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NATURAL DISASTERS AND POLITICAL PARTICIPATION: THE CASE OF JAPAN AND THE 2011 TRIPLE DISASTER

Abstract
What effect do natural disasters have on political participation? Some argue that natural disasters decrease political participation because of the way they reduce individual and group resources. Others argue that they stimulate political participation by creating new social norms. Previous studies have been limited both by their focus on a specific disaster type and a lack of regional variation. This article advances the literature by assessing the effect of the 2011 triple disaster in Japan on political participation at both the individual and district level. Drawing on multiple sources of data, I use a difference-in-differences identification strategy to show that the 2011 triple disaster in Japan resulted in a 6 percent increase in participation in political groups in regions heavily affected by the disaster, and a 2.5 percent increase in voter turnout in districts in prefectures that were significantly affected by the disaster. The results also show that the effect at the individual level is largely confined to individuals with large social networks, suggesting that the effect of natural disasters on political participation is a combination of their direct and indirect impact on variables that operate through different subpopulations. Directions for future studies are suggested.

Keywords
Japan, natural disasters, political participation, social networks

INTRODUCTION

Natural disasters have an awful power. Aside from the immediate human, infrastructural, and economic damage they inflict, they can also have political repercussions that reverberate beyond the initial disaster period. Intuitively, one would expect natural disasters to depress political participation. As people rebuild their lives in the aftermath of a natural disaster, politics would seem to be a low priority. However, there is an emerging strand of literature that identifies a positive relationship between a variety of types of traumatic events and political participation (Bateson 2012; Blattman 2009; Sinclair, Hall, and Alvarez 2011; Fair et al. 2017). Natural disasters in particular have been found to occasion an increase in voting and political participation in highly affected areas (Sinclair, Hall, and Alvarez 2011; Fair et al. 2017). While scholars have proposed a number of theories that explain how natural disasters might elevate political participation, underlying mechanisms have rarely been directly assessed. Other studies cast doubt on this relationship entirely, finding empirical support for the negative impact that one would naturally expect (e.g., Rudolph and Kuhn 2018). What is the causal effect of natural disasters on political participation, and what might be underlying this relationship?

In this article I advance the literature by assessing the effect of the 2011 triple disaster on participation in political groups and voter turnout in Japan. The 2011 triple disaster constitutes a critical test case for theories that link the experience of natural disasters to increased political participation, owing to the sheer magnitude of the disaster. The large exogenous shock of the triple disaster generates conditionally as-if random variation that we can leverage to study disasters' effect on political participation. First, using a difference-in-differences causal identification strategy, I show that participation in political groups nearly doubled in regions highly affected by the disaster, and that electoral districts in affected prefectures experienced a 2.5 percentage point increase in turnout relative to non-highly affected prefectures.

The second finding of this article speaks to the debate about how disasters affect political participation. Examining heterogeneous treatment effects, I find that the impact of the disaster at the individual level is contingent on an individual's social network size: the participation-inducing effect of the disaster is observed only for individuals with large social networks. These results suggest that the positive impact of natural disasters may be a combination of direct and intervening variables that operate through different subpopulations, rather than uniformly. Future studies should seek to further specify how these intervening variables interact in subpopulations to produce this result.

TRAUMATIC EVENTS AND POLITICAL PARTICIPATION

Studies of the effect of natural disasters on political participation fall into a broader class of investigations into the relationship between traumatic events and political participation. Research in this area has caught the attention of scholars of political behavior largely because of counterintuitive findings regarding the relationship between the experience of traumatic events and participation. Whereas one would naturally expect traumatic experiences to reduce political participation, a number of studies show that they have the opposite effect. Blattman (2009), for instance, drawing on interviews with individuals abducted by Ugandan rebel leader Joseph Kony, shows that former abductees were more likely to be politically active later on in life. This suggests that, rather than frightening victims into political acquiescence, the experience of violence can spur positive behavioral changes that last long after the experience has ended.

Similarly, Voors et al. (2010) find that exposure to political violence results in increased levels of altruism, and Bateson (2012) finds that crime victimization has a strongly positive effect on political participation. Natural disasters in particular have begun to receive growing attention from scholars of political behavior over the past decade because of the way they can be used to understand the immediate short-term effects of exogenous shocks on voting behavior and political participation. Furthermore, a natural disaster is one type of traumatic event that can exert a substantial effect on politics and society, so another benefit of studying natural disasters is that by doing so we can gain analytical leverage on other types of traumatic events.

Natural disaster studies are theoretically important because some of their findings call into question long-standing theoretical models of political participation. As Brady (1999) notes, much of the literature argues that participation is largely a function of individual resources, such as income and education, and these resources are thought to be the major predictor of political participation (e.g., Verba, Scholzman, and Brady 1995). Because natural disasters reduce individuals' resources, it follows that they should also be expected to reduce political participation among affected individuals.

Some evidence supports these expectations. Kosec and Mo (2015), for instance, show that natural disasters decrease aspiration, which is strongly associated with political participation. Similarly, Rudolph and Kuhn (2018) find that the costs of voting resulting from flooding in Germany far outweighs any increase in political engagement linked to effective flood management. Bodet, Thomas, and Tessier (2016) find similar results in Canada. Additionally, there is evidence that disasters can result in political instability (Dal Bo and Dal Bo 2011; Dube and Vargas 2013), and that they are associated with a decrease in trust in government (Hommerich, 2012).

However, a growing number of studies find that natural disasters have a positive effect on political participation. Flooding caused by Hurricane Katrina in the United States has been found to have contributed to a rise in voter turnout in highly affected areas (Sinclair, Hall, and Alvarez 2011) as well as increased community participation (Han 2009). Fair et al. (2017) find similar results for flooding in Pakistan. A related vein of research identifies a relationship between candidates' share of the vote and their handling of natural disasters. Bechtel and Hainmueller (2011), for example, find evidence of increased voting for incumbents who were perceived to have successfully managed damaged caused by flooding in Germany, whereas Cole, Healy, and Werker (2012) find the obverse result in India, where incumbents experienced a decrease in their share of the vote in highly affected areas, owing to the way they were perceived to have mismanaged flooding there. To varying degrees, these types of studies call into question the existing understanding of turnout and political participation, in so far as they suggest that the relationship between individual resources and political participation is less straightforward than previously thought. Specifically, they suggest that the effect of a sudden reduction in individual resources can have the opposite of the effect expected by traditional theories of participation. If such effects are in fact generalizable phenomena and not coincidental products of a particular cases, then an updating of existing theories is warranted. A first step in evaluating the extent to which this is true is to examine the mechanism underlying these counterintuitive findings and see if it truly represents a divergence from existing theories.

