of transient displacements, $\hat{u} = u|\mathbf{d}_*$. We will assume that an appropriate covariance functions and corresponding hyperparameters for X , T , and C_v have already been chosen. If κ is kept finite, then we can follow Rasmussen and Williams (2006) to find that \hat{u} is distributed as $\mathcal{N}(\mu_{\hat{u}}, C_{\hat{u}})$, where

$$\mu_{\hat{u}}(z) = C_u(z, \mathbf{y}) (\Sigma + \kappa^2 \mathbf{P}\mathbf{P}^T)^{-1} \mathbf{d}_* \quad (6)$$

and

$$C_{\hat{u}}(z, z') = C_u(z, z') - C_u(z, \mathbf{y}) (\Sigma + \kappa^2 \mathbf{P}\mathbf{P}^T)^{-1} C_u(\mathbf{y}, z'). \quad (7)$$

However, we are interested in the limit as $\kappa \rightarrow \infty$, and the form for eq. (6) and eq. (7) is not suitable for evaluating this limit. We can use the partitioned matrix inversion identity (e.g., Press et al., 2007) to rewrite eq. (6) and eq. (7) as

$$\mu_{\hat{u}}(z) = [C_u(z, \mathbf{y}) \quad \mathbf{0}] \begin{bmatrix} \Sigma & \mathbf{P} \\ \mathbf{P}^T & -\kappa^{-2} \mathbf{I} \end{bmatrix}^{-1} \begin{bmatrix} \mathbf{d}_* \\ \mathbf{0} \end{bmatrix} \quad (8)$$

and

$$C_{\hat{u}}(z, z') = C_u(z, z') - [C_u(z, \mathbf{y}) \quad \mathbf{0}] \begin{bmatrix} \Sigma & \mathbf{P} \\ \mathbf{P}^T & -\kappa^{-2} \mathbf{I} \end{bmatrix}^{-1} \begin{bmatrix} C_u(\mathbf{y}, z') \\ \mathbf{0} \end{bmatrix}. \quad (9)$$

Taking the limit as $\kappa \rightarrow \infty$, we get the solution for the mean and covariance of \hat{u} ,

$$\mu_{\hat{u}}(z) = [C_u(z, \mathbf{y}) \quad \mathbf{0}] \begin{bmatrix} \Sigma & \mathbf{P} \\ \mathbf{P}^T & \mathbf{0} \end{bmatrix}^{-1} \begin{bmatrix} \mathbf{d}_* \\ \mathbf{0} \end{bmatrix} \quad (10)$$

and

$$C_{\hat{u}}(z, z') = C_u(z, z') - [C_u(z, \mathbf{y}) \quad \mathbf{0}] \begin{bmatrix} \Sigma & \mathbf{P} \\ \mathbf{P}^T & \mathbf{0} \end{bmatrix}^{-1} \begin{bmatrix} C_u(\mathbf{y}, z') \\ \mathbf{0} \end{bmatrix}. \quad (11)$$

The posterior transients displacements, $\hat{\mathbf{u}}$, are found by evaluating eq. (10) and (11) for each displacement component. We differentiate the Gaussian process $\hat{\mathbf{u}}$ (see Abrahamsen (1997)) to form an estimate of transient strain rate, $\hat{\epsilon}$, which is itself a spatially and temporally continuous Gaussian process. The components of the transient strain rate tensor, $\hat{\epsilon}_{ij}$, are given by

$$\hat{\epsilon}_{ij}(z) = \frac{1}{2} \frac{\partial}{\partial t} \left(\frac{\partial \hat{u}_i(z)}{\partial x_j} + \frac{\partial \hat{u}_j(z)}{\partial x_i} \right). \quad (12)$$

which are distributed as $\mathcal{N}(\mu_{\hat{\epsilon}_{ij}}, C_{\hat{\epsilon}_{ij}})$, where

$$\mu_{\hat{\epsilon}_{ij}}(z) = \frac{1}{2} \frac{\partial}{\partial t} \left(\frac{\partial \mu_{\hat{u}_i}(z)}{\partial x_j} + \frac{\partial \mu_{\hat{u}_j}(z)}{\partial x_i} \right) \quad (13)$$

and

$$C_{\hat{\epsilon}_{ij}}(z, z') = \frac{1}{4} \frac{\partial^2}{\partial t \partial t'} \left(\frac{\partial^2 C_{\hat{u}_i}(z, z')}{\partial x_j \partial x'_j} + \frac{\partial^2 C_{\hat{u}_j}(z, z')}{\partial x_i \partial x'_i} \right). \quad (14)$$

2.1 Outlier detection

In our formulation for estimating transient strain rates, we have assumed that noise in the data vector is normally distributed. This is not an appropriate assumption for GNSS displacement timeseries which are prone to more outliers than would be predicted for normally distributed noise. It follows that proposed methods for analyzing GNSS timeseries should be robust against outliers (e.g., Blewitt et al., 2016). In order to make our estimates of transient strain more robust, we automatically identify and remove outliers in the GNSS data as a pre-processing step.

Our method for detecting outliers is based on the data editing algorithm described in Gibbs (2011). We calculate the residuals between the observations and a best fitting model, and data with residuals that are anomalously large are identified as outliers. We treat \mathbf{d}_* as a sample of \mathbf{d} and assume that there is no temporally correlated noise (i.e., $v = 0$). The best fitting model for \mathbf{d}_* is considered to be the expected value of the random vector $u(\mathbf{y}) + \mathbf{P}\mathbf{a}$ after conditioning it with non-outlier observations. We still consider u to have a separable covariance function as in eq. (2), and the choice for X and T does not need to be the same as that used in Section X. Since outliers are determined based on how well a spatially and temporally dependent model fits the data, we are able to identify anomalous observations which may not be immediately apparent based on inspection of individual timeseries.

To begin the algorithm, we let Ω be the index set of non-outliers in \mathbf{d}_* and initiate it with all n indices. This algorithm is iterative, and for each iteration we calculate the residual vector

$$\mathbf{r} = \frac{\mathbf{d}_* - \mathbb{E}[(u(\mathbf{y}) + \mathbf{P}\mathbf{a})|\tilde{\mathbf{d}}_*]}{\sigma} \quad (15)$$

$$= \frac{1}{\sigma} \left(\mathbf{d}_* - \begin{bmatrix} C_u(\mathbf{y}, \tilde{\mathbf{y}}) & \mathbf{P} \end{bmatrix} \begin{bmatrix} C_u(\tilde{\mathbf{y}}, \tilde{\mathbf{y}}) + \text{diag}(\tilde{\sigma}^2) & \tilde{\mathbf{P}} \\ \tilde{\mathbf{P}}^T & \mathbf{0} \end{bmatrix}^{-1} \begin{bmatrix} \tilde{\mathbf{d}}_* \\ \mathbf{0} \end{bmatrix} \right), \quad (16)$$

where the tilde indicates that only elements corresponding to indices in Ω are retained (e.g., $\tilde{\mathbf{y}} = \{y_i\}_{i \in \Omega}$). We then update Ω to be

$$\Omega = \{i : |r_i| < \eta \tilde{r}_{rms}\}, \quad r_i \in \mathbf{r} \quad (17)$$

where \tilde{r}_{rms} is the root-mean-square of $\tilde{\mathbf{r}}$ and η is an outlier tolerance. We use $\eta = 4$ in this study, which seems to accurately identify outliers without unnecessarily decimating the data. Iterations continue until the new Ω is equal to the previous Ω .

