Casting roles, casting votes:

Lessons from Sesame Street on media representation, racial biases, and voting

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Abstract

Evidence on the media's potential to reduce prejudice is limited. *Sesame Street*'s positive representation of minority characters and working women was distinctive in the media landscape of 1969. Using age and technological variation in broadcast reception, we show that *Sesame Street* reduced prejudice decades later and impacted voting, a consequential outcome. Exposed cohorts in high coverage counties are 4.2 ppts more likely to vote, have lower measures of racial biases, and report more votes for minority and women candidates to the U.S. House by 8.1 and 5.8 ppts respectively. On ballots featuring white men, turnout gains are split between parties.

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1 Introduction

Can child mass media shape prejudices and racial biases in adulthood? Can it impact who we elect as representatives as adults? Women and minority groups in the U.S. have been underrepresented in media roles or have been represented through portrayals of negative and damaging stereotypes. Concern has been drawn in the popular press and scientific community to the long-run impact of these patterns. As a result, much research on mass media has focused on its detrimental effects. Yet media technology does not intrinsically dictate harmful content. Indeed, we know little about mass media's potential to *reduce* prejudice in the long-run, particularly when exposure occurs in childhood. Media instruments, such as television, can present great opportunity, with the potential to impact large populations, increasing their exposure to, knowledge of, and respect for social groups. This is particularly true for young children. For many young kids, *Sesame Street*, which took the nation by storm when it started airing in 1969, was a window into a very different America than what they saw in their daily life and in other television programs.

In this paper, we study the effects of exposure to *Sesame Street*, a children's educational television show that portrayed an inclusive, egalitarian and diverse America. We do so using a difference-in-differences strategy that compares cohort differences between high and low broadcast coverage counties, as cohorts who were in their preschool years in 1969, when the show started airing, watched substantially more *Sesame Street* than slightly older cohorts.[1] Using 2006 to 2020 data from a large election survey and an online survey of racial biases we explore these novel and socially important questions, addressing an area where rigorous empirical evidence is lacking. We examine impacts on voting behavior as elections are a consequential context where millions of Americans choose between political candidates, expressing preferences that may be influenced by candidate demographics. We find that exposure to *Sesame Street* impacted elections in two distinct ways. First, *Sesame Street* coverage increased political knowledge and engagement later in life, increasing voter turnout. Second, *Sesame Street* shaped voter preferences in adulthood, many decades later. Younger cohorts in high coverage counties report more votes for minority and women candidates to the U.S. House of Representatives. As a result, they report more votes for Democrats who are more likely to be women and minorities. In elections featuring two white men, turnout gains are split between parties. We find little evidence that the show changed policy views or political identities in a way that favored one party over another. Instead, we find that *Sesame Street* coverage reduced

[1] This strategy was first used by Kearney and Levine (2019) to investigate *Sesame Street's* impacts on education.

measures of white racial bias against Blacks, and increased interest in gender biases, suggesting that a long-run reduction in biases and prejudice explains the differences in candidate choice. These findings provide needed causal evidence on the impacts of race and gender representation in child media, a topic of concern to parents and policy makers but one for which empirical evidence is scarce. We find that child mass media programming can positively impact long-run patterns of prejudice and bias, demonstrating that representation in mass media is important, not just as role models for underrepresented groups, but also for other audience members whose biases and prejudices are shaped by such representation. Finally, we show that these changes carry important implications, namely for the vote shares received by underrepresented candidates in elections.

It is hard to overstate the cultural impact *Sesame Street* had in the United States. First airing in November 1969, the show rapidly became extremely popular. It is estimated that in the early 1970s over half of kids between the ages of 2 and 5 watched the show in the previous week if they had the technological capability to do so (Kearney and Levine (2019)). For many kids, the show became an hour-long daily routine.[2] The show was strikingly different from contemporaneous children's programming.[3] In addition to its novel focus on educational content, *Sesame Street* portrayed a thriving racially diverse urban community, modeled off of New York's African-American neighborhoods. Its human cast was exceptionally diverse from the outset, and all human characters, including women, held jobs.[4] This choice of cast and setting was deliberate. The show was designed to educate young kids, particularly from lower income groups (Cooney (1967)). Input was sought from experts and practitioners in education, child development, psychology and the arts, including Dr. Chester Pierce, a psychiatrist and expert on the psychiatric consequences of racism, and the effects of television's portrayals of minorities (Harrington (2019)). Dr. Pierce played an important role in defining the affective skills that the show sought to encourage in its viewers such as improving children's self-image and racial tolerance (Long (1973), Fisch et al. (1999), Lesser (1974)). Thus *Sesame Street* portrayed an urban integrated community in a positive light, with Black actors cast as role models and figures of authority, working married women, friendly and

[2]Early *Sesame Street* episodes lasted 60 minutes and had 130 episodes per season. Re-runs were also common.

[3]*Mr. Rodger's neighborhood*, began in 1968. Early cast members were European-American, except for Officer Clemmons, an African-American police officer. He appeared in two episodes in August 1968 and in 15% of episodes between 1969 and 1972. *Captain Kangaroo* was a popular children's show from 1955 to 1984 that aired on CBS, which had different broadcasting patterns. Mr. Baxter, its first regular Black character, made limited appearances starting in 1968.

[4]Human characters in season 1 were Susan, an African-American nurse; her spouse Gordon, an African-American veteran and history teacher; Bob, a European-American music teacher; and Mr. Hooper, a European-American shopkeeper. David, an African-American law student; and Linda, a deaf European-American librarian joined in season 2. Season 3 added Luis, the Mexican-American owner of the Fix-It shop; and Maria, a Puerto Rican librarian.

egalitarian interactions between cast members of different races and between adults, children, and Muppets.[5] For many kids, *Sesame Street* was their first glimpse at a diverse America, allowing them to form strong bonds with fictional characters whose race did not always match their own.[6]

Recognition that biases and prejudices carry important social and economic costs motivates an extensive literature in child development, psychology, and economics that seeks to understand and reduce prejudice (Lang and Spitzer (2020), Rutland and Killen (2015), Paluck et al. (2021), Bertrand and Duflo (2017)). Much of this work is guided by contact theory, the idea that prejudices can be reduced by interacting and cooperating with equal status and common goals (Allport et al. (1954)). Field and natural-experiments have generally found that interpersonal contact in schools or through sporting activities reduces in-group bias in the short and medium-run (Paluck et al. (2019), Corno et al. (2022), Boisjoly et al. (2006), Carrell et al. (2019), Finseraas et al. (2019), Rao (2019), Mousa (2020), Lowe (2021)). Only a few recent papers have examined the long-run effects of interpersonal childhood contact finding impacts on political preferences and neighborhood choice (Billings et al. (2021), Brown et al. (2021), Merlino et al. (2022)). In this paper we show that positive contact via children's media can also induce long-run effects on biases.

Whether media based contact with celebrity role-models or out-group characters can impact prejudices has been less studied, especially in the long-run (Schiappa et al. (2005), Paluck et al. (2021)). A few papers show that famous minority role models, such as President Obama or the soccer star Mohamed Salah, reduced indicators of prejudice in adults in the short-run (Plant et al. (2009), Marble et al. (2021)). Though a large literature has studied role-model impacts on the aspirations and education of underrepresented groups (Dee (2005), Bettinger and Long (2005), Eble and Hu (2020)), evidence on how minority role models impact majority biases is limited (Kearney and Levine (2020), Bertrand and Duflo (2017)).[7] A growing literature is exploring the media's role in worsening racial and ethnic tensions (Wang (2021), Ang (2023), Müller and Schwarz. (2023), Adena et al. (2015), Yanagizawa-Drott (2014)). Few papers explore media's potential to improve these outcomes. Paluck (2009) and Blouin and Mukand (2019) show that radio soap operas in

[5]These groundbreaking choices were controversial. In May of 1970, the Mississippi State Commission for Educational Television voted to ban the airing of *Sesame Street* on the state's educational television station. Public backlash against this decision led to the show's rapid reinstatement (Greene (2019)).

[6]Upon the death of Emilio Delgado, a long time cast member, memorial Twitter comments often expressed that he was their first exposure to Hispanic culture. Loretta Long who played Susan, recalled of a 1970 cast trip to Jackson Mississippi that *"Little white kids would reach out to kiss me or 'Gordon,' the other Black character, and you could see their mothers were uneasy. But they'd loosen up, because how can you hate someone who makes your child so happy?"*

[7]Beaman et al. (2009) see less stereotyping of women by men in Indian villages that randomly had to elect women.

Rwanda reduced inter-group prejudice; Armand et al. (nd) shows that U.S. radio broadcasts in the 1940s of *The Adventures of Superman* increased racial tolerance; and Jensen and Oster (2009) show the important effects TV had on gender attitudes in India. This paper contributes to this space by showing that representation in children's television reduced biases and impacted voting in the very long-run.

Though not previously documented, there are reasons to believe *Sesame Street* could have these impacts. Biases are thought to be more malleable in early childhood as children's responsiveness to race and gender is still evolving and can be shaped by children's environment (Beelmann and Heinemann (2014), Rutland and Killen (2015), Hughes et al. (2023), Heron-Delaney et al. (2011), Raabe and Beelmann (2011)).[8] Indeed, evidence suggests that shows like *Sesame Street* can reduce race and gender prejudices in the short-run, hinting at the possibility of long-run effects (Cole et al. (2003), Mares and Pan (2013), Vittrup and Holden (2011)).[9] Because children may be especially sensitive to media content, representations of race and gender in child media are a contentious topic for policy makers and parents. Yet despite frequent discussions in the popular press on how minority and gender representations in children's media impacts kids, this paper is the first to causally identify long-run impacts. Existing economic research on children's media is limited, and has generally focused on education and human capital outcomes. In their examination of *Sesame Street*'s impacts on academic and labor market outcomes, Kearney and Levine (2019) find that the show generated important improvements in students' grade-for-age during their school years, though long-term effects on educational attainment and wages were small and inconclusive.[10] Other work on child media's human capital impacts has found a range of

[8]While newborns exhibit no differences, by 3 mo. infants look more at own- than other-race faces and can better distinguish between them by 9 mo. (Kelly et al. (2005), Kelly et al. (2007)). Showing infants images of other-race faces attenuates the emergence of own-race preferences (Heron-Delaney et al. (2011)). Most children correctly identify race by ages 2 to 3 (Raabe and Beelmann (2011)). By ages 3 to 5, explicit in-group favoritism in white children is common, peaking at ages 6 or 7 and declining thereafter, likely as children respond to social norms against racial bias (Raabe and Beelmann (2011)). In a large meta-analysis, Raabe and Beelmann (2011) find that the development of explicit prejudices in majority-group children correlates with environments where contact with the minority group is low. See Hughes et al. (2023) for an overview of the developmental literature on race.

[9]Cole et al. (2003) look at the impact of two *Sesame Street* inspired shows designed by the Children's Television Workshop (CTW) for broadcast in Israel and Palestine. They find improved out-group attitudes for Israeli-Jewish preschoolers after 4 months of exposure. In an international meta-analysis of *Sesame Street* using CTW generated data, Mares and Pan (2013) find positive short-run impacts on attitudes towards social differences. Vittrup and Holden (2011) find positive effects on children's out-group attitudes shortly after having children watch videos that featured positive relationships between a racially diverse cast. Two of the five videos used were excerpts from *Sesame Street*.

[10]Early academic studies in psychology and education as well as impact evaluations commissioned by governments and its producers generally conclude that *Sesame Street* was effective at teaching its targeted curriculum (Fisch et al. (1999), Fisch and Truglio (2014) and Murphy (1991)), and had positive effects on pro-social behaviors (Paulson (1974), Leifer (1975), Bankart and Anderson (1979), Zielinska and Chambers (1995)).

effects, likely due to differences in content and substitutions in time use.[11] In this paper we show that children's media has long-run effects on outcomes other than human capital by impacting bias formation through representations of minorities and gender roles. Our findings complement the work by Adukia et al. (2023) who document how minorities and females are underrepresented in prize winning children's books and show that representation in the children's books consumed in local communities correlates with local politics. Adukia et al. (2023) substantially increased our understanding of the media context in which children develop. The natural experiment we study, examining the launch of one of the most widely watched children's television shows, allows us to consider the implications of the patterns they document as we can identify the causal effects of exposure to representation on prejudice and voting in later life.

Finally, these findings also contribute to the literature examining the media's impacts on voter preferences and turnout (Campante et al. (2022)). Turnout effects likely depend on whether a media source increases or reduces exposure to political information (DellaVigna and Kaplan (2007), Oberholzer-Gee and Waldfogel (2009), Sørensen (2019), Gentzkow (2006), Falck et al. (2014), Ellingsen and Hernæs (2018)). Existing work has focused on adult news media with the exception of Durante et al. (2019) who find that light entertainment TV in Italy in the 1980s increased populist voting, especially for those exposed in childhood. They hypothesize exposure reduced political knowledge. In contrast, we document for the first time an increase in political knowledge and voter turnout as a result of exposure to children's media. Our contrasting findings highlight the importance of considering media content, even in the analysis of children's media.

The paper is organized as follows. Section 2 presents the data. Section 3 details our empirical strategy. Section 4 documents our first finding, that *Sesame Street* coverage increased electoral participation. Section 5 examines our second result, showing how *Sesame Street* coverage shaped respondents' preferences over candidates. Section 6 examines which candidate attributes explain changes in voter behaviors. Section 7 shows impacts on racial biases. Section 8 concludes.

2 Data

Reported voting behavior and ballot composition: This paper uses post-election survey responses on political behavior and public opinion collected by the Cooperative Congressional Election Study

[11] These include positive human capital impacts (Gentzkow and Shapiro (2008), Riley (2019)), negative impacts (Hernæs et al. (2019), Durante et al. (2019)) and ambiguous correlations in the cross section (Fiorini and Keane (2014)).

(CCES) in election years from 2006 to 2020.[12] In addition to self-reported voting behavior, the CCES also provides validated voter registration and voter turnout by matching respondents to the Catalist database of registered voters.[13] The main sample consists of 57,221 survey responses from non-immigrant citizens born between 1959 and 1968, who faced 2,957 distinct *major party ballots* for their U.S. House Representative.[14] *Major party ballots* are defined as a ballot where the two front-runners are a Democrat and a Republican, both of whom receive over 5% of their district's vote.[15] The 3,025 distinct major party ballots thus defined constitute 87% of ballots for the U.S. House of Representatives between 2006 and 2020. Race and gender demographic information was collected for the two front-running candidates and are observed for both candidates on over 99% of these ballots.[16] Of the 5,914 candidacies in our data, 522 (8.8%) are election runs by Black candidates, 320 (5.4%) are by non-Black Hispanic candidates, 4,891 (82.7%) are by non-Hispanic white candidates, and 181 (3.1%) are by candidates of other races (East Asian, South Asian, Native American, Middle Eastern, Pacific Islander...). 1,288 candidacies (21.8%) are election runs by women. 689 of the major party ballots feature a non-Hispanic white candidate running against a minority candidate and 1,034 feature a woman running against a man. Table A2 of the appendix provides, for the major party ballots in our sample, a full picture of ballot composition by major party candidate demographics and the number of survey responses associated with each type of ballot. Table A2 also reports summary statistics for election results by ballot composition. On average, 16% of these U.S. House elections are won with less than a 10 point margin of victory. In addition to the CCES data, we also examine data from the Current Population Survey (CPS) voting and registration supplement which also records self-reported turnout and registration.