EXPLAINING THE RELATIONSHIP BETWEEN NATURAL DISASTERS AND POLITICAL PARTICIPATION: DIRECT VERSUS INDIRECT

Explanations of the positive relationship between disasters and collective behavior can be broadly divided into two types. One possibility is that the experience of natural disasters causes a direct psychological impact on affected individuals, much akin to the individual psychological impact of other traumatic events on individual motivation to participate found by scholars like Blattman (2009) and Bateson (2012). In support of this explanation are a number of studies that show a direct effect on various types of cooperative behavior. Yamamura (2016), for instance, shows that community participation increased in areas significantly affected by the Hanshin-Awaji earthquake in Japan. Similarly, Cassar, Healy, and Kessler (2017) find an increase in social trust in areas heavily affected by natural disasters in Thailand, and Hawkins and Maurer (2010) find that bonding and bridging social capital helped ameliorated the impact of Hurricane Katrina in the United States.

More generally, natural disasters have been shown to increase pro-social behavior (Bardo 1978; Toya and Skidmore 2012). Thus, it is reasonable to expect this pattern to hold for political participation, as well, since an increase in pro-social attitudes might aid in overcoming the collective action dilemma. That is, by creating a pressing need to work together provide a public good, such as shelter or a means of gauging local radiation levels, natural disasters might result in a new norm of cooperative behavior that carries over into the political realm. Such an explanation finds some support in the work of Han (2009), who shows that non-political personal commitments and personal bonds, such as those formed amid the flooding of New Orleans in 2005 developed into long-term political commitments.

Another possibility is that natural disasters have a more indirect and largely coincidental effect on political participation. Fair et al. (2017), for instance, argue that flooding in Pakistan resulted in increased political engagement largely because of the way the government response to the disaster caused citizens to become more aware of the importance of government in their lives. Similarly, Sinclair, Hall, and Alvarez (2011) argue that the increased media attention to parts of New Orleans heavily affected by the disaster made residents of these areas more interested in and attentive to politics, thereby motivating them to vote. In both of these cases, it was less the experience of the disaster itself than the effect of other variables associated with the disaster that resulted in the increase in participation. These types of explanations point to a far more contingent and coincidental relationship than suggested by the former mechanism; one that depends largely on factors unrelated to the disaster should thus be expected to exhibit more variability than it would if it depended solely on the socio-psychological response to the disaster itself. Moreover, such explanations largely conform with existing theories of participation.

THE SOCIAL CONTEXT OF NATURAL DISASTERS AND EFFECT HETEROGENEITY

One way to begin to discriminate between these competing explanations is to look for heterogeneity in the treatment effect. If it is true that natural disasters affect political participation through their direct impact on social cooperation, then we should expect their effect to be more or less uniformly distributed across all affected individuals. If, however, disasters exert their effect through some exogenous variable, we would expect the increase in participation to be concentrated only among some subset of affected individuals. Accordingly, a first step in deciding between these explanations is to look for effect heterogeneity at the individual level conditional on a theoretically relevant variable that is correlated with the outcome variable but not the primary treatment variable.

Social network size is a suitable variable for this purpose. Social network size refers the number of people one interacts with on a regular basis (Wasserman and Faust 1994), and it is well-known to be positively associated a number of forms of political participation (Huckfeldt, Mendez, and Osborn 2004; Ikeda 2012; Mutz 2002). The participation-inducing effect of social network size is theorized to be a consequence of the way in which having a large number of contacts increases overall network heterogeneity,

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which, in this context, refers to the amount of variance in the political views of one's set of contacts (Nir 2005; Hu, Lin, and Cui 2015). Network heterogeneity, in turn, has been shown to increase participation through multiple pathways (Nir 2005, 2011; Barnidge et al. 2018). For instance, those with heterogeneous social networks are more likely to be exposed to competing viewpoints than those with more homogenous networks, and such exposure is in turn thought to motivate individuals to resolve these viewpoints through information seeking, a behavior that is widely known to be positively associated with political participation (Scheufele et al. 2006; Kim and Chen 2015).

We can leverage this relationship between social network size and participation to identify heterogeneity in the effect of natural disasters and political participation in the following manner. First, we observe that if natural disasters have a direct effect, such as by establishing a new shared social norm of participation, then we should expect the increase in participation to occur for individuals with both large and small networks, precisely because it should be a shared norm and not the result of stimuli to specific subunits of the treatment group. If, however, the effect is dependent on upon unobserved intervening variables, then we should expect the treatment to only have an effect on a specific subset of treated units. In this present case we might expect well-connected individuals to be more likely to experience the effect of such intervening stimuli by virtue of their social connectedness. To give an example, if the effect of natural disasters on political participation is a result of learning from government intervention, as suggested by Fair et al. (2017), individuals with large social networks might be more likely to experience this effect because they have a higher likelihood of learning about the effectiveness of government intervention through their many social contacts.

THE CASE OF THE 2011 JAPAN TRIPLE DISASTER

On March 11, 2011 a magnitude 9.0 struck off the coast of the Tohoku region of Japan. In addition to a series of large aftershocks, the quake also caused multiple tsunamis that struck up and down the eastern coast of Japan (Kazama and Noda 2012). It is estimated that the disaster resulted in 15,782 deaths, the full or partial destruction of 240,332 houses, and a total of about 16–25 trillion Yen (about $143–$224 billion USD) in structural damage (Kazama and Noda 2012, 783). Most of the causalities were concentrated in three prefectures located in the Tohoku region of Japan: Miyagi, Fukushima, and Iwate. Hence, the earthquake that caused the triple disaster is sometimes referred to as the "Tohoku Earthquake" (Mori, Takahashi, and the 2011 Tohoku Earthquake Tsunami Joint Survey Group 2012). However, significant damage and injuries were also reported in other prefectures, such as Ibaraki prefecture, Chiba prefecture, and Kanagawa prefecture, where the earthquake caused the largest amount of liquefaction recorded anywhere in the world (Kazama and Noda 2012, 790). The disaster also caused the meltdown of a nuclear reactor in Fukushima prefecture, resulting in radiation poisoning in areas around the nuclear plant, as well as widespread public concern about the spread of radiation in the adjacent prefectures of Chiba, Tochigi, and Saitama (Samuels 2013; Novikova 2016, 57). The combined earthquake, tsunamis, and nuclear meltdown are often referred to as a triple disaster.

