Governor partisanship explains the adoption of statewide mandates to wear face coverings

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Abstract. Public mask use has emerged as a key tool in response to COVID-19. We develop and document a classification of statewide mask mandates that reveals variation in their scope and timing. Some U.S. states quickly mandated the wearing of face coverings in most public spaces, whereas others issued narrow mandates or no mandate at all. We consider how differences in COVID-19 epidemiological indicators, state capacity, and partisan politics affect when states adopted broad mask mandates. The most important predictor is whether a state is led by a Republican governor. These states were much slower to adopt mandates, if they did so at all. COVID-19 indicators such as confirmed cases or deaths per million are much less important predictors of statewide mask mandates. This finding highlights a key challenge to public efforts to increase mask-wearing, widely believed to be one of the most effective tools for preventing the spread of SARS-CoV-2 while restoring economic activity.

Keywords: masks; COVID-19; U.S. states; public policy; partisanship

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Introduction

Public mask wearing is now widely viewed as a low-cost and effective means for reducing SARS-CoV-2 virus transmission (1; 2; 3). However, it was not until April 3, more than a month after the first reported case of the novel coronavirus in the US, that the CDC formally recommended mask wearing to the general public (4). Across the U.S., voluntary adherence to the CDC's mask recommendation has been uneven. Unlike some other societies, mask wearing in response to contagion is not a cultural norm in the U.S. (5). The absence of such a norm or a national mask mandate has resulted in considerable policy variation across states (6). As with other non-pharmaceutical interventions (NPIs), such as business and school closings and stay-at-home directives (7), many U.S. states did not broadly require that citizens wear masks across a range of indoor public spaces, even as the scientific and public heath case for mask wearing grew stronger. There was also considerable variation in how quickly states adopted mandates, among those that did.

At first glance, this variation appears to fall sharply along political party lines. For example, sixteen of the 17 states that have not adopted broad mandates are led by Republican governors. It is also the case that most of the early-adopting states were led by Democratic governors. But it is possible that first impressions fail to account for other factors. The mounting scientific evidence that masks are an effective means for slowing the spread of SARS-CoV-2 makes it more important to understand the most important drivers of state COVID-19 responses. Using originally collected data on mask mandates across states, we examine how differences in COVID-19 indicators, state capacity, and partisan politics may have affected the speed of mask mandate adoption. Specifically, we recorded when a state issued a mask mandate (if it did), developed a three-point scale to classify the breadth of each mandate, and performed an event history analysis to explore variations in the timing of the broadest mandates that require individuals to wear masks while indoors in public spaces.

Controlling for the seven-day moving average of reported COVID-19 deaths and state citizen ideology, we find the governor's party affiliation is the most important predictor of state differences in the timing of indoor public mask mandates. The marginal effect of a having Republican governor instead of a Democrat was a 29.9 day (95% CI: 24.6 to 35.2) delay in the announcement of broad state-wide mask mandates. This effect is far larger than the effect of any other variables examined and is robust to many different sensitivity analyses testing a large number of possible confounders.

Data

We collected data on all statewide directives mandating masks issued over the period April 1 to August 13. We consider a public mask mandate to be any policy that requires individuals to wear masks or other mouth and nose coverings when they are outside their places of residence. We include only mandates which apply to all individuals within a given setting, allowing exceptions for individuals with certain medical conditions or for young children. Our data thus do not include mandates which only require the use of masks or other personal protective equipment by specific employees as part of business operations.

To further capture variation across mask mandates applying to the general public, we create a typology with three ordered categories that encompass all state-wide public mask mandates issued over this period:

Limited mandate (Level 1). Policies in this category involve limited mask mandates applying only to specific public settings. For example, mask mandates at this level might apply only to transportation services (e.g., issued by Vermont on May 1, augmented to a Level 3 policy on July 24) (8; 9), to retail establishments (e.g., issued by Alaska on April 22 and ended on May 22) (10), or to large gatherings where social distancing is not possible (e.g., issued by New Hampshire on August 11)(11). A common example of a limited mandate is one which applies only to people visiting government buildings (e.g., issued by Utah on June 26 and South Carolina on August 3) (12; 13).

Broad indoor mandate (Level 2). Policies in this category constitute broad mask mandates requiring the use of masks or cloth face coverings by the public across most or all sectors of public activity indoors or in enclosed spaces. Mandates in this category may also include requirements that members of the public wear masks while waiting in line to enter an indoor space, or while using or waiting for shared transportation. For example, Minnesota's mask mandate (issued July 22) requires people over five years of age who are medically able to wear facial coverings or masks "in an indoor business or public indoor space, including when waiting outdoors to enter an indoor business or public indoor space, and when riding on public transportation, in a taxi, in a ride-sharing vehicle, or in a vehicle that is being used for business purposes" (14).

Broad indoor and outdoor mandate (Level 3). Policies in this most comprehensive category mandate the use of face coverings by the public across all public indoor spaces and in outdoor settings, though exceptions may be made for outdoor mask wearing where social distancing is possible. For instance, New York issued a mask policy on April 15 mandating all individuals who are medically able and over two years of age to wear a mask when in a public place and unable to maintain social distancing (15). Washington state's mask mandate, issued June 24, requires that "every person… wear a face covering that covers their nose and mouth when in any indoor or outdoor setting" (16).

