Antitrust Deregulation and the Politics of the American Knowledge Economy

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April 14, 2022

Abstract

Theory suggests antitrust policy is a low salience issue that responds to ideological changes among political elites and not to changes in preferences among the electorate. I argue that the Democratic Party's reluctance to increase antitrust enforcement intensity after 1992 had an electoral connection in that, as the Party turned towards the knowledge economy, it increasingly aligned its policy positions with the views of affluent service professionals who did not favor more antitrust intervention. I find some evidence to support this hypothesis in three exploratory analyses using observational data. First, I show that, in 1981, the federal government began challenging large mergers and acquisitions at much lower rates and that later Democratic presidents accommodated that policy change. Second, using a panel study of top income shares at the state level, I show that antitrust deregulation allowed service professionals in several consistently Democratic states to generate large economic returns. Third, using multilevel regression and post-stratification on survey data, I find that younger and more affluent individuals and those in states that are the most exposed to the service economy support regulation of mergers and acquisitions at significantly lower levels, levels normally observed in states that support Republican presidential candidates by large margins.

Introduction

After a long nascence, antitrust policy has sprung onto the public agenda. A year-long investigation into the nature of competition in digital markets has led a House antitrust subcommittee to issue a 450-page report equating the managers of today's technology firms with the oil and railroad barons of the late 19th century (Majority Staff, 2020). Members of Congress have introduced at least seven major pieces of legislation seeking to reform the nation's system of antitrust enforcement. And

President Biden has appointed vocal critics of the existing system, including Jonathan Kanter, Tim Wu, and Lina Khan, to prominent positions in his administration.

Those who support the reform movement have rooted their demands in a specific narrative about the evolution of antitrust enforcement, the reasons why it has been so lax in recent decades, and what the repercussions are for the American political economy. In this perspective, the current regime arose when top antitrust officials serving under President Reagan embraced many teachings of the law and economics movement associated with the University of Chicago, and these ideological shifts in turn motivated a more relaxed posture towards potentially anti-competitive activity like mergers and acquisitions involving large companies (Khan, 2017; Stoller, 2019; Lynn, 2010). Democratic presidents after Reagan did not revert to a more aggressive antitrust posture simply because they embraced this "neoliberal" ideology (Khan and Vaheesan, 2017; Stoller, 2019). And, as with the first Gilded Age, the corporate behemoths that have emerged from this system of lax enforcement have exercised their market power in ways that disadvantage workers and consumers but increase corporate profits, thereby exacerbating economic inequality (Baker and Salop, 2015; Khan and Vaheesan, 2017).

Though mostly raised by normative scholars, some of these claims are empirical in nature, but testing has been impeded in some cases by a lack of data and in some cases by a lack of interest. First, whether antitrust officials under President Reagan dramatically reset antitrust priorities is an open question. Some contend that antitrust officials routinely respond to pressures and constraints arising from multiple political institutions including presidents, Congress, and the courts (Wood and Anderson, 1993). Others agree that that law and economics movement caused a sea change in enforcement priorities but contend that the shift began before Reagan was elected and arose from the replacement of many lawyers with economists inside the antitrust bureaucracy (Eisner and Meier, 1990). Still others have looked for patterns of change within the time period from 1980 to the

present, which is the period in which good data on mergers and acquisitions exists (Macher and Mayo, 2020; Coate, 2018; Baker and Shapiro, 2008; Leary, 2002; Coate, 2000). But these studies do not allow us to compare the contemporary system of antitrust enforcement to that which prevailed before 1981.

Second, the question of whether lax antitrust enforcement has contributed to rising income inequality is mostly based on case studies (Khan and Vaheesan, 2017) and is conventionally understood as arising from the increase market power that producers have over consumers in accordance with neoclassical economic theory (Baker and Salop, 2015; Khan and Vaheesan, 2017). An alternative perspectives suggests, in contrast, that large mergers and acquisitions mechanically increase inequality by bestowing great financial rewards on the service professionals—mostly lawyers and bankers—who implement them while laying off many blue-collar workers (Short, n.d.), and that the surge in this form of business activity is therefore emblematic of increasing financialization in the American political economy more broadly (Krippner, 2011). In this alternative perspective, lax antitrust enforcement has created both geographic (urban versus rural) and demographic (service professionals versus blue-collar workers) cleavages that have the potential to disrupt the status quo. But if this is true, the relevant political question becomes: why has the Democratic Party has accommodated lax antitrust enforcement if it has been exacerbating economic inequality for more than 40 years and if intensifying antitrust enforcement would garner electoral benefits, like increasing support among blue collar workers? Existing studies which establish an empirical relationship between financialization and economic inequality, like Philippon and Reshef (2012), do not address this question.

Finally, as suggested above, reformers emphasize ideological and behavioral shifts among political elites, especially among judges and antitrust agency staff, when explaining the durability of the current antitrust system. But no work, of which I am aware, has sought to look for evidence of a

broader electoral connection beneath the Democratic Party's shift on antitrust enforcement. Theory and some data suggests such a connection should exist. For starters, the technology companies that are currently under scrutiny and the service professionals on which those companies intensely rely are core target constituencies for the Democratic Party and have been since the Party first turned towards the knowledge economy in the early 1980s (Iversen and Soskice, 2019; Short, 2022). Recent survey work also suggests that tech entrepreneurs are in fact unique in their political preferences, especially when it comes to government regulation of business behavior (Broockman, Ferenstein, and Malhotra, 2018). Similarly, if today's system of lax antitrust enforcement has been propelled by the law and economics movement, we might expect those who have been exposed to this ideology, namely lawyers, to be uniquely supportive of its tenets. And there are reasons to suspect that these differences could feed back into the political system, as legal professionals are known to donate huge sums to political campaigns (more than $320 milion in 2020) by strong Democratic margins (garnering 81 of legal professional donations in 2020) (Open Secrets, n.d.).

In this paper, I implement a series of exploratory analyses using observational data to test the plausibility of these three claims. First, I emulate the technique developed by Phillippon (2015) to develop a consistent time series showing merger enforcement intensity—the number of mergers and acquisitions (M&A) challenged by antitrust officials divided by the total number of large M&A deals consummated—for every year from 1955 to 2015, and then fit a Bayesian changepoint model to the data to determine if there are any critical junctures. Consistent with reformist assertions, the evidence suggests that President Reagan dramatically reset agency priorities and that subsequent Democratic presidents have done little to shift antitrust enforcement back to post-war levels. Next, I develop a model to test whether M&A mechanically influences economic inequality and whether exposure to the service economy predicts heterogeneous effects across states. Estimating this model with panel data, I find that M&A activity does exacerbate economic inequality, that the effect is

confined to only a few states, and that some strongly Democratic states that are well integrated into the knowledge and service economies, including California, New York, Massachusetts, and Washington, are among those that have a uniquely strong economic interest in relaxed antitrust oversight. Given that certain demographic groups and certain regions have an economic interest in lax antitrust enforcement, I then analyze public opinion data to look for some evidence of demographic and regional differences in attitudes towards antitrust enforcement. Analysis of prior polling indicates that younger and more affluent voters are significantly less likely to support doing more to regulate M&A and, at the state level, unionization and exposure to the service economy play a more important role than presidential partisanship in explaining regional variation in attitudes towards antitrust enforcement.

These tests do not reject the reformist assertion that a certain kind of ideology, arising from the law and economics movement, partly sustains today's lax system of antitrust enforcement. They do provide some tentative evidence suggesting that this ideology may be shared more broadly than frequently assumed among the kinds of younger and more affluent professionals that the Democratic Party has consistently courted for over forty years. Whether this remains true today requires new polling on antitrust policy preferences while controlling, at a minimum, for industry and occupation, which is a subject of ongoing work. But existing data also suggests that the Democratic Party's restraint on antitrust issues may arise from a kind of policy feedback inherent to the realignment associated with the knowledge economy transition. Though elected Democrats may have been reluctant participants in initiating the policy changes that have caused increasing financialization in the American political economy, those policies—including lax antitrust enforcement—seem to bestow great economic rewards to the service professionals that reside in several strongly Democratic states, service professionals who are known to support Democratic political campaigns by large margins. Whether those constituencies can be persuaded to set aside

their own individual or regional economic interests to support the reform movement remains to be seen. To the extent they cannot, these demographic and regional cleavages may complicate the path to reform in ways not anticipated by the reformist narrative with its exclusive emphasis on the role of ideology in shaping administrative action among small groups of elite actors.

