LETTER

Population matters when modeling hurricane fatalities

Jung et al. (1) find gender bias leads to misperception of hurricane risk in both experimental and historical evidence. We affirm the rich literature on gender bias. However, we argue that Jung et al.'s empirical analysis suffers from endogeneity. Once addressed, we find the previous results to be of questionable robustness. Although gender bias may exist in limited-information experiments, historical evidence does not indicate gender bias in hurricane fatalities.

Damages do not determine deaths, but rather both are simultaneously determined by multiple factors, including hurricane characteristics and the (omitted) underlying population and vulnerability (2), which lead to endogeneity, or correlation between damages and the error term that can bias estimated coefficients. This result necessitates either instrumentation or elimination of "damages" from the specification's right-hand side. Additionally, the negative binomial estimator assumes the hazard potential is identical for each observation. By using "deaths" as the dependent variable, Jung et al.'s (1) study assumes the treated population in each case is identical which, as Pielke et al. (3) show, is not true.

We mend the endogenous specification in three ways. First, we removed the endogenous damages. Second, we included underlying population at risk as explanatory variables, testing both annual US population data and the average population density of the five coastal counties surrounding the point of landfall. (No annual county-level data exist back to 1950. We calculated the 2000 county-to-country population density ratio and, assuming it is constant across time, we scaled the ratio by annual US population density. We recommend future work refine this assumption.) Third, we normalized deaths by (*i*) dividing deaths by real damages (nominal damages taken from the International Catastrophe Insurance Managers; deflator information taken from the Bureau of Economic Analysis; population taken from the US Census) and (*ii*) dividing deaths by total US population in the year of landfall (we normalized deaths by the five-county population and found similar results). After a log-transformation, these measures are approximately normally distributed so ordinary least squares is appropriate.

Table 1 presents our results. The first "all deaths" column replicates the author's main result. The next three "all deaths" columns present negative binomial results controlling for underlying population at risk. Conditional on population, we find a reversal of Jung et al.'s (1) finding: storm name femininity is now protective. However, the joint impact of femininity and interactions is not statistically different from zero. The last two columns present models of the deaths—US population and deaths—damages normalizations. Both specifications find no increase in fatalities for more feminine-sounding storms.

The experiments in Jung et al.'s study (1) are interesting but the motivational facts are of questionable robustness. We establish this finding by controlling for population and correcting for endogeneity. Further research on the subject of hurricane naming is therefore warranted and encouraged.

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1 Jung K, Shavitt S, Viswanathan M, Hilbe JM (2014) Female hurricanes are deadlier than male hurricanes. *Proc Natl Acad Sci USA* 111(24):8782–8787.
2 Malter D (2014) Female hurricanes are not deadlier than male hurricanes. *Proc Natl Acad Sci USA* 111(34):E3496.
3 Pielke RA, Jr, et al. (2008) Normalized hurricane damage in the United States: 1900–2005. *Nat Hazards Rev* 9(1):29–42.
4 Jung K, Shavitt S, Viswanathan M, Hilbe JM (2014) Reply to Christensen and Christensen and to Malter: Pitfalls of erroneous analyses of hurricanes names. *Proc Natl Acad Sci USA* 111(34):E3499–E3500.

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Table 1. Regression results

Dependent variable	All deaths	All deaths	All deaths	All deaths	Ln (Deaths − US population ratio)	Ln (Deaths − real damage ratio)
Pressure	−0.552***	−0.693***	−0.645***	−0.666***	−0.752***	0.777***
	(0.147)	(0.137)	(0.161)	(0.134)	(0.150)	(0.149)
Normalized damages	0.863***					
	(0.333)					
Gender Index (female = 10, male = 0)	0.172	0.453***	0.287*	0.387**	0.00673	0.126
	(0.118)	(0.166)	(0.167)	(0.162)	(0.142)	(0.133)
US population		0.00865		0.0216		−0.716***
		(0.137)		(0.143)		(0.134)
County population density			0.365**	0.306*	0.542*	−0.0612
			(0.158)	(0.175)	(0.301)	(0.202)
Gender Index × Pressure	0.395**	0.162	0.268*	0.152	0.0405	−0.272**
	(0.157)	(0.115)	(0.146)	(0.113)	(0.128)	(0.125)
Gender Index × Normalized damages	0.705***					
	(0.246)					
Gender Index × US population		−0.348**		−0.336**		−0.304**
		(0.143)		(0.145)		(0.133)
Gender Index × County population density			−0.548***	−0.480***	−0.563*	−0.0424
			(0.169)	(0.170)	(0.309)	(0.212)
Constant	2.476***	2.585***	2.782***	2.637***	−17.20***	−4.286***
	(0.125)	(0.182)	(0.203)	(0.183)	(0.152)	(0.143)
Observations	92	92	92	92	82	82
Estimator	NB	NB	NB	NB	OLS	OLS
F test: all gender parameters sum to zero (two-sided *P* value)	0.00112	0.264	0.980	0.400	0.195	3.62e-05

SEs are in parentheses. ***$P < 0.01$, **$P < 0.05$, *$P < 0.1$. All specifications show robust (White) SEs as Jung et al. (4) suggest. The results are qualitatively the same without using robust SEs.