Because Level 1 reflects a very limited mask mandate from the perspective of preventing transmission of the novel coronavirus, we concentrate our analysis on adoption of mandates at Level 2 or 3: mandates that at a minimum include a broad requirement to wear masks indoors in public spaces. For these policies, we coded both the dates on which statewide policies were issued in each state at each level, as well as the date of enactment of each policy. Because our objective is to better understand the factors that influenced Governors' decisions to implement mask mandates, we focus on the dates the policies were issued. (If our objective were instead to study the effects of mask mandates, the date of enactment might be more appropriate. Notably, we include a sensitivity analysis using dates of enactment that does not find substantive or statistical differences in the factors predicting mask mandate adoption.)

The top panel of Figure 1 shows when broad statewide mandates requiring masks in indoor public spaces (Level 2 or higher mandates) were adopted across the US, starting in April 2020. The bottom panel shows when the broadest dual indoor-outdoor mandates were adopted (Level 3). These adoptions occurred in two phases: from the middle of April to the end of May, eleven states adopted Level 2 or higher mandates; most (8) were Level 3 mandates. The second phase began in mid- to late-June, and continued into early August. In this later phase, an additional 22 states adopted mask mandates of at least Level 2 or higher, bringing the total number of states with broad mandates to 33. Most of these (17) were also Level 3 mandates (for a total of 25 Level 3 mandates). Four of these 17 (Maryland, Michigan, New Jersey, and Oregon) had already adopted Level 2 mandates in April. Thus, by 12 August 2020, two-thirds of states, containing at least 76% of the U.S. population, had a statewide mask mandate requiring masks in indoor settings. Half of states, containing 63% of the population, further required masks to be worn outdoors statewide.

Results

We use Cox proportional hazards models to explore how different factors influenced the timing of broad statewide mask mandates across the fifty U.S. states. These factors

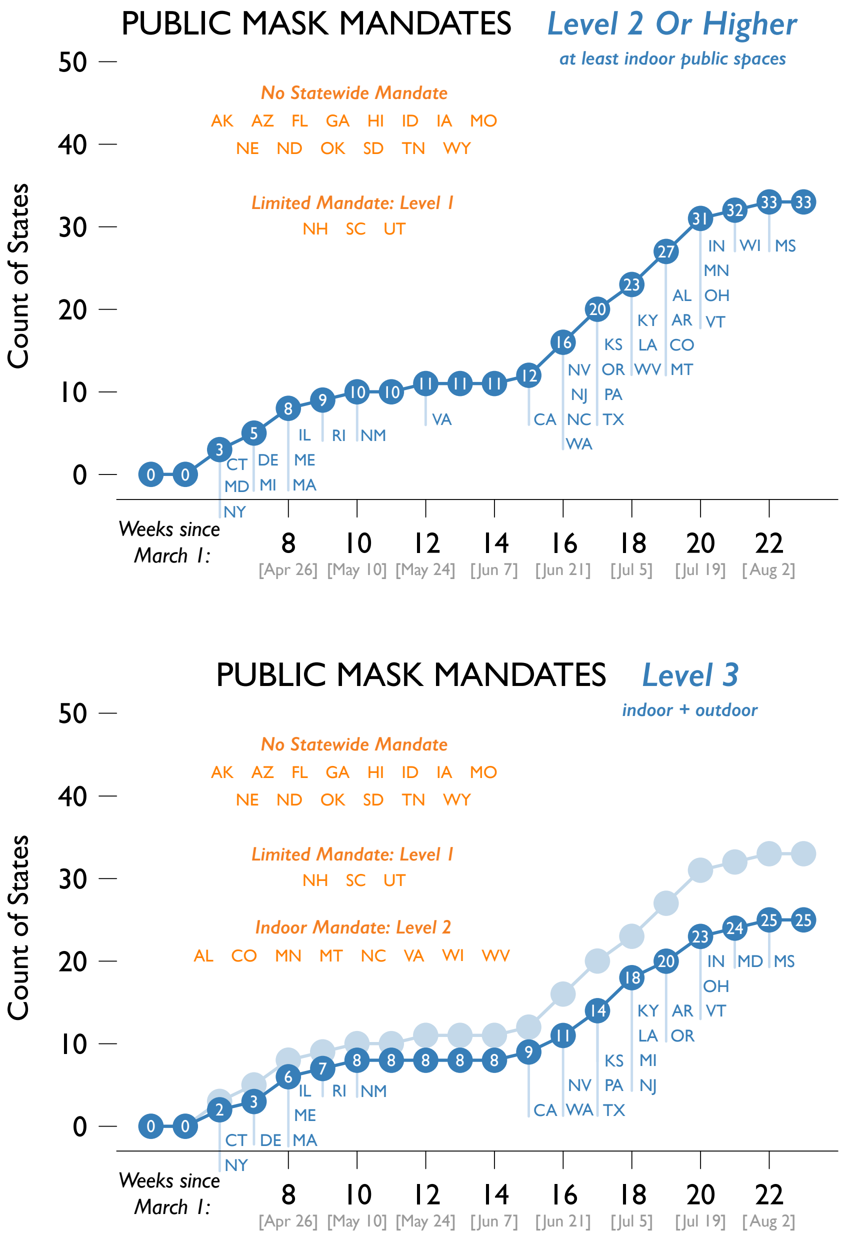

Figure 1. *Adoption of broad statewide mask mandates, 1 April 2020 to 12 August 2020.* All states adopting broad mandates (Level 2 or Level 3) maintained them at least through 12 August 2020. Alaska and Hawaii adopted limited mask mandates (Level 1) but later ended those mandates. Sources: Authors' original data collection (17). Data available at `http://covid19statepolicy.org`.