1 Testing Political Theories of Antitrust Policy Change

A core plank in the antitrust reform platform is that President Reagan dramatically reduced government oversight of proposed mergers and acquisitions involving large companies and that subsequent Democratic presidents failed to unwind this drastic change in administrative priorities. At least two prior studies by political scientists seem to reject this hypothesis, but each relies on interrupted time series regressions with thirty or fewer observations and seven or more predictors (Eisner and Meier, 1990; Wood and Anderson, 1993), which makes it difficult to assess the true significance of these authors' findings. Both studies also sought to explain changes either in pure counts of enforcement activity and resources (the number of cases filed, the dollar amount of budget allocations, etc.) even though this empirical strategy has serious drawbacks. More recent studies in the legal and economic literature have therefore tended to measure antitrust enforcement intensity as a ratio of administrative action (merger challenges) relative to some measure of economic activity (the number of proposed mergers involving large companies) (Macher and Mayo, 2020; Coate, 2018; Baker and Shapiro, 2008; Leary, 2002; Coate, 2000). But these studies focus almost exclusively on the post-Reagan period and therefore say little about whether the changes observed during the Reagan administration deviated in any meaningful sense from those observed in earlier decades.

To overcome these data limitations, I first developed a unique dataset showing merger enforcement intensity—or the number of mergers challenged by the Department of Justice Antitrust Division

(DOJAD) as a share of the number of proposed mergers involving large companies—for every year from 1955 to 2015. Specifically, I followed Phillippon (2015) and Baker and Wurgler (2000, pp. 2250) in using historical time series known to be correlated with M&A activity to impute the number of large mergers consummated in each year from 1955 to 1981 and then merged these measures with comprehensive SDC Platinum data on M&A from 1982 to 2015. I then followed a similar procedure to make the historical series of DOJAD merger cases from 1955 to 1997 reported in Gallo et al. (2000) consistent with the official DOJ workload statistics reported from 1970 to 2015. I then took the ratio of DOJAD challenges relative to the number of large mergers, and normalized the resulting time series to the high point observed in 1961.

Figure 1 illustrates why it is essential to interpret regulatory activity relative to the amount of economic activity that is meant to be regulated, a quantity which is arguably endogenous to the chosen stringency of regulatory scrutiny. Panel A shows the number of cases the DOJAD filed from 1955 to 2015 challenging a merger. As shown, the numbers vary but there is no easily discernible trend and one could easily conclude that, during the Clinton presidency, antitrust enforcement intensity significantly increased. Panel B shows the number of large mergers executed in each year over the same time frame. It shows that the amount of M&A activity exploded around 1980, most likely in anticipation of relaxed regulatory scrutiny, and continued to grow dramatically during the Clinton presidency. Panel C shows the relative enforcement rates obtained by taking the number of cases from Panel A, dividing by the number of large deals from Panel B, and normalizing the numbers so that they portray the likelihood of a challenge relative to the high enforcement rates of 1961. It provides strong evidence of a critical juncture in 1981. Assume, for example, that the average relative enforcement rates are roughly 50 percent and 1 percent before and after 1981 based on the plot. Before Reagan was elected, if the number of deals executed increased by 10 percent relative to the number of deals executed in 1961, firms could expect the number of DOJAD

challenges to be about 55 percent of the number of cases filed in 1961. After Reagan was elected, firms could expect DOJAD challenges to be about 1.1 percent of the number of cases filed in 1961 for the same relative increase in M&A activity.

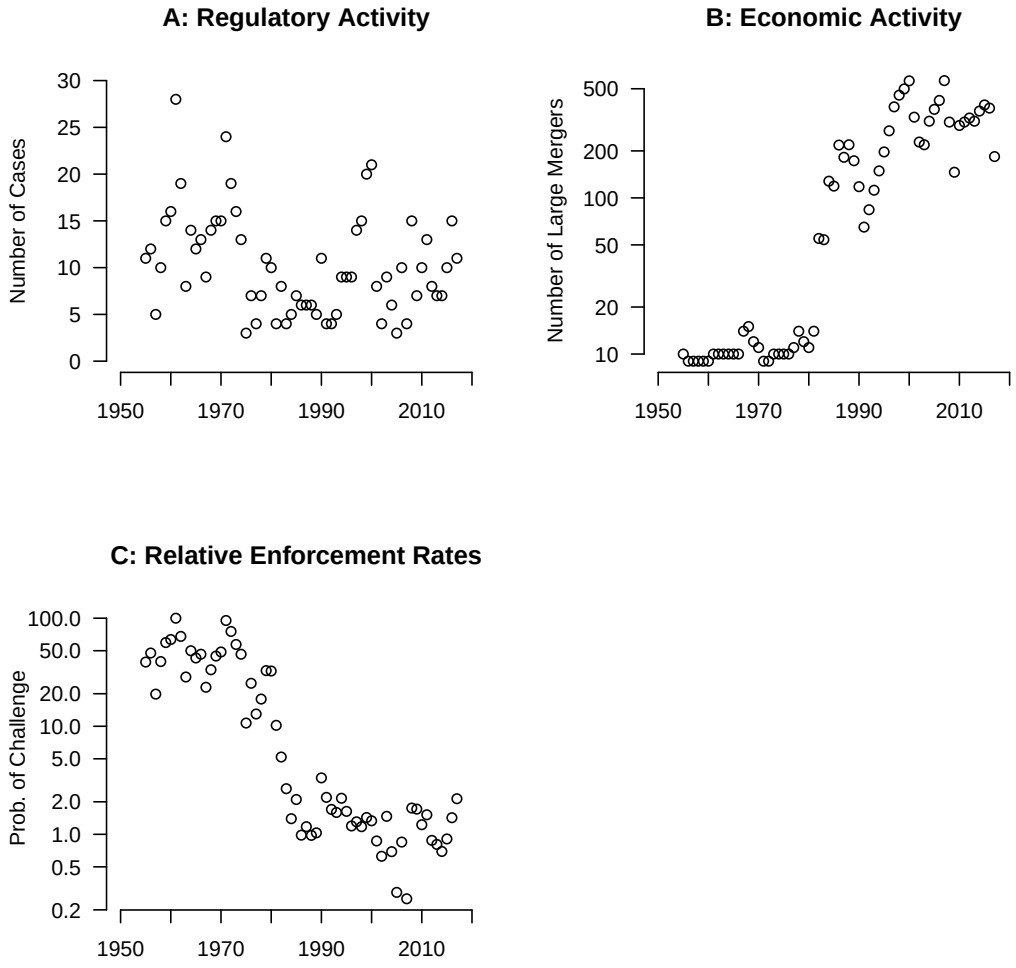

Figure 1: Panel A shows the number of DOJAD cases challenging a merger from 1955 to 2015. Panel B shows the number of large mergers in each year over the same time frame, based on imputed estimates from 1955 to 1981 as described in the text. Panel C shows the likelihood of facing a regulatory challenge, which is obtained by dividing the number of cases from Panel A by the number of mergers in Panel B, and normalizing relative to the maximum in 1961.

The analysis above supports the conventional view that antitrust policy changed dramatically in a short period of time around the early 1980s. To determine precisely when the change took place I

estimated a Bayesian changepoint model (Carlin, Gelfand, and Smith, 1992) of the following form:

$$y_t = \phi y_{t-1} + a_1 x_t + b_2 z_t \text{ for t} \in (2 : k) \tag{1}$$

$$y_t = \phi y_{t-1} + a_2 x_t + b_2 z_t \text{ for t} \in (k+1 : T) \tag{2}$$

where k is a potential change point, y_t is relative enforcement rates, x_t is the DOJAD's annual budget appropriation, and z_t is presidential party affiliation. The model is first order autoregressive under the assumption that agency officials and bureaucrats will base the current year's priorities on the last year's priorities with some modifications. In terms of specifying distributional assumptions, the change point k was drawn from a uniform distribution on the interval $(2, T)$ where T in this case is 61 (for the 61 years between 1955 and 2015), and the autoregressive coefficient, ϕ, was drawn from a uniform distribution on the interval (0,1). The remaining coefficients on budget restraints (a_1 and a_2) and presidential party (b_1 and b_2) were given uninformative normal priors.

The results obtained by estimating equations 1-2 using the JAGS Gibbs sampler in R are shown in column (1) of Table 1. Because budget appropriations increase (almost) monotonically over the entire period, the negative coefficients a_1 and a_2 suggest that relative enforcement rates trended down before and after the change point, despite increasing budgets, though the trend is more negative after the change point. More importantly, the 95 percent confidence intervals for the coefficients measuring the effect of presidential party affiliation, b_1 and b_2, effectively surround zero. As a result, we cannot reject the null hypothesis that the partisanship of the sitting president has no bearing whatsoever on the likelihood of challenging large mergers in either period after we control for the substantial change in enforcement priorities at the change point. Because those coefficients are effectively zero, I dropped them from the model and ran a second simulation to (slightly) improve precision on the estimate of the change point, k. The results are shown in column

(2) of Table 1. The 95 percent confidence interval for k is (26.0, 27.9) with a mean of 26.7. This suggests the change point is precisely estimated and took place in the 27th year, or 1981.