The response to the disaster was immense. Samuels (2013) notes that minutes after the disaster then-Prime Minister Naoto Kan mobilized over 100,000 troops of the Japanese Self Defense Forces (SDF) to the "highest level" for the purpose of carrying out search and rescue operations, and the combined efforts of the Japanese government, SDF, and local authorities resulted in the successful evacuation of 270,000 people living in affected areas (Samuels 2013, 10). In the ensuing days and weeks, international aid poured into Japan, with the Chinese, Taiwanese, and Korean governments—among others—and NGOs providing over \$200,000,000 worth of humanitarian relief funds for those in affected prefectures, in addition to providing search and rescue teams. It is estimated that the Japanese business community alone received over 122.4 billion Yen for humanitarian relief (Samuels 2013, 18).

In many respects, the Japanese government's response to the disaster was effective, but in the days and weeks after the disaster criticism of the government was not in short supply. In particular, critics of the government noted that it suffered from "systemic shortfalls," such as the lack of a central emergency management agency, the absence of a comprehensive emergency response plan, and limited interagency communication (Samuels 2013, 9). As Samuels (2013) points out, part of the problem lay in the way Prime Minister Naoto Kan handled the situation, such as by creating ad hoc emergency tasks forces that were perceived to have undermined the authority of existing career civil servants (Samuels 2013, 10). The sudden appearance of a number of such task forces was in part responsible for perceived delays in the government response, as demonstrated by the government's five-hour delay in announcing the explosion of the Number One reactor at Fukushima (Samuels 2013, 12).

Another example is the discrepancy between the Japanese government's radius of mandatory evacuation resulting from the nuclear meltdown and that of the US Nuclear Regulatory commission, which set a far wider radius than did Japan. At the local level, many of the communities affected by the disaster were insufficiently incorporated into local governance structures, leaving affected individuals "cut off from first responders," and insufficiently accounted for in government reconstruction plans (Samuels 2013, 40). It was this perceived unreliability of government aid and information that led to the creation of citizens groups, such as the radiation watch group identified by Novikova (2016).

A number of authors have argued that the post-disaster increase in political participation observed by Mōri (2015), Aldrich (2012) and Ogawa (2015), among others, is a result of this type of community participation. Aldrich (2015), for instance, argues that increased levels of non-traditional political participation following the disaster are the result of social capital formed as a by-product of citizens' attempts to respond to the crisis (Aldrich 2015, 146). Similarly, Choate (2011) finds that the increase in political participation is a result of the sense of community fostered by the post-disaster community groups that emerged in response to the government's slow and sometimes ineffectual handling of the crisis. All of this echoes the causal logic outlined by scholars like Han (2009), Cassar, Healy, and Kessler (2017), and Yamamura (2016), who see disasters as resulting in social changes that can set the groundwork for political participation. Yet, while these accounts of the link between the triple disaster and the post-disaster rise in participation are suggestive, scholars have yet to establish this link empirically.

HYPOTHESES

The 2011 triple disaster was horrific, but it gives political participation scholars an opportunity to learn about the causal effects of disasters, as well as to explain the apparent post-disaster rise in certain kinds of political participation. We can think of the disaster as a conditionally as-if randomly assigned treatment. This assumption permits the identification of the effect of the disaster on political participation. This also represents a critical case in that, if there really is a relationship between natural disasters and participation, we should certainly expect to find one here given the immensity of the damage it inflicted. In what follows, I accomplish three tasks. First, I assess the effect of the triple disaster on participation in the Tohoku region of Japan, the most heavily and directly affected region in the country. Following the authors above who have argued that the disaster caused an increase in participation in political groups, I hypothesize that Tohoku will exhibit a post-disaster increase in average participation in political groups relative to non-affected regions.

> ***H1***: Tohoku will experience a relative increase in participation in political groups after the Triple Disaster.

Secondly, while some authors argue for a direct effect of the triple disaster on political participation via its effect on norms of participation, a large body of literature finds an indirect effect. As discussed above, if the effect is direct we should not observe heterogeneity conditional on social network size, whereas an indirect effect would be more likely to result in a heterogeneous treatment effect. Since there does not exist a preponderance of evidence on either side, I pose the following research question:

> ***RQ1***: Is the effect of the disaster identified in H1 conditional on social network size?

Third, I assess the effect of the disaster on the larger set of prefectures both directly and indirectly affected by the triple disaster. While participation in political groups is a good indicator of political participation, it is not the only one. Another indicator is voter turnout. Participation in political groups tends to be highly correlated with turnout, so if the proposed causal relationship between experiencing natural disasters and political participation proposed above is correct, we should also expect to see a rise in voter turnout in directly affected prefectures. Furthermore, we should also expect the disaster to have an indirect effect on voting in neighboring prefectures.

The reasoning here is as follows. Extant studies show that the disaster had an indirect effect on neighboring prefectures, largely owing to conflicting government estimates of the radius of dangerous radioactive particles (Recknagel 2011; Novikova 2016). As Novikova (2016) points out, this uncertainty spurred citizens in neighboring prefectures to organize radiation watch groups. It is reasonable to expect that this sort of spontaneous citizen activity would also have provided a stimulus to political activity, given that participation in non-political groups is a main pathway to political participation in Japan (Lee 2016). Accordingly, for this analysis I operationally define political participation as voter turnout, and I examine the effect of the disaster on voter turnout in the larger set of prefectures that were directly or indirectly affected by the disaster as identified in Kazama and Noda (2012) and Novikova (2016). These prefectures are Miyagi, Iwate, Fukushima, Ibaraki, Kanagawa, Tochigi, Chiba, and Saitama. I hypothesize that districts in these prefectures will experience an increase in voter turnout relative to districts in non-affected prefectures after experiencing the disaster.