Race and gender biases: Data on race and gender biases is available from Harvard's *Project Im-*

[12]Survey questions vary from year to year. Kuriwaki (2021) harmonized key variables and generated harmonized weights for the cumulative 2006-2020 dataset that are used in all estimations.

[13]Registration and voting is validated by matching the CCES data through Catalist LLC., a political data vendor. Their matching procedure is validated in Ansolabehere and Hersh (2012). They find that Catalist correctly identifies 94% of voters, 96% of non-voters, and a respondent's race in 99.9% of cases.

[14]Responses from Washington DC and responses missing information on either 1969 *Sesame Street* coverage, or ballot demographics are omitted. In election years, respondents are contacted both prior to, and shortly after, election day. The main sample uses responses from individuals who complete the post-election survey. There is no evidence that exposure to *Sesame Street* coverage is associated with post-election survey completion (see appendix table A1).

[15]Election statistics are published by the clerk of the U.S. House of Representatives and are available from MIT Election Data and Science Lab (2021).

[16]Candidates were categorized into four mutually exclusive racio-ethnic categories (Black, non-Black Hispanic, other, non-Hispanic white) and two gender categories (man, woman). Demographics were compiled using three existing sources. Information for candidacies between 2006 and 2014 is from Fraga and Hassell (2021); data for the 2018 and 2020 elections is from OpenSecrets; and data for women candidates was made available by the Center for American Women and Politics (CAWP). Data on 2,359 additional candidacies was compiled by research assistants.

plicit, a research platform that makes online implicit association tests (IATs) available to the public. Anyone can log on to their website, take an IAT, and complete the accompanying survey which also includes questions on explicit biases.[17] Implicit association tests (IATs) are psychological tests designed to measure implicit associations respondents have about individual characteristics. For this project, we focus on the race IAT and the gender-career IAT data.[18] IATs have been used in psychology, and in economics, as a way of measuring implicit attitudes (Glover et al. (2017), Carlana (2019), Corno et al. (2022), Lowes et al. (2015)). Our main datasets consists of the 350,080 race IAT scores and the 84,479 gender-career IAT scores of U.S. resident citizens born between 1959 and 1968, along with their survey responses.[19]

***Sesame Street* coverage:** The predicted share of households in a county that could watch *Sesame Street* in their homes in 1969 was estimated by Kearney and Levine (2019). These predicted coverage rates were calculated using data from the 1968–1969 edition of the trade publication, *TV Factbook*. This lists all commercial and non-commercial broadcasting television stations, their technical and geographical specifications (UHF or VHF signal, signal power, location, and height of the tower) and, for commercial stations only, the coverage rates for surrounding counties.[20] Using this information, the authors estimate the empirical relationship between a station's technical specifications and a county's coverage rate using the commercial station sample. This relationship

[17]Because test takers self select into the sample, *Project Implicit* data is not representative and comparisons to a representative sample are not available (Ratliff and Smith (2021)). Potential impacts on selection into taking an IAT test are discussed in section 3.2.

[18]When completing the race IAT, respondents are sequentially presented with images of Black and white individuals, and words that have positive or negative connotations. For the gender-career IAT respondents are presented with gendered names, and words associated with career or family. Respondents complete a series of rapid sorting exercises grouping together the images, or names, with words. The IAT is meant to measure the strength of a respondent's association between individual characteristics and word connotations, as sorting is easier when associated items are sorted together. IAT measures have been the subject of increased scrutiny in the psychology literature (Ratliff and Smith (2021)). There is generally agreement that these measures are relevant and predictive, particularly for socially sensitive topics (Greenwald et al. (2009), Bertrand and Duflo (2017)), aggregate regional measures (Hehman et al. (2019)), and voting outcomes (Gawronski et al. (2015), Jost (2019)). Disagreements have centered on the validity of the *implicit* bias construct, and if it differs from *explicit* biases, and whether the low test-retest reliability is due to high measurement error, or the construct itself being time variant (Gawronski (2019), Schimmack (2021), Connor and Evers (2020)).

[19]Respondents that report having taken three or more IATs are dropped from the sample. In survey years where respondents report their age rather than their year of birth, birth years are coded as the $\textit{survey year} - \textit{age}$ for surveys taken between July and December, or the $\textit{survey year} - (\textit{age} + 1)$ for surveys taken between January and June.

[20]Particularly important to coverage rates was the broadcast signal type: UHF versus VHF. VHF signals travel further and are less impacted by physical obstacles. Televisions at the time were generally equipped to receive VHF signals. However demand for expanded channel offerings and the removal of legal barriers by the Federal Communications Commission led to a growth in UHF broadcasting channels, with newer television sets equipped to receive both UHF and VHF signals. Uptake of this new technology was costly though. Television sets were found in 95 percent of households, but a UHF signal could only be received in 54 percent of these. More details on the construction of predicted coverage rates are available in Kearney and Levine (2019) and its appendix.

is then applied to the non-commercial stations that broadcast *Sesame Street* to generate a simulated estimate of their coverage rates. Finally, counties are assigned the simulated coverage rate associated with the best simulated signal received from surrounding towers. Figure 1a maps the Kearney and Levine (2019) estimates of simulated coverage levels in each county. Figure 1b shows the distribution of coverage rates assigned to CCES survey respondents based on their county of residence. Note that because this simulated estimate of the coverage rate is estimated using the average relationship between broadcast technology factors and signal receipt, it is not a function of the quality of receiving television sets in the county which is likely correlated with household wealth. Though national ratings data is not available for the early years of *Sesame Street*, these estimated coverage rates can be validated using the limited Nielsen ratings that are available for 28 metropolitan areas (Haydon (1973)). They strongly predict *Sesame Street* ratings, with viewership in the past week increasing by 0.59 percent in the 2 to 5 year old demographic for each percentage point increase in predicted coverage. Consistent with the 69.4% national coverage rate that was estimated for *Sesame Street* in 1970, these measures generate a national predicted coverage rate of 65% with a standard deviation of 20 percentage points.

3 Empirical strategy

To identify the effects of early childhood exposure to *Sesame Street* coverage on voting, we employ the same identification strategy as Kearney and Levine (2019). We limit our analysis to individuals born between 1959 and 1968 who were between the ages of 1 and 10 when *Sesame Street* began airing in 1969.[21] We compare the outcomes of cohorts born between 1959 and 1963, who would have been attending elementary school when the show started airing, to slightly younger individuals in the same county, born between 1964 and 1968, who would have been 5 and younger. Preschool attendance was uncommon in this period with only 9% of 3 year olds and 19% of 4 year olds attending preschool in the 1970 census. For 5 year olds, 57% attended kindergarten with 88% of kindergartners enrolled in half day programs (Kearney and Levine (2019)). Not only were preschool aged children the targeted demographic of the show's producers, they would generally have had the opportunity to watch *Sesame Street* if the household could receive the broadcast signal. Indeed, our limited viewership data from 28 metropolitan areas confirms that 2 to 5 year olds accounted for 71% of the show's viewership, while 19% were between 6 and 11, and 9.7% were older than 12.

[21] These respondents would have been between their late thirties and early sixties when they were surveyed in the election years between 2006 and 2020 or when they completed their IATs.

In addition to the variation in cohort exposure, differences in broadcast coverage rates between counties generates a second dimension of exposure variation for identification. We identify likely treated respondents as those from cohorts that were under the age of 6 when the show first aired *and* who live in high broadcast coverage counties. Likely untreated respondents are those who were 6 or older in 1969 as well as those who were younger but who live in counties with a poor broadcast signal. The main empirical models used in estimations using the CCES data on voting related outcomes are specified below,

$$Y_{icjdy}^{CCES} = \beta_0 + \beta_1(preschool69_i * SSCov_j) + \beta_2 X_i + \gamma_{scy} + \delta_{jdy} + \epsilon_i, \tag{1}$$

$$Y_{icjdy}^{CCES} = \lambda_0 + \sum_{c=59, c\neq 63}^{68} \lambda_{1c}(Cohort_{ic} * SSCov_j) + \lambda_2 X_i + \gamma_{scy} + \delta_{jdy} + \epsilon_i. \tag{2}$$

Y^{CCES} is the CCES outcome variable of interest for individual i, in cohort c, residing in county j, and voting in congressional district d in year y. The indicator variable $preschool69_i$ is set to 1 if the individual would have been below the age of 6 when *Sesame Street* first aired. This is interacted with the variable $SSCov_j$, the predicted level of *Sesame Street* coverage in county j. Figures also present disaggregated cohort level estimates of λ_{1c}, as specified in equation 2 where $Cohort_{ic}$ is an indicator variable set to one if the respondent was born in cohort c.

Our preferred specification includes X_i, controls for an individual's gender and race. For the CCES data, γ_{scy} is a (*state* $\times$ *cohort* $\times$ *year*) fixed effect which captures electoral behaviors and preferences in a particular election year that would impact all respondents in a state from the same birth cohort as well as the impact of any time-varying state level policy shocks or events that may have impacted particular cohorts within a state over their lifetime. δ_{jdy} is a (*county* $\times$ *congressional district* $\times$ *year*) fixed effect which captures the electoral behaviors and preferences of respondents that are constant across cohorts residing in the same county and voting in the same congressional district election. β_1 is the coefficient of interest. It identifies the difference in the voting patterns on a congressional district ballot between the preschool age cohorts in high *Sesame Street* coverage counties, as compared to slightly older cohorts in their county voting on the same ballot, while controlling for the voting patterns of same age cohorts in their state who were surveyed in the same year. This specification differs slightly from that used in Kearney and Levine (2019). The fixed effects are augmented to control for congressional districts, and survey years, to capture all aspects of a particular congressional election that affect all respondents in the congressional district (the presence of third party candidates, idiosyncratic shocks, up-ballot elections, etc.). For non-voting outcomes, controlling for voting patterns in a particular election is not necessary, and not always possible as congressional districts are not observed in other datasets.[22] For these outcomes we estimate β_1^{KL} using the same fixed effects as Kearney and Levine (2019): a (*state* $\times$ *cohort*) fixed effect, γ_{sc}^{KL}, and a *county* fixed effect, δ_j^{KL}.

Robustness of key results to the inclusion of additional controls is explored in appendix table A3. Given the educational nature of the show, we estimate specifications that control for a respondent's education, and income. We also estimate specifications that include interactions between cohort indicators and county characteristics as local socio-economic conditions that correlate with *Sesame Street* coverage might have impacted cohorts differently. The inclusion of additional controls does not alter our main estimates meaningfully. Key results are also robust to using a very simple difference-in-differences specification, where we regress outcomes on the $preschool69_i \times SSCov_j$ interaction; the $preschool69_i$ indicator; and $SSCov_j$; with race and gender controls but no fixed effects. For key outcome variables, we report all three of these coefficients in panel b of appendix table A3, providing insight into how these outcomes correlate with the preschool cohort indicator and *Sesame Street* coverage—correlations that are otherwise unobserved, as they are absorbed by the fixed effects in our main specification. Finally, permutation tests for key results are presented in appendix figure A1. Placebo coefficients for all key results are normally distributed around 0 and none are as large in magnitude as our main coefficients.

A measure of coverage, not viewership. It is important to note that our treatment measures a county's exposure to *Sesame Street* broadcast coverage, not actual viewership. Because our ratings data is very limited, we cannot estimate the impact of actual viewership. Rather, our estimates are "intent-to-treat" estimates, identifying the policy relevant effect of making the show more accessible in a county rather than the impact of watching the show on an individual.[23] For interpretation, we follow the approach used in Kearney and Levine (2019). As seen in figure 1b, the theoretical comparison of counties with 0% and 100% coverage is out of sample. We will interpret the effects of a 30 percentage point increase in the predicted coverage rate of a county, roughly the gap between the 25th percentile (48.5%) and the 75th percentile (81%) of the predicted coverage distribution. This can be thought of as the equivalent of moving from a typical county

[22]For turnout and registration outcomes in the CPS data, we estimate β_1^{CPS} using δ_{jy}^{CPS}, a $(county \times year)$ fixed effect as congressional districts are not observed in the CPS voting data.

[23]Though of great interest, micro-data on viewership, and even national county level ratings data is unavailable, limiting our ability to evaluate selection into viewership. A very limited examination of heterogeneity in ratings using the 28 metro areas where viewership data is available is reported with heterogeneity results in appendix section A2.

with weak coverage, like Los Angeles county which had a predicted coverage rate of 49%, to a typical high coverage county, like Sacramento county, which had an estimated coverage rate of 79%.

Unobserved time substitutions. It is also important to highlight that treatment in our context is composed of both increased exposure to *Sesame Street* coverage, but also a time substitution away from other activities. These alternate activities could have been viewership of other TV programs, or entirely different uses of children's time.[24] While we frame our interpretation as the impacts of *Sesame Street* coverage, our estimates could also be picking up the effects of reduced exposure to other television programming in which minorities and women were not represented in a respectful or egalitarian manner.

3.1 Migration: Selection and attenuation bias

Ideally we would observe respondents' counties of residence in childhood and be able to assign them to $SScov_j^{child}$ to estimate β_1^{child}. As this is unobserved in the CCES and IAT data, β_1 estimates β_1^{adult} as it is estimated using $SScov_j^{adult}$, the 1969 coverage in respondents' county of residence when surveyed as adults, between 2006 and 2020. Substantial out-of-county migration between childhood and adulthood will generate attrition bias, potentially biasing our results towards zero. The possibility that exposure affected out-of-county migration, and destination choice, is also a concern. We explore these issues below.