It should be noted that this algorithm does not identify jumps in GNSS time series, which are another common defect. Some, but not all, jumps can be automatically removed by parsing station logs and identifying dates of antenna changes. However, it is still necessary to manually inspect and remove jumps of unknown origin. That being said, this outlier detection algorithm significantly reduces the effort needed to manually clean GNSS data.

3 Application to Cascadia Slow Slip Events

In this section we estimate transient strain from GNSS data in Cascadia. In particular, we are interested in identifying transient strain resulting from slow slip events (SSEs) (e.g., Dragert et al., 2001). Our results illuminate where elastic strain energy, which can eventually be released seismically, is accumulating during SSEs. The estimated strain rates also reveal how strain from SSEs contributes to strain accumulated over seismic cycles. We demonstrate that transient strain rates estimated with the method described above are an effective tool for detecting SSEs and potentially other geophysical signal in GNSS data. In Cascadia, SSEs can be detected by monitoring for associated seismic tremor (Rogers and Dragert, 2003), which is actively being done by the Pacific Northwest Seismic Network (PNSN) (Wech, 2010). We verify that transient strain rates estimated with the method described above are indeed identifying strain from SSEs by comparing $\dot{\epsilon}$ to the tremor records. We envision that GNSS derived transient strain rates could be useful for monitoring subduction zones which have SSEs but no associated tremor (Schwartz and Rokosky, 2007).

We use publicly available continuous GNSS data from the University Navstar Consortium (UNAVCO) (Herring et al., 2016), which can be found at www.unavco.org. We limit the dataset to the stations and time ranges which are pertinent to seven of the most recent SSEs in the Puget Sound region; the earliest SSE began in the August 2010, and the most recent SSE began in the February of 2017. We consider these most recent SSEs because the station coverage is sufficiently dense for us to make well constrained inferences of optimal prior models. The positions of GNSS stations used to constrain transient strain rates are shown in Figure 1.

3.1 Noise model

Before we determine the transient strain rates, we must establish a prior for the transient deformation, u , and the noise, v . In this section we discuss our choice for the noise covariance function C_v . There have been numerous studies on temporally correlated noise in GNSS data (e.g., Zhang et al., 1997; Mao et al., 1999; Williams et al., 2004; Langbein, 2008). In these studies, temporally correlated noise was described with some combination of Brownian motion (also referred to as a Weiner process or a random walk), a first-order Gauss-Markov (FOGM) process, and/or flicker noise. There is some physical justification for using Brownian motion as a noise model because it accurately describes the power spectrum of motion resulting from instability in geodetic monuments (e.g., Wyatt, 1982, 1989; Langbein and Johnson, 1997). Here we describe the time dependence of v as a FOGM

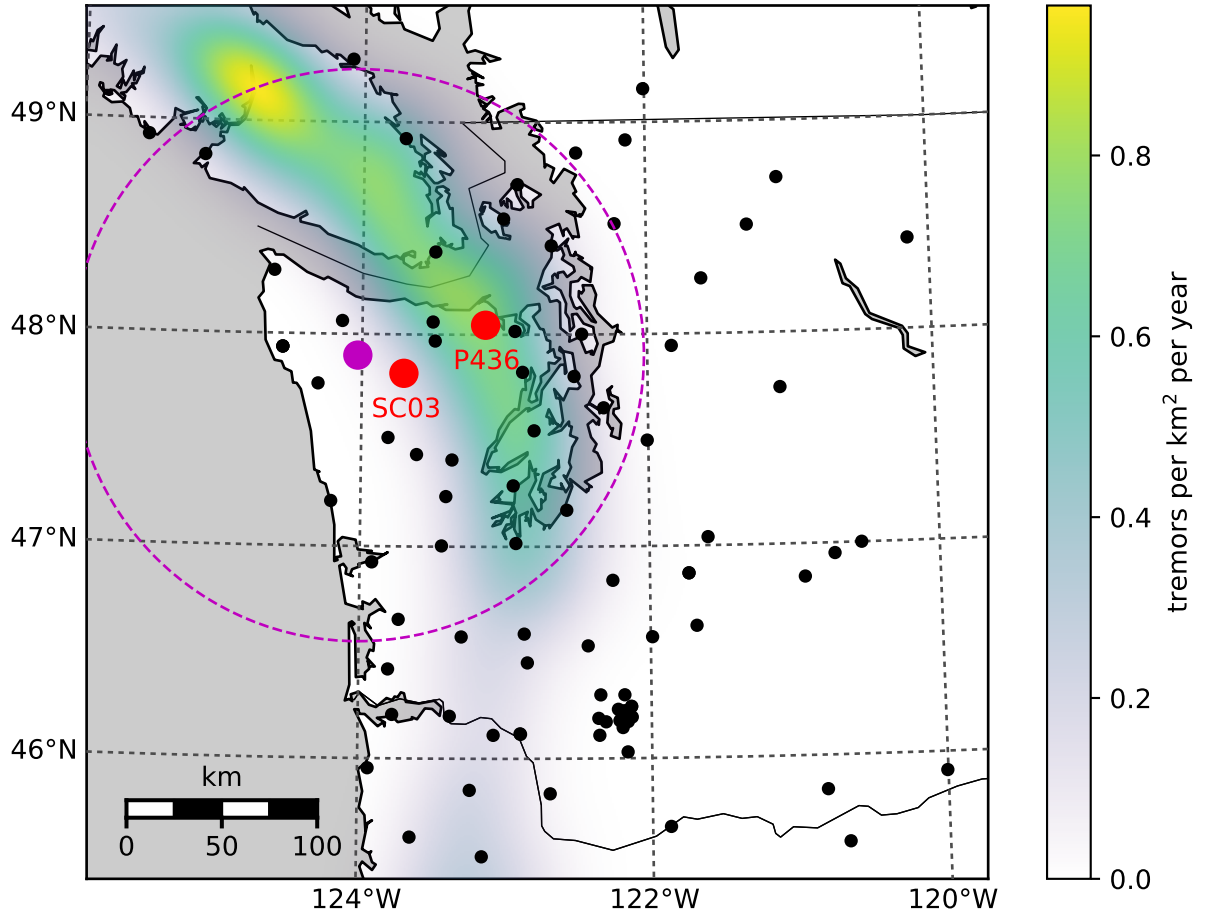


Figure 1: Positions of continuous GNSS stations used to estimate transient strain rates. The colored regions indicates the density of detected seismic tremor as determined by (Wech, 2010). The red dots show the positions of GNSS stations mentioned in this paper. The magenta dot indicates the location of the strain rates shown in Figure 7 and 8. The magenta circle demarcates the spatial extent of the tremors shown in Figure 8

process and consider v to be spatially uncorrelated. A FOGM process is a solution to the stochastic differential equation

$$\dot{v}(t) + \alpha v(t) = \beta w(t), \quad (18)$$

where $w(t)$ is white noise with unit variance. The FOGM process degenerates to the commonly used Brownian motion noise model under the condition that $\alpha = 0$ and $v(0) = 0$, which is evident from the definition of Brownian motion (e.g., Papoulis, 1991). Our noise model which satisfies eq. (18) is a Gaussian process with zero mean and covariance function

$$C_v((x, t), (x', t')) = \frac{\beta^2}{2\alpha} \exp(-\alpha|t - t'|) \delta(\|x - x'\|_2). \quad (19)$$