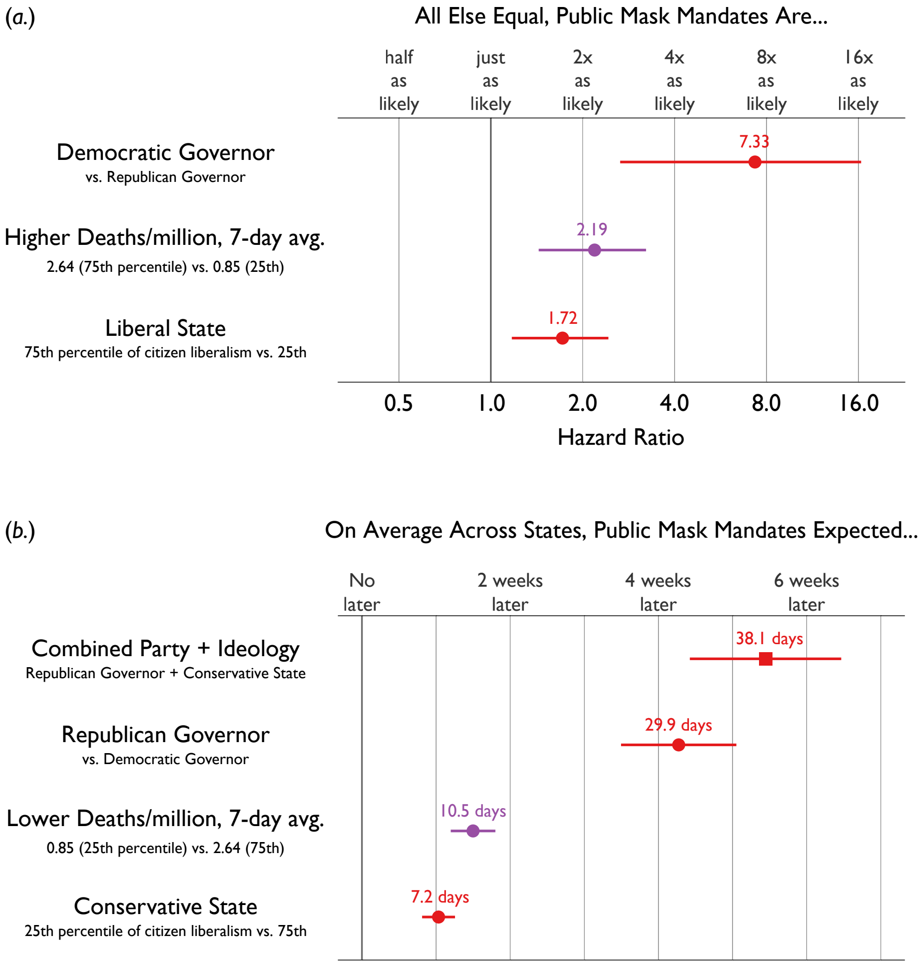

Figure 2. *Relative probability (a.) and expected delay (b.) of adopting at least a Level 2 mandate for public masks, by factor.* The top panel shows on a log scale the estimated hazard ratios obtained from a Cox proportional hazards model on mask mandates adopted by the fifty states, April 1 – August 12, 2020. Red circles mark the hazard ratios for political covariates, and purple circles indicate hazard ratios for other covariates. The bottom panel shows on a linear scale the estimated average marginal effects obtained by post-estimation simulation from the model. The red square marks combined effect of partisanship and ideology, red circles indicate the independent effects of governor party and citizen ideology, and purple circles indicate average marginal effects for other covariates. Horizontal lines are 95% confidence intervals. Solid symbols indicate significance at the 0.05 level.

include COVID-19 indicators, state capacity, and partisan politics. Figure 2 reports the results from our baseline model, which controls for the log of COVID-19 deaths per million population reported in the state as seven-day moving averages, the ideological orientation of each state's citizenry, and the party of the governor (18; 19; 20). These results are reported both using traditional hazard ratios (top panel of Figure 2) and as average marginal effects across all fifty states, expressed as the average expected days of delay associated with each factor (bottom panel of Figure 2).

By far, the most powerful predictor of broad mask mandate adoption and timing is the political party of the Governor. Holding constant state ideology and the rate of deaths per million, at any given time Democratic governors are 7.33 times (95% CI: 2.68 to 16.17 times) more likely to adopt a mask mandate than are Republican governors. We can also use the estimated Cox model to predict the total expected delay hypothetically associated with having a Republican governor (rather than a Democratic governor) in each state, while leaving state ideology and daily deaths per million at their observed values for each state-day. We find that averaged across the fifty states, the marginal effect of having a Republican governor is a 29.9 day delay in adopting a broad indoor mask mandate (95% CI: 24.6 to 35.2 days).

The party of the governor is not the only political variable that influences the likelihood of adoption. Holding constant the party of the governor, states with more liberal citizens adopt mandates earlier than states with more conservative citizens. For example, states at the 75th percentile of citizen ideology (more liberal) are 1.72 times more likely to adopt mask mandates at a given time than more conservative states at the 25th percentile of citizen ideology (95% CI: 1.18 to 2.40 times). The marginal effect of this inter-quartile difference in citizen ideology is a 7.2 day delay of indoor mask mandates in more conservative states (95% CI: 5.8 to 8.7 days).