H2: Districts in prefectures affected by the disaster will experience a relative increase in voter turnout.

DATA AND VARIABLES

I conduct two main analyses, the first of which is a pair of individual-level analyses, and the second of which is an analysis of district-level turnout data. In the first analysis, I analyze data from the 2005 and 2008 Japanese General Social Survey (JGSS), the 2010 and 2012 East Asian Social Survey (EASS). These two surveys contain identical measures of all of the variables under consideration, except where otherwise indicated. Here, the unit of analysis is the individual, and the outcome variable is a dummy variable that indicates whether or not a respondent participated in a political group. The treatment is given by a binary variable that indicates whether or not a respondent resides in Tohoku. This variable takes on a value of 1 if the respondent resides in Tohoku (treated group, N = 293) and a 0 otherwise (control group, N = 2042).

After demonstrating the effect of the disaster on participation in political groups, I then assess heterogeneity in this effect by assessing the extent to which the effect of the disaster is conditional on social network size. Social network size is given by an ordinal variable that is constructed by summing together two items on the JGSS and EASS, one that asks the respondent to select the number of family members they interact with on a daily basis, and another that asks the respondent the number of non-family members they interact with on a daily basis. Both of the items require the respondent to identify a range, so responses are recoded as ordinal variables that range from 0–6. Accordingly, this procedure results in a social network size index that ranges from 0 to 12, where a value of 0 indicates that a respondent does not interact with any people, and a value of 12 indicates that a respondent interacts with at least 200 people on a daily basis. The average value for this variable across regional groups and years is 3.412 (SD: 1.844).

I include two sets of controls. The first is a set of time-varying covariates that previous research has shown to affect political participation: political efficacy, political sophistication, and community participation. Political efficacy is included because it is possible that the change in participation might be due to an increase in the perceived potential to affect political change through participation, rather than the triple disaster. This explanation has been used to explain participation in Japan in previous studies (Almeida and Stearns 1998). Political efficacy is given by a survey item that asks respondents to state the extent to which they feel capable of influencing the government. The resulting ordinal variables ranges from 1 to 6, where a value of 1 indicates that a respondent does not feel capable of influencing the government at all, and a value of 6 indicates that a respondent feels highly capable of influencing the government.

Previous studies of the effect of natural disasters on political participation have also shown that they can increase awareness of politics (e.g., Sinclair, Hall, and Alvarez 2011). Political sophistication is used as a proxy for political awareness. Political sophistication is defined as the extent to which a respondent feels capable of understanding politics, and it is measured by an item that asks respondents to state the extent to which they do not feel that politics is too complicated to understand. The resulting ordinal variable has a minimum value of 1, which indicates that the respondent strongly feels that politics is too complicated to understand, and a maximum of 4, which indicates that the

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respondent feels confident in their ability to understand politics.[1] Community participation is included because increased political participation may be a consequence of increases in social capital, and community group participation is a commonly used indicator of social capital (e.g., Yamamura 2016). This is measured by a dummy variable that indicates whether or not a respondent participates in a non-political community organization.

Controls are also added for a number of demographic variables. Normally, a difference-in-differences analysis controls for time invariant factors, but in the case of a natural disaster of this magnitude, it is possible that observed changes in the dependent variable could be due to compositional changes in one of the regions. In order to assess this possibility balance checks were conducted for pre-disaster Tohoku and post-disaster Tohoku. These checks consisted of difference in means tests for age, sex, education, income. Age is a numerical variable that indicates a respondent's age in years. Sex is indicated by a variable that takes on a value of 1 for Male and 2 for Female. Income is given by an ordinal variable that ranges from 1 to 19, where 19 indicates the highest level of income. Education is indicated by an ordinal variable that ranges from 1 to 7, where 7 indicates the highest level of education (graduate degree). The results of the difference in means tests are shown in Table 1. Looking at Table 1, we see that the only figure that approaches statistical significance is the difference between average age in pre- and post-disaster Tohoku, such that the average age appears to have decreased somewhat in post-disaster Tohoku. Since age has been shown to be a strong predictor of political participation in Japan (Lee 2016), I control for respondents' age.

Another concern we might have is that of pre-treatment covariate balance between the treatment and control groups. Ideally, treatment and control groups should be roughly similar, especially in covariates thought to predict the outcome variable. In order to assess balance, difference in means tests were conducted for pre-treatment Tohoku and non-Tohoku. The results are shown in Table 2. The two groups do not show statistically significant differences in the covariates shown by previous studies to have a strong impact on political participation in Japan: age and sex. However, average household income and education level appear to be somewhat lower in Tohoku than in the rest of the country. I thus add a control for income, as well as for sex. Available data do not contain comparable measures of education, so I am not able to add a separate control for education, though the income variable should account for some income variation. Moreover, Tohoku did not experience a change in average education level, so this difference should not be expected to generate bias.

TABLE 1 Difference in Means Test for Pre and Post-Disaster Tohoku

	Pre-Disaster Tohoku Average	Post-Disaster Tohoku Average	P-Value
Age	56.856	53.755	0.049
Sex (Female)	1.525	1.541	0.759
Income	8.925	8.914	0.971
Education	4.206	4.236	0.761

TABLE 2 Difference in Means Tests for Pre-Treatment Regions

	Tohoku Average	Non-Tohoku Average	P-Value
Age	53.755	54.241	0.667
Sex (Female)	1.541	1.502	0.269
Income	8.914	9.463	0.012
Education	4.236	4.490	0.0001

A final concern I address for the first set of analyses is random assignment of treatment. An implicit assumption of the difference-in-differences analysis that will be employed here is exogeneity of the treatment. In this case the treatment is operationalized as residing in a region affected by the 2011 triple disaster.