We constrain the hyperparameters for v , α and β , with a set of 38 continuous GNSS stations in Washington and Oregon which are east of 121° W. These stations are sufficiently far from the subduction zone that they are unlikely to contain transient signal. We clean the data for these stations by removing jumps at times of equipment changes, and we remove outliers that have been detected with the algorithm described in Section 2.1. We then find α and β for each station timeseries with the Restricted Maximum Likelihood (REML) method (e.g., Harville, 1974; Cressie, 1993). The REML method finds the hyperparameters, which we collectively refer to as θ , that maximize the likelihood function

$$\mathcal{L}(\theta) = \left(\frac{|P^T P|}{(2\pi)^{n-6m} |\Sigma(\theta)| |P^T \Sigma P|} \right)^{\frac{1}{2}} e^{-\frac{1}{2} d_*^T K(\theta) d_*}, \quad (20)$$

where

$$K(\theta) = \Sigma(\theta)^{-1} - \Sigma(\theta)^{-1} P (P^T \Sigma(\theta)^{-1} P)^{-1} P^T \Sigma(\theta)^{-1}. \quad (21)$$

For this section, d_* is used to denote the displacement time series for just one station, and thus $m = 1$. We are also assuming C_u is zero when estimating the noise hyperparameters for this section. Harville (1974) showed that choosing the hyperparameters which maximize eq. (20) is equivalent to choosing the hyperparameters which maximize the probability of drawing d_* from d . We use the REML method over the maximum likelihood (ML) method (e.g., Langbein and Johnson, 1997) because the REML method properly accounts for the improper prior that we assigned to α . The distribution of inferred α and β are shown in Figure 2. The amplitude of FOGM noise for the east and north components is notable low and are clustered around 0.5 mm/yr^{0.5}. The corresponding estimates of α tend to cluster around 0 yr⁻¹, suggesting that Brownian motion would have also been an appropriate noise model. The amplitude of FOGM noise for the vertical component is significantly larger with a median value of 13.8 mm/yr^{0.5}. The inferred values for alpha are higher for the vertical component with a median value of 1.87 yr⁻¹. In Figure X, we use the median values of α and β to generate two random samples of FOGM noise. The samples span seven years and over these seven years the east and north samples drift by about 1 mm. In the context of detecting SSEs, which produce several mm's of surface displacement on the timescale of weeks, the estimated FOGM noise for the east and north component is negligible. In contrast, the estimated FOGM noise for the vertical component is significantly larger than the signal we would expect from SSEs. We suspect that the higher amplitude for the FOGM model in the vertical component is accommodating for deficiencies in our rather simple seasonal model. Based on this analysis, we henceforth ignore temporally correlated noise in the east and north component because of its low amplitude, and we do not estimate u for vertical deformation because of the low signal to noise ratio.

Another significant source of noise in GNSS data is common mode error (e.g., Wdowinski et al., 1997; Dong et al., 2006), which is noise that is highly spatially correlated. When not accounted for, common mode error manifests itself as spatially uniform undulations in \hat{u} . However, we are primarily interested in estimating strain which is insensitive to common mode error. We therefore do not include common mode error in our noise model.

For these reasons, we henceforth make the simplifying assumption that temporally and spatially correlated noise is negligible and $v = 0$ for the easting and northing component of GNSS data.

3.2 Transient displacement model

We next discuss our prior model for transient displacements, and, specifically, our choice for the covariance functions $X(x, x')$ and $T(t, t')$. For X , we use the squared exponential (SE) covariance function,

$$X(x, x') = \exp\left(\frac{-\|x - x'\|_2^2}{2\lambda^2}\right). \quad (22)$$

The SE covariance function is commonly used in Kriging (e.g, Cressie, 1993) and Gaussian process regression (e.g., Rasmussen and Williams, 2006). The SE is a valid (i.e. positive definite) covariance function for any number of spatial dimensions, and it describes an isotropic Gaussian processes with realizations that are infinitely