Researchers and policy-makers use several metrics to track SARS-CoV-2 transmission, and governors have access to daily data on COVID-19 measures including confirmed cases, deaths, and positive test result rates. In our models, daily deaths per million consistently dominates measures of new cases per million and test positivity rates as a factor associated with the timing of broad statewide mask mandates. Nevertheless, the effect of daily deaths is much weaker than the effect of governors' party affiliation. We find that a state at the 75th percentile for daily COVID-19 deaths per million population is 2.19 times more likely to adopt a mask mandate than a state at the 25th percentile (95% CI: 1.45 to 3.19). Our model suggests a state with a lower rate of daily COVID-19 deaths will adopt mask mandates 10.5 days later than a state with a higher daily death rate (95% CI: 8.5 to 12.5 days).

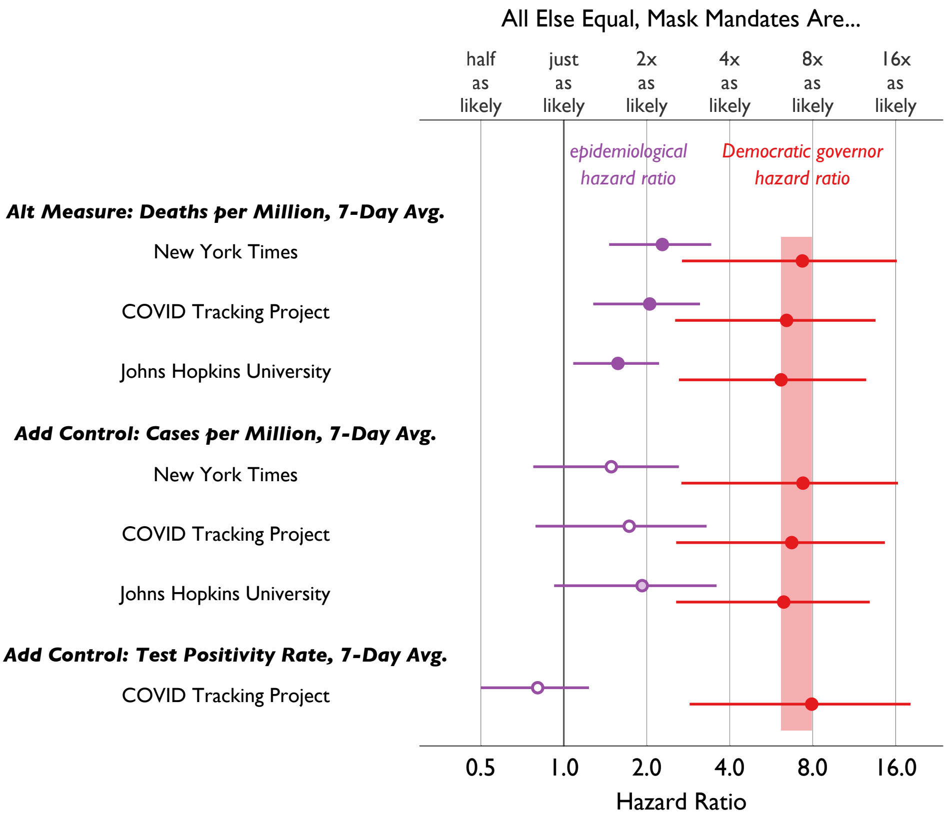

Figure 3. *Sensitivity of results to alternative COVID-19 epidemiological indicators.* Estimated hazard ratios of mask mandate adoption (Level 2 or higher) for various epidemiological indicators (in purple) and for Democratic governors (in red) from a series of Cox proportional hazards models adding the epidemiological covariate listed at the left of the plot. Horizontal lines are 95% confidence intervals. Solid symbols indicate significance at the 0.05 level; shaded symbols indicate significance at the 0.1 level. Axes are log scaled.

As Republican governors and conservative citizens often go together, the relative importance of politics on mask mandate adoption is even greater. When combined, the expected delay in adopting at least an indoor mask mandate for a state with both a Republican governor and a conservative citizenry is 38.1 days (95% CI: 31.1 to 45.1 days) when compared to a Democratic governor in a liberal state. The majority of this delay is attributable to the party of the executive, highlighting the importance of state-level political leadership in fighting the virus.

We conducted several additional analyses to test the robustness of these findings. First, we considered the possibility that our results were sensitive to either the source of daily COVID-19 data used in the model or the set of COVID-19 indicators used for each state-day. Our baseline model used data reported by the New York Times on daily COVID-19 deaths for each state (18). Figure 3 reports results from a series of models that use alternative sources of daily death counts (21; 22). As the top of the figure makes clear, the gap between the effect of governor partisanship and the effect of deaths per million remains at least as large as in the baseline model across the alternative indicators.

While deaths are perhaps the most politically salient consequence of the pandemic, they are the least timely indicator of the severity of transmission in a given place and time, operating at a lag of approximately two or more weeks from the time of infection (23; 24). We therefore consider models adding controls for more timely indicators of the spread of SARS-CoV-2: the number of confirmed COVID-19 cases per million reported in each state each day and the rate of test positivity (in both cases, as seven-day moving averages). States taking prompt action to stem the spread of the virus should arguably be responsive to these indicators. Although the effect of higher rates of case growth is in the expected direction of possibly encouraging mask mandates, the relationship is not statistically significant in a model that controls for the count of deaths. (This pattern holds regardless of the data source used for confirmed cases.) The rate of positive tests in a state had no relationship with mandate timing once deaths per million is controlled. In all models, the partisan effect was unchanged.