	(1)	(2)
a_1	-0.06	-0.06
	(-0.1, -0.02)	(-0.09, -0.03)
a_2	-0.21	-0.2
	(-0.29, -0.13)	(-0.27, -0.13)
b_1	-0.07	–
	(-0.48, 0.35)	–
b_2	0.15	–
	(-0.21, 0.51)	–
k	26.73	26.7
	(26, 27.98)	(26, 27.88)
ϕ	0.4	0.41
	(0.17, 0.62)	(0.19, 0.63)
N	61	61

Table 1: This table shows the mean estimated coefficients and 95 percent confidence intervals (in parentheses) obtained from estimating Equations 1-2 using the runjags library in R (Denwood, 2016). The dependent variable in each model is DOJAD enforcement rates relative to 1961, as explained in the text. The parameter ϕ and the pairs (a_1, a_2), and (b_1, b_2) respectively capture dependence on the prior year's enforcement rates as well as budget appropriations and presidential party affiliation before and after the change point, k.

In sum, when we use an appropriate metric of DOJAD enforcement priorities that is consistent across several deacdes, we do see evidence of a critical juncture in antitrust enforcement priorities starting with Reagan's presidency in 1981. But we also observe that, even after Democrats regained control of the presidency in 1993 and again in 2009, they did little to return the DOJAD to its pre-1981 policy of challenging almost half of all large mergers that were proposed in each year, even though this kind of policy shift can be (and was) implemented without Congressional approval. The DOJAD did revise the merger guidelines in 1992, in 1997, and again in 2010. But the data shows that these revisions tinkered at the margins, at best. These findings support reformist contentions but also raise new questions about why Democratic presidents after Reagan chose not to reinvigorate antitrust enforcement. I explore this question in the remaining sections.

2 Explaining Lax Enforcement Under Democrats After Reagan

For some policy observers, the empirical results described above may come as little surprise. In a new book, Stoller (2019) argues, for example, that key political appointees to the DOJAD in Democratic administrations of the 1960s and 1970s began to embrace Chicago School "neoliberal" ideology on antitrust issues so that, by the time of Reagan's victory in 1980, liberal support for more aggressive enforcement within the bureaucracy had all but vanished. In this respect, Stoller (2019) adds descriptive heft to a similar theory articulated earlier by Eisner and Meier (1990) and Eisner (1991) and echoed by Khan (2018).

One potential drawback of these theories is that, in placing so much emphasis on ideological shifts among political elites, the authors fail to explore the possibility that some core Democratic constituencies may not want the government to regulate M&A more aggressively. In other words, existing scholarship has done little to entertain the hypothesis that there is an electoral connection supporting the Democratic Party's shift in attitudes about antitrust enforcement. One possibility is that Democratic Party leaders embraced certain "neoliberal" policies, including a more relaxed antitrust posture, to better align themselves with a targeted constituency like the younger more affluent urban professionals who were contributing to knowledge economy development (Geismer, 2015; Iversen and Soskice, 2019). Alternatively, that constituency might be much narrower and be dominated by the leaders of high tech businesses who were known to support lax antitrust enforcement (Short, 2022). Another possibility is that, in a kind of feedback mechanism, the move towards lax antitrust enforcement shifted the economic and political terrain and thereby made it more difficult for Party leaders to revert to more robust enforcement. In this scenario, even if the initial shift to lax enforcement occurred under a Republican president, if liberal service professionals who work in knowledge economy hubs benefited from that policy, their own attitudes

towards antitrust enforcement may have changed over time, and Party leaders may have shifted accordingly (Rodden, 2019). Through either mechanism—by targeting a new constituency with idiosyncratic preferences from the outset or by responding after the fact to induced changes in preferences—Democratic officials may have partly located their accommodation with lax antitrust in the demands of certain constituencies or regions.

Rigorous testing of these hypotheses is not straightforward. Archival records for business lobbyists might reveal more about the posture of business leaders towards antitrust enforcement in the 1980s and 1990s, work which is ongoing. The archives for prominent Democratic Party leaders, like Bill Clinton, might also contain evidence of intent to cater policy to new constituencies. But it is much more likely that, at least in documentary evidence, these political considerations would remain subtext while the official narrative would focus on legitimating the policy shift through ideological appeals to prove that the change is sound or justified by economic theory (Trumbull, 2012). In fact, today's antitrust reformers may over emphasize ideology simply because it is much more likely to be observed in the historical record. Similarly, rigorous testing of whether knowledge economy and service professionals are materially different in their attitudes towards antitrust enforcement than ordinary Democratic voters would require new surveys with previously unasked questions about demographics, like occupation, targeted to heavily sample within groups, like venture capitalists or inventors, who are seldom the focus of public opinion surveys (but see Broockman, Ferenstein, and Malhotra, 2018). This work is also ongoing, but the effect of the reform movement itself, which has quite publicly signalled major challenges to the existing policy regime among progressives, will muddle the interpretation of these new surveys. To answer the historical question of whether Democratic Party leaders perceived an electoral connection in 1993 and 2009 we must therefore appeal to observational data, however imperfect.

In the sections that follow, I engaged in two such analyses. First, as described in depth in Appendix

B, I develop a model for testing whether cyclical shocks in M&A activity exacerbate state-level income inequality and whether state-level exposure to the service economy, as proxied by the state's share of GDP in legal services, significantly predicts heterogeneity in that relationship. When estimated separately for the pre- and post-1981 period, this model reveals whether service professionals have a concrete economic interest in lax antitrust enforcement and whether the regional expression of those interests has changed over time. In section 2.1, I estimate the model in the pre- and post-1981 periods and describe the results.

Second, after finding that merger waves do exacerbate income inequality, I then test the hypothesis that the states where these economic effects are most acutely felt have historically opposed more aggressive policing of M&A at lower levels than would be expected based on aggregate measures of liberalism like two-party Democratic presidential voteshare. I test this contention by fitting a multilevel model, with both individual and regional variables, to historical polling data about antitrust enforcement and by then imputing state-level margins of support for doing more to regulate M&A using Census data, following the technique developed by Park, Gelman, and Bafumi (2004). This analysis explores whether there once existed a plausible cleavage, demographic or geographic, among Democrats that might explains the Party's shift towards accommodating lax enforcement. In section 2.2, I introduce this model and describe the results.

2.1 Antitrust and the Professionals Who Service the Knowledge Economy

The American public has, for some time, believed that M&A lead to concrete economic benefits for white collar professionals, like lawyers and bankers, who implement these transactions while imposing great economic costs on the blue-collar employees of the companies involved. In a 1990 survey, for example, 80 percent of respondents said that corporate mergers and takeovers help the lawyers and bankers who arrange them, while 62 percent said that the deals hurt the employees of

the companies involved (Shulman, 1990). But to my knowledge, no prior work has sought to test whether there is a plausible causal relationship between M&A activity and income inequality or quantify the size of the effect or look for changes in the size of that effect over time. In fact, even those who contend that such a relationship exists tend to understand that relationship through the lens of neoclassical economic theory, which teaches that inequality generally arises somewhat diffusely, over long periods of time, from the power producers have to charge higher prices to consumers (Khan and Vaheesan, 2017). This framing overlooks an equally plausible mechanism, consistent with public opinion, in which these transactions have immediate differential impacts on white and blue collar workers (Short, n.d.).

To test for a plausibly causal effect between relative growth in M&A activity and income inequality according to this mechanism, I leverage the fact that in 1981, each of the fifty states plus the District of Columbia was differentially suited to take advantage of the explosion in financial activity that came with antitrust deregulation. My identification strategy is based on the econometric model for heterogeneous treatments developed by Card (1992) and others (Rajan and Zingales, 1998; Baker, Bloom, and Davis, 2016; Angrist and Pischke, 2008) and is motivated and described in greater detail in Appendix B. The main specification is:

$$Is_{it} = \alpha_i + \beta_t + \delta\xi_{i,t-1} + \gamma(\xi_{i,t-1} \times X_t) + \epsilon_t \tag{3}$$

where Is_{it} is a top income share in state i and year t, α_i and β_t are state and year fixed effects, X_t is the annual amount of large M&A activity relative to total national income, and $\xi_{i,t-1}$ represents exposure to the service economy. The coefficient on the interaction term, γ, captures the main effect of interest.

To estimate the model, I first followed a procedure similar to that used in section 1 to estimate the

total dollar value of global M&A completed in each year where the transaction value was more than $100 million in 1948 dollars, which is approaching $1 billion in today's dollars. I then use the time series showing M&A activity relative to total income, $X_t = log(\frac{M_t}{I_t})$, as capturing a common national shock to which each state is differentially exposed. I estimate each state's exposure to these shocks and to the service economy more generally, $\xi_{i,t-1}$, using each state's share of total GDP earned in the prior year in legal services, much in the way that Card (1992) estimated exposure to a new federal minimum wage with the share of teenage workers likely to be affected. For example, if in 1986, the total amount of income generated in legal services was $10 billion and New York generated $1 billion of the total, then its exposure to the service economy in 1987 would be 10 percent or 0.1.

Table 2 presents the results. The main finding is that growth in M&A activity relative to total income accurately predicts some fluctuations in fiscal income inequality across both periods, but the effect size for the top 0.1 percent of the income distribution has more than doubled since 1981, even after we control for the possible effect due to capital gains in the second period.[1] This strongly suggests that the income of affluent professionals at the top of the income distribution is sensitive to changes in M&A activity.