We find no evidence that younger cohorts residing in counties that had high *Sesame Street* coverage differ in their probability of being life-long residents of a city as compared to older cohorts in the same county. Problematic selection bias could arise if exposure to *Sesame Street* coverage impacted respondents' inter-county mobility between 1969 and their survey date. Though we do not observe county of residence in 1969, in many waves CCES respondents report how long they have lived in their current *city*. We identify city residents who have lived in their city since they were five or younger and estimate both $\hat{\beta}_1^{KLadult}$ and the simplified specification on this indicator. In both cases, we fail to reject the null of no effect as reported in column 1 of panel a, table A4. We also estimate $\hat{\beta}_1^{KLadult}$ on indicators for being non-Hispanic white (column 4) and female (column 5). We find no evidence that younger cohorts residing in counties that had high

[24] Data on children's time-use for this period is not available limiting our ability to examine patterns in time substitution when the show began airing.

Sesame Street coverage differ in their demographics from older cohorts in the same county.

We find no evidence that younger cohorts residing in counties that had high *Sesame Street* coverage differ in their probability of facing diverse congressional ballots as compared to older cohorts in the same county. We estimate both $\hat{\beta}_1^{KLadult}$ and the simplified specification on the indicator for facing a minority-white ballot and the indicator for facing a woman-man ballot. For both specifications and both indicators we fail to reject the null of no difference between the treated and untreated cohorts as reported in column 2 and 3 of panel a, table A4. All respondents residing in counties that had high *Sesame Street* coverage in 1969 are more likely to face such ballots, as indicated by the positive coefficients on $SScov_j^{adult}$ in the simplified specification. This is not surprising given that these counties tend to be more urban (see figure 1a) and less white, as indicated by the negative coefficient on $SScov_j^{adult}$ in column 4, table A4. However, as compared to slightly older cohorts in their county, younger cohorts in high coverage counties are no more likely to reside in parts of the county that face a diverse congressional ballot.

We find no evidence that *Sesame Street* exposure impacted the probability of residing in one's childhood state or county. To estimate the effects of $SScov_j^{child}$ on migration, we turn to restricted-use data from the Panel Survey of Income Dynamics (PSID). We identify 1,204 original panel members born between 1959 and 1968 whose county of residence is observed in both 1969 and 2019.[25] Using this sample, we estimate $\hat{\beta}_1^{KLchild}$, reported in panel b of table A4, on indicators for residing in 2019 in the same state (column 2) and county (column 3) as in 1969. In both cases, we fail to reject the null of no effect. There is no evidence that younger cohorts from high coverage counties differ in their likelihood of residing in their childhood county or state as compared to slightly older cohorts from their county.

We find no evidence that *Sesame Street* exposure impacted the probability of residing in a county that had high *Sesame Street* coverage as an adult. Even if the show did not impact the propensity to move, it may have impacted destination choice. Problematic selection would arise if younger and older cohorts from high coverage counties reside as adults in counties that differ in their 1969 coverage rates. To check this, we estimate $\hat{\beta}^{KLchild}$ on $SSCov_j^{adult}$. Results, reported in panel b of

[25] The original household sample included 4,312 members from these cohorts. In column 1 of panel b table A4 we estimate $\hat{\beta}^{KLchild}$ on an indicator for attrition between the original PSID sample and 2019. We find no evidence that exposure to *Sesame Street* coverage impacted attrition in the PSID.

table A4, column 4, fail to reject the null of no such difference. Critically, we can rule out that a 30 percentage point increase in the coverage rate impacted the 1969 *Sesame Street* coverage rate in adult counties of residence outside of a [-6.4, 5] percentage point range with a 95% confidence level.

Coverage rates between destination and sending counties are correlated for county-to-county migrants, counteracting attenuation bias. Only 47.3% of the observed PSID sample reside in their childhood county in 2019, raising the issue of attenuation bias.[26] Nevertheless, the correlation between the $SSCov_j^{adult}$ and $SSCov_j^{child}$ in the PSID is high at 0.547 (p-value$<$0.001) as reported in column 1, row 1 of table A5. The severity of attenuation bias is reduced, as even among inter-county migrants, these coverage rates are correlated (column 2). For the 30% of inter-county migration that occurs between neighboring counties the correlation between adult and childhood county coverage rates is high at 0.789 (column 3). We also observe a positive correlation of 0.101 (p-value=0.06) between $SSCov_j^{adult}$ and $SSCov_j^{child}$ for long distance movers (column 4). These patterns in the PSID are consistent with general US mobility data, reported in row 2 of table A5.[27]

Overall, we find no evidence that the show impacted inter-county migration, and we find no evidence that exposed cohorts differentially select into high or low coverage adult counties of residence. Nevertheless, data limitations due to the small size of the PSID sample prevent us from completely ruling out some impact on migration and destination choice. Accordingly, we interpret our estimates of β_1 most accurately as the difference between younger and older cohorts residing in counties that had high *Sesame Street* coverage in 1969. The balance of evidence suggests observed differences are the result of exposure to the show and that, to the extent that migration is affecting our estimates, it is primarily attenuating them. In table A6 we report estimates of our main effects on the small sub-sample of CCES respondents who report living in the same city since before age 6. Consistent with attenuation bias being the main impact of migration on our estimates, we generally find larger effects for this sub-sample, though the small sample size generates imprecise estimates.

[26] 9% of the CCES sample are city never-movers though this survey question will not capture respondents who moved between cities within a county, and respondents who moved, but then returned, to their childhood city.

[27] Using census estimates of 2016-2020 county-to-county moves for each county pair in our main sample, we regress the destination county's coverage on the sending county's coverage, weighting the regression by the pair's total county-to-county movers. Results in the second row of table A5 show very similar correlation estimates.

3.2 Selection into the *Project implicit* data

A particular concern with analysis of IAT scores collected by *Project Implicit* is selection into the sample. IAT tests are publicly available online through the *Project Implicit* website and any individual can log on and takes a test. Exposure to *Sesame Street* coverage may have affected the probability individuals select into the sample, biasing estimates.[28] To check for selection, we calculate $ShareSS_j = \frac{N_{j,1964 \leq Cohort_i \leq 1968}}{N_j}$, the share of our sample observations in each county j coming from treated cohorts. We then regress $ShareSS_j$ on the county's predicted coverage rate, $SSCov_j$, to check if a larger share of a county's IAT test takers are from treated cohorts in counties that had higher levels of *Sesame Street* coverage. Results are reported in appendix table A7. Column 1 presents estimates using all the race IAT test takers, column 2 using only non-Hispanic white race IAT test takers, and column 3 presents estimates for the gender-career IAT test takers. For the race IAT, both estimates in columns 1 and 2 are small and not statistically significant (p-value$\in [0.31, 0.84]$). There is no evidence that exposure to *Sesame Street* coverage changed selection into the race IAT data. In contrast, *Sesame Street* coverage increased the likelihood of taking the (less widely used) gender-career IAT test.[29] Given this selection, our analysis in section 7 will focus on the race IAT data. Nevertheless, it should be noted that this impact on selection into taking the gender-career IAT test suggest the show may have increased awareness of gender and career biases.

4 *Sesame Street* coverage increased electoral participation

In this section, we document our first finding: that younger cohorts in high *Sesame Street* coverage counties are more likely to turnout to vote and have an active voter registration, likely due to an increase we observe in political interest and knowledge. Examination of our second set of results on voter preferences for candidate demographics is presented in section 5.

Younger cohorts in high *Sesame Street* coverage counties vote more. The $\hat{\beta}_1$ estimate of 0.139 (p-value $<$ 0.001) on verified general election turnout, reported in the first row of table 1, panel a, column 2, implies that the probability of younger cohort members turning out to vote increases by 4.2 percentage points, a 6.6% increase, if they reside in a county where the coverage rate was

[28] If more biased individuals are less likely to take IAT tests, selection will bias estimates towards zero.

[29] The positive and statistically significant estimate of 0.034 implies that a 30 percentage point increase in the *Sesame Street* coverage rate increased the share of test takers from treated cohorts in a county by 1.0 percentage point.

30 percentage points higher. This result is highly statistically significant and robust to alternative specifications.[30] Figure 2a plots the cohort level $\hat{\lambda}_{1y}$ estimates, confirming a discontinuity in validated turnout between the 1964–1968 cohorts and the 1959–1963 cohorts. The estimated $\hat{\beta}_1$ and $\hat{\lambda}_{1y}$'s are similar when using the indicator for self-reported election turnout, plotted in figure 2a and reported in the second row of table 1, panel a. Critically, self-reported voter turnout is also observable in the CPS voting and registration supplement.[31] The estimated coefficient of 0.108 (p-value = 0.051) on self-reported turnout in the CPS, reported in column 8, is also positive and statistically significant. We cannot reject equality (p-value= 0.77) with the CCES point estimate of 0.092 (p-value = 0.06) estimated on a comparable sample.[32]

It is important to note how similar the estimates using verified and self-reported turnout are, even though there is a large gap between the 85% of respondents who self-report voting and the 63% for whom turnout is validated. Recent work on the CCES data finds that this gap is due to overreporting turnout to surveyors (Enamorado and Imai (2019)). One may suspect that exposure to *Sesame Street*'s pro-social messaging could (ambiguously) affect the propensity to misrepresent one's turnout by either increasing the desire to supply a socially desirable response or by decreasing the propensity to lie to a surveyor. To check this, we generate an indicator for inconsistent validated and self-reported turnout records. Estimated $\hat{\beta}_1$'s on this indicator are reported in the third row of table 1, panel a. We see no impact on the probability of having inconsistent verified and self-reported turnout records. There is no evidence that younger cohorts in high *Sesame Street* coverage counties differentially misrepresent their behavior to surveyors.

To better understand why *Sesame Street* coverage impacts turnout, we examine respondents' reasons for not voting.[33] Results are presented in appendix table A8. Younger cohorts in high *Sesame Street* coverage counties are less likely to report not voting because they dislike the candidates and because they are not registered. These effects are robust across specifications and explain about

[30]It is robust to using the same specification as Kearney and Levine (2019) (column 3), to the simplified specification, and to alternative specifications that include added controls (appendix table A3, column 1). Dropping data from 2006 and 2008, years when the quality of turnout validation was poorer (Grimmer et al. (2018)) give a similar estimate of 0.117 (p-value = 0.018). Figure A1 plots the coefficients from 1000 permutation tests that randomly assign cohorts (panel a) or counties (panel e). Placebo coefficients are normally distributed around 0 and none are as large in magnitude.

[31]Self-reported turnout is also observed in the Performance of American Elections and the American National Election Studies but these smaller datasets contain too few observations from the 1959–1968 birth cohorts for estimation.

[32]The CPS voting data is coded to be comparable to our CCES sample: we drop non-citizens and naturalized Americans, as well as responses from proxy respondents. The CPS does not report congressional districts and county codes are only available for some CPS counties. For comparability, in column 6 we limit the CCES data to observations from counties identified in the CPS and estimate β_1^{CPS} using (county × year) and (state × cohort × year) fixed effects.

[33]This question is only administered to self-reported non-voters and thus does not shed light on the reasons for non-turnout for respondents who misreport their voting behavior.

half of the effect on reported non-turnout. The other reported estimates are less clear but taken jointly suggest an increased interest in participating in elections, increased knowledge about the voting process, and an increased willingness to incur non-pecuniary costs to vote.

Younger cohorts in high *Sesame Street* coverage counties register to vote more. The $\hat{\beta}_1$ estimate of 0.091 (p-value = 0.017) on having a verified active registration record, reported in the first row of table 1, panel b, column 2, implies that for a 30 percentage point increase in the coverage rate, the probability of having a verified active registration record is 2.7 percentage points higher, a 3.7% increase. This effect is statistically significant, and robust to alternative specifications (column 3). Figure 2b plots the cohort level $\hat{\lambda}_{1y}$ estimates and confirms a discontinuity in validated registrations between younger and older cohorts. While still present, the estimate of $\hat{\beta}_1$ on self-reported voter registration is smaller in magnitude by about half with little visual evidence of a discontinuity between treated and untreated cohorts. Self-reported registration status is also collected in the CPS data where we also observe a small positive effect, albeit statistically insignificant (column 8). Importantly, we cannot reject equality between the CPS and CCES estimates (p-value= 0.74) when estimated with the same specification and subset of counties.

It is worth noting that a substantial share of respondents misreport their registration status. 94% of respondents self-report being registered to vote while only 74% have a verified active registration. The estimated coefficient of -0.066 (p-value= 0.07) on an indicator of inconsistency, in the third row of table 1, panel b, implies that younger cohorts in high coverage counties are less likely to misreport their registration status. Is this misreporting driven by misinformation or misrepresentation? Turnout results showed no evidence of an impact on prosocial misrepresentation. Furthermore, evidence discussed below shows that younger cohorts in high coverage counties are more politically informed suggesting that *Sesame Street* coverage increased respondents' knowledge of their voter registration status.

Heterogeneity in this increase in electoral participation is examined in the appendix. Table A9 evaluates differences by respondent demographics. We cannot reject the null of no heterogeneity in estimated effects. There is no evidence that these turnout effects are driven by a specific demographic group. Furthermore, appendix table A10 shows that turnout increases hold for different ballot compositions, regardless of candidate demographics, and are also observed during midterm elections which are not affected by the demographic composition of the presidential ballot

($\hat{\beta}_1 \in [0.121, 0.178]$ across the sub-samples). Consequently, this effect on electoral participation should be thought of as distinct from the effect on preferences over candidate demographics that we document in section 5.

How did *Sesame Street* increase electoral participation? Below we explore several explanations.

We find no evidence of a difference in educational attainment and income. Given the findings in Kearney and Levine (2019) we start by considering income and education channels. The CCES survey collects coarse information on respondents' education levels and their household income. Treatment effects on years of education and family income are reported in appendix table A11. Like Kearney and Levine (2019) we cannot reject the null of no effect on educational attainment (in years) and family income. Though the coefficients are positive at 0.167 (p-value= 0.49) and 3.5 (p-value= 0.36) respectively, they are not statistically significant. The same estimation applied to the IAT data also fails to reject the null of no effect at 0.034 (p-value= 0.56) on educational attainment. Furthermore, estimates on our main outcomes reported in appendix table A3 that include controls for educational attainment and family income do not substantially differ from out main estimates. Increased educational attainment and family income are not the main mechanisms behind our results.

Younger cohorts in higher *Sesame Street* coverage counties have greater political knowledge and interest. Estimates in table 2, panel a, show they are more likely to recognize the names of their senators and representatives, with an estimate of 0.179 (p-value = 0.012) out of 3 names. They report a greater interest in politics with an estimate of 0.155 (p-value = 0.048) on a 4 point scale and are less likely to report accessing no news media in the past 24 hours with an estimate of -0.029 (p-value= 0.15), though this last result is not statistically significant.