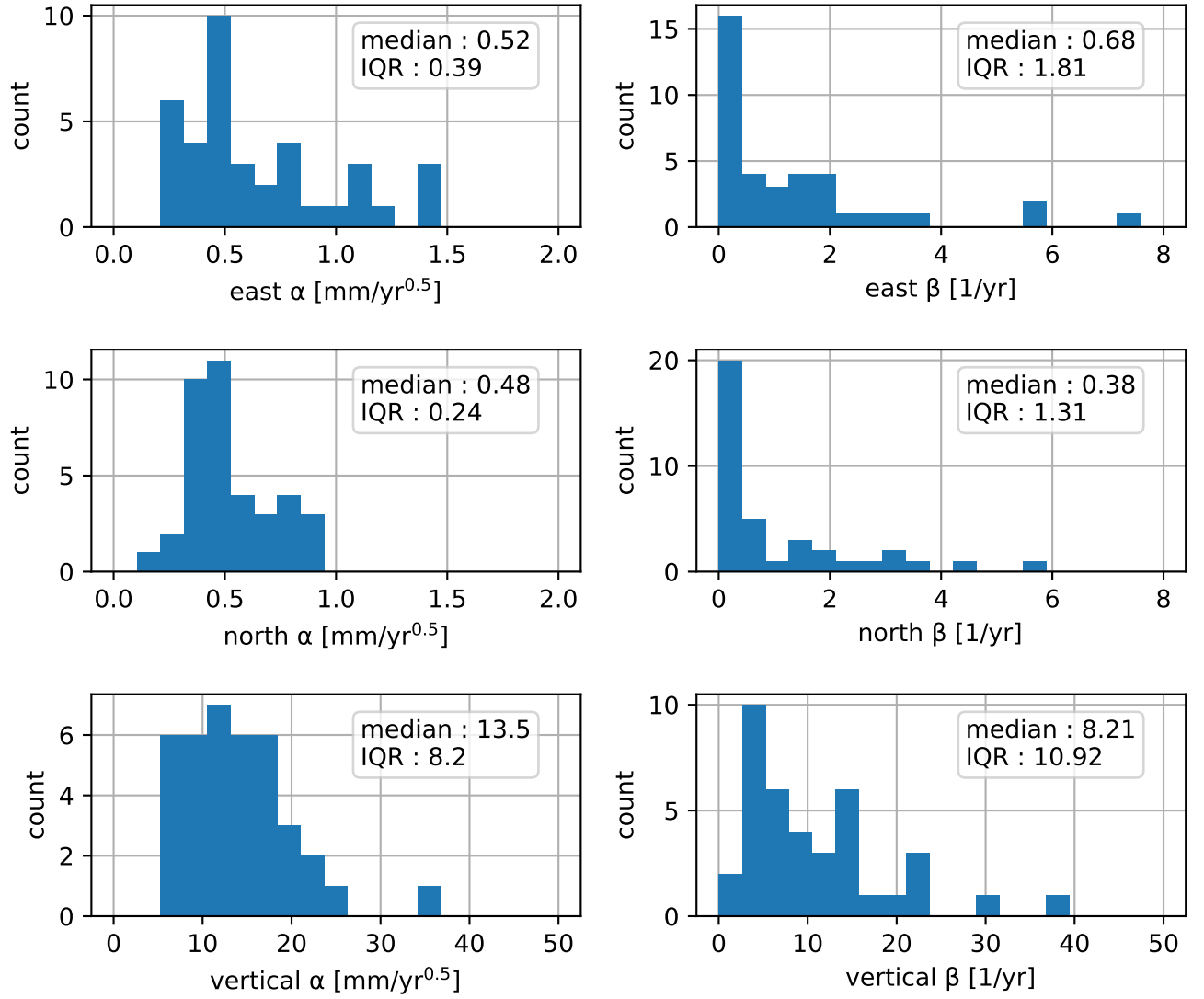


Figure 2: Distribution of estimated FOGM hyperparameters from eq. (19). Hyperparameters are independently estimated for each of the stations shown in Figure X and for each displacement component. “IQR” is the inter-quartile range.

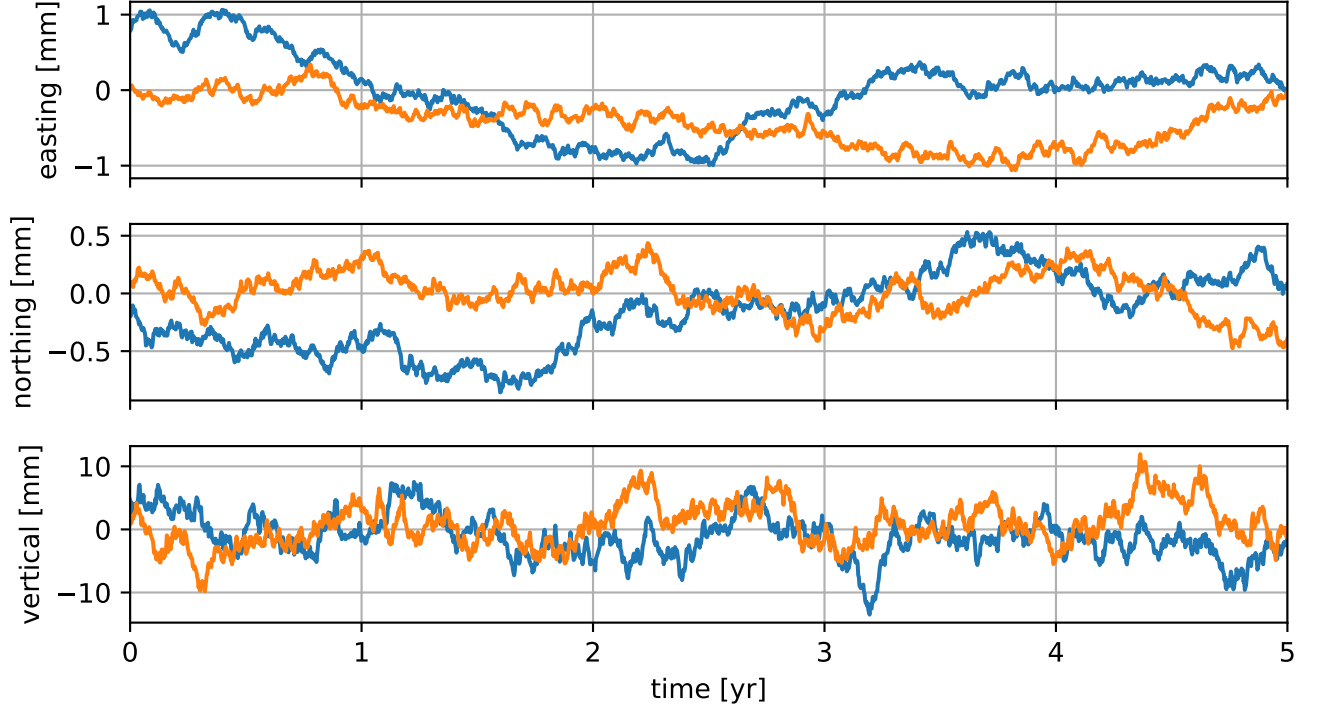


Figure 3: Two FOGM noise samples, where the hyperparameters have been set to the median values from Figure 2.

differentiable. In terms of geodetic applications, Kato et al. (1998) and El-Fiky and Kato (1999) demonstrated that the SE is an appropriate covariance model for describing long-term tectonic strain rates in Japan.

We consider three potential covariance functions to describe the temporal covariance of u . First, we consider the one-dimensional SE covariance function,

$$T(t, t') = \phi^2 \exp\left(\frac{-|t - t'|^2}{\theta^2}\right), \quad (23)$$

simply because the SE covariance function is versatile and commonly used for non-parametric modeling. Note that T includes the hyperparameter ϕ , which serves to scale the covariance function C_u . Second, we consider integrated Brownian motion (IBM). IBM has zero mean and its covariance function can be found by integrating the covariance function for Brownian motion as

$$T(t, t') = \int_0^t \int_0^{t'} \phi^2 \min(\tau, \tau') d\tau' d\tau \quad (24)$$

$$= \frac{\phi^2}{2} \min(t, t')^2 \left(\max(t, t') - \frac{1}{3} \min(t, t') \right), \quad t, t' \geq 0. \quad (25)$$

IBM has been used, in the context of Kalman filtering, as a non-parametric model for the time dependence of geophysical signals (e.g., Segall and Mathews, 1997; McGuire and Segall, 2003; Ohtani et al., 2010; Hines and Hetland, 2016). It should be emphasized $t = 0$ is a reference time at which the Gaussian process is exactly zero. For some geophysical signals, it is appropriate to have this reference time. For example, if we are trying to identify postseismic deformation then t should be zero at the time of the earthquake. However, if we are interesting in detecting transient events, where there is no known start time, then IBM may not be an appropriate prior and an isotropic Gaussian process should be preferred. In the following analysis, we make the quite arbitrary choice that t is zero on the first epoch of \mathbf{d}_* . Our third option for T is the Wendland class of covariance functions (Wendland, 2005). Wendland covariance functions have compact support and hence their corresponding covariance matrices will be sparse. In our analysis, we exploit this sparsity with the CHOLMOD software package (Chen et al., 2008). Wendland functions are positive definite in \mathbb{R}^d , and they describes an isotropic Gaussian process with realizations that can be differentiated k times. The form of the covariance function depends on the choice of d and k . To describe the temporal covariance of u , we only require that $d = 1$