One concern with the approach above is that it may not capture all of the important differences between the states in their exposure to the service economy. As an exploratory exercise, I also estimated a Bayesian equivalent of 3 (see Appendix B for more information) in which I allowed the coefficients, δ_i and γ_i, to vary by state i. This muddies the water for statistical inference, as γ_i now captures some combination of an effect and un-modeled heterogeneity in service economy exposure. But as an exploratory exercise, the results are nevertheless interesting and help identify

[1] This control is included to account for the possibility that income accruing to shareholders, rather than service professionals, are generating the results.

	top 0.1 percent share of fiscal income		
	1963-1980	1981-2015	1989-2015
	(1)	(2)	(3)
state share of legal services (t-1) (δ)	-9.504^{***}	-17.302	-4.030
	(2.571)	(26.463)	(21.589)
state share of legal services (t-1) $\times X_t$ (γ)	3.060^{***}	11.063^{***}	7.251^{***}
	(0.471)	(2.327)	(1.856)
capital gains			32.509^{***}
			(7.693)
state share of capital gains (t-1)			18.100^{***}
			(4.441)
state share of capital gains (t-1) $\times X_t$			38.938^{***}
			(9.660)
Constant			
N	816	1,785	1,272
R^2	0.531	0.630	0.610
Adjusted R^2	0.521	0.622	0.601
Residual Std. Error	9.771 (df = 798)	33.113 (df = 1748)	35.588 (df = 1242)
F Statistic	53.209^{***} (df = 17; 798)	82.620^{***} (df = 36; 1748)	66.918^{***} (df = 29; 1242)

$^{*}p < .1$; $^{**}p < .05$; $^{***}p < .01$

Table 2: This table shows the main results of estimating equation 3. All three models include state and year fixed effects (estimates not shown). The dependent variable in each model is the top 0.1 percent's share of fiscal income. The state's share of legal services in period $t-1$ is used as a measure of exposure to the service economy (δ in Equation 3). The main effect of interest (γ in Equation 3) is the interaction between this measure of exposure and X_t, which measures M&A activity relative to total fiscal income in year t, on the log scale (see text). Models (1) and (2) cover different time periods, 1963-1980 and 1981-2015 respectively. In model (3), I added controls for inflation adjusted capital gains at the state level plus each state's exposure to capital markets (its share of capital gains in the prior year) and that exposure interacted with X_t. It covers the time period for which this data was available (1989-2015).

states that may be acutely dependent on generating economic benefits from antitrust deregulation.

Table 3 presents the results. The first key finding, shown in column (1), is that growth in M&A activity did not significantly increase fiscal income inequality from 1963-1980 anywhere except New York. This is consistent with New York being the center of legal and financial services for M&A activity before deregulation. The second key finding is that, in the period of antitrust deregulation from 1981-2015, growth in M&A activity significantly increased fiscal income inequality in only eight states plus the District of Columbia. Most of these states tend to favor Democratic presidential candidates by relatively strong margins, and at least four of them—California, Massachusetts, New York, and Washington—are considered knowledge economy leaders. In 1980, few might have predicted that antitrust deregulation would financially benefit anyone outside of a select group of professionals in New York. But the explosion in M&A activity that antitrust deregulation wrought has fundamentally altered the economic landscape so that today, groups of affluent professionals who reside in at least six consistently Democratic states plus the District of Columbia have incomes that are intensely tied to the lax antitrust regime that Reagan inaugurated.

The key takeaway, here, is that it is possible to identify specific constituencies and regions that have benefited from antitrust deregulation, and that the constituencies and regions that have benefited from this policy shift tend to favor Democratic presidential candidates by large margins (Open Secrets, n.d.). In this sense, the popular perception that M&A deregulation worked for the benefit of lawyers and bankers but to the detriment of employees has at least some basis in fact. And while it has always been true that M&A activity exacerbates income inequality, the magnitude of that effect grew substantially after deregulation and its geographic location spread well beyond New York City. In the next section, I explore the possible impact of these economic shifts on public support for having the federal government do more to regulate M&A.

state	1963-1980	1981-2015
California	3.15	15.3
	(-0.5, 6.3)	(8.28, 22.63)
Connecticut	3.37	52.66
	(-1.42, 8.68)	(9.14, 100.13)
District of Columbia	2.24	49.44
	(-3.67, 6.74)	(28.33, 70.46)
Florida	3.31	25.12
	(-1.4, 8.3)	(10.2, 39.92)
Massachusetts	3.28	56.55
	(-1.27, 8.27)	(29.5, 84.49)
Nevada	3.14	55.72
	(-1.94, 8.5)	(-0.09, 117.8)
New Jersey	3.37	27.12
	(-1.01, 8.18)	(3.39, 52.07)
New York	2.86	9.63
	(0.22, 5.32)	(4.02, 15.26)
Washington	2.99	47.75
	(-2.48, 7.88)	(14.21, 83.45)

Table 3: This table shows the main results of estimating a Bayesian form of equation 3 that allows for heterogeneity in δ_i and γ_i using the runjags library in R (Denwood, 2016). The results for states that do not have a significant effect in any period are not shown.

2.2 The Political Geography of Public Support for Antitrust Re-regulation

While antitrust issues do not have a prominent place in public opinion surveys, some questions have been asked consistently enough over time to enable a basic understanding of how views on antitrust regulation have evolved in the post-Reagan period. In this section, I focus on 16 polls that asked respondents whether the federal government should make it easier or harder for companies to merge. I ignore differences in question wording and consider responses to fall into one of three categories: support for doing more, taking no position or opposed to doing more, and either refusing to answer the question or responding "I don't know."

The aggregate annual trends (where outcomes of multiple polls in the same year have been averaged) are shown in Figure 2. The results suggest that Reagan took advantage of a unique policy window in 1981 when the public was largely opposed to more federal regulation. Public attitudes quickly swung in the other direction as antitrust policy changed, with those favoring more regulation exceeding 60 percent through most of the late 1980s. Since the late 1980s, the gap between those who favor and oppose more regulation has drastically narrowed (from about 40 to about 10 percent), though overall support has been somewhat consistently between 45 and 55 percent.

For five of these sixteen polls, raw data is available for further analysis. To determine what economic and demographic cleavages have influenced public opinion towards antitrust regulation since 1981, I fit the following multilevel model using these raw data sets:

$$Pr(y_{jt} = 1) = logit^{-1}(stateyear_{i[j],t} + inc_j + educ_j + age_j + d * hisp_j + race_j + e * sex_j) \quad (4)$$

$$stateyear_{i[j],t} = state_{i[j]} + year_t + a * partisan_{i[j],t} + b * union_{i[j],t} + c * exposure_{i[j],t-1} \quad (5)$$

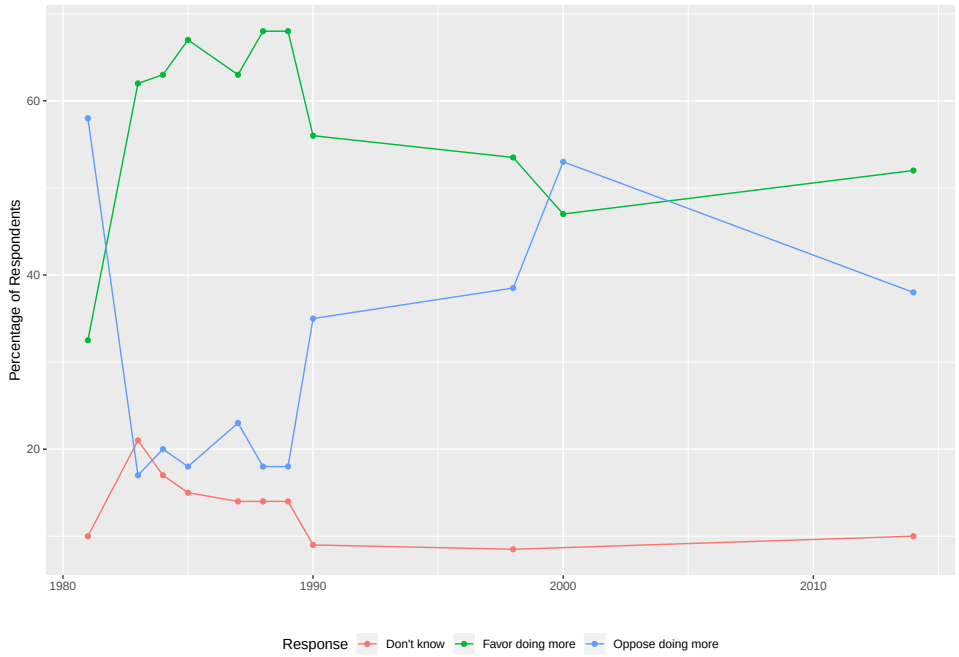

Figure 2: This figure shows the percentage of respondents who favor (green) or oppose (blue) doing more to regulate M&A, or who either refused to answer the question or did not know how to respond (red). The data includes 16 separate polls but response percentages in the same year are averaged to produce 11 annual observations since 1981. *Source*: iPOLL Databank provided by the Roper Center for Public Opinion Research.