Younger cohorts in higher *Sesame Street* coverage counties are more likely to report moderate political identities, but do not differ in strong political identification or active political engagement. With respect to political identification, a more complex picture emerges in table 2, panel b. Younger cohorts in high coverage counties are more likely to report identifying with a political ideology and major political party. Political ideology is measured in both the IAT and CCES data in which we estimate treatment effects of 0.037 (p-value $<$ 0.001) and 0.024 (p-value=

0.53) respectively. Though the CCES estimate on having a political ideology is not statistically significant, we do observe increased identification with major political parties in the CCES data with an estimate of 0.055 (p-value= 0.08). All of these effects are driven by respondents who express weak preferences towards a political party or identity.[34] These indicators of political affiliation are a common form of political engagement. The majority of respondents indicate that they identify with a political party and ideology. We do not observe a treatment effect on reporting a strong party identification, or rarer more involved forms of political engagement, examined in panel c, associated with individuals highly engaged in the political process such as voting in primary elections, donating money to political campaigns, putting up political signs, attending political meetings and working for a candidate or campaign. For all of these indicators, we are unable to reject the null of no treatment effect (p-values $\in [0.19, 0.77]$).[35] Consistent with this, table A13 shows that estimates of β_1 on electoral participation are larger for marginally engaged respondents, as proxied by the fact that their primary election turnout is not verified.

How do these findings on political participation compare to the existing literature? Much causal work on voter turnout has focused on interventions conducted shortly before elections. In a meta-analysis of RCTs, Green and Gerber (2019) estimate an average treatment effect of +4 percentage points as a result of door-to-door canvasing, similar to our estimated 4.2 percentage point effect of residing in a high coverage county. More comparable to our treatment are findings reported in Sondheimer and Green (2010) that examines the causal-effects on long-run validated voter turnout from three well-known early childhood education interventions.[36] They report treatment and control differences in validated turnout ranging from +2 to +8.8 percentage points. Cohodes and Feigenbaum (2021) also find a +6 percentage point impact on turnout from improvements in educational quality at the high school level. With regards to mechanisms, our results suggest that the impacts of *Sesame Street* coverage on registration and turnout stem from an increase in the political engagement and knowledge of marginal voters. This is consistent with meta-analyses of the literature on voter turnout that highlight party identification, political interest, and political knowledge as consistently linked to individual turnout (Smets and Van Ham (2013)). Previous work has also identified increases in political interest, knowledge and participation following the

[34]These respondents describe their preferences as "leaning towards", "not strong preference for", "slightly" or "moderate" as opposed to "strong" or "very".

[35]We also find no evidence of a difference in measures of non-political civic engagement as proxied by blood donation, union membership, and having been part of the military, reported in appendix table A12.

[36]These are the Perry preschool experiment; the I Have a Dream natural experiment; and the STAR experiment.

introduction of public television in Norway in the 1960s (Sørensen (2019)).

5 *Sesame Street* coverage shaped who receives reported votes

In this section, we examine how *Sesame Street* coverage impacted voter preferences between candidates for the U.S. House of Representatives. In particular, we examine whether exposure to the show's working women, and the egalitarian and respectful interactions among the racially integrated cast increased reported voting for minority and women candidates.

5.1 Younger cohorts in high *Sesame Street* coverage counties report more votes for minority candidates

We begin by restricting the sample to the 13,120 respondents facing one of the 689 ballots that featured a minority candidate running against a non-Hispanic white candidate.[37] We build four mutually exclusive indicators defined for all respondents: reporting a vote for the minority candidate, the white candidate, a third party candidate or not voting. The share of respondents in each of these categories is reported in the first column of table 3, panel a. Estimates of β_1 using each indicator as an outcome are reported in column 2 of table 3, panel a. Younger cohorts in high *Sesame Street* coverage counties are substantially more likely to report voting for a minority candidate compared to older cohorts in these counties. The $\hat{\beta}_1$ estimate of 0.271 (p-value < 0.001) reported in the first row of table 3, panel a, implies that for a 30 percentage point increase in the coverage rate, the probability a voter reports a vote for a minority candidate is 8.1 percentage points higher, a 20% increase. This effect is highly statistically significant, and robust to alternative specifications (column 3 and appendix table A3, column 2).[38] About half of this effect can be attributed to reductions in reported non-voting, as indicated by the negative and statistically significant coefficient reported in the fourth row. The other half is driven by reduced reported voting for white candidates, as indicated by the -0.144 (p-value$=$ 0.09) estimate reported in the second row of table 3, panel a. That reported voting for white candidates is lower for younger cohorts, despite their higher turnout, signals that we should think of this effect on preferences over candidates as distinct from the impacts on electoral participation discussed in section 4.

[37] Note that estimates in appendix table A4 show that the 1969 *Sesame Street* coverage rates in this subset of counties was 9.03 percentage points higher on average. However, we find no evidence that younger cohorts differ in their likelihood of facing this type of ballot, mitigating concerns of a treatment effect on selection into this subsample.

[38] Figure A1 plots the coefficients from 1000 permutation tests that randomly assign cohorts (panel b) or counties (panel f). Placebo coefficients are normally distributed around 0 and none are as large in magnitude.

Figure 3a plots the cohort level $\hat{\lambda}_{1y}$ estimates for the minority candidate indicator, the white candidate indicator and the non-voting indicator, confirming a discontinuity in voter behavior between the elementary aged and preschool aged cohorts. Similar impacts are found when the sample is limited to respondents who report voting for a major party, and validated voters, as reported in appendix table A14. Table A6 also suggests a substantially larger effect at 0.848 (p-value < 0.001) for respondents who have lived in their city since before the age of 6, as compared to 0.216 (p-value $= 0.039$) for those who moved after age 6, a difference that is statistically significant (p-value $= 0.004$). This pattern is consistent with migration generating attrition bias though it should be interpreted with caution given the small subsample. Potential impacts on election outcomes are discussed in appendix A1.

In columns 4 and 5 of table 3 we split the sample to estimate the effect separately for white and minority respondents. We find no evidence of differential effects in reported voting for minority candidates across these groups (p-value$=$ 0.79), though the small sample size generates imprecise estimates for the minority sample. For white respondents, the effect is driven by reduced voting for white candidates while the turnout effect plays a larger role for minority respondents.

5.2 Younger cohorts in high *Sesame Street* coverage counties report more votes for women candidates

We select respondents voting on one of the 1,034 ballots that feature a woman running against a man.[39] As above, we construct four mutually exclusive indicators representing reported votes. Estimates of β_1 on each of these indicators are reported in column 2 of table 3, panel b. Younger cohorts residing in high *Sesame Street* coverage counties are more likely to report voting for a woman candidate. The $\hat{\beta}_1$ estimate of 0.192 (p-value < 0.001) reported in the first row of table 3, panel b, implies that for a 30 percentage point increase in the coverage rate, the probability of reporting a vote for a woman candidate is 5.8 percentage points higher, a 15% increase. This effect is highly statistically significant, and robust to alternate specifications (column 3 and appendix table A3, column 3).[40] Most of the gains by women candidates come from a reduction in the share of reported non-voting, as indicated by the negative and significant coefficient of comparable magnitude reported in row 4 of panel b. Exposure to *Sesame Street* coverage had little effect on

[39]This subset of respondents live in counties that had higher 1969 *Sesame Street* coverage rates, by 11.3 percentage points on average. However, we find no evidence that younger cohorts differ in their likelihood of facing this type of ballot, mitigating concerns of a treatment effect on selection into this subsample (appendix table A4).

[40]Figure A1 plots the coefficients from 1000 permutation tests that randomly assign cohorts (panel c) or counties (panel g). Placebo coefficients are normally distributed around 0 and none are as large in magnitude.

reported voting for men candidates as demonstrated in row 2 of panel b. Reported voting for third party candidates is lower for treated respondents though only a very small share of respondents report third party votes. Figure 3b plots the cohort level $\hat{\lambda}_{1y}$ estimates for the woman candidate indicator, man candidate indicator, and non-voting indicator, confirming a discontinuity in voter behavior between treated and untreated cohorts. Possible implications on election outcomes are discussed in appendix A1.

In columns 4 and 5, we split the sample to estimate the effect separately for female and male respondents. We find no statistically significant differences in reported voting for women across these groups (p-value= 0.79), though suggestively the turnout effect is more important for male voters while female voters switch their reported votes from men to women candidates.

Note that this result is not as clear cut as the effect on voting for minority candidates. When the sample is limited to validated voters reporting a major party vote, in appendix table A14, we cannot reject the null of no effect, though the direction of the coefficient suggests a small positive effect on women candidates' vote share. We also do not find any evidence of a larger effect for respondents who have lived in their city since before the age of 6 (appendix table A6) though the small samples complicates this comparison.

5.3 Younger cohorts in high *Sesame Street* coverage counties report more votes for Democratic candidates.

As in the preceding sections, we build four mutually exclusive indicators, representing respondents' reported votes: whether they report voting for the Democratic candidate, the Republican candidate, a third party candidate or not voting. Estimates of β_1 on each of these indicators, are reported in column 2 of table 3, panel c. Younger cohorts in high *Sesame Street* coverage counties are more likely to report voting for a Democratic candidate. The $\hat{\beta}_1$ estimate of 0.100 (p-value = 0.007) reported in the first row of panel c implies that for a 30 percentage point increase in the coverage rate, the probability a voter reports a Democratic vote is 3.0 percentage points higher, a 7.5% increase. The $\hat{\beta}_1^{KL}$ estimate (column 3) is of a slightly smaller magnitude, but still statistically significant at conventional levels (p-value=0.07). Most of these Democratic gains come from a reduction in reported non-voting, as indicated by the negative and significant coefficients reported in the fourth row. Estimates on reported voting for Republican candidates are positive but smaller in magnitude. Reported voting for third party candidates is lower for treated respondents though only a very small share report third party votes. When the sample is limited to validated voters voting for major party candidates, in appendix table A14, we cannot reject the null of no effect, though the coefficient suggests a small positive effect favoring Democratic candidates.

It is worth highlighting that the impact of *Sesame Street* coverage on reported votes is not driven by a treatment effect on the propensity to misrepresent one's vote. Exposure to *Sesame Street*'s prosocial messaging could have impacted the probability respondents misrepresent their voting behavior, a concern given that we only observe how respondents report voting, not their actual vote. The data suggests this is not driving our results. First, in section 4 we compared reported and verified turnout data and found no treatment effect on respondents' propensity to misrepresent their turnout behavior to surveyors, a common form of misrepresentation in polling data. Second, columns 3 and 4 of table A14 compare estimates on the reported voting of validated voters and non-validated voters who report voting for a major party candidate. Non-validated voters who report a major party vote are respondents for whom the issue of prosocial misrepresentation may be particularly pronounced. In all three panels, we fail to reject the null that treatment effects are equal across these two types of respondents (p-value=$\in [0.65, 0.77]$), rejecting the idea that respondents prone to misrepresentation are driving our results.

6 Disentangling effects on correlated candidate characteristics

The results presented in section 5 show that exposure to *Sesame Street* coverage increased the likelihood respondents report voting for minority, women and Democratic candidates. What candidate attributes voters are responding to is unclear. Many candidate attributes correlate with one another. Democratic candidates are more demographically diverse. 69.3% of women candidacies and 67.8% of minority candidacies are Democratic candidacies. Minority candidacies are also more likely to be women at 33% versus 19.4% of white candidacies. Below we consider which of these observable candidate characteristics voters are responding to. We show that these changes in reported votes are driven by candidate demographics rather than their party.

It is important to note that as we cannot control for all candidate characteristics, we cannot rule out that being a minority or woman candidate correlates with unobservable candidate characteristics that may be generating this voter response. In this section we explore potential mechanisms, ruling out potential confounds. In section 7 we provide evidence that younger cohorts in high *Sesame Street* coverage counties have reduced measures of biases, the most likely explanation for the observed impacts on reported voting.

6.1 Candidates from underrepresented backgrounds drive the increase in reported voting for Democrats

Younger cohorts in high coverage counties report more votes for Democrats. Is this due to different political views or a response to candidate characteristics that correlate with candidate party?

Table 4 re-estimates the effects on reported vote by party for sub-samples of ballots with different demographic compositions. Column 1 limits the sample to the 44% of respondents voting on ballots where both candidates are white men. The turnout effect is still apparent with the -0.093 (p-value= 0.10) estimate on not voting, but the gains in turnout are split between Democrat and Republican candidates, favoring Republicans. Indeed, conditional on voting for a major party, the coefficient on reporting a Democratic vote is negative and not statistically significant at -0.011 (p-value= 0.88), as reported in column 2. On ballots featuring two white men, we cannot reject that voting for both parties was comparably affected by the turnout effect of *Sesame Street* coverage.

In contrast, we see a strong treatment effect favoring Democrats on the sub-sample of ballots where the Democratic candidate is a woman, in column 3, at 0.183 (p-value = 0.002); and where the Democratic candidate is a minority in column 5, at 0.235 (p-value = 0.009). Overall, the increase in reported voting for Democrats is driven by women and minority Democratic candidacies.

6.2 Voting effects for minority candidates holds regardless of party or gender

Younger cohorts in high *Sesame Street* coverage counties report more votes for minority candidates than slightly older cohorts in their counties. In table 5, panel a, we consider whether candidate characteristics that correlate with a candidate's minority identity, namely political party and gender, explain this pattern.[41] Columns 1 and 2 of table 5, panel a, estimate β_1 separately for ballots where the minority candidate is running as the Democrat (column 1) and the Republican (column 2). In both sub-samples, $\hat{\beta}_1$ is positive and statistically significant implying increased reported voting for minority candidates regardless of their party, with suggestively larger effects for Republican minority candidates (p-value= 0.16). In columns 3 and 4 of table 5, panel a, we estimate β_1 separately on ballots where the minority candidate is a woman (column 3) and a man (column 4).

[41] Minority candidates are more likely to run as Democrats. Of the 689 ballots featuring a minority candidate running against a white candidate, the minority candidate is running as a Democrat on 76% of the ballots. Minority candidates are also more likely to be women. 33% of minority candidacies are woman candidacies as opposed to 19.4% of white candidacies.

In both sub-samples, $\hat{\beta}_1$ is positive and statistically significant. Thus the increase in reported voting for the minority candidate holds for both minority men and women candidates, and is possibly larger for women minority candidates (p-value= 0.14).