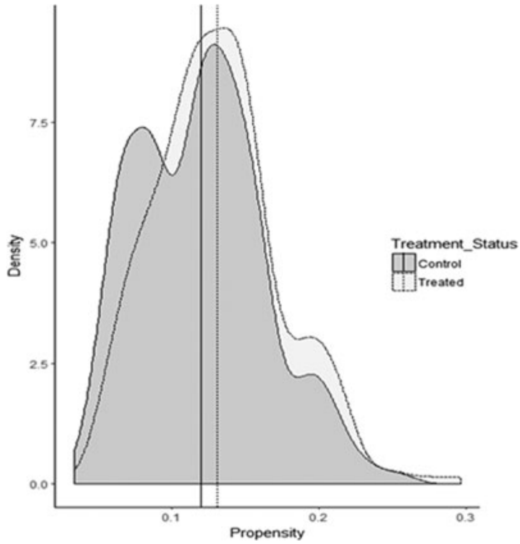
One way to assess exogeneity of treatment assignment is to see if the treated and non-treated groups differ in their propensity to receive the treatment conditional on some set of covariates thought to predict the outcome variable. Prima facie, it seems unlikely that Tohoku is somehow predisposed to experience natural disasters, since earthquakes are common throughout Japan, and its 54 nuclear reactors are spread fairly evenly across the four main islands (World Nuclear Association 2018). With regard to the propensity to experience earthquakes, there does not appear to be any evidence that Tohoku was more likely to experience earthquakes. In fact, the epicenter of the Tohoku earthquake is in a location considered by scientists to have had a low likelihood of major seismic activity (Boyle 2013). Last, people who live near the coast are in some sense predisposed to experience a tsunami, but Japan is an island nation, and about 80% of its inhabitants live in a coastal region (Hinrichsen 1998). So, it seems unlikely that Tohoku residents are somehow more likely to experience tsunamis as opposed to Japanese inhabitants of other coastal regions.

On the whole, then, it seems very unlikely that treatment assignment is endogenous. To be sure, a propensity score analysis was conducted. This consisted of regressing the treatment dummy on age, sex, income, and education using logistic regression. Treatment propensity is defined as the predicted probability of receiving the treatment, and it is given by the predicted probabilities extracted from a logistic regression. The distribution of propensity scores for the treatment and control groups are shown in Figure 1. The means of the treated and control groups are indicated by vertical lines. Figure 1 shows considerable overlap between the two distributions, and the means of the two groups are fairly close, further supporting the assumption of exogeneity of treatment assignment.

In the second analysis the unit of analysis shifts to electoral district. Here, I analyze district-level voter turnout from the 2003, 2005, 2009, 2012, and 2014 House of Representatives elections for the PR tier. The turnout data come from Steven Reed's "PR by SM" data set, which contains PR vote data by single member electoral district. Prefecture covariates were obtained from the Japanese Ministry of Internal Affairs and Communications (MIAC).

The primary outcome variable for this analysis is district-level PR voter turnout, which is defined as the number of voters divided by the number of registered voters in each district in each election year. The treatment is indicated by a dummy variable that takes on

FIGURE 1 Propensity Score Distributions for Treatment and Control Groups



the value of 1 if a district belongs in one of the prefectures that were significantly affected by the disaster—as defined above— and a value of 0 otherwise.[2] Treated prefectures are defined as prefectures that were significantly affected by the disaster, both directly and indirectly. I choose this approach because the effect of this particular disaster was not limited to districts in which the disaster had a scientifically measurable impact, such as in prefectures surrounding Fukushima where citizens were not certain of the extent of radiation, but nevertheless acted on the assumption that radiation contamination could occur at any moment (Novikova 2016, 72). For every time period except 2003–2005 I add controls for per capita income, population change, and the percent of the population that is between age 15 and 65.[3] Again, these controls might plausibly predict political participation. I also control for population change in order to further take into account compositional changes.

METHOD

I employ a difference-in-differences identification strategy in both sets of analyses. Difference-in-differences (DD) is a causal identification strategy that identifies the effect of a given treatment by finding the difference in the outcome variable between the treated and control groups in two or more time periods. In order for the DD approach to work we must assume that there is no large difference in the outcome variable between the treatment and control groups in the pre-treatment period (pre-2011 in this case), otherwise known as the parallel trends assumption. This assumption assures that the treatment and control groups are in fact comparable, and that observed differences in the outcome variable are not due to trends unrelated to the treatment (see Supplementary

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Material for full mathematical derivation). Here, the average treatment effect on the treated (ATT) is estimated with OLS through the following interaction model, where α_3 gives the ATT, and $\mathbf{X}_{it}$ is a matrix of time-varying confounders:

$$Y_{it} = \alpha_0 + \alpha_1 D_{it} + \alpha_2 T_{it} + \alpha_3 D_{it} T_{it} + \mathbf{X}_{it}\gamma + \varepsilon_{it} \tag{1}$$

I show heterogeneity in the individual-level effect of the earthquake as follows. First, the data is split into two groups: a small social network group and a large social network group. Then, the OLS interaction models shown in equation (1) are estimated for each subset. I operationally define a large social network individual as one with a score of 5 or higher on the 12-point social network size index.

In the second analysis, a difference-in-differences analysis is conducted on district-level turnout data. The analysis for this section is essentially the same as the first, with the exception that the unit of analysis is electoral district, and the outcome variable is PR voter turnout. PR turnout is chosen over SMD turnout because the characteristics of the particular candidates run in a given district present a further identification challenge, whereas all voters face the same set of parties in the PR tier. The analysis is conducted on matched prefectures in order to ensure maximum comparability between treatment and control groups, thereby reducing sensitivity of the results to model specification. Prefectures are matched on the following 2010 prefecture-level covariates using rescaled Euclidean distance: population density, the percent of the population aged 15 to 65, per-capita income, unemployment, net population change, the percent of workers employed in primary industries, the percent of workers employed in secondary industries, and the percent of workers employed in tertiary industries. These covariates are chosen in order to ensure maximum similarity between treated and control regions. Balance checks were performed in order to ensure that the matching procedure was successful. The results of these tests, shown in Table 3, suggest that there are no significant differences between the matched prefecture pairs on any observable covariates. The matched data set contains electoral data for 150 unique districts (73 treated and 77 control) for each of the five years, yielding a total of 750 observations (300 for each pair of years). For this analysis the set of years is $t \in \{2003, 2005, 2009, 2012, 2014\}$. The data for this analysis is estimated with the