In the regional model of Equation 5, $partisan_{i[j],t}$ in state i and year t is the average Democratic presidential vote share in the two elections prior to year t, $union_{i[j],t}$ captures the share of each state's private sector labor force that is a union member, and $exposure_{i[j],t}$ captures each state's exposure to the service economy as above (the share of total income in legal services earned by each state in the prior year). Along with time invariant aspects of each state ($state_i$) and average effects across all states in each year ($year_t$), the estimated coefficients on these three variables (a, b, and c) influence average levels of support for more regulation in the regional model. The term $i[j]$ denotes the state i that individual j belongs to. The individual model of Equation 4 includes six basic demographic variables that were reported in each poll. Because gender and hispanic heritage are binary variables, the effects of those characteristics are modeled as simple regression coefficients (d and e) (Park, Gelman, and Bafumi, 2004, p. 377). The binary dependent variable y_{jt} represents

whether individual j in polling year t favors more regulation of M&A. All of the regression and multilevel coefficients were given uninformative normal priors.

The results from estimating the model are shown in Figure 3. The results from the individual level model suggest that the main demographic cleavages are not race or gender. Though women are slightly more likely to support more regulation than men, and whites are slightly more opposed to more regulation than other racial groups, the differences are somewhat small compared to the other effects seen in the model, most notably age and income. Adults aged 18-34 are significantly more opposed to regulation than adults who are 65 or older for example, which suggests that one of the main cleavages in public opinion about antitrust policy is generational. Similarly, those who have a household income less than $50,000 are significantly more likely to support more regulation than those who make $100,000 per year or more.

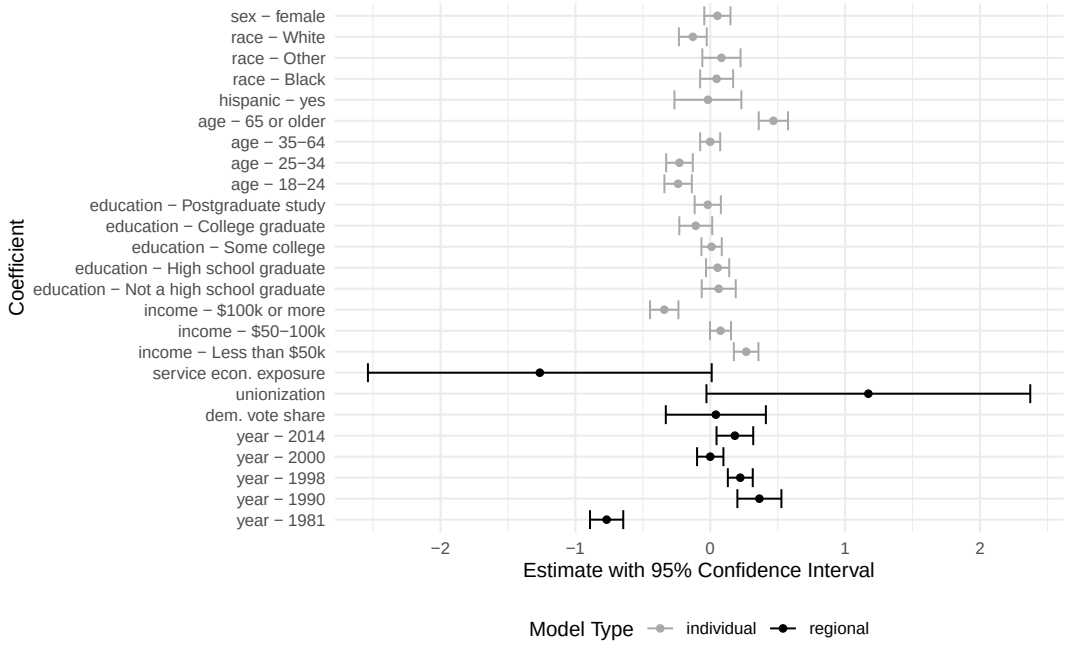

Figure 3: This figure shows the average estimated coefficient (dots) and 95 confidence intervals (bars) for both the individual characteristics (grey) and regional characteristics (black) included in the multilevel model (see Equations 4 and 5).

The results from the regional model are also revealing, in that exposure to the service economy and

union membership have roughly equal and substantively large opposite effects on average levels of support (though the coefficients are only marginally precise), while state partisanship plays an insignificant role in predicting aggregate antitrust attitudes. The difference in these estimated coefficients is meaningful because partisanship, unionization, and service economy exposure are weakly correlated (they have Pearson correlation coefficients ranging from 0.12-0.25). As a result, it is unlikely the effect of partisanship is being absorbed, for example, in the estimates for the other two variables. Both sets of findings (from the individual and regional models) are consistent with the view that political realignment has caused Democratic officials to base antitrust policy priorities on the preferences of younger and more affluent voters who are integrated into the knowledge economy, even if those priorities are somewhat at odds with the views of older and less affluent voters in states with higher levels of union membership.

To see how these effect sizes aggregate into state-level differences in support for having the federal government do more to regulate mergers, I implemented post-stratification using IPUMS-USA census data (Ruggles et al., 2019) from the census year nearest to the polling year. In short, the multilevel model described above allows me to compute the predicted probability of support within 720 separate demographic categories (3 income levels, 5 education levels, 4 age brackets, 3 race categories, and 2 categories each for gender and hispanic heritage) for each of the fifty states plus the District of Columbia (51 state categories) for a total of 36,720 categories in each polling year. Those predicted probabilities are then effectively weighted by the share of the state's population in each of the census categories (Park, Gelman, and Bafumi, 2004, p. 376) to impute average levels of support for each state and year.

If political realignment has influenced Democratic preferences towards antitrust regulation, we would expect the states with concrete economic interests in lax regulation, by virtue of their advanced position in the service and knowledge economies, to have materially different preferences

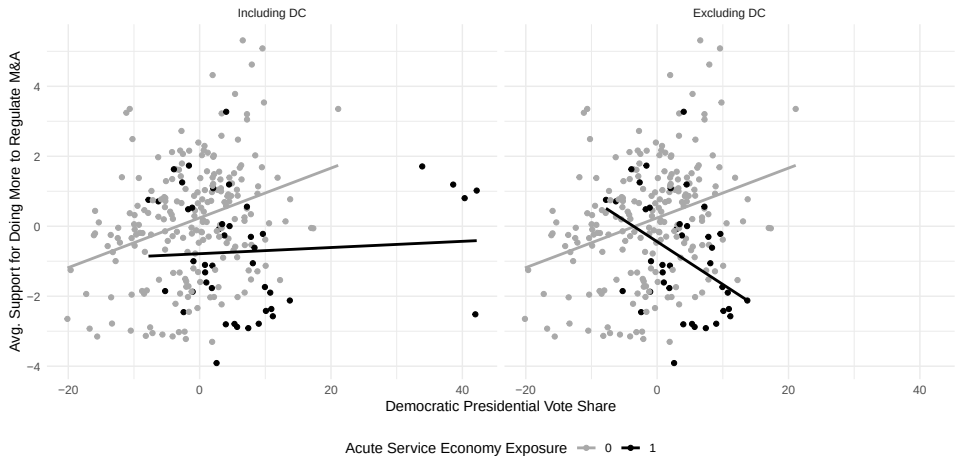

Figure 4: These figures show the predicted levels of support for having the government do more to regulate M&A at the state level as a function of average presidential vote share in the two elections prior to the polling year. In each figure, the data is pooled for all five polling years (1981, 1990, 1998, 2000, and 2014) and both variables have been de-meaned within each polling year. The figure on the left includes the District of Columbia while the figure on the right excludes it as an outlier. Each figure shows the relationship for the states with acute exposure to the service economy (black, see text) and all other states (grey). Regression lines are included as a visual aid but are not reported.

for more regulation. Figures 4 suggests that this is in fact the case. As illustrated, presidential partisanship explains some variation in attitudes towards antitrust enforcement: moving from one extreme to the other opens up a difference of about 3 percentage points in average levels of support for doing more to regulate M&A across states. But when the District of Columbia is excluded as an outlier (right), a comparable support gap exists between the service economy states and other states who support Democratic presidential candidates at 10 percentage points above the national average. Party identification is known to be a strong and significant predictor of differences in attitudes towards regulation of big business generally and of antitrust enforcement more specifically (individual year regression results not shown), so it is no surprise that regional differences can be partly explained by partisanship. What is surprising is the drastic cleavage observed between states that tend to favor Democratic presidents, with the service economy states behaving "as if" they were extremely conservative on this issue. As illustrated, service economy states that favor Democratic presidential candidates by about 5 percentage points had about the

same average levels of support for regulating M&A more aggressively as non-service economy states that favor Republican presidential candidates by 20 percentage points or more.