6.3 Voting effects for women candidates hold for Democratic women only, with larger effects for minority women

Younger cohorts in high *Sesame Street* coverage counties report more votes for women candidates. In table 5, panel b, we consider if this is a response to candidate characteristics that correlate with a candidate's gender, namely political party and minority identity.[42] Columns 1 and 2 of table 5, panel b, estimate β_1 on ballots where the woman candidate is running as a Democrat (column 1) and Republican (column 2) separately. Treated respondents are no more likely to report voting for Republican women than for the Democratic men they run against, though estimates are imprecise as there are only 268 such ballots. In contrast, the $\hat{\beta}_1$ for Democratic women is large and statistically significant suggesting the main effect is driven by increased reported voting for women running as Democrats, though we cannot reject equality between the two coefficients. Columns 3 and 4 of table 5, panel b, estimate β_1 on ballots where the woman candidate is white (column 3) or a minority (column 4). In both sub-samples, $\hat{\beta}_1$ is positive and statistically significant implying increased reported voting for women candidates holds for both white and minority women, though the magnitude of the estimate is larger for minority women, a difference that is statistically significant (p-value= 0.06). Overall, results in panels a and b suggest that reported voting for minority women candidates is doubly affected in high *Sesame Street* coverage counties.

6.4 No clear impacts on preferences for policies or incumbents

Beyond party affiliation, other characteristics may correlate with candidate demographics. Below we examine impacts on support for specific policies, and preferences for incumbent representatives.

Younger cohorts in high *Sesame Street* coverage counties do not differ in their policy preferences in ways consistent with party platforms. Exposure to *Sesame Street* coverage may have shaped treated respondents' policy preferences towards policies that are more likely to be espoused by

[42]Women candidates are more likely to run as Democrats. Of the 1,034 ballots featuring a woman candidate running against a man, the woman is running as the Democrat on 74% of the ballots. Women candidates are also more likely to be minorities. 26.2% of candidacies by women are also minority candidacies as opposed to 20.5% of candidacies by men.

minority and female candidates. To investigate this mechanism, we examine CCES respondents' views on specific policies. Appendix table A15 presents aggregated indices of responses to broad policy topics.[43] Overall, we cannot rule out that younger cohorts in high *Sesame Street* coverage counties support environmental policies or abortion rights at the same rate as older cohorts in their counties. Treatment effects are observed signaling increased support for same-sex marriage, an effect consistent with increased tolerance for diversity more generally. In contrast, we observe reduced support for less restrictive immigration policies. The CCES survey and the IAT survey also include opinion questions on race relations in the US, and race related policies. Panel b presents results for the CESS data and panel c for the IAT data.[44] For questions on race related policies and race relations in the US, estimated effects are generally small, not statistically significant, and when comparable, inconsistent across data sources and question topics.[45] Overall, impacts on policy views are mixed or null and do not clearly align with any particular political platform.

Younger cohorts in high *Sesame Street* coverage counties identify more with political ideologies and parties, without clearly favoring one end of the political spectrum. Table 6, panel a, examines treatment effects on respondents' reported political ideology in both the CCES and IAT data. Younger respondents in high coverage counties are more likely to report identifying with a liberal or a conservative ideology as opposed to being moderate or neutral, as reported in column 2 for the CCES data and column 5 for the IAT data. However, the shift towards having a political ideology is split between liberal and conservative. When we restrict the sample to individuals expressing a political ideology, in columns 3 and 6, we cannot reject the null of no effect on identifying with a liberal ideology (p-value$\in [0.13, 0.63]$). A similar pattern is observed for party identification, a question that is only asked in the CCES data. Results reported in panel b show a reduction in identifying as an independent but conditional on having a party identity, we cannot reject the null of no effect on identifying with the Democratic party (p-value=0.21).

Younger cohorts in high *Sesame Street* coverage counties do not differ in their voting for incumbents. As younger cohorts in high *Sesame Street* coverage counties demonstrate greater political

[43]Administered questions vary across surveys. Selected questions used to construct indices were administered in multiple survey years. Disaggregated estimates are reported in table A16.

[44]Estimates for all questions estimated separately are presented in appendix table A17.

[45]For instance, while we detect statistically significant positive effect on support for affirmative action policies in the IAT survey, the coefficient for the same topic in the CCES survey is negative. We also detect a statistically significant reduction in agreement with statements expressing belief in structural racial barriers for the small sample of Black respondents in the CCES, but not for the full sample of all respondents.

knowledge, they may differ in their propensity to vote for incumbent candidates. We explore this in appendix table A18. There is no evidence of a difference in reported voting for incumbent candidates. Turnout effects are split. Respondents report increased voting for both incumbent and non-incumbent candidates. When we restrict the sample to major party voters, we cannot reject the null of no effect on reported voting for the incumbent candidate (p-value= 0.87).

Overall, the balance of evidence presented in this section suggests that the observed changes in reported voting behavior occur in response to candidate demographics. This conclusion is further corroborated by analysis of the *Project Implicit* data below confirming impacts on exposed cohorts' measures of racial biases.

7 Younger cohorts in high *Sesame Street* coverage counties have reduced measures of racial biases

In this section we show that the observed differences in voter preferences between younger and older cohorts in high coverage counties are the result of a difference in racial biases. *Sesame Street*'s portrayal of positive minority role models in an integrated cohesive community reduced long-run racial biases, increasing the probability of reporting a vote for diverse candidates. Given the evidence of selection into the gender-career IAT discussed in section 3.2, we focus our analysis on results reported in panel a of table 7 that examine effects on the race IAT. It should be noted however that the selection pattern is itself suggestive that exposure to *Sesame Street* coverage also increased interest in issues surrounding gender-career biases, though we cannot make conclusive statements on how it impacted gender-career IAT scores, reported in panel b of table 7.[46]

Younger cohorts in high *Sesame Street* coverage counties have lower long-run measures of white racial bias towards Blacks. Column 2 of table 7 presents $\hat{\beta}_1^{KL}$ estimates on standardized race IAT scores. Estimates in row 1 reveal a negative and statistically significant effect on this measure of implicit bias for non-Hispanic white test takers.[47] The $\hat{\beta}_1^{KL}$ estimate of -0.067 sds (p-value = 0.007) reported in row 1 of column 2 implies that for a 30 percentage point increase in the coverage rate,

[46]If less biased test takers select into taking the test, estimates will be biased towards zero. Given the selection patterns detected, estimates of $\hat{\beta}_1^{KL}$ on the gender-career IAT scores are, unsurprisingly, small and not statistically significant (panel b of table 7).

[47]Figure A1 plots the coefficients from 1000 permutation tests that randomly assign cohorts (panel d) or counties (panel h). Placebo coefficients are normally distributed around 0 and none are as large in magnitude.

standardized IAT scores decrease by -0.02 standard deviations.[48] Like Corno et al. (2022), and consistent with contact theory, we also find an opposite sign effect for Blacks (row 2) though it is imprecisely estimated given the small number of Black test takers. Figure 4 plots the cohort level $\hat{\lambda}_{1y}^{KL}$ estimates on the IAT scores of white test takers, confirming a discontinuity in IAT scores between the elementary aged and preschool aged cohorts.

In addition to IAT test scores measuring implicit biases, the IAT survey also asks respondents explicit questions on preferences and warmth felt towards European and African Americans, which we examine in table 8. Column 2 pools results for non-Hispanic white and Black test takers, results for non-Hispanic white test takers are presented in column 3, and for Black test takers in column 4. Overall, white test takers rarely express explicit in-group preferences. When asked whether they prefer European Americans over African Americans, 60% respond that they like the two groups equally. Similarly, when asked separately about the warmth they feel towards European Americans and African Americans on a 10 point scale, 66% report the same value for both groups (appendix figures A2a and A2b).[49] In this context, it is not surprising that we detect no discernible effect in rows 1 and 3 on the explicit preferences expressed by white test takers (p-value $\in [0.17, 0.51]$). Estimates of $\hat{\beta}_1^{KL}$ for white test takers are small in magnitude and not statistically significant. However, when we limit the sample to test takers who do not report the same value for both groups, in rows 2 and 4, the magnitude of the effects become larger, and a small effect signaling reduced in-group preferences on warmth questions becomes marginally statistically significant (p-value$=$ 0.09).[50] Reporting no preferences across racial groups is less common for African Americans, with 36% reporting that they like European and African Americans equally and assigning the same level of warmth towards these two groups on the 10 point scale (figures A2a and A2b). Effects for this group are less muted and display a similar pattern of reduced in-group preferences for exposed cohorts. Overall, these patterns suggest a muted reduction in explicit measures of in-group preferences that is more pronounced when restricting the data to respondents who are willing to express variation in their warmth towards individuals based on their race.

[48]For reference, Corno et al. (2022) find a -0.63 standard deviation effect on the race IAT of white freshman college students in South Africa measured at the end of their freshman year after being randomly allocated to a shared dorm room with a Black student.

[49]Expressing racial preferences, particularly for white Americans, is stigmatized in contemporary U.S. culture. As such, the pool of test takers reporting no preferences likely includes truthful responses as well as socially-desirable misrepresentations that reduce the signal received from measures of explicit racial preferences.

[50]The -0.192 point estimate suggests that for a 30 ppt increase in the coverage rate, the difference in expressed warmth for European versus African Americans for these white non-zero test takers shrinks by -0.06 points, a 0.017 sd decrease.

Non-incumbent minority candidates gain most from their electorate's exposure to *Sesame Street* coverage. If racial biases are driving the change in voting patterns, we would expect to observe larger effects for lesser known minority candidates. For known candidates, voters can base decisions on observed behaviors, choices, and viewpoints, potentially reducing the influence of racial biases. Consistent with this hypothesis, we find that non-incumbent minority candidates gain most from an electorate's exposure to *Sesame Street*. Table 9 presents heterogeneity results by the incumbency status of the minority candidate. Column 2 shows estimates on incumbent minority candidates while columns 1 and 3 do so for non-incumbent minority candidates that face a non-incumbent (column 1) or an incumbent (column 3). For non-incumbent minority candidates, the impact of *Sesame Street* is large at 0.313 (p-value= 0.23) and 0.363 (p-value = 0.003) respectively for columns 1 and 3.[51] For incumbent minority candidates, while still positive, the estimate is substantially smaller at 0.070 (p-value= 0.62) and not statistically significant. The difference between these two groups approaches statistical significance at the 10% level (p-value= 0.11) when comparing impacts reported in columns 2 and 3.

Given these findings, it is interesting to consider how the impacts of *Sesame Street* may have differed based on the social and racial context in which it aired. A priori it is not clear whether media "contact" with minority role models would have a larger effect in areas where opportunities for other forms of contact were high or low, making this an interesting question for empirical research. Where contact opportunities were few, or where pre-existing racial stereotypes were strong, the show may have filled a vacuum allowing for a large effect. Alternatively, effects may have been larger in areas where preschool age exposure to the show could later be reinforced through other interactions. This question is of academic interest as it would shed light on contact theory mechanisms. Unfortunately, the lack of national county level viewership data complicates interpretation of any heterogeneity analysis based on county level characteristics. This is because the county characteristics we examine may have affected both viewership probability, as parents may have been less likely to allow their children to watch the program, as well as how the show impacted its viewers. Nevertheless, heterogeneity in effects by county characteristics are presented and discussed in appendix A2. Overall, we find no evidence of significant heterogeneity across a number of relevant variables that help characterize the social and racial environment of counties.

[51] Estimates in column 1 are imprecise due to the small number of white-minority ballots with no incumbents.

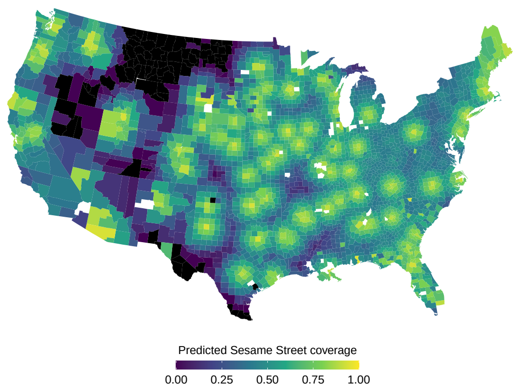
8 Conclusion

Can child mass media shape prejudices and racial biases in adulthood? Can it impact who we elect as our representatives as adults? The importance of representation in the media has increasingly been recognized and discussed in the popular press. How gender and race are approached in child media has also been the subject of recent politicized debate. Yet despite this flood of attention, there is little causal evidence on how positive non-stereotyped representation affects social and economic outcomes.

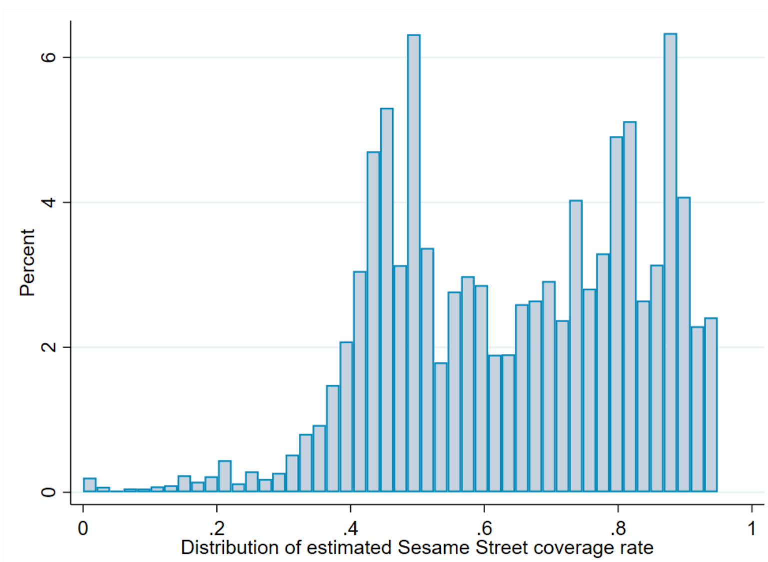
This paper helps inform these debates. In this paper, we show that preschool-age exposure to *Sesame Street* coverage, and its portrayal of an inclusive, egalitarian and diverse America, had long-run effects on adult voting patterns and racial biases. We find that decades later, exposed cohorts were more interested and engaged in the political process, registering and turning out to vote at higher rates than slightly older cohorts in their same county. When voting, these cohorts are more likely to report voting for minority and women candidates. Voting for Democratic candidates increased because of the increase in voting for diverse candidates. When we limit the sample to ballots featuring two white men, turnout gains are split between both parties, consistent with other evidence that the show increased moderate political engagement across the political spectrum. Overall, the evidence suggests that the change in voting behavior stems from a change in voter preferences for candidate demographics, consistent with the reductions in racial bias measures of white IAT test takers we observe for these cohorts.

Our results suggest multiple new directions for future research. Research in economics on the effects of children's media is limited. This is particularly true for long-run effects on outcomes other than education. These results show that child media can have long-run impacts on consequential adult decisions, suggesting that other long-run impacts could be important as well. This paper also demonstrates that exposure to underrepresented role models has important effects on majority group members, a relationship that has not been the focus of work in the past. Finally, this paper shows that mass media can reduce prejudices and biases, revealing an important yet understudied behavioral lever for efforts in prejudice reduction.