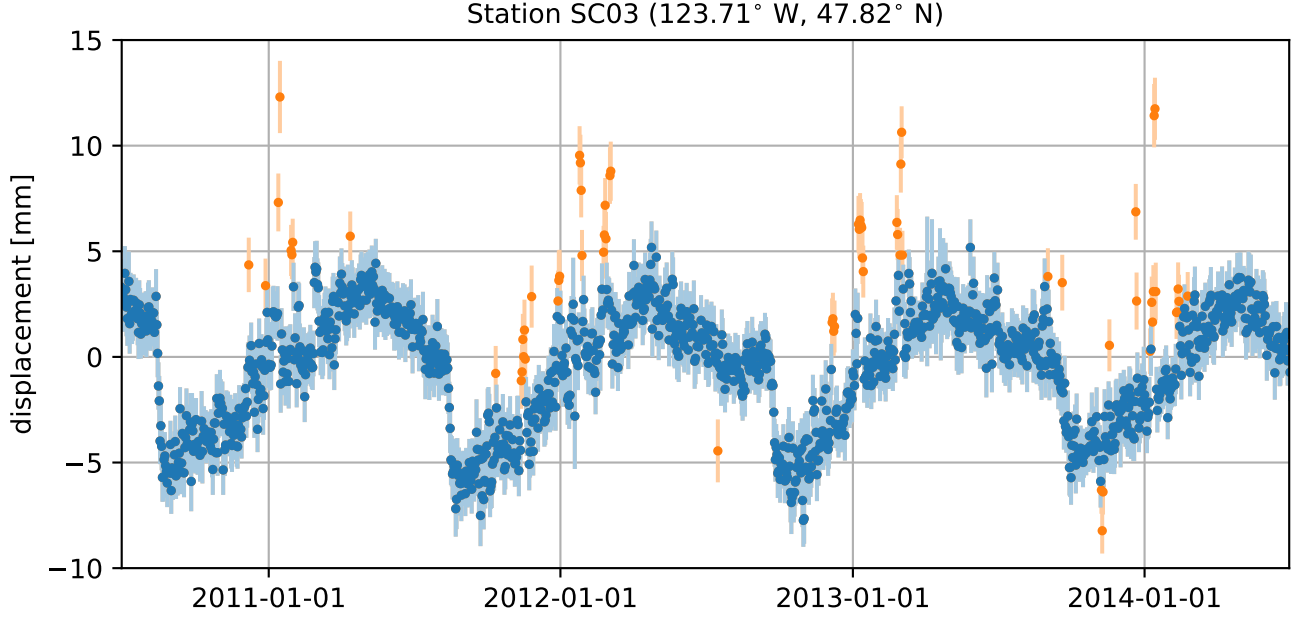


Figure 4: Detrended easting component of GNSS displacement data for station SC03, which is located on Mt. Olympus in Washington. The orange markers indicate outliers which were automatically detected using the algorithm from Section 2. The error bars show one standard deviation uncertainty.

and $k = 1$. The corresponding Wendland covariance function is

$$T(t, t') = \phi^2 \left(1 - \frac{|t - t'|}{\theta} \right)_+^3 \left(\frac{3|t - t'|}{\theta} + 1 \right), \quad (26)$$

where

$$(t)_+ = \begin{cases} t, & t > 0 \\ 0, & \text{otherwise.} \end{cases} \quad (27)$$

We next want to determine appropriate hyperparameters for X and each of the three candidate functions for T . First, we again clean the GNSS datasets by removing offsets at times of equipment changes and removing outliers with the method describe in Section 2.1. For the outlier detection algorithm, our prior model, u , is chosen to have a length-scale and time-scale which is able to approximately describe SSE displacements. We use the SE covariance function for X with $\lambda = 100$ km, and we use the Wendland covariance function for T , due to its computational efficiency, with $\theta = 0.1$ yr and $\phi = 1$ mm. The outlier detection algorithm is particularly effective at removing outliers for stations at high elevation, which can be adversely affected by ice or snow during the winter (Lisowski et al., 2008). As an example, Figure 4 shows the outliers which were automatically detected for station SC03, which is located on Mount Olympus in Washington. After cleaning the data, we divide it into subsets which are four months long and centered on the time of an SSE. The times of seven SSEs are determined using tremor records from Wech (2010). We find the optimal hyperparameters for T and X for each subset of data. The REML method is once again used to determine the hyperparameters for T and X which maximize the probability of drawing \mathbf{d}_* from \mathbf{d} , where \mathbf{d}_* now refers to a data subset. We chose to make each data subsets four months long because it is long enough to encompass the SSE, and it is short enough to still be computationally tractable. However, four months is too short to resolve the sinusoids in \mathbf{d} , and they are omitted from \mathbf{d} for the REML analysis. The estimated hyperparameters for u are summarized in Table 1. Based on the interquartile ranges (IQR), the hyperparameters for the SE and Wendland model do not vary significantly between SSEs. This suggests that the median of estimated hyperparameters should be an appropriate prior model for all Cascadia SSEs. For the IBM model, there are several anomalously large values for λ and ϕ , which is they they have large IQRs.

Next we identify which covariance function for T best describes the SSEs. One approach is to compare the REML likelihood for each covariance function, similar to the analysis in Langbein (2004). In Table 1, we summarize how the log REML likelihoods for the Wendland and IBM covariance functions compare to the SE covariance function. Based on the differences in log REML likelihoods, the data is substantially more likely to come form the SE or Wendland covariance functions than from the IBM covariance function. The REML likelihoods do not definitively indicate whether the SE or Wendland covariance function is preferable. As another

T	direction	λ	ϕ	θ	diff. log(REML)
SE	east	92 ± 25 km	0.62 ± 0.11 mm	0.026 ± 0.011 yr	-
SE	north	91 ± 53 km	0.43 ± 0.05 mm	0.030 ± 0.017 yr	-
Wendland	east	95 ± 30 km	0.66 ± 0.15 mm	0.093 ± 0.044 yr	0.78 ± 0.87
Wendland	north	92 ± 57 km	0.46 ± 0.10 mm	0.116 ± 0.057 yr	0.08 ± 0.58
IBM	east	110 ± 130 km	290 ± 420 mm/yr ^{1.5}	-	-16.4 ± 7.8
IBM	north	150 ± 560 km	110 ± 250 mm/yr ^{1.5}	-	-10.1 ± 2.3

Table 1: Optimal hyperparameters for transient displacements determined with the REML method. The temporal covariance function is indicated by the “ T ” column, and the spatial covariance function, X , is the squared exponential in all cases. The hyperparameters are estimated for each of the seven SSEs considered in this study, and the tabulated values indicate the median and interquartile ranges of estimates. The “diff log(REML)” column compares the log REML likelihood to the log REML likelihood for the SE model. Positive values indicate that observations are more consistent with the SE model.