In addition to alternative measures of public health indicators, we consider a series of additional control variables, none of which alter our findings regarding the effect of partisan governors (Figure 4). First, we add a third measure of partisan politics, either Trump's vote share in the state in the 2016 presidential election or the percentage of people in the state who watch Fox News regularly (25; 26). Neither helps explain mask mandate timing in models that also control for governor party and citizen ideology. This may indicate that direct effects of these factors cannot be isolated, or that their impact on timing is mediated through governors and through their conservative audiences. Next, we consider the possibility that states adopt mask mandates either in imitation of policies adopted by other states or in reaction to the spread of the virus in neighboring states. We find that, controlling for governor party, citizen ideology, and the daily death rate within a state, neither the adoption of mask mandates by neighboring states nor the average death rate in neighboring states is associated with the timing of mandates. We also control for the rate of mask mandate adoption in "peer states":

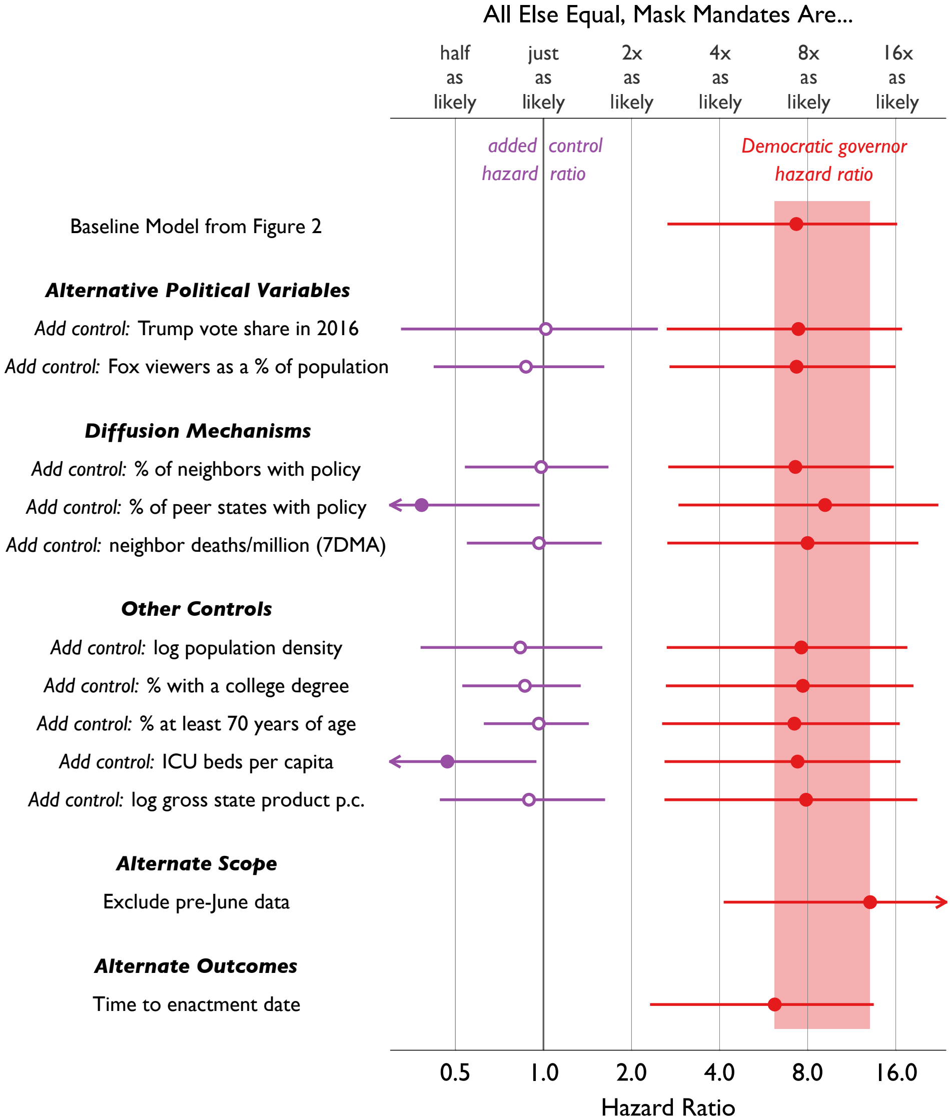

Figure 4. *Democratic governors' greater propensity to enact Mask Mandates is highly robust.* Estimated hazard ratios of mask mandate adoption (Level 2 or higher) for effect of Democratic governors from a series of Cox proportional hazards models including various added controls or alternative outcome measures. Horizontal lines are 95% confidence intervals. Solid symbols indicate significance at the 0.05 level. Arrows indicate confidence intervals that extend outside the plotting range. Axes are log scaled.

other states (often *not* neighbors) identified using network analysis as the innovators which a given state most often imitates across a variety of policy areas (27). Somewhat puzzlingly, whether peer states have adopted mandates is negatively associated with mandate adoption once our baseline controls are included, a result we suspect is spurious (and which does not alter our main findings).