9 Figures



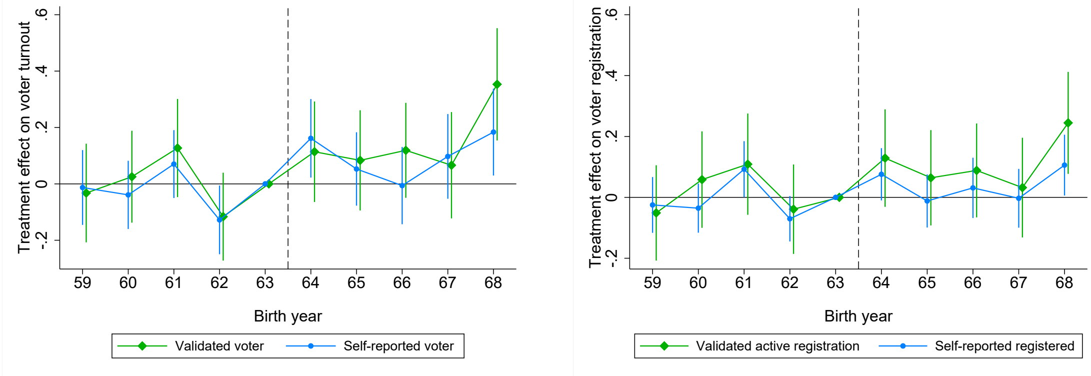
(a) Estimated county coverage rates of *Sesame Street* in 1969



(b) Distribution of estimated coverage rates for CCES respondents

Figure 1. *Sesame Street* coverage

Notes: Estimated coverage is simulated using the average relationship between the broadcast technology of signal towers and signal receipt. In figure a, the color gradient represents the predicted share of households that could watch *Sesame Street* in that county in 1969 as calculated in Kearney and Levine (2019). Figure b plots the distribution of estimated *Sesame Street* coverage rates for our sample of CCES respondents.

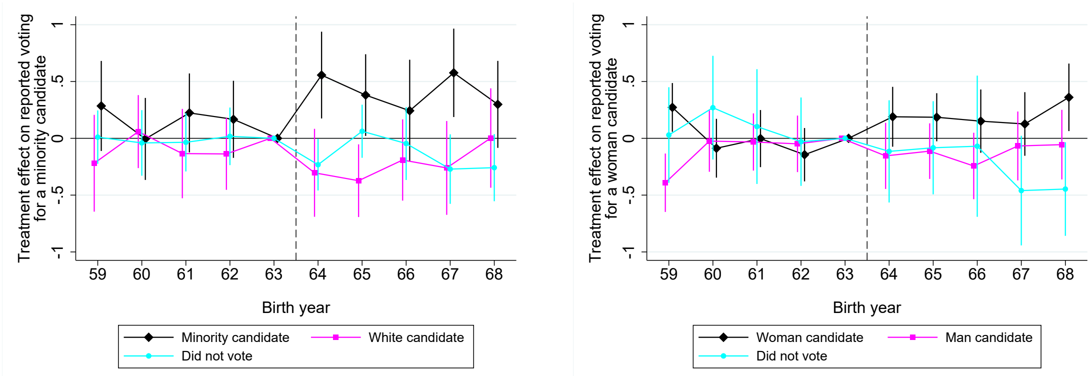


(a) Self-reported and validated turnout

(b) Self-reported and validated registration

Figure 2. Voter turnout and registration is higher for younger cohorts in high *Sesame Street* coverage counties

Notes: These figures plot the $\hat{\lambda}_{1c}$ estimates from the interaction between birth year indicators and the county's predicted *Sesame Street* coverage in 1969 as specified in equation 2. Controls for race and gender as well as ($state \times cohort \times year$) and ($county \times congressional\ district \times year$) fixed effects are included. The outcome variables are indicators variables. The 1963 cohort is the omitted category. Estimates include survey weights. 95% confidence intervals are depicted using standard errors clustered at the county level.



(a) Reported voting on minority-white ballots

(b) Reported voting on woman-man ballots

Figure 3. Reported voting for minority and women candidates is higher for younger cohorts in high *Sesame Street* coverage counties

Notes: These figures plot the $\hat{\lambda}_{1c}$ estimates from the interaction between birth year indicators and the county's predicted *Sesame Street* coverage in 1969 as specified in equation 2. Controls for race and gender as well as ($state \times cohort \times year$) and ($county \times congressional\ district \times year$) fixed effects are included. In each figure, coefficients for three indicators are plotted: reported voting for the minority (figure a) or woman (figure b) candidate; reported voting for the white (figure a) or man (figure b) candidate; reporting not voting. Estimates on reported voting for third party candidates are omitted for clarity. The sample is limited to respondents who face minority-white ballots (figure a) or woman-man ballots (figure b). The 1963 cohort is the omitted category. Estimates include survey weights. 95% confidence intervals are depicted using standard errors clustered at the county level.

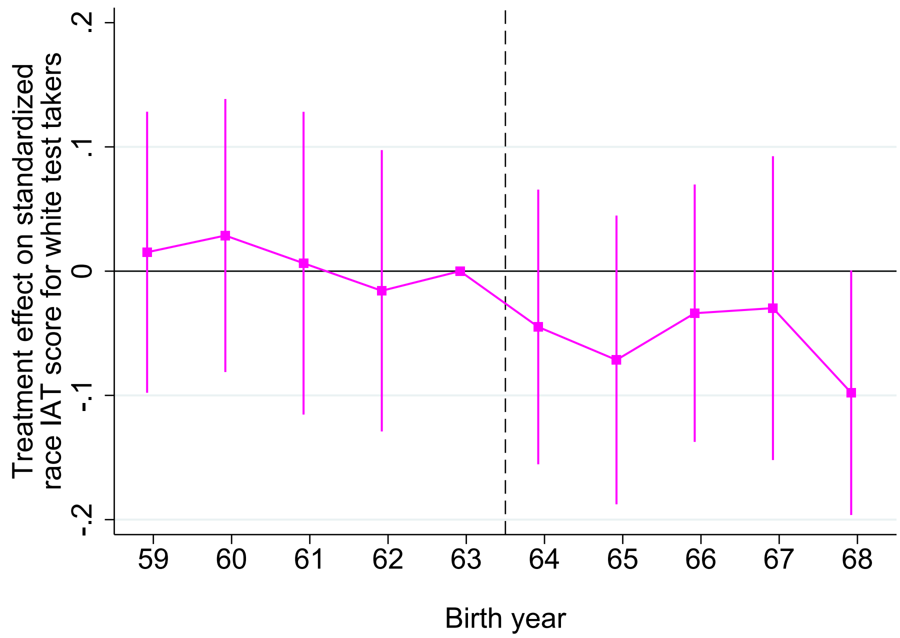


Figure 4. Measures of implicit bias against Blacks is lower for younger white test takers in high *Sesame Street* coverage counties

Notes: These figures plot the $\hat{\lambda}_{1c}^{KL}$ estimates from the interaction between birth year indicators and the county's predicted *Sesame Street* coverage in 1969. Controls for race and gender as well as (*state*×*cohort*) and (*county*) fixed effects are included. The 1963 cohort is the omitted category. Larger positive values indicate stronger Black-bad and white-good associations. 95% confidence intervals are depicted using standard errors clustered at the county level.

10 Tables

Table 1: Voter turnout and registration is higher for younger cohorts in high *Sesame Street* coverage counties

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
	Full CCES			**CCES with CPS counties only**			**CPS**	
Dependent indicator variable	**Dependent mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	**Dependent mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{CPS}$	**Dependent mean**	$\hat{\beta}_1^{CPS}$
Panel a: ***Sesame Street*** **increased self-reported and validated voter turnout**								
Verified general election turnout	0.634	0.139***	0.109***	0.643	0.074	0.096*		
		(0.040)	(0.040)		(0.059)	(0.058)		
		[51,723]	[56,850]		[28,788]	[29,230]		
Self-reported election turnout	0.851	0.120***	0.118***	0.862	0.087*	0.092*	0.716	0.108*
		(0.032)	(0.033)		(0.047)	(0.048)		(0.055)
		[51,723]	[56,850]		[28,788]	[29,230]		[23,016]
Inconsistent self-reported voting status with validation	0.229	-0.034	-0.018	0.230	0.013	-0.005		
		(0.034)	(0.034)		(0.052)	(0.050)		
		[51,723]	[56,850]		[28,788]	[29,230]		
Panel b: ***Sesame Street*** **increased voter registration and knowledge of registration status**								
Verified active voter registration	0.744	0.091**	0.081**	0.747	0.081	0.093*		
		(0.038)	(0.038)		(0.054)	(0.054)		
		[47,604]	[52,030]		[26,454]	[26,832]		
Self-reported voter registration	0.944	0.044*	0.049*	0.948	-0.001	0.005	0.865	0.020
		(0.024)	(0.028)		(0.038)	(0.038)		(0.043)
		[47,353]	[51,777]		[26,338]	[26,720]		[22,939]
Inconsistent self-reported registration status with validation	0.210	-0.066*	-0.062*	0.209	-0.060	-0.078		
		(0.037)	(0.033)		(0.059)	(0.059)		
		[47,353]	[51,777]		[26,338]	[26,720]		
Controls: Gender and race		Yes	Yes		Yes	Yes		Yes
FE: County		.	Yes		.	.		.
FE: State x cohort		.	Yes		.	.		.
FE: County x cong. district x year		Yes	No		Yes	No		NA
FE: State x cohort x year		Yes	No		Yes	Yes		Yes
FE: County x year		.	No		.	Yes		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates are estimated using the same specification as Kearney and Levine (2019), with $(county)$ and $(state \times cohort)$ fixed effects. $\hat{\beta}_1^{CPS}$ estimates control for $(county \times year)$ and $(state \times cohort \times year)$ as congressional districts are not available in the CPS voting data. Columns 1-3 use the full CCES sample. As county of residence is only observable for some counties in the CPS data, columns 4-5 estimate effects in the CCES using only respondents living in counties observed in the CPS. CPS voting data is cleaned and coded to be comparable to our CCES sample: non-citizens and naturalized Americans are dropped, as well as responses reported by proxy respondents. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table 2: Younger cohorts in high *Sesame Street* coverage counties are more politically informed with increased rates of moderate political identities

	(1)	(2)	(3)	(4)	(5)
	CCES			Race IAT	
Dependent variable	**Mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	**Mean**	$\hat{\beta}_1^{KL}$
Panel a: Political knowledge					
Recognized names of elected representatives (out of 3)	2.6	0.179** (0.071) [50,836]	0.139** (0.070) [55,939]		
Interest in government and public affairs (scale 1-4)	3.4	0.155** (0.079) [46,835]	0.165** (0.078) [51,248]		
Accessed no news media in the past 24 hours	0.056	-0.029 (0.020) [41,890]	-0.018 (0.021) [45,616]		
Panel b: Political identification					
Has a political ideology	0.615	0.024 (0.039)	0.015 (0.038)	0.804	0.037*** (0.009)
..... Weakly identifies with a political ideology	0.396	0.017 (0.041)	0.034 (0.038)	0.599	0.027** (0.012)
..... Strongly identifies with a political ideology	0.220	0.007 (0.032)	-0.018 (0.028)	0.205	0.010 (0.010)
N		[51,575]	[56,707]		[319,563]
Identifies with a major political party	0.839	0.055* (0.031)	0.069** (0.033)		
..... Weakly identifies with a major political party	0.418	0.073* (0.043)	0.091** (0.040)		
..... Strongly identifies with a major political party	0.421	-0.018 (0.040)	-0.022 (0.039)		
N		[51,527]	[56,659]		
Panel c: Involved political engagement					
Verified congressional primary turnout	0.347	0.026 (0.042) [47,604]	0.014 (0.041) [52,030]		
Donated money to a political campaign or organization	0.241	0.011 (0.038) [43,637]	0.016 (0.033) [47,405]		
Put up a political sign	0.198	0.033 (0.038) [43,637]	0.045 (0.035) [47,405]		
Attended a political meeting	0.134	-0.041 (0.031) [43,637]	-0.006 (0.028) [47,405]		
Worked for a candidate or campaign	0.061	-0.020 (0.021) [43,637]	-0.008 (0.018) [47,405]		
Reports having run for office	0.028	0.012 (0.013) [43,491]	0.014 (0.013) [47,267]		
Controls: Gender and race		Yes	Yes		Yes
FE: County		.	Yes		Yes
FE: State x cohort		.	Yes		Yes
FE: County x cong. district x year		Yes	No		No
FE: State x cohort x year		Yes	No		No

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing outcome and control variables are omitted. Most dependent variables are indicator variables. Self-report of interest in and public affairs is an index variable ranging from 1 (Hardly at all) to 4 (Most of the time). Recognized names of elected representatives takes on values from 0 to 3 indicating whether the respondent recognizes the name of their current U.S. House Representative, and both U.S. Senators. All estimates using CCES data employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table 3: Younger cohorts in high *Sesame Street* coverage counties report more votes for minority, women, and Democratic candidates

	(1)	(2)	(3)	(4)	*p-value of difference*	(5)
	Share	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$			
Panel a: Reported voting on minority-white ballots				**Respondent is**		
				White		**Minority**
Minority	0.409	0.271***	0.195***	0.312***	—0.790—	0.383
		(0.082)	(0.074)	(0.095)		(0.260)
White	0.398	-0.144*	-0.083	-0.266***	—0.289—	0.022
		(0.085)	(0.070)	(0.092)		(0.268)
Third party	0.028	0.004	-0.006	-0.009	—0.918—	-0.016
		(0.021)	(0.017)	(0.028)		(0.063)
Not voting	0.165	-0.132*	-0.106*	-0.037	—0.050—	-0.389**
		(0.069)	(0.063)	(0.077)		(0.169)
N		[11,811]	[12,819]	[8,055]		[2,469]
Panel b: Reported voting on woman-man ballots						
				Female		**Male**
Woman	0.396	0.192***	0.192***	0.214**	—0.794—	0.170
		(0.058)	(0.050)	(0.106)		(0.129)
Man	0.417	-0.031	-0.004	-0.164*	—0.127—	0.055
		(0.062)	(0.059)	(0.094)		(0.109)
Third party	0.025	-0.041**	-0.056***	-0.032	—0.874—	-0.039
		(0.017)	(0.020)	(0.021)		(0.037)
Not voting	0.162	-0.121***	-0.132***	-0.017	—0.164—	-0.186*
		(0.046)	(0.047)	(0.070)		(0.100)
N		[19,034]	[20,602]	[8,952]		[7,565]
Panel c: Reported voting by candidate party						
Democrat	0.398	0.100***	0.065*			
		(0.037)	(0.036)			
Republican	0.407	0.033	0.059*			
		(0.037)	(0.036)			
Third party	0.027	-0.021*	-0.021*			
		(0.011)	(0.011)			
Not voting	0.168	-0.113***	-0.103***			
		(0.033)	(0.034)			
N		[51,723]	[56,850]			
Controls: Gender and race		Yes	Yes	Yes		Yes
FE: County		.	Yes	.		.
FE: State x cohort		.	Yes	.		.
FE: County x cong. dist. x year		Yes	No	Yes		Yes
FE: State x cohort x year		Yes	No	Yes		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Unless otherwise specified, estimates of $\hat{\beta}_1$ are reported as indicated by the listed fixed effects. The sample is limited to respondents voting in U.S. House elections that feature a Democratic candidate and a Republican candidate (all panels); one of whom is a minority and the other is white (panel a); or one of whom is a man and the other is a woman (panel b). Each observation in the sample has one of the mutually exclusive voting behaviors listed set to 1 and all others set to 0. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Values between columns 4-5 give the p-value on the interaction term of the equivalent fully interacted specification testing for heterogeneity between the two groups. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * $p<0.1$, ** $p<0.05$ and *** $p<0.01$.