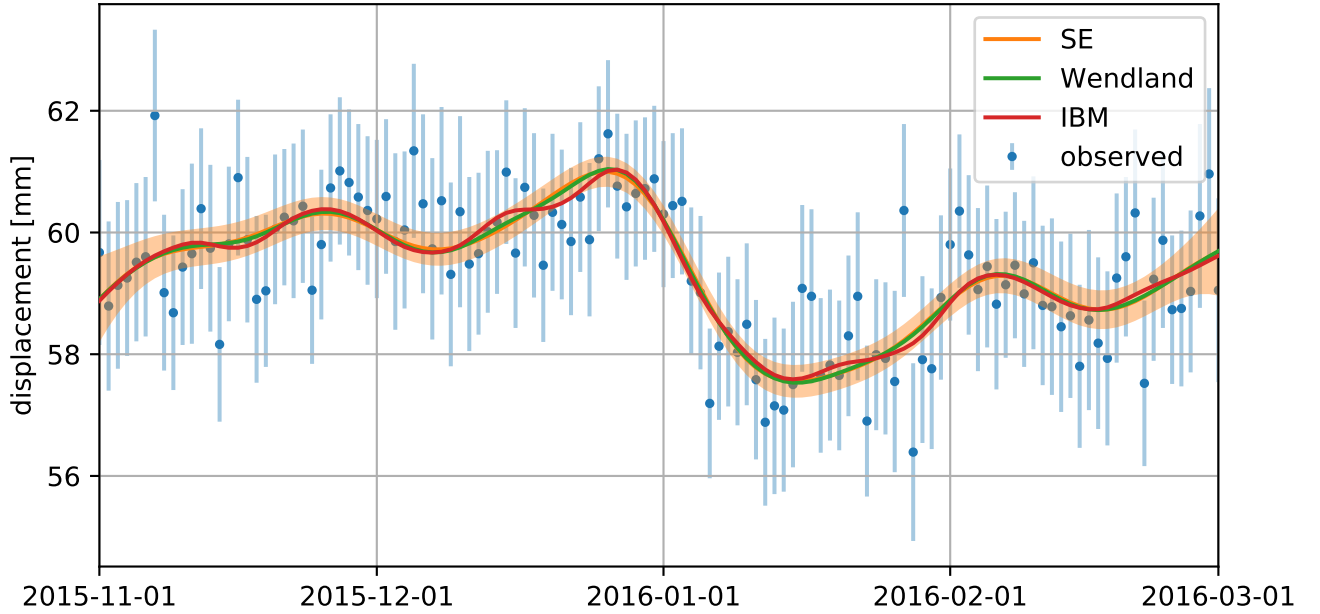


Figure 5: Fit.

test to determine the most appropriate covariance function for T , we compare the observations to the predicted displacements for each covariance function. let $\hat{\mathbf{d}} = (u(\mathbf{y}) + \mathbf{P}\mathbf{a})|_{\mathbf{d}_*}$ be the data prediction vector. It can be shown that $\hat{\mathbf{d}}$ is normally distributed with mean

$$\mu_{\hat{\mathbf{d}}} = \begin{bmatrix} C_u(\mathbf{y}, \mathbf{y}) & \mathbf{P} \end{bmatrix} \begin{bmatrix} \Sigma & \mathbf{P} \\ \mathbf{P}^T & \mathbf{0} \end{bmatrix}^{-1} \begin{bmatrix} \mathbf{d}_* \\ \mathbf{0} \end{bmatrix} \quad (28)$$

and covariance

$$\mathbf{C}_{\hat{\mathbf{d}}} = C_u(\mathbf{y}, \mathbf{y}) - \begin{bmatrix} C_u(\mathbf{y}, \mathbf{y}) & \mathbf{P} \end{bmatrix} \begin{bmatrix} \Sigma & \mathbf{P} \\ \mathbf{P}^T & \mathbf{0} \end{bmatrix}^{-1} \begin{bmatrix} C_u(\mathbf{y}, \mathbf{y}) \\ \mathbf{P}^T \end{bmatrix}. \quad (29)$$

We compute $\hat{\mathbf{d}}$ using SE, Wendland, and IBM covariance functions for T and the median hyperparameters from Table 1. Figure 5 compares the easting component of \mathbf{d}_* to $\hat{\mathbf{d}}$ for the Winter 2016 SSE. For most stations, the data prediction vector appears to appropriately describe displacements throughout the SSE, regardless of the choice of T . The prediction for the IBM model contains slightly more high frequency, and perhaps spurious, features. The predictions for the Wendland and SE covariance functions are nearly indistinguishable. In our estimates of transient strain discussed in the next section, we ultimately settle on the Wendland covariance function for T and use the median values from Table 1 for the hyperparameters. We choose the Wendland covariance function over the SE covariance function because of its computational advantages noted above.

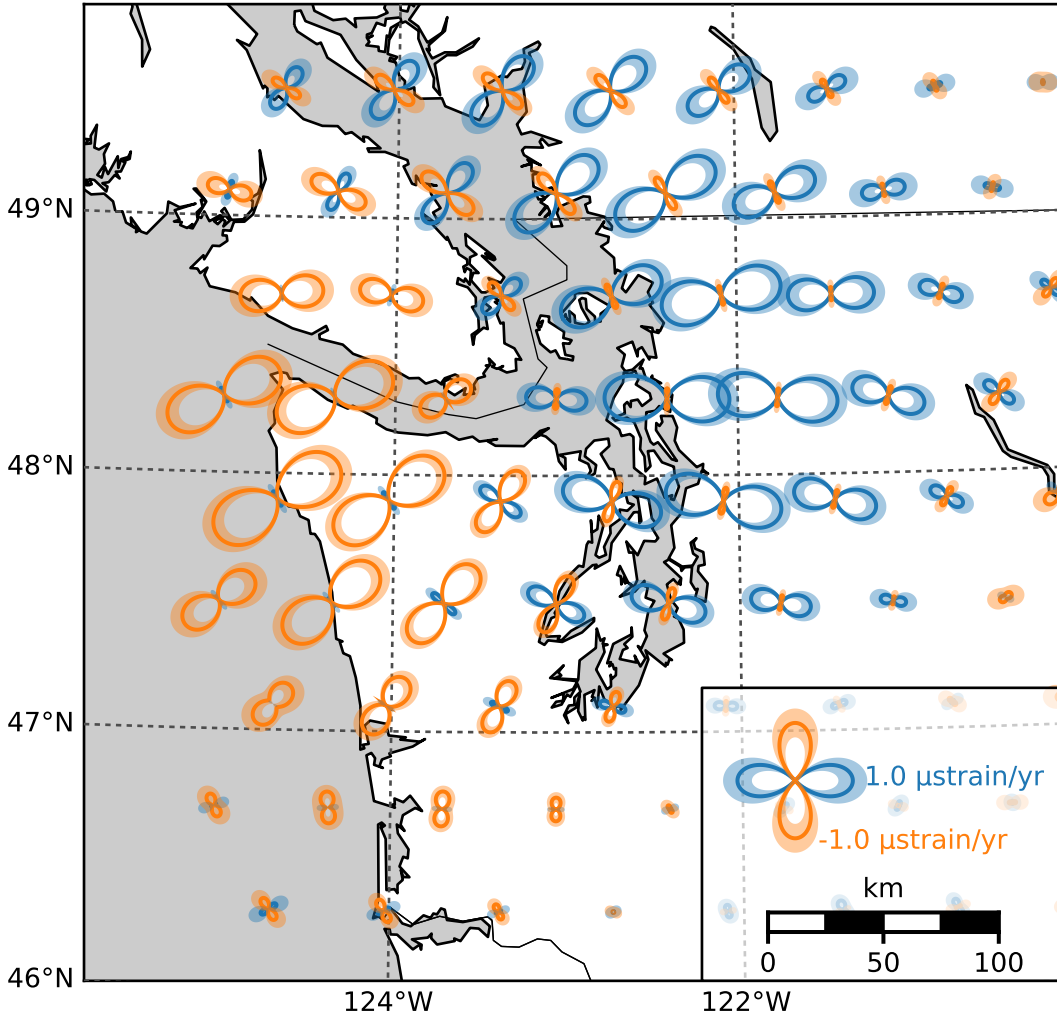


Figure 6: Strain map.