In sum, affluent service professionals in a few states have historically had a concrete economic interest in lax antitrust enforcement and the younger and more affluent workers who live in these regions have historically expressed much higher levels of opposition to increasing antitrust enforcement. In fact, when regional and demographic differences in attitudes towards antitrust enforcement are aggregated into state-level estimates of support for doing more, the states that are most acutely exposed to the service economy support increasing antitrust enforcement intensity at about the same average level found in states that tend to support Republican presidents by overwhelming margins. Though not definitive in establishing an electoral connection, this suggests that there was, historically, an alignment between economic interests and political preferences on antitrust policy, and that this alignment cleaved Democratic leaning states into two camps: those where increasing liberalism predictably leads to more support for antitrust intervention and those where increasing liberalism counter-intuitively leads to more *opposition* to antitrust intervention. To the extent that elected Democrats prioritized the demands of this latter group, their accommodation of lax antitrust enforcement may have been rooted in popular demands.

Conclusion

The reformers who are calling for a massive overhaul of the nation's antitrust laws have raised many important empirical claims about how and why antitrust enforcement has remained so lax in recent decades. Though data limitations and a historical lack of interest in antitrust policy among political scientists make it difficult to rigorously assess some of these claims, observational data does provide some perspective on when antitrust policy changed and why elected Democrats

have historically accommodated that change. Specifically, after developing a consistent measure of merger enforcement intensity from 1955-2015, it becomes clear that President Reagan dramatically reset antitrust enforcement priorities during his first year in office and that later Democratic presidents did not materially strengthen antitrust policing of M&A. This critical juncture in antitrust policy delivered significant economic benefits to many relatively affluent service professionals, and some evidence suggests that those benefits have been confined to a handful of states that tend to favor Democratic presidential candidates, some of which are leading the knowledge economy transition. These economic changes, in turn, seem to have motivated significant differences in political behavior, as the states that have reaped the largest economic benefits from lax antitrust enforcement have historically behaved as if they are quite conservative when it comes to public support for increasing antitrust intervention.

Taken together, the results support the reformist contention that President Reagan initiated a sea change in antitrust priorities. But they also suggest that the new equilibrium in antitrust policy is not purely ideological and may have been supported, at least in part, by the idiosyncratic preferences of the service and knowledge economy professionals that have become increasingly important constituencies for the Democratic Party. The results do not rule out ideology among political elites as a strong explanatory factor. The hard constraints imposed upon antitrust officials by Supreme Court Justices, including liberal Justices who have embraced the law and economics movement, suggest that elite ideology has played a crucial role in supporting the new equilibrium. But when representatives for organized labor appeared before Congress in the late 1980s and early 1990s to argue that large mergers and acquisitions were destroying blue collar jobs and being used to dissolve collective bargaining agreements (Fallick and Hassett, 1996), members of the Democratic Party should have theoretically perceived a political opportunity to resuscitate antitrust enforcement. That they did not suggests that Democratic Party leaders, concerned with

winning elections, likely perceived any loss of blue-collar votes to be outweighed by some other electoral benefit. The data that exists, though not conclusive, suggests that Democratic Party leaders may have believed there was more to gain from aligning policy with the views of white collar professionals than there was to lose from alienating blue collar workers.

More broadly, the results suggest that policy feedback may, in some settings, interact with partisan realignment to support the status quo in federal policy-making. Iversen and Soskice (2019) contend, for example, that federal policies supporting the knowledge economy transition were driven by the public demands of middle-class voters, but Short (2022) finds that bi-cameralism, in a setting of divided power, effectively muted the electoral connection and led to policies that responded more intensely to the demands of organized business interests, especially high tech business interests. Short (2022) also concedes, however, that an electoral connection may still explain policy shifts that took place within the Democratic Party as the Party committed itself to knowledge economy development. The current paper suggests that such a connection may partly explain the dramatic shift within the Democratic Party on antitrust enforcement. But it also suggests that prior policy (antitrust deregulation) induced economic changes (large economic benefits for service professionals in certain regions) that made it harder for the Party to rescuscitate antitrust enforcement, as the benefits of prior policy increasingly accrued to a constituency at the center of the Party's realignment.

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Appendix: Antitrust Deregulation and the Politics of the American Knowledge Economy

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April 14, 2022

Appendix A - Estimating Historical Merger Enforcement Intensity

To develop a historical time series reflecting merger enforcement intensity before and after 1980, I first used a generalized additive model to estimate the number of large deals executed from 1955 to 2015, and then estimated the probability of a challenge in each year by dividing the number of merger cases the Department of Justice Antitrust Division (DOJAD) filed by the total number of large deals. I consider a deal to be "large" and vulnerable to regulatory change using the transaction value, or the purchase price. This is consistent with the regulatory view which has historically focused on the value of acquired assets in individual deals and in the aggregate. The FTC, for example, tracked all mergers in manufacturing and mining from 1948 to 1978 and tabulated them according to the value of assets acquired, with a "large merger" series including all deals with acquired assets worth more than $100 million in nominal (not adjusted) dollars. Here I use a lower transaction value, $50 instead of a $100 million, to get more variation in early years, but importantly, I also adjust the threshold for inflation to capture deals of comparable size across years. The model

therefore estimates the number of deals in each year where the deal value exceeded $50 million in 1948 dollars, or about $400 million in today's dollars.

I describe the procedure for estimating the number of large deals in each year at length in the replication file and provide only a brief summary here. In essence, I use transaction level data on M&A in recent decades to estimate the number of large deals executed each quarter using known correlates of aggregate M&A activity, and then use the model to impute the number of large deals in prior years. The SDC Platinum database published by Thompson Reuters provides comprehensive data on mergers and acquisitions from 1982 to the present, including the transaction date and value. The Center for Research on Securities Prices (CRSP) also provides comprehensive data on publicly traded companies that have been delisted because of a merger from 1955 to 2015 including the date of delisting. I use the acquired company's stock market valuation one month prior to the delisting as a proxy for deal value. The Federal Trade Commission publishes the number of deals involving publicly traded companies in mining and manufacturing from 1955 to 1979, and that series can be extended using SDC Platinum Data with appropriate filters. Similarly, aggregate M&A activity is known to be highly correlated with other macroeconomic statistics like aggregate amounts of corporate debt (Baker and Wurgler (2000), pp. 2250). The Flow of Funds (FOF) database published by the Federal Reserve provides quarterly measures of corporate debt from 1955 to the present. I follow Phillippon (2015) in first fitting a model estimating quarterly counts of deals above the inflation adjusted threshold from 1982 to 2015 using SDC data as the dependent variable and CRSP data and FOF data as predictors. I then impute expected deal counts from 1955 to 1981 using the longer time span of the predictors. Last, I sum the quarterly counts to obtain annual measures.

I follow a similar procedure for estimating DOJAD enforcement activity. The official DOJ workload statistics contain the total number of merger cases filed from 1970 to 2015, but are not available for earlier years. I therefore use comparable measures published by Gallo et al. (2000) from 1955 to

1997 to estimate a linear model, and impute official measures from 1955 to 1969.

Given annual estimates of the number of merger cases filed and the number of large deals from 1955 to 2015, I calculate enforcement rates or percentages by taking the ratio. By this measure, enforcement rates reached their peak in 1961 when the DOJAD filed 28 cases but only 10 large deals were consummated. To obtain a measure resembling a likelihood of challenge, I normalize these estimates by dividing by this maximum value (2.8). The resulting metric estimates the likelihood of facing a DOJAD challenge relative to the high enforcement rates of 1961.

Appendix B - Inequality Model

Descriptive Facts About M&A and Economic Inequality

The main model reported in the paper was born of two basic observations, shown in Figures 1 and 2 respectively. In Figure 1, the horizontal axis captures the logarithm of the ratio of total large M&A value and total fiscal income in each year from 1962 to 2015. This essentially captures the aggregate amount of large deals in each year relative to total income. The vertical axis captures the top 1 percent share of fiscal income in the same year at the federal level (nationwide). The grey points correspond to the period prior to antitrust deregulation, and we observe no significant relationship between relative M&A activity and income inequality. In contrast, after antitrust deregulation, from 1981 to 2015, there appears to be a strong positive correlation between income inequality and the amount of M&A activity.