Table 4: No evidence that younger cohorts in high *Sesame Street* coverage counties differ in their party preferences on ballots featuring two white men

	(1)	(2)	(3)	(4)	(5)	(6)
	Both candidates are white men		**Democrat is a woman**		**Democrat is a minority**	
	All	**Major party voters**	**All**	**Major party voters**	**All**	**Major party voters**
Democrat	0.014	-0.011	0.183***	0.156**	0.235***	0.217**
	(0.059)	(0.073)	(0.060)	(0.068)	(0.089)	(0.094)
Republican	0.087		-0.026		-0.099	
	(0.064)		(0.062)		(0.090)	
Third party	-0.008		-0.022		-0.004	
	(0.018)		(0.016)		(0.022)	
Not voting	-0.093*		-0.135***		-0.132*	
	(0.056)		(0.049)		(0.075)	
N	[22,143]	[17,071]	[16,782]	[13,307]	[8,946]	[6,916]
Controls: Gender and race	Yes	Yes	Yes	Yes	Yes	Yes
FE: County x cong. dist. x year	Yes	Yes	Yes	Yes	Yes	Yes
FE: State x cohort x year	Yes	Yes	Yes	Yes	Yes	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents and are estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. The sample is limited to respondents voting in U.S. House elections that feature a Democratic candidate and a Republican candidate. Each observation in the sample has one of the mutually exclusive voting behaviors listed set to 1 and all others set to 0. Estimates in even columns limit the sample to respondents who report voting for a major party candidate. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table 5: Younger cohorts in high *Sesame Street* coverage counties report more votes for minority candidates of both parties, and women Democrats

	(1)	*p-value of difference*	(2)	(3)	*p-value of difference*	(4)
Panel a: Reported voting when minority candidates are						
	Minority candidate is			**Minority candidate is**		
	Democrat		**Republican**	**Woman**		**Man**
Minority	0.235*** (0.089)	—0.161—	0.569** (0.232)	0.514*** (0.191)	—0.138—	0.190* (0.106)
White	-0.099 (0.090)	—0.364—	-0.317 (0.232)	-0.138 (0.179)	—0.916—	-0.160 (0.114)
Third party	-0.004 (0.022)	—0.614—	-0.028 (0.043)	-0.072* (0.037)	—0.007—	0.047* (0.024)
Not voting	-0.132* (0.075)	—0.598—	-0.225 (0.168)	-0.304*** (0.109)	—0.100—	-0.077 (0.086)
N	[8,946]		[2,655]	[4,282]		[7,285]
Panel b: Reported voting when women candidates are						
	Woman candidate is			**Woman candidate is**		
	Democrat		**Republican**	**White**		**Minority**
Woman	0.228*** (0.065)	—0.221—	0.036 (0.149)	0.138** (0.066)	—0.056—	0.467*** (0.162)
Man	-0.042 (0.072)	—0.580—	0.054 (0.167)	0.012 (0.072)	—0.247—	-0.189 (0.162)
Third party	-0.021 (0.017)	—0.369—	-0.066 (0.049)	-0.033* (0.018)	—0.590—	-0.060 (0.049)
Not voting	-0.164*** (0.052)	—0.233—	-0.025 (0.110)	-0.117** (0.055)	—0.428—	-0.217* (0.117)
N	[14,169]		[4,473]	[14,340]		[4,371]
Controls: Gender and race	Yes		Yes	Yes		Yes
FE: County x cong. dist. x year	Yes		Yes	Yes		Yes
FE: State x cohort x year	Yes		Yes	Yes		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents and are estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. The sample is limited to respondents voting in U.S. House elections that feature a minority and white candidate (panel a); or a man and woman candidate (panel b). Each observation in the sample has one of the mutually exclusive voting behaviors listed set to 1 and all others set to 0. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Values between columns 1-2 and 3-4 give the p-value on the interaction term of the equivalent fully interacted specification testing for heterogeneity between the two groups. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table 6: Younger cohorts in high *Sesame Street* coverage counties report more political and party identification

	(1)	(2)	(3)	(4)	(5)	(6)
	Full CESS			**Full IAT**		
	Share	$\hat{\beta}_1$	**CCES if lib. or cons. dem. or rep.**	**Share**	$\hat{\beta}_1^{KL}$	**IAT if lib. or cons.**
Panel a: Impacts on political ideology						
Liberal	0.239	0.012 (0.033)	0.084 (0.056)	0.516	0.025** (0.012)	-0.006 (0.013)
Moderate/Neutral	0.324	-0.030 (0.039)		0.196	-0.037*** (0.009)	
Conservative	0.376	0.012 (0.040)		0.288	0.012 (0.011)	
Not sure	0.061	0.006 (0.024)				
N		[51,575]	[29,549]		[319,563]	[256,923]
Panel b: Impacts on party identity						
Democrat	0.429	0.071* (0.038)	0.057 (0.046)			
Independent	0.139	-0.058** (0.028)				
Republican	0.410	-0.016 (0.044)				
Not Sure	0.022	0.003 (0.013)				
N		[51,527]	[42,430]			
Controls: Gender and race		Yes	Yes		Yes	Yes
FE: County		.	.		Yes	Yes
FE: State x cohort		.	.		Yes	Yes
FE: County x cong. district x year		Yes	Yes		.	.
FE: State x cohort x year		Yes	Yes		.	.

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Each observation in the sample has one of the mutually exclusive political identities listed set to 1 and all others set to 0. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Estimates using the CCES data employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table 7: Younger white test takers in high *Sesame Street* coverage counties have lower measures of implicit bias against Blacks

	(1)	(2)	(3)	(4)
Dependent variable and sample	**Mean score**	$\hat{\beta}_1^{KL}$		
Panel a: Impacts on the race IAT				
Race IAT score of white test takers	0.145	-0.067***	-0.067***	-0.051*
		(0.025)	(0.025)	(0.028)
		[261,189]	[261,189]	[252,098]
Race IAT score of Black test takers	-0.800	0.065	0.052	0.097
		(0.079)	(0.078)	(0.098)
		[43,397]	[43,397]	[38,601]
Race IAT score of Hispanic/other/unreported test takers	-0.063	0.041	0.043	-0.104
		(0.065)	(0.064)	(0.074)
		[44,356]	[44,356]	[38,288]
Panel b: Impacts on the gender-career IAT				
Gender-career IAT score of all test takers	0.107	0.012	0.016	0.046
		(0.047)	(0.047)	(0.060)
		[83,687]	[83,687]	[76,329]
Gender-career IAT score of female test takers	0.210	-0.023	-0.012	0.022
		(0.059)	(0.059)	(0.077)
		[54,826]	[54,826]	[47,638]
Gender-career IAT score of male test takers	-0.089	0.091	0.092	-0.023
		(0.096)	(0.096)	(0.142)
		[28,280]	[28,280]	[23,152]
Controls: Gender and race		Yes	Yes	Yes
FE: County		Yes	Yes	.
FE: State x cohort		Yes	Yes	.
FE: County x year		No	No	Yes
FE: State x cohort x year		No	No	Yes
Controls: Education level		No	Yes	Yes

Note: Each coefficient is the result of a separate regression using the indicated controls and fixed effects. In panel a, standardized race IAT scores are the dependent variable. Larger positive values indicate stronger Black-bad and white-good associations. In panel b, standardized gender-career IAT scores are the dependent variable. Larger positive values indicate stronger man-career and woman-family associations. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table 8: Younger cohorts in high *Sesame Street* coverage counties exhibit muted differences in explicit racial preferences reported on the race IAT survey

	(1)	(2)	(3)	(4)
	White and Black test takers		White test takers	Black test takers
Dependent variable and sub sample	**Mean**	$\hat{\beta}_1^{KL}$	$\hat{\beta}_1^{KL}$	$\hat{\beta}_1^{KL}$
Reported preference for in-group over out-group Americans (1 to 7) (1-Strongly prefers out-group, 4-likes equally, 7-strongly prefers in-group)	4.6	-0.028 (0.024)	-0.016 (0.025)	-0.175* (0.105)
		[288,344]	[247,539]	[40,376]
..........Reported preference for non-equal respondents	5.3	-0.045 (0.039)	-0.031 (0.042)	-0.130 (0.136)
		[123,781]	[97,944]	[25,443]
Difference in warmth towards in-group and out-group Americans (-10 to 10) (-10-Strongly prefers out-group, 0-likes equally, 10-strongly prefers in-group)	0.584	-0.061 (0.040)	-0.056 (0.041)	-0.090 (0.178)
		[297,737]	[255,188]	[42,109]
..........Difference in warmth for non-zero respondents	1.5	-0.231** (0.095)	-0.192* (0.112)	-0.378* (0.200)
N		[112,997]	[85,991]	[26,614]
Controls: Gender and race		Yes	Yes	Yes
FE: County		Yes	Yes	Yes
FE: State x cohort		Yes	Yes	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Difference in warmth is the difference in the warmth reported for the respondent's in-group and respondent's out-group on 10 point scales. The distribution of these indicators in plotted in appendix figure A2. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table 9: Non-incumbent minority candidates gain more from electorate exposure to *Sesame Street* coverage

	(1)	*p-value of difference*	(2)	*p-value of difference*	(3)
	Ballots with no incumbents		**Minority candidate is incumbent**		**Minority candidate is non-incumbent**
Minority	0.313 (0.258)	—0.400—	0.070 (0.140)	—0.111—	0.363*** (0.120)
White	-0.318 (0.233)	—0.216—	0.015 (0.141)	—0.597—	-0.085 (0.125)
Third party	0.126* (0.073)	—0.119—	0.005 (0.030)	—0.051—	-0.077** (0.030)
Not voting	-0.121 (0.185)	—0.886—	-0.090 (0.126)	—0.507—	-0.200* (0.109)
N	[1,844]		[4,270]		[5,114]
Controls: Gender and race	Yes		Yes		Yes
FE: County x cong. district x year	Yes		Yes		Yes
FE: State x cohort x year	Yes		Yes		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 with controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. The sample is limited to respondents voting in U.S. house white-minority ballots. Each observation in the sample has one of the mutually exclusive voting behaviors listed set to 1 and all others set to 0. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Values between columns 1-2 and 2-3 give the p-value on the interaction term of the equivalent fully interacted specification testing for heterogeneity between the two groups. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

References

Adena, M., Enikolopov, R., Petrova, M., Santarosa, V., and Zhuravskaya, E. (2015). Radio and the rise of the Nazis in prewar Germany. The Quarterly Journal of Economics, 130(4):1885–1939.

Adukia, A., Eble, A., Harrison, E., Runesha, H. B., and Szasz, T. (2023). What we teach about race and gender: Representation in images and text of children's books. The Quarterly Journal of Economics, 138(4):2225–2285.

Allport, G. W., Clark, K., and Pettigrew, T. (1954). The nature of prejudice.

Ang, D. (2023). The birth of a nation: Media and racial hate. American Economic Review, 113(6):1424–1460.

Ansolabehere, S. and Hersh, E. (2012). Validation: What big data reveal about survey misreporting and the real electorate. Political Analysis, 20(4):437–459.

Armand, A., Atwell, P., Gomes, J. F., and Schenk, Y. (n.d.). It's a bird, it's a plane, it's Superman! Using mass media to fight intolerance. Technical report.

Bankart, C. P. and Anderson, C. C. (1979). Short-term effects of prosocial television viewing on play of preschool boys and girls. Psychological Reports, 44(3):935–941.

Beaman, L., Chattopadhyay, R., Duflo, E., Pande, R., and Topalova, P. (2009). Powerful women: Does exposure reduce bias? The Quarterly Journal of Economics, 124(4):1497–1540.

Beelmann, A. and Heinemann, K. S. (2014). Preventing prejudice and improving intergroup attitudes: A meta-analysis of child and adolescent training programs. Journal of Applied Developmental Psychology, 35(1):10–24.

Bertrand, M. and Duflo, E. (2017). Field experiments on discrimination. Handbook of economic field experiments, 1:309–393.

Bettinger, E. P. and Long, B. T. (2005). Do faculty serve as role models? the impact of instructor gender on female students. American Economic Review, 95(2):152–157.

Billings, S. B., Chyn, E., and Haggag, K. (2021). The long-run effects of school racial diversity on political identity. American Economic Review: Insights, 3(3):267–84.

Blouin, A. and Mukand, S. W. (2019). Erasing ethnicity? Propaganda, nation building, and identity in Rwanda. Journal of Political Economy, 127(3):1008–1062.

Boisjoly, J., Duncan, G. J., Kremer, M., Levy, D. M., and Eccles, J. (2006). Empathy or antipathy? The impact of diversity. American Economic Review, 96(5):1890–1905.

Brown, J. R., Enos, R. D., Feigenbaum, J., and Mazumder, S. (2021). Childhood cross-ethnic exposure predicts political behavior seven decades later: Evidence from linked administrative data. Science Advances, 7(24):eabe8432.

Campante, F., Durante, R., and Tesei, A. (2022). Media and social capital. Annual Review of Economics, 14(1):69–91.

Carlana, M. (2019). Implicit stereotypes: Evidence from teachers' gender bias. The Quarterly Journal of Economics, 134(3):1163–1224.

Carrell, S. E., Hoekstra, M., and West, J. E. (2019). The impact of college diversity on behavior toward minorities. American Economic Journal: Economic Policy, 11(4):159–82.

Clubb, J. M., Flanigan, W. H., and Zingale, N. H. (1987). Electoral data for counties in the United States: Presidential and congressional races, 1840-1972. Inter-university Consortium for Political and Social Research.