Other controls which fail to explain mandate timing when added to the model include the percentage of state residents above the age of 70 or in possession of a college degree (28), as well as the log of population density (29) and the log of gross state product per capita (a reasonable non-finding given the minimal economic consequences of a mask mandate, in contrast to many other non-pharmaceutical interventions) (30). We consider one last control: the (pre-epidemic) count of ICU beds in each state per capita, which if low might add urgency to state policies to combat the pandemic (31). We find the opposite: states with more ICU beds per capita are more likely to adopt mask mandates. It may be that states that are more generally prepared to address health care needs are also more likely to implement preventive services such as mandates. In any event, inclusion of this control does not alter our main findings.

Finally, we consider changes to the model outcome and scope. The first change is simple: instead of measuring time to the issuance of mask mandates, we model the time to the enactment dates contained in those mandates; our results are unchanged.

The second modification is more noteworthy and divides the data into two periods. In the first period, the months of April and May, states that adopted mask mandates did so either before they eased social distancing mandates, or concurrent with efforts to ease social distancing and re-open business sectors. For most states, the second period, June and July, followed substantial easing of social distancing policies and saw rising numbers of cases starting in mid-June (32). In the first period, states may have issued mask mandates as a preventative policy layer to mitigate transmission risks associated with easing social distancing restrictions (33). By the second period, the benefits of wearing non-medical masks against SARS-CoV-2 transmission were better understood (34; 35). Despite partisan resistance to mask mandates on the part of Republican voters and President Trump, one could imagine governors of both parties coalescing in June and July around mask mandates as the least costly intervention to protect fragile state economies and create a path to normal social interactions (1; 2). Yet when we restrict our analysis to start on June 1 instead of April 1, we find substantively similar relationships – and an even stronger partisan governor effect (hazard ratio of 13.08, 95%CI: 4.18 to 31.34), suggesting mask mandates became *more* partisan in the summer of 2020.

Discussion

Masks are an important tool in the strategies of countries that have slowed the spread of COVID-19 (35; 34). Near-universal mask wearing reduces the risk implicit in returning to aspects of normal life, and may be especially important for protecting essential workers who are not able to limit their exposure (35; 36; 34).

In some countries, mask wearing is a well-established cultural norm (5). This was not the case in the U.S. prior to COVID-19. By early summer 2020, governors of both parties surely recognized the pandemic's threat to their states and were also well aware that it was transmitted via aerosols. One might therefore expect these leaders to be eager to encourage mask wearing as an alternative to the steep social and economic costs implicit in prolonged social distancing measures such as stay-at-home orders. The rapid progression of the pandemic also suggests that mandates, rather than public education campaigns, would be the preferred approach to ensuring mask compliance. Why, then, were so many Republican governors reluctant to promote a relatively low-cost and effective intervention?

Our event history analysis cannot speak to motives. The most likely explanation, we believe, is that the absence of a mask wearing norm in the U.S. opened the door to reactionary responses. From the beginning of the pandemic, President Trump has seemed more concerned about the pandemic's threat to the economy than to public health, and may see mask wearing as a prominent public reminder of a problem that he was consistently trying to minimize. The President remains opposed to a national mandate (37) and has mocked those who do wear masks (38). His behavior promoted partisan division on mask wearing. Republican identifiers are now much less likely than Democratic identifiers to say that they wear masks all or most of the time (53% vs. 76% in August 2020) (39). In addition, many Americans, and especially Republicans, resisted mask wearing as a sign of weakness or "unmanly" behavior (40; 41; 42), perhaps based on the mistaken assumption that self-protection is the primary objective of mask wearing. This political dynamic may help to explain why President Trump steadfastly refused to wear a mask despite widespread urging that he set a public example (43), and why he publicly mocked Democratic Presidential nominee Joe Biden for wearing one.

The U.S. is as polarized politically as it has ever been, including across and within state governments (44; 45; 46). A plausible hypothesis is that many Republican governors delayed imposing mask mandates, not because they believed them to be ineffective or unnecessary, but because they were unpopular with Republican voters, who continue to support Trump by wide margins (47). A Republican governor who mandated masks risked being portrayed as weak, threatening their base of support and possibly the support of their party's national leader. Democratic antipathy toward Trump and his cavalier rejection of the recommendations of his own government, a generally positive view of mask wearing among constituents (48; 49; 39), and widespread mask wearing by Democratic leaders (including Biden) made it much easier politically for Democratic governors to support mask mandates.

Although Republican governors were generally more resistant to mandates, some eventually succumbed to the reality of rising cases and deaths in their states (including Greg Abbott of Texas, Kay Ivey of Alabama, and Tate Reeves of Mississippi). Others, however, continued to reject the recommendations of public health officials in the face of rampant cases and deaths in their states (including Brian Kemp in Georgia, Ron DeSantis in Florida, and Doug Ducey in Arizona). By the end of our study period, 33 states required masks indoors (Level 2 or higher), so that three-quarters of Americans now live under state-wide indoor mask mandates. The remaining resistance is highly partisan: thirteen of the 14 states that have yet to adopt broad mandates are led by Republican governors.

Mask wearing can help to reduce transmission of the coronavirus. Widespread compliance with mask mandates in many localities across the U.S. suggests that a valuable norm may be developing that will have longer term positive consequences for pandemic response (50). However, at the state-level, the politicization of this important intervention has delayed the adoption of broad, consistent mandates on an affordable and effective behavior to reduce coronavirus spread.