TABLE 3 Difference in Means Tests for Matched Prefecture Covariates

	Treated Prefectures	Control Prefectures	P-Value
Population Density	1,716.000	2,010.170	0.780
Per Capita Income	2,742.500	2,709.750	0.810
% of Population Age 15–65	64.060	63	0.310
Unemployment Rate	6.650	6.670	0.950
Net Population Change	−0.150	−0.100	0.600
Share of Workers in Primary Industry	5.170	4.450	0.690
Share of Workers in Secondary Industry	24.840	24.510	0.890
Share of Workers in Tertiary Industry	65.640	65.280	0.880

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following fixed effects model:

$$Y_{it} = \delta_t + \alpha D_{it} + \mathbf{X}_{it}\beta + \varepsilon_{it} \tag{2}$$

Here, δ_t is a time period dummy, α is the estimate of the ATT, and $\mathbf{X}_{it}$ is a matrix of time-varying control variables. A fixed effects model is chosen for this analysis because the data track turnout in the same districts over time, and hence constitute panel data and not repeated cross-sectional data as in the first set of analyses. Equation 2 is estimated separately for pairs of consecutive time periods (i.e., 2003–2005, 2009–2012, 2009–2014).[4] In order to test empirically the parallel trends assumption, placebo DDs are performed for the 2003–2005 and 2005–2009 time periods. All models in this section are run on first-differenced data, and they are estimated with district-clustered standard errors in order to control for serial correlation.

RESULTS

The discussion of the results proceeds as follows. First, the results for the test of the individual-level effect of the triple disaster are presented. These are followed up by the test of heterogeneity in this effect. Then, the results for the test of the effect of the disaster on prefecture-level turnout are presented. These results are followed by a discussion of the implications for theories of political participation.

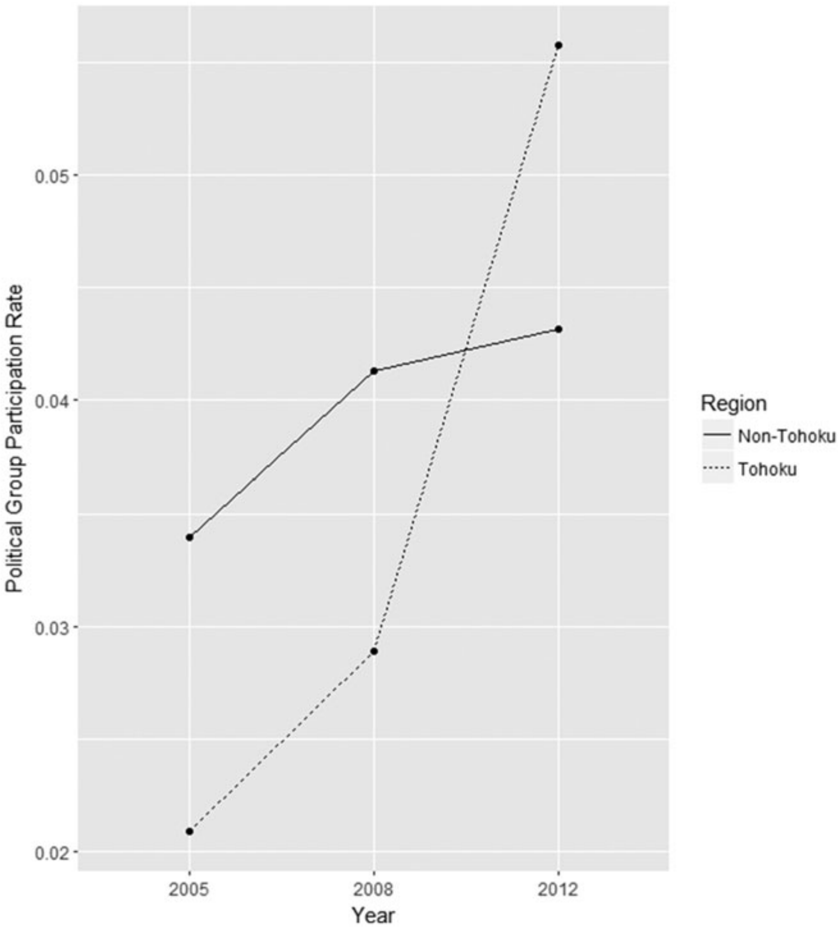
RESULTS OF THE INDIVIDUAL-LEVEL ANALYSIS

Figure 2 shows the fraction of respondents in each group reporting that they participated in a political group. The figure clearly shows a parallel trend in participation prior to the disaster. After the disaster, however, participation in political groups increases fairly dramatically in Tohoku. This provides visual support for the parallel trends assumption. In other words, it seems likely that the trend observed in the pre-treatment period would have continued absent the treatment, which suggests that the two groups are otherwise similar.

Table 4 shows the results for the test of H1. Model 1 in Table 4 quantifies the results shown in Figure 2; this is the placebo DD for the pre-treatment period 2005–2008. The coefficient on the Post-Period*Tohoku interaction term is very close to zero, and it does not reach statistical significance. This is consistent with the parallel trends assumption, since it suggests that Tohoku and the rest of Japan would have proceeded on a similar trajectory had the disaster not occurred.

Model 2 in Table 4 tests the main hypothesis that the triple disaster had a positive effect on political participation amongst residents of the Tohoku region (H1). The results for Model 2 show that participation in political groups in post-disaster Tohoku increased by about 6.3 percentage points relative to the control group. This is an almost two-fold increase in participation in political groups over the baseline participation rate of 2.8 percent in 2008. The 95 percent confidence interval for the effect of the disaster is quite large, ranging from about 0.4 to about 12 percentage points, but the coefficient is statistically significant at the 0.05 level. Note that this effect obtains even when the control variables are included in the regression, suggesting that the effect of the disaster is independent of these time varying confounders.