If such a relationship is causal, it might explain why income inequality occasionally declines. Figure 2 shows the time series for the top 1 percent fiscal income share from 1975 to 2015. The fact that has garnered the most attention, of course, is the significant increase in the top 1 percent share from about 10 percent to over 20 percent since 1980. But there are also fluctuations around that

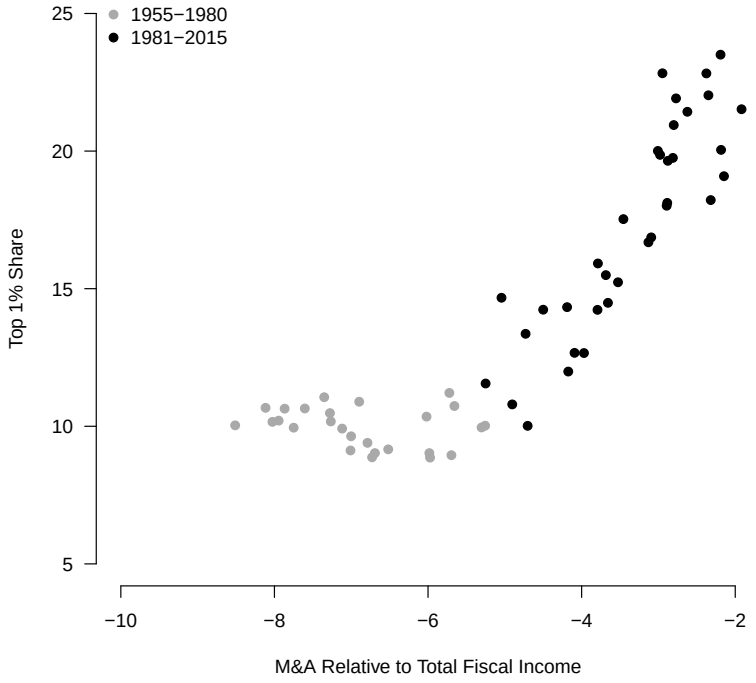

Figure 1: The horizontal axis shows $log(a_t/b_t)$ where a_t represents the total dollar value of large mergers in year t and b_t represents total fiscal income in the same year. This captures the amount of money that changes hands due to mergers and acquisitions as a share of total income on the log scale. The vertical axis shows the top 1 percent share of fiscal income. As indicated in the legend, grey points are for the period 1962-1980 while black points are for the period 1981-2015.

trend that deserve explanation. The grey boxes represent periods since 1981 in which M&A activity grew at least 25 percent slower than total income over the prior year, periods which overlap with economic recessions. In these periods, income inequality can substantially decline. Moreover, the effect seems to be immediate. Given these observation, the challenge was to develop a theory as to how and why changes in the growth of this kind of deregulated financial activity immediately translates into changes in income inequality in the same year.

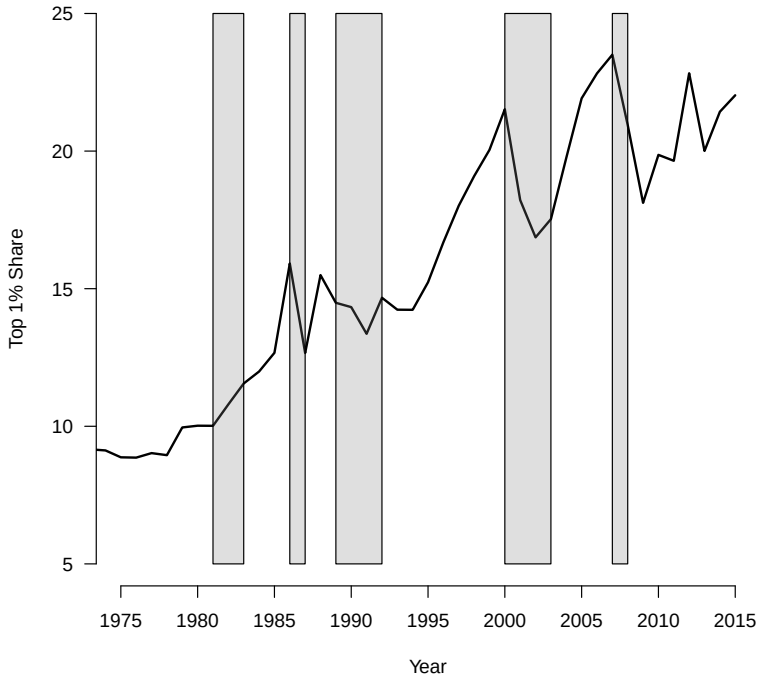

Figure 2: This figure shows the time series of top 1 percent income shares in the US from 1975-2015. The grey bands represent periods since 1981 in which M&A activity grew at least 10 percent slower than total income over the prior year. These downturns appear to be correlated with periods in which fiscal income inequality significantly declines after 1980.

Potential Mechanisms by Which M&A Exacerbates Inequality

Theoretically, large mergers and acquisitions create a number of opportunities for top earners to increase their share of income from labor. First, large deals are often quite complex, and the attorneys for both corporate entities often need to perform a significant amount of due diligence. According to the data from the World Inequality Databse, the threshold income for the 99th percentile in the United States in 2014 (based on pre-tax factor income) was about $445,000 in constant 2016 dollars. Only the most senior associate attorneys would likely surpass this threshold income. A seventh-year associate at a large law firm might earn a salary of about $300,000 per year today, but it would take a fairly generous bonus (by law firm standards) for an associate at that

level to make it into the top percentile of the national income distribution. Many law firm partners, however, easily fall within the top percentile. Average compensation for law firm partners was about $716,000 per year in 2014 (Lowe, 2016). Holding all else equal, if either the number of large deals or the size and complexity of a given number of deals increases, total law firm profits from M&A activity should also increase, as should per partner income.

Investment banks also profit from rising levels of M&A activity. Associate salaries for investment bankers are generally much lower than for attorneys, though bonuses often make up a much larger portion of their overall compensation. Some fourth-year associates at firms that pay above market could conceivably make it into the top percentile in years with high numbers of large transactions (Butcher, 2016). More senior directors and partners will quite often fall within the top percentile even in bad years. Investment banks are generally compensated based on a commission calculated as a percentage of the transaction size, and so as the size of mergers and acquisitions increases, compensation for both non-equity directors (with bonuses tied to performance) and equity partners (with compensation tied to firm profits) should increase.

Mergers and acquisitions are pursued for a variety of business reasons but, as suggested above, one of the primary justifications for relaxing antitrust scrutiny of mergers and acquisitions around 1981 was to create a "takeover market" to oust executives who do not meet shareholder expectations. Deals undertaken for this purpose necessarily create opportunities to renegotiate executive compensation, and while it is possible that compensation could be adjusted downward, that outcome is not likely for two reasons. First, only large firms can reasonably finance large acquisitions and such acquisitions typically increase the merged firm's size, even if the overarching goal is divestiture. Since executive compensation appears to be correlated with firm size (Gabaix, Landier, and Sauvagnat, 2013), executives considered for new management are likely to fall within the top percentile of the income distribution and can expect increased compensation after the merger is completed. Also,

executives may have simply become better at extracting firm profits for personal gain since the 1970s (Bebchuck and Fried, 2004). Changes in federal regulation of executive compensation after 1980 may have contributed to this process (Davis and Thompson, 1994; Wallace and Ferris, 2006; Hacker and Pierson, 2010; Lazonick and Mazzucato, 2013).

Each of these mechanisms is a plausible channel through which those who reside within the top of the labor income distribution in any given year might further increase their share of labor income. But mergers and acquisitions could also exacerbate capital income inequality. Many mergers are undertaken with the intent of increasing the overall market valuation of the merging companies. If mergers and acquisitions do actually increase shareholder value, and if the top percentile of the capital income distribution disproportionately benefit from that increase in stock market value, then M&A activity might also exacerbate capital income inequality. Also, many private equity and hedge fund managers charge a relatively menial management fee, which the IRS recognizes as labor income, but recognize substantial income from the profits generated by the mergers they manage, and that income is technically taxed as capital income (the "carried interest" provision). This is an additional mechanism by which M&A activity might exacerbate capital income inequality.

The following model for the relationship between aggregate income inequality and merger activity allows us to explore whether any of these mechanisms are facially plausible. First, assume that the amount of income earned by the top 1 percent of a given income distribution (fiscal, labor, or capital) in year t, represented by i_t, depends on the nominal value of M&A activity above a given cutoff in year t, represented by M_t. If those in the top 1 percent earn a time-constant percentage of total M&A activity, β_2, this relationship could be modeled as:

$$i_t = \beta_0 + \beta_1 t + \beta_2 M_t + \mu_t \tag{1}$$

where the β_1 coefficient captures the slope of any linear exogenous time trend.

Second, since income inequality and not income levels are the primary concern, divide both variables of interest, i_t and M_t, by total income (fiscal, labor, or capital), I_t, and take the logarithm of our main predictor to allow for diminishing marginal returns.[1] Letting $X_t = log(\frac{M_t}{I_t})$, the relationship of interest then becomes:

$$Is_t = \frac{i_t}{I_t} = \beta_0 + \beta_1 t + \beta_2 X_t + \mu_t \tag{2}$$

In this specification, Is_t is the top 1 percent income share in year t and the coefficient, β_2, captures the relationship between top income shares and merger activity relative to different measures of total income (fiscal income or the pre-tax labor or capital component of national income).