Methods Appendix

We estimate an event history model to predict the timing of announced mask mandates across U.S. states from April 1, 2020 to August 12, 2020. Specifically, we model the likelihood that a state will implement a mask mandate of at least Level 2 (broadly requiring face coverings indoors) as a function of time in days with a Cox proportional hazards model, clustering standard errors by state. All states are at risk of adopting a mandate starting on April 1, and remain at risk until they adopt a mandate at either Level 2 or Level 3. In this model, the baseline hazard rate non-parametrically captures the effects of purely national trends – such as the common tendency of states to adopt mask mandates due to the national resurgence of new COVID-19 cases and deaths, or as a result of new scientific findings regarding the effectiveness of masks in reducing coronavirus transmission. This leaves only cross-state variation in the timing of mask mandates to be explained by covariates.

Our primary specification, reported in Table 1, includes two time-invariant covariates – the ideological orientation of each state's citizenry and the party of the governor (19; 20). We also control for a time-varying covariate, the daily reported COVID-19 deaths per million population in each state using data from the New York Times (on COVID-19 deaths) (18) and the US Census (on population) (28). Deaths per million enter the model as a seven-day average (to smooth over differential rates of reporting over weekends and weekdays) and logged (to allow for diminishing marginal effects of rising COVID-19 deaths and to mitigate the influence of outliers, which in some cases likely reflect idiosyncratic reporting delays).

Logging this term improves model fit (concordance increases from 0.821 to 0.833), but poses the problem of how to deal with seven-day averages over periods with no reported deaths. A common but flawed solution is to add a small "fudge" factor (e.g., 0.01, or 1, etc.) to cases of zero deaths to ensure the log of deaths per million is always defined; however, this technique produces different results depending on the (arbitrary) amount added. This is an underappreciated but unsurprising problem, as the range of plausible adjustments covers several orders of magnitude. While differences across plausible "fudge" factors do not affect our substantive or statistical conclusions enough to change our findings, a non-arbitrary solution is preferable. Instead, we rely on the data to suggest the appropriate treatment of zeros by including an additional covariate indicating cases of exact zero values of the moving average of deaths. In turn, before logging the moving average of deaths, we replace zeros with ones, ensuring (without loss of generality) that the zero cases drop out of the log term. The results from this zero-adjusted log specification are similar to those from models that use a "fudge" factor, but arguably less arbitrary and more data-driven.

The hazard ratios associated with each covariate in our primary model are reported in Table 1. The Schoenfeld residuals for each covariate shows no evidence of violation of proportionality, supporting the proportional hazard assumption. For continuous covariates, we show the hazard ratio associated with an interquartile shift in the covariate, as recommended by Harrell (51). These are the hazard ratios reported in the top panel of Figure 2 in the main text. Following the approach of Adolph et al. (7), we contextualize these findings by computing the average marginal effect of each covariate averaged across the fifty states (52), expressed as the expected days of delay associated with each covariate, averaged across the fifty states with all other covariates taking on their observed values day by day for each state. These quantities are shown in the bot-

Table 1. Cox proportional hazards model of state-level mask mandates, Level 2 or higher, 1 April to 12 August 2020.

Covariate	Counterfactuals *pre*	*post*	hazard rate	95% CI *lower*	*upper*
Democratic governor	0	1	7.33	2.68	16.17
Citizen ideology	40.1	54.4	1.72	1.18	2.40
log(Daily deaths/million, 7-day moving average)	0.85	2.64	2.19	1.45	3.19
Daily deaths/million is exactly zero	No	Yes	0.39	0.08	1.76
Total state-policy-days at risk				4818	
Total state-policies at risk				50	
Total events				33	
AIC				189.2	
Concordance index (Harrell's *c*)				0.83	

Each row shows the hazard ratio for (the counterfactual change in) the covariate listed at the left. To simplify comparison across covariates with different scales of measurement, hazard ratios for the interquartile range are shown for continuous covariates. Covariates with both 95 confidence limits above 1.0 significantly increase the chance of adopting a statewide mask mandate. Standard errors used to compute confidence intervals are clustered by state. The concordance index shows the proportion of all pairs of states for which the model correctly predicts which state will adopt a mask mandate first. Schoenfeld residuals show no evidence of violation of proportionality for any covariate. The Efron method is used to resolve ties.

tom of Figure 2 in the main text. Analyses were performed in `R` (version 4.0.2) using the `survival` and `coxed` (52) packages. All visualizations were constructed using the `tile` package (53).

In addition to the primary model reported in the Table 1 and Figure 2, we consider a series of sensitivity analyses. Throughout these analyses, we attempt to keep each estimated model parsimonious, as including too many covariates is a particular concern for event history models with small numbers of observed events (54). The first sensitivity analyses reported in Figure 3 simply replace the New York Times death data used in the primary model with data from alternative sources (the COVID Tracking Project or Johns Hopkins University) (21; 22). However, most of the sensitivity analyses retain the covariates of the primary model and serially add a single additional covariate.