3.3 Transient Strain Rates

Having established a noise model and a prior for transient displacements, we use the cleaned GNSS dataset described in Section 3.2 to calculate transient strain rates in the Puget Sound region. We calculate transient strain rate on each day from January 1, 2010 to March 15, 2017. The strain rates are estimates on a grid of points spanning the study area. In Figure 6, we show the transient strain rates on January 1, 2016, which coincides with the height of an SSE. We have included a video showing the map view of strain rates over time as supplementary material. The strain rates shown in Figure 6 are generally similar to the strain rates for the other six SSEs considered in this study. The SSEs cause compression in the Olympic Peninsula and extension east of Puget Sound. For comparison, long-term strain rates indicate trench perpendicular compression throughout this study region (Murray and Lisowski, 2000; McCaffrey et al., 2007, 2013). The SSEs are thus concentrating tectonically accumulated strain energy trench-ward, and pushing the subduction zone closer to failure. Similar conclusions have been drawn based on fault slip models (e.g., Dragert et al., 2001; Wech et al., 2009; Schmidt and Gao, 2010), which reveal that SSEs are occurring downdip of the seismogenic zone. The SSEs, which are thrust events, migrate stress updip and push the locked portion of the fault closer to failure. A key difference between the strain inferred here and strain derived from fault slip models is that our estimated strain rates are not based on an assumed physical model. In contrast, fault slip models can be biased by errors in the assumed fault geometry or lithospheric rheology. Moreover, the degrees of freedom in fault slip models usually cannot be constrained by GNSS data alone, and it is necessary to impose regularization which further biases the results. Since our estimated strain rates lack such systematic errors, we can be more confident that our solution is unbiased and has meaningful uncertainties.

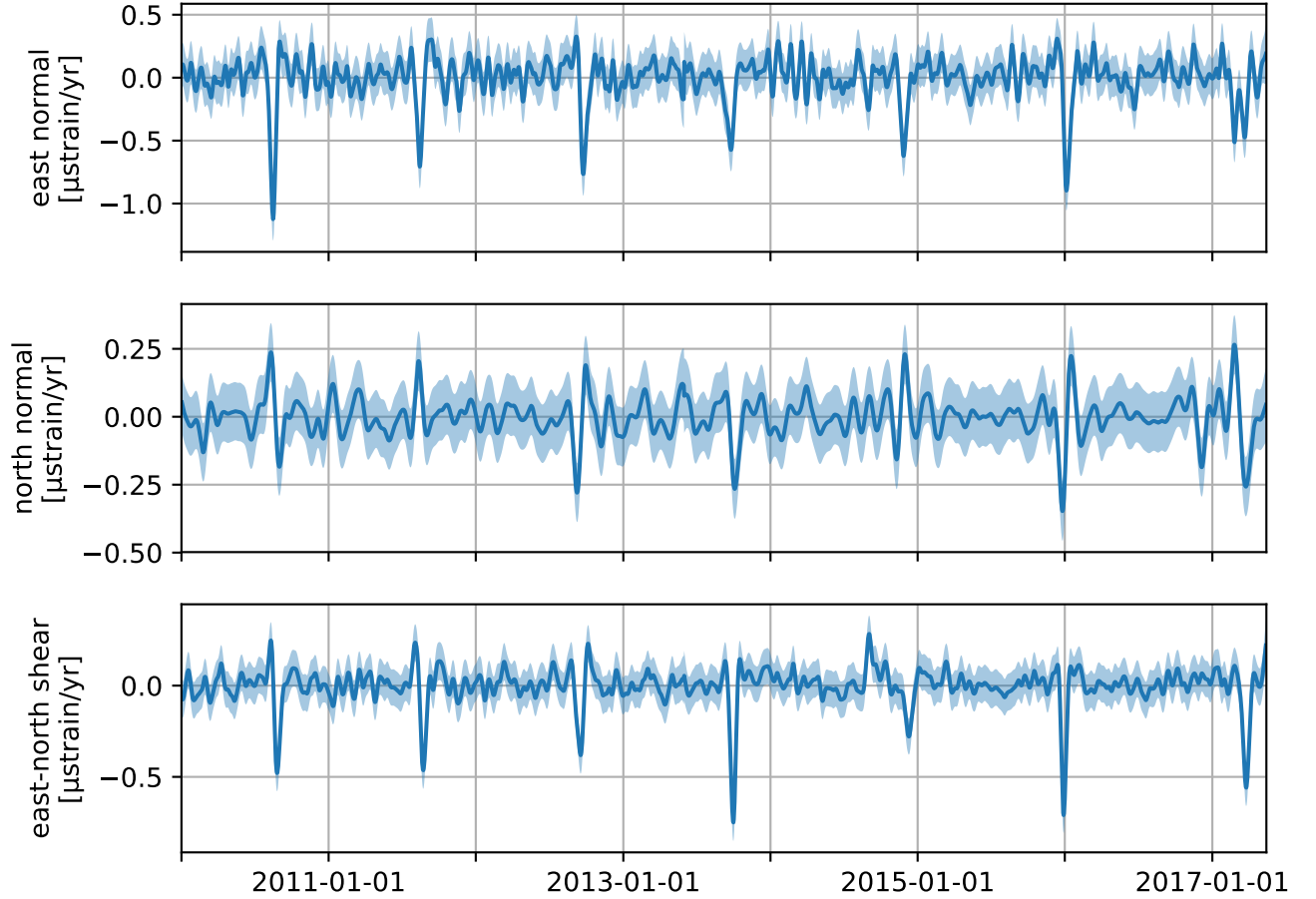


Figure 7: Strain time series.

In Figure 7 we show the time dependence of estimated transient strain rates at a position on the Olympic Peninsula, where strain rate magnitudes from SSEs are largest. To verify that the estimated strain rates are accurately identifying geophysical signal, we compare the strain rate magnitude to the frequency of seismic tremor (Figure 8). The shown strain rate magnitude has been normalized by the estimated uncertainty and can be considered a signal-to-noise ratio. Estimated strain rate magnitudes greater than ~ 3 can then be interpreted as a detected transient event. For each detected transient event there is a corresponding peak in seismic tremor. We are also able to clearly identify transient strain associated with a more subtle inter-SSE event in August 2014. In between SSEs the normalized strain rate magnitude is consistently between 0 and 2, indicating that no anomalous non-SSE events are being detected, at least at this location.