FIGURE 2 Participation in Political Groups in Treatment and Control Groups Across Time



ASSESSING HETEROGENEITY IN THE INDIVIDUAL-LEVEL EFFECT OF THE 2011 TRIPLE DISASTER ON POLITICAL PARTICIPATION

Having identified the effect of the 2011 Japan triple disaster on political participation (H1), I now turn to the test of heterogeneity in its effect (RQ1). The results for the test of RQ1 are shown in Table 5. The results for Model 1 in Table 5 show a roughly 37 percentage point increase in participation in political groups for those with large social networks ([95% CI: [0.034, 0.700]). This figure is statistically significant at the 0.05 level.

By contrast, Model 2 in Table 5 shows about a one percent increase in participation in political groups for those with smaller social networks, but it does not reach statistical significance. The effect size in the large network group less that of the small network group is 0.356 (SE: 0.172), and the 95 percent confidence interval of this difference ranges from 0.694 to 0.018, so this difference is statistically significant at the 0.05 level. These results support the argument for an indirect effect in so far as they

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TABLE 4 Results for Placebo Test and Main Effects

	Dependent Variable: Participation in a Political Group	
	2005–2008	2008–2012
Model	(1)	(2)
Post Period	0.007 (0.006)	0.026*** (0.010)
Tohoku	−0.010 (0.010)	−0.014 (0.012)
Tohoku*Post-Period	−0.004 (0.014)	0.063** (0.030)
Age	0.007*** (0.002)	0.009*** (0.003)
Sex (Female)	−0.028*** (0.005)	−0.006 (0.008)
Income	0.002*** (0.001)	0.005*** (0.002)
Community Participation	0.101*** (0.017)	0.088*** (0.023)
Efficacy		0.002 (0.004)
Political Sophistication		−0.023*** (0.005)
Constant	−0.020 (0.011)	−0.005 (0.030)
Observations	4,914	2,390

*p < 0.10; **p < 0.05; ***p < 0.01
Note: Robust standard errors are shown in parentheses.

suggest that the political participation-inducing effect of the disaster is largely concentrated amongst individuals with large social networks, rather than a widely shared norm.

ASSESSING ALTERNATIVE EXPLANATIONS FOR INDIVIDUAL-LEVEL EFFECT

Scholarship on political participation in Japan suggests a number of alternative explanations for the results observed above. One possibility is that the post-disaster increase in political participation is a result of changes in structural variables. Japan scholars have often looked to structural theories to explain fluctuations in Japanese political

TABLE 5 Results for Split Models (H2)

	Dependent Variable: Participation in a Political Group	
	Large Social Network Group	Small Social Network Group
Model	(1)	(2)
Post Period	0.007 (0.036)	0.010 (0.008)
Tohoku	−0.058* (0.035)	−0.009 (0.015)
Tohoku*Post-Period	0.370** (0.171)	0.013 (0.022)
Community Participation	0.172* (0.089)	0.135*** (0.024)
Age	−0.003 (0.010)	0.013*** (0.002)
Sex (Female)	−0.050 (0.037)	−0.023*** (0.008)
Income	−0.0001 (0.008)	0.003** (0.001)
Constant	0.079 (0.100)	−0.055*** (0.020)
Observations	159	3,026

*p < 0.10; **p < 0.05; ***p < 0.01
Note: Robust standard errors are shown in parentheses.

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participation. Almeida and Stearns (1998), for instance, use the political opportunity model developed by McAdam (1982), Tarrow (1994), and others to explain changes in participation in Japan. In particular, they argue that Japan's Large Grassroots Environmental Movements (LGEM) gained steam when fractures in the central government, like policy disagreement between the Ministry of Health and Welfare (MHW) and the Ministry of International Trade and Industry (MITI), energized local grassroots movements by signaling the vulnerability of the growth-oriented bureaucratic consensus, creating an incentive to participate because it was believed that citizens movements would be able to enact policy change by exploiting these divisions.

Conversely, after achieving a fair degree of political success, the LGEMs declined as a result of changes in Japan's political opportunity structure, such as the central government's decision to move civil litigation to local courts, the end of intra-bureaucratic conflict, and changes in the government's stance towards pollution legislation (Almeida and Stearns 1998, 52). However, if this sort of dynamic is what is driving the effect of the 2011 triple disaster on political participation, we would have expected the efficacy variable in Table 4 to be both positive and statistically significant, since the theory suggests that openings in the political opportunity structure will increase political efficacy. However, while the efficacy variable is positive, it does not come close to conventional levels of significance.

Another possible explanation is that the effect is due to the influx of humanitarian relief funds and government aid discussed above. These relief funds might have temporarily provided citizens with more resources, thus increasing their propensity to engage in voting or participation in political groups. Yet, this argument lacks plausibility for a number of reasons. First, while it comports with certain theoretical expectations, it does not comport with the dynamics of Japanese political participation discussed above; namely, that increased wealth is not associated with increased participation (Lee 2016). Second, since I have controlled for self-reported wealth, the results of the analysis should be robust to fluctuations in individual resources.

A final explanation suggested by the literature is that the rise in political participation will have been occasioned by an increase in community participation. Looking at Table 5, we see that the coefficient on community participation is indeed positive and statistically significant. Yet, the results of the analysis are robust to this specification, suggesting that social network size is exerting an effect that is independent or distinct from that of community participation.

ANALYSIS OF DISTRICT-LEVEL VOTER TURNOUT

I now turn to H2, which concerns the effect of the disaster on the larger set of prefectures identified in the literature as having been directly or indirectly affected by the 2011 triple disaster. Table 6 presents the results of this second set of analyses. Recall that this set of analyses was performed on district level turnout data, so the outcome variable is now district-level voter turnout, and the treatment variable is a binary variable indicating whether or not a district belongs in a prefecture that was highly affected by the disaster. Model 1 and Model 2 in Table 6 test the parallel trends assumption required to identify the effect of the disaster. The results for both Model 1 and Model 2 strongly support the parallel trends assumption. For both models, the coefficient on the treatment variable ("Disaster")

TABLE 6 Results for Matched Turnout Data (H3)

	Dependent Variable: PR Turnout			
	2003–2005	2005–2009	2009–2012	2009–2014
Model	(1)	(2)	(3)	(4)
Post Period	0.100*** (0.008)	0.002 (0.015)	−0.176*** (0.025)	−0.287*** (0.023)
Disaster	−0.007 (-0.005)	0.005 (0.003)	0.022*** (0.006)	0.025*** (0.004)
Per Capita Income		−0.00003** (0.00002)	0.00004** (0.00002)	0.0001*** (0.00002)
Population Change		−0.028 (0.068)	0.510** (0.200)	−0.156 (0.143)
Population Aged 15–65		0.002 (0.003)	−0.047*** (0.006)	−0.033*** (0.006)
Observations	300	300	300	300

*p < 0.10; **p < 0.05; ***p < 0.01

Note: District clustered robust standard errors are shown in parentheses.

is very close to zero and is far from being statistically significant, meaning that it is likely that the two groups would have continued on similar paths had the disaster not occurred.