As a final robustness check, we report a complementary analysis focusing on the time to adoption of mandates requiring masks both indoors and outdoors (Level 3 mandates). The results of this analysis are reported in Table 2 and Figure 5. Only half of the states had adopted so broad a mask directive by August 12, and while the hazard ratio associated with a higher moving-average of deaths per million was little changed from the more inclusive model of both Level 2 and Level 3 mandates (2.03 versus 2.19), the haz-

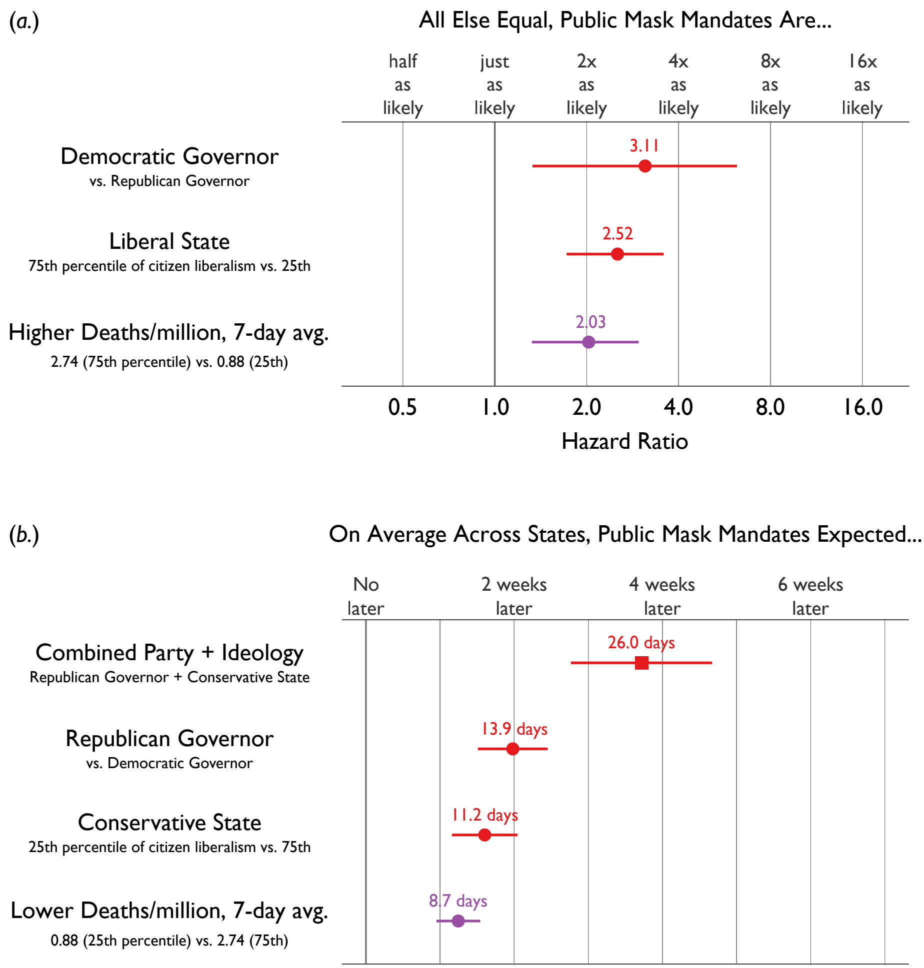

Figure 5. *Relative probability (a.) and expected delay (b.) of adopting a Level 3 mandate for public masks, by factor.* The top panel shows on a log scale the estimated hazard ratios obtained from a Cox proportional hazards model on mask mandates adopted by the fifty states, April 1 – August 12, 2020. Red circles mark the hazard ratios for political covariates, and purple circles indicate hazard ratios for other covariates. The bottom panel shows on a linear scale the estimated average marginal effects obtained by post-estimation simulation from the model. The red square marks combined effect of partisanship and ideology, red circles indicate the independent effects of governor party and citizen ideology, and purple circles indicate average marginal effects for other covariates. Horizontal lines are 95% confidence intervals. Solid symbols indicate significance at the 0.05 level.

Table 2. Cox proportional hazards model of state-level mask mandates, Level 3 only, 1 April to 12 August 2020.

Covariate	Counterfactuals *pre*	*post*	hazard rate	95% CI *lower*	*upper*
Democratic governor	0	1	3.11	1.34	6.15
Citizen ideology	41.7	56.3	2.52	1.74	3.54
log(Daily deaths/million, 7-day moving average)	0.88	2.74	2.03	1.34	2.93
Daily deaths/million is exactly zero	No	Yes	0.37	0.10	1.30
Total state-policy-days at risk				5307	
Total state-policies at risk				50	
Total events				25	
AIC				153.1	
Concordance index (Harrell's *c*)				0.82	

Each row shows the hazard ratio for (the counterfactual change in) the covariate listed at the left. To simplify comparison across covariates with different scales of measurement, hazard ratios for the interquartile range are shown for continuous covariates. Covariates with both 95 confidence limits above 1.0 significantly increase the chance of adopting a statewide mask mandate. Standard errors used to compute confidence intervals are clustered by state. The concordance index shows the proportion of all pairs of states for which the model correctly predicts which state will adopt a mask mandate first. Schoenfeld residuals show no evidence of violation of proportionality for any covariate. The Efron method is used to resolve ties.

ard rate associated with the party of the governor shrank (from 7.33 to 3.11) and the hazard rate associated with conservative citizen ideology grew (from 1.72 to 2.52). In all cases, these results remained significant at the 0.05 level and associated with substantively noteworthy average marginal effects. States with higher rates of daily deaths per million could be expected to adopt combined indoor-outdoor mask mandates 8.7 (95% CI: 6.8 to 10.7) days later than states with low rates of daily deaths. The expected delay associated with Republican governors was 13.9 days (95% CI: 10.7 to 17.0), while more states with more conservative citizens could be expected to adopt Level 3 mandates 11.2 (95% CI: 8.2 to 10.7) days later than states with liberal citizens. The combined delay for states with Republican governors and conservative citizens was 26.0 days (95% CI 19.5 to 32.6).

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