4 Other Potential Applications

In the previous section we demonstrated that transient strain rates can be used to detect SSEs or other transient geophysical phenomena. Another potential application would be to use the GNSS derived transient strain rates to develop noise models for borehole strain meters (BSMs). The plate Boundary Observatory maintains 43 BSMs in the Pacific Northwest, and it has been demonstrated that BSMs are able to record transient geophysical events such as SSEs (e.g., Dragert and Wang, 2011). However, noise in BSM data is not well understood, which complicates the use of BSM data for identifying geophysical signal. The noise consists, in part, of a long-term decay resulting from the instrument equilibrating with the surrounding rock (Gladwin et al., 1987). Typically, this noise is dealt with in an ad-hoc manner by fitting and removing exponentials and low order polynomials. If we use GNSS derived strain rates as an approximation of the signal in BSM data, then we could perform a more rigorous quantification of BSM noise.

We also note that there is potential for a more thorough analysis of the spatio-temporal noise in GNSS data, v , than what was performed in Section 3.1. In this study we have not explored the spatial covariance of v , since it would not have significantly affected our estimated strain rates. If, for example, we were interested in inferring fault slip from GNSS displacements then it may be necessary to account for the spatial covariance of

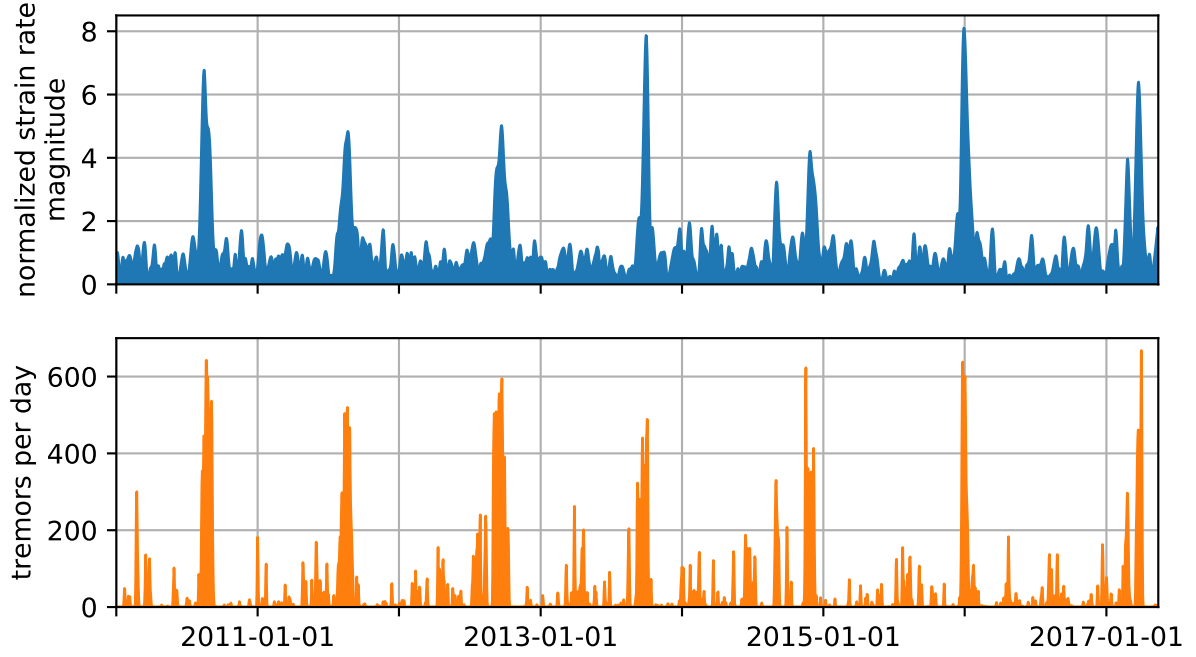


Figure 8: Strain magnitude time series.

the v .

In future studies, we may explore more sophisticated seasonal models than what is used in eq. (3). As suggested earlier, we may want to quantify the spatial covariance of seasonal deformation. We may also want to explore whether the semiannual terms in eq. (3) are necessary or if higher frequency terms are warranted. The periodic Gaussian process (Mackay, 1998) is an alternative model for seasonal deformation and is well suited for exploring the roughness of seasonal deformation. The periodic Gaussian process has zero mean and covariance function,

$$T(t, t') = \phi \exp \left(\frac{-\sin(\pi|t - t'|)^2}{2\theta^2} \right). \quad (30)$$

Realizations have annual periodicity and the roughness is controlled by θ . Decreasing θ has the same effect as including higher frequency sinusoids in the seasonal model. The optimal value for θ can be found with the REML method as described in Section 3.1.

Numerical Notes/scaling. Use partitioned matrix and Cholmod to solve for strain.

5 Conclusion

In this paper we describe how Gaussian processes can be used to identifying transient strain from GNSS data. Our approach is Bayesian and we use a spatio-temporal stochastic model to describe the prior for transient deformation and the noise in GNSS data. Our estimated transient strain rates are complete with meaningful estimates of uncertainties, which allows us to objectively discern geophysical signal from noise. Other commonly used methods for estimate long-term and transient strain assume a parametric form for the underlying signal, and errors in this assumption will result in a biased solution. In contrast our stochastic prior model is chosen with maximum likelihood methods, which is an objective and data-driven approach. The lack of subjectivity in our method gives us confidence that the results are unbiased.

We have demonstrated with GNSS data from Cascadia that Gaussian processes are effective for identifying transient strain. We also envision that Gaussian processes can be an effective tool to handle other geophysical things.

used to tackle other several other closely related applications

Since our method lacks subjectivity, the obto be most consistent with the data maximum likelihood methods. that is, constrained by the observed data. and is less likely the results and lead to this and We prefer a stochastic prior model over a parametric prior model such

The uncertainties on the estimate of transient strain can be trusted, provided that we have chosen appropriate stochastic models for the prior and noise. These models are objectively chosen with on maximum likelihoods methods, and

A stochastic prior model is preferable to a As a result, we obtain sensible

The material presented in this paper is primarily focused on GPS data; however, we speculate that the RBF-FD scheme can be of particular use in denoising borehole strain meter (BSM) data. The Plate Boundary Observatory has deployed X BSMs along the Western United States. BSM data contains low frequency drift resulting from relaxation of the borehole (Gladwin et al., 1987), which can obscure the geophysical signal of interest. We suggest that the RBF-FD scheme may be useful in denoising BSM data. Since the RBF-FD scheme provides a straight-forward mapping from GPS displacements to strain at any target locations, it is possible to use GPS derived strains as *a priori* information for strain at BSM sites. GPS derived strains could then aid in discerning tectonic signal from noise in BSM data.

Additional uses for strain: transient detection, prior for strain meters,

Additional potential applications of the RBF-FD method: regularizing inverse problems,

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