By contrast, the results for the post-disaster models show a clear rise in turnout in districts in treated prefectures relative to districts in the control group. Looking at Model 3 in Table 6, we see a roughly 2.2 percentage point (95% CI: [0.01, 0.03]) increase in turnout in districts located in prefectures significantly affected by the disaster in the year after the disaster (2012). The results for Model 4 show that this effect remains more or less constant in 2014. On average, in 2014 districts in treated regions experience a roughly 2.5 percentage point increase in voter turnout relative to the control districts (95% CI: [0.017, 0.032]). Overall, these results strongly support H2, though the effect size is somewhat smaller than the effect size identified in other studies, such as the 3 percentage point effect identified by Sinclair, Hall, and Alvarez (2011), and is about 40 percent of the average change in turnout across the years in the sample (0.064) and about 1/3 of the standard deviation of turnout across all districts and years.

ROBUSTNESS CHECK: ENTROPY BALANCING

In order to assess the robustness of the results shown in Table 6, I conduct the same analysis on entropy balanced data. Entropy balancing is a data pre-processing method for reducing model dependence of causal estimates. It is thought to be an improvement over propensity score matching in that it retains a reweighted version of the full data set, hence does not entirely discard data. Following the procedures described in Hainmueller (2012), a vector of weights was obtained that was then used to estimate a weighted least-squares version of the model shown in equation 2. The results of this estimation are shown in Table 7. Overall, the results do not diverge significantly from those derived from the analysis of the matched data. Looking at Table 7 we see that the only major difference is that the estimate of the treatment effect obtained from the analysis performed on the entropy balanced data is more than a full percentage point lower than that obtained from the matched data. In contrast to the 2.5 percent increase in turnout shown in Table 6,

TABLE 7 Regression Results for Entropy Balanced Data

	Dependent Variable: PR Turnout			
	2003–2005	2005–2009	2009–2012	2009–2014
Year	0.090 (0.002)	0.007 (0.014)	−0.119**	−0.211**
Disaster	0.005 (0.003)	−0.005 (0.003)	0.012** (0.003)	0.012** (0.004)
Per Capita Income		−0.00002** (0.00002)	−0.00001 (0.00001)	0.00003** (0.00001)
Population Change		0.009 (0.033)	0.201* (0.087)	−0.296** (0.106)
Population Aged 15–65		0.001	−0.016** (0.003)	−0.003 (0.004)
Observations	600	600	600	595
Adjusted R-Squared	0.827	−0.975	0.841	0.878
F Statistic	2,384.742**	134.632**	1,901.954**	3,465.620**

*$p < 0.05$; **$p < 0.01$

Table 7 shows a roughly 1.2 percentage point increase in voter turnout in treated districts (95% CI: [0.004, 0.018]). Moreover, there is very little overlap between the 95 percent confidence intervals of the two estimates, suggesting that this difference is marginally statistically significant.

To put this figure on context, this effect is less than 1/5 of the effect size reported by Fair et al. (2017) in Pakistan, and about 2/5 of the 3 percentage point increase in participation found by Sinclair, Hall, and Alvarez (2011). Also, considering that the average overall change in turnout between each pair of years is about 6 percentage points, it seems fair to say that both of the estimates are relatively small.

CONCLUSION

This study does indeed find support for a positive relationship between natural disasters and political participation. However, the results of the analyses also suggest that this relationship has less to do with the direct socio-psychological effect of the disaster than with other potential factors: it shows that the increase in political participation observed for those in the treatment group obtains only for treated individuals with large social networks, and this result is robust to indicators of a direct socio-psychological effect. While it is not possible to identify this variable using the present data, the literature suggests a number of possibilities worthy of further exploration. For instance, it could be that that the observed treatment effect heterogeneity is a function of access to government resources, perhaps because well-connected individuals are more likely to obtain such resources. Another possibility is that it is due to network effects, like post-hoc information seeking. I leave it to future researchers to assess these and other explanations.

Secondly, while the analysis does show that the triple disaster increased voter turnout in affected districts, the effect is much smaller than that found in previous studies. Further, this estimate varies considerably by data pre-processing method. This instability of the effect estimate lends some support to the work of Bodet, Thomas, and Tessier

(2016), who find that disaster effect estimates are highly sensitive to estimation procedure. So, while natural disasters do appear to have a positive effect on political participation, researchers should scrutinize this effect carefully. Future studies ought to bring more attention to intervening variables in order to precisely identify the relationship between natural disasters and political participation. Accounting for the distribution of specific types of government intervention or prior local government administrative capacity might be useful in this regard. Finally, it is difficult to generalize based on a single case study. Scholars would do well to engage in comparative studies of the relationship between natural disasters and participation in order to better account for the role of context in moderating this relationship.

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CONFLICTS OF INTEREST

Matthew Jenkins declares none.

SUPPLEMENTARY MATERIAL

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NOTES

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1. While this is not a perfect measure of political awareness, it should act as a reasonably good proxy, since we would expect news reading and other information gathering activities to be highly correlated with the feeling that one understands politics.

2. A prefecture is a subnational administrative jurisdiction, roughly comparable to a US state.

3. Ideally, one would control for all matching variables. However, sufficient data is only available for these covariates.

4. Note that this is exactly equivalent to the interaction model in equation 1 since only data from two elections years is analyzed at a time.

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