As an exploratory exercise, I take the first difference of this specification to eliminate exogenous time trends and allow the difference in $log(\frac{M_t}{I_t})$ to interact with a dummy variable, D_t, representing the policy interventions associated with the start of the Reagan presidency. The specification then becomes:

$$\Delta Is_t = \beta_1 + \beta_2 \Delta log(\frac{M_t}{I_t}) + \beta_3 D_t + \beta_4 \Delta log(\frac{M_t}{I_t}) \times D_t + \epsilon_t \tag{3}$$

where Δ indicates differencing between time periods (i.e. $\Delta Is_t = Is_t - Is_{t-1}$) and the residuals, ϵ_t, are equal to $\Delta\mu_t$. Taking the logarithm of the main predictor gives the main independent variable of interest the qualities of an elasticity in Equation 3 because:

[1] Investment banks are known to step down their proposed commission percentage as the final acquisition price exceeds certain thresholds. Lawyer and executive compensation may also behave similarly.

$$log(\frac{M_t}{I_t}) - log(\frac{M_{t-1}}{I_{t-1}}) = log(\frac{M_t/M_{t-1}}{I_t/I_{t-1}}). \quad (4)$$

As a result, the model essentially captures the correlation between changes in top income shares and the extent to which the rate of growth in merger activity exceeds the rate of income growth. The model specified in Equation 3 can be estimated with OLS regression given data on annual top income shares and total income, available at the World Inequality Database, and the amount of annual merger activity. The annual amount of merger activity relative to income can be estimated following the basic procedure reported above and in the paper.

Table 1 shows the results. One key finding is the statistically significant coefficient on the interaction term, $\Delta \log(M_t/I_t) \times D_t$, when the dependent variable is inequality in fiscal income (column 1). This suggests that the relationship between fiscal income inequality and the growth in M&A activity relative to fiscal income changed by an order of magnitude after 1981. Another key finding is that the immediate part of this effect operates predominately through labor income, as the coefficient on the interaction term remains significant when the dependent variable is inequality in labor income (column 2) but is not when the dependent variable is inequality in capital income (column 3). The benefit accruing to top earners in years with substantial amounts of M&A activity might snowball into higher capital income shares in later years through a delayed effect (Saez and Zucman, 2016). But the immediate effect is observed in the labor income distribution. A third key finding is that the effect size is roughly three times larger for fiscal income inequality (column 1), which includes capital gains than for labor income inequality (column 2) which does not. This may reflect the influence of the "carried interest" provision which allows fund managers who engage in high levels of M&A to tax the substantial portion of their income as a capital gain rather than labor income (Saez and Zucman, 2016, p. 545), at a substantial tax savings.

	Fiscal income share	Labor income share	Capital income share
	(1)	(2)	(3)
$\Delta \log(M_t / I_t)$	0.160*	0.012	0.015
	(0.083)	(0.024)	(0.212)
D_t	0.284	0.098	0.792**
	(0.321)	(0.113)	(0.356)
$\Delta \log(M_t / I_t) \times D_t$	1.567***	0.502**	0.302
	(0.581)	(0.201)	(0.360)
Constant	−0.045	0.071*	−0.534*
	(0.136)	(0.037)	(0.311)
N	53	48	48
R^2	0.183	0.144	0.136
Adjusted R^2	0.133	0.085	0.078
Residual Std. Error	1.348 (df = 49)	0.498 (df = 44)	1.024 (df = 44)
F Statistic	3.659** (df = 3; 49)	2.457* (df = 3; 44)	2.317* (df = 3; 44)

*p < .1; **p < .05; ***p < .01

Table 1: This table shows the results of estimating Equation 3. The first term, $\Delta \log(M_t / I_t)$, captures growth in merger activity relative to growth in total income. The second term, D_t, is a dummy variable equal to 0 before 1981 and 1 after. The last term is an interaction between these two independent variables which captures a change in the relationship between merger activity and various forms of income inequality after 1981.

Final Published Model

The final model used in the paper essentially extrapolates from equation 2 to include variation at the state level and over time. Instead of interacting the measure of relative merger activity with a dummy variable representing policy intervention, separate regressions are run for the before and after time period and the difference in effect sizes is apparent from the reported coefficients. The key innovation is based on the fact that, in both periods, each of the states is differentially positioned to capture income from merger activity. I argue in the paper that this lagged exposure, $\xi_{i,t-1}$ essentially captures the extent to which each state i is integrated into (or exposed to) the service economy in period t, and it can be explicitly modeled in a variety of ways.

The model estimated in the paper is:

$$Is_{it} = \alpha_i + \beta_t + \delta\xi_{i,t-1} + \gamma(\xi_{i,t-1} \times X_t) + \epsilon_{it} \tag{5}$$

This is effectively the panel equivalent of equation 2, except that the model now includes state and

year fixed effects in lieu of a single intercept and a time trend, and the main measure of service economy exposure, $\xi_{i,t-1}$, is included as a control because it changes over time. Because X_t is co-linear with the year fixed effects, β_t, it does not enter into the model as a predictor. When using first differences to eliminate the state fixed effects and to reduce autocorrelation in the residuals (see replication file), this reduces to:

$$\Delta Is_{it} = \beta_t + \delta\Delta\xi_{i,t-1} + \gamma(\Delta(\xi_{i,t-1} \times X_t)) + \epsilon_{it} \tag{6}$$

The estimated coefficient on that control, δ, captures the extent to which changes in service economy exposure are associated with income inequality. In the paper, this measure of exposure actually has a significantly negative coefficient in the first period, which suggests that, holding all else equal, as states became more exposed to the service economy from 1963-1980, they tended to see income inequality decline. The coefficient remains negative in the second period (1981-2015) but is imprecisely estimated, and so we cannot reject the null hypothesis of no relationship between service economy exposure and income inequality at conventional confidence levels.

To estimate this equation in the paper, I use top top 0.1 percent fiscal income shares as the main dependent variable. Though labor income statistics might be preferable to fiscal income statistics, they are yet not available in long time series at the state level. I therefore use fiscal income statistics but control for a possible effect through capital income by controlling for state level capital gains and the interaction between a state's exposure to capital markets (its share of capital gains in the prior year) and relative growth in M&A activity.

I use each state's share of total income in legal services in the prior year as my main measure of exposure because it is highly correlated with the share of income in securities and commodities

brokerage (which covers investment banking), but is more granular[2] and is consistently reported across the shift from SIC to NAICS codes in 1997. It is also well correlated with a measure of exposure based on the advisory fees earned from reported mergers, as shown in the replication file.

I estimate equation 6 using first differencing instead of demeaning to deal with state fixed effects because the estimates are more conservative and because it substantially reduces autocorrelation in the residuals. I also use the square root of the number of tax units in state i and year t as weights to reduce the observable heteroskedasticity that is likely due to measurement error surrounding income statistics from smaller states (less than 2 million tax units) (Jayaratne and Strahan, 1996, pp. 649). Standard errors are clustered at the state level.

The results reported in the paper are robust to variations in functional form, to the use of the top 1 percent of fiscal income as a dependent variable, to the use of a lagged average of exposure that smooths fluctuations within states, to the use of alternative measures of exposure based on M&A advisory fees or shares of income in securities and brokerage services, and to the inclusion of controls for capital gains at the state level, exposure to capital markets, and capital gains earned by way of M&A activity.

An Exploratory Bayesian Model with Heterogeneous Effects

It is possible that the chosen measure of exposure does not capture all of the heterogeneity among states in their ability to extract income from financial activity like M&A. For example, the dependent variable is a top income share at the state-year level, but the income threshold for achieving an income in the top 1 or 0.1 percent varies substantially by state while the salaries for the investment bankers and lawyers who work on these deals tend to be set according to national pay scales. At the same time, data on these thresholds are not as reliable as the estimated top shares, and even if

[2]In the state level GDP data published by the Bureau of Economic Analysis, the finance and insurance industry code is at the three-digit code level (523x) while the legal services code is at the four digit level (5411).

accurate measures were readily available, it is not a simple matter to account for variations in these thresholds in determining a state's exposure in the model above.

As an exploratory exercise, I therefore used the also fit the following hierarchical model:

$$\Delta Is_{it} \sim \mathcal{N}(\beta_t + \delta_i \Delta\xi_{i,t-1} + \gamma_i(\Delta(\xi_{i,t-1} \times X_t)), \sigma_i^2) \tag{7}$$

$$\beta_t \sim \mathcal{N}(\mu_\beta, \sigma_{\beta^2}) \tag{8}$$

$$\delta_i \sim \mathcal{N}(\mu_\delta, \sigma_{\delta^2}) \tag{9}$$

$$\gamma_i \sim \mathcal{N}(\mu_\gamma, \sigma_{\gamma^2}). \tag{10}$$

As shown in the replication file, we can also allow for correlation between δ_i and γ_i, so that

$$\begin{pmatrix} \delta_i \\ \gamma_i \end{pmatrix} \sim N\left(\begin{pmatrix} \mu_\delta \\ \mu_\gamma \end{pmatrix}, \begin{pmatrix} \sigma_\delta^2 & \rho\sigma_\delta\sigma_\gamma \\ \rho\sigma_\delta\sigma_\gamma & \sigma_\gamma^2 \end{pmatrix}\right)$$

and $\rho \sim \mathcal{U}(-1, 1)$, but there is little evidence of any correlation and the model converges more readily if the coefficients are modeled as being independent.

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