Cohodes, S. and Feigenbaum, J. J. (2021). Why does education increase voting? evidence from boston's charter schools. Technical report, National Bureau of Economic Research.

Cole, C., Arafat, C., Tidhar, C., Tafesh, W. Z., Fox, N., Killen, M., Ardila-Rey, A., Leavitt, L., Lesser, G., Richman, B., et al. (2003). The educational impact of Rechov Sumsum/Shara'a Simsim: A Sesame Street television series to promote respect and understanding among children living in Israel, the West Bank, and Gaza. International Journal of Behavioral Development, 27(5):409–422.

Connor, P. and Evers, E. R. (2020). The bias of individuals (in crowds): Why implicit bias is probably a noisily measured individual-level construct. Perspectives on Psychological Science, 15(6):1329–1345.

Cooney, J. G. (1967). The potential uses of television in preschool education.

Corno, L., La Ferrara, E., and Burns, J. (2022). Interaction, stereotypes, and performance: Evidence from South Africa. American Economic Review, 112(12):3848–75.

Dee, T. S. (2005). A teacher like me: Does race, ethnicity, or gender matter? American Economic Review, 95(2):158–165.

DellaVigna, S. and Kaplan, E. (2007). The Fox News effect: Media bias and voting. The Quarterly Journal of Economics, 122(3):1187–1234.

Durante, R., Pinotti, P., and Tesei, A. (2019). The political legacy of entertainment TV. American Economic Review, 109(7):2497–2530.

Eble, A. and Hu, F. (2020). Child beliefs, societal beliefs, and teacher-student identity match.

Economics of Education Review, 77:101994.

Ellingsen, S. and Hernæs, Ø. (2018). The impact of commercial television on turnout and public policy: Evidence from norwegian local politics. Journal of Public Economics, 159:1–15.

Enamorado, T. and Imai, K. (2019). Validating self-reported turnout by linking public opinion surveys with administrative records. Public Opinion Quarterly, 83(4):723–748.

Falck, O., Gold, R., and Heblich, S. (2014). E-lections: Voting behavior and the internet. American Economic Review, 104(7):2238–65.

Finseraas, H., Hanson, T., Johnsen, Å. A., Kotsadam, A., and Torsvik, G. (2019). Trust, ethnic diversity, and personal contact: A field experiment. Journal of Public Economics, 173:72–84.

Fiorini, M. and Keane, M. P. (2014). How the allocation of children's time affects cognitive and noncognitive development. Journal of Labor Economics, 32(4):787–836.

Fisch, S. M. and Truglio, R. T. (2014). G is for growing: Thirty years of research on children and Sesame Street.

Fisch, S. M., Truglio, R. T., and Cole, C. F. (1999). The impact of Sesame Street on preschool children: A review and synthesis of 30 years' research. Media Psychology, 1(2):165–190.

Fraga, B. L. and Hassell, H. J. (2021). Are minority and women candidates penalized by party politics? Race, gender, and access to party support. Political Research Quarterly, 74(3):540–555.

Gawronski, B. (2019). Six lessons for a cogent science of implicit bias and its criticism. Perspectives on Psychological Science, 14(4):574–595.

Gawronski, B., Galdi, S., and Arcuri, L. (2015). What can political psychology learn from implicit measures? Empirical evidence and new directions. Political Psychology, 36(1):1–17.

Gentzkow, M. (2006). Television and voter turnout. The Quarterly Journal of Economics, 121(3):931–972.

Gentzkow, M. and Shapiro, J. M. (2008). Preschool television viewing and adolescent test scores: Historical evidence from the Coleman study. The Quarterly Journal of Economics, 123(1):279–323.

Glover, D., Pallais, A., and Pariente, W. (2017). Discrimination as a self-fulfilling prophecy: Evidence from french grocery stores. The Quarterly Journal of Economics, 132(3):1219–1260.

Green, D. P. and Gerber, A. S. (2019). Get out the vote: How to increase voter turnout. Brookings Institution Press.

Greene, B. (2019). The unmistakable black roots of 'Sesame Street'.

Greenwald, A. G., Poehlman, T. A., Uhlmann, E. L., and Banaji, M. R. (2009). Understanding and using the Implicit Association Test: III. Meta-analysis of predictive validity. Journal of Personality and Social Psychology, 97(1):17.

Grimmer, J., Hersh, E., Meredith, M., Mummolo, J., and Nall, C. (2018). Obstacles to estimating voter id laws' effect on turnout. The Journal of Politics, 80(3):1045–1051.

Harrington, A. (2019). Psychiatry, racism, and the birth of 'Sesame Street'.

Haydon, M. (1973). Draft report on audience data. Commissioned by CTW to study audience for Sesame Street and Electric Company. Technical report, Box 44, Folder 32. Children's Television Workshop Collection. Special Collections in Mass Media and Culture. University of Maryland Libraries.

Hehman, E., Calanchini, J., Flake, J. K., and Leitner, J. B. (2019). Establishing construct validity evidence for regional measures of explicit and implicit racial bias. Journal of Experimental Psychology: General, 148(6):1022.

Hernæs, Ø., Markussen, S., and Røed, K. (2019). Television, cognitive ability, and high school completion. Journal of Human Resources, 54(2):371–400.

Heron-Delaney, M., Anzures, G., Herbert, J. S., Quinn, P. C., Slater, A. M., Tanaka, J. W., Lee, K., and Pascalis, O. (2011). Perceptual training prevents the emergence of the other race effect during infancy. PloS one, 6(5):e19858.

Hughes, D., Cox, B., and Das, S. (2023). Growing up, learning race: An integration of research on cognitive mechanisms and socialization in context. Annual Review of Developmental Psychology, 5(1):137–167.

Jensen, R. and Oster, E. (2009). The power of TV: Cable television and women's status in India. The Quarterly Journal of Economics, 124(3):1057–1094.

Jost, J. T. (2019). The IAT is dead, long live the IAT: Context-sensitive measures of implicit attitudes are indispensable to social and political psychology. Current Directions in Psychological Science, 28(1):10–19.

Kearney, M. S. and Levine, P. B. (2019). Early childhood education by television: Lessons from Sesame Street. American Economic Journal: Applied Economics, 11(1):318–50.

Kearney, M. S. and Levine, P. B. (2020). Role models, mentors, and media influences. The Future of Children, 30(1):83–106.

Kelly, D. J., Quinn, P. C., Slater, A. M., Lee, K., Ge, L., and Pascalis, O. (2007). The other-race effect develops during infancy: Evidence of perceptual narrowing. Psychological Science, 18(12):1084–1089.

Kelly, D. J., Quinn, P. C., Slater, A. M., Lee, K., Gibson, A., Smith, M., Ge, L., and Pascalis, O. (2005). Three-month-olds, but not newborns, prefer own-race faces. Developmental Science, 8(6):F31–F36.

Kuriwaki, S. (2021). Cumulative CCES Common Content (2006-2020). Harvard Dataverse, 4:2018.

Lang, K. and Spitzer, A. K.-L. (2020). Race discrimination: An economic perspective. Journal of Economic Perspectives, 34(2):68–89.

Leifer, A. D. (1975). How to encourage socially-valued behavior.

Lesser, G. S. (1974). Children and television: Lessons from Sesame Street.

Logan, J. R. (2017). The American Communities Project: Dataset of United States schools desegregation court cases and school demographic data. Brown University.

Long, L. M. (1973). Sesame Street: A space age approach to education for space age children. University of Massachusetts Amherst.

Lowe, M. (2021). Types of contact: A field experiment on collaborative and adversarial caste integration. American Economic Review, 111(6):1807–1844.

Lowes, S., Nunn, N., Robinson, J. A., and Weigel, J. (2015). Understanding ethnic identity in Africa: Evidence from the implicit association test (IAT). American Economic Review, 105(5):340–345.

Marble, W., Mousa, S., Siegel, A. A., et al. (2021). Can exposure to celebrities reduce prejudice? the effect of mohamed salah on islamophobic behaviors and attitudes. American Political Science Review, 115(4):1111–1128.

Mares, M.-L. and Pan, Z. (2013). Effects of Sesame Street: A meta-analysis of children's learning in 15 countries. Journal of Applied Developmental Psychology, 34(3):140–151.

Merlino, L. P., Steinhardt, M. F., and Wren-Lewis, L. (2022). The long run impact of childhood interracial contact on residential segregation. Technical report, IZA Discussion Papers.

MIT Election Data and Science Lab (2021). U.S. House 1976–2020.

Mousa, S. (2020). Building social cohesion between christians and muslims through soccer in post-ISIS Iraq. Science, 369(6505):866–870.

Murphy, R. T. (1991). Educational effectiveness of Sesame Street: A review of the first twenty years of research 1969–1989. ETS Research Report Series, 1991(2):i–30.

Müller, K. and Schwarz., C. (2023). From hashtag to hate crime: Twitter and anti-minority sentiment. American Economic Journal: Applied Economics, 15(3):270–312.

Oberholzer-Gee, F. and Waldfogel, J. (2009). Media markets and localism: Does local news en espanol boost hispanic voter turnout? American Economic Review, 99(5):2120–2128.

Paluck, E. L. (2009). Reducing intergroup prejudice and conflict using the media: A field experiment in Rwanda. Journal of Personality and Social Psychology, 96(3):574.

Paluck, E. L., Green, S. A., and Green, D. P. (2019). The contact hypothesis re-evaluated. Behavioural Public Policy, 3(2):129–158.

Paluck, E. L., Porat, R., Clark, C. S., and Green, D. P. (2021). Prejudice reduction: Progress and challenges. Annual Review of Psychology, 72:533–560.

Paulson, F. L. (1974). Teaching cooperation on television. AV Communication Review, 22(3):229–246.

Plant, E. A., Devine, P. G., Cox, W. T., Columb, C., Miller, S. L., Goplen, J., and Peruche, B. M. (2009). The obama effect: Decreasing implicit prejudice and stereotyping. Journal of Experimental Social Psychology, 45(4):961–964.

Raabe, T. and Beelmann, A. (2011). Development of ethnic, racial, and national prejudice in childhood and adolescence: A multinational meta-analysis of age differences. Child Development, 82(6):1715–1737.

Rao, G. (2019). Familiarity does not breed contempt: Generosity, discrimination and diversity in Delhi schools. American Economic Review, 109(3):774–809.

Ratliff, K. A. and Smith, C. T. (2021). Lessons from two decades of Project Implicit. In The Cambridge handbook of implicit bias and racism, page 4. Cambridge University Press.

Riley, E. (2019). Role models in movies: The impact of Queen of Katwe on students' educational attainment. The Review of Economics and Statistics, pages 1–48.

Rutland, A. and Killen, M. (2015). A developmental science approach to reducing prejudice and social exclusion: Intergroup processes, social-cognitive development, and moral reasoning. Social Issues and Policy Review, 9(1):121–154.

Schiappa, E., Gregg, P. B., and Hewes, D. E. (2005). The parasocial contact hypothesis. Communication Monographs, 72(1):92–115.

Schimmack, U. (2021). The Implicit Association Test: A method in search of a construct. Perspectives on Psychological Science, 16(2):396–414.

Smets, K. and Van Ham, C. (2013). The embarrassment of riches? a meta-analysis of individual-level research on voter turnout. Electoral Studies, 32(2):344–359.

Sondheimer, R. M. and Green, D. P. (2010). Using experiments to estimate the effects of education on voter turnout. American Journal of Political Science, 54(1):174–189.

Sørensen, R. J. (2019). The impact of state television on voter turnout. British Journal of Political

Science, 49(1):257–278.

Vittrup, B. and Holden, G. W. (2011). Exploring the impact of educational television and parent–child discussions on children's racial attitudes. Analyses of Social Issues and Public Policy, 11(1):82–104.

Wang, T. (2021). Media, pulpit, and populist persuasion: evidence from father Coughlin. American Economic Review, 111(9):3064–92.

Yanagizawa-Drott, D. (2014). Propaganda and conflict: Evidence from the Rwandan genocide. The Quarterly Journal of Economics, 129(4):1947–1994.

Zielinska, I. E. and Chambers, B. (1995). Using group viewing of television to teach preschool children social skills. Journal of Educational Television, 21(2):85–99.

Appendix

Table A1: Younger cohorts in high *Sesame Street* coverage counties do not differ in their post-election survey response rates

	(1)	(2)	(3)
Dependent variable	**Dependent variable mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$
Took the post-election survey	0.888	-0.004 (0.024) [59,251]	-0.011 (0.024) [64,335]
Controls: Gender and race		No	No
FE: County		.	Yes
FE: State x cohort		.	Yes
FE: County x cong. district x year		Yes	No
FE: State x cohort x year		Yes	No

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Educational attainment (in years) and reported family income are continuous variables built from binned response options (6 and 12 bins respectively). Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A2: Demographic composition of candidates on major party ballots

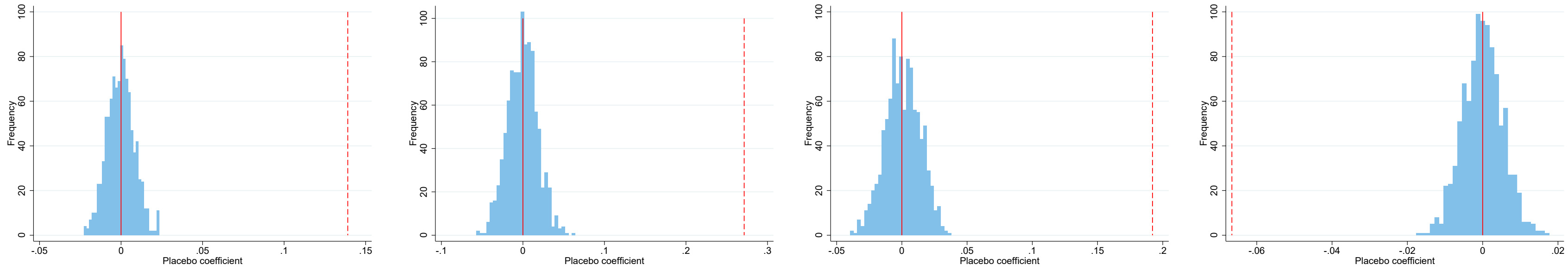
Race		Gender						Election outcomes				
								Democratic margin (ppts)			Share	
Democrat	Republican	Democrat	Republican	Respondents	Share	Ballots	Share	Mean	Median	Std. Dev.	Close (±10)	Very close (±5)
White	White	Man	Man	25,304	0.442	1,333	0.451	-6.10	-12.35	30.05	0.152	0.075
		Man	Woman	3,642	0.064	179	0.061	-0.83	1.34	28.47	0.218	0.117
		Woman	Man	11,101	0.194	516	0.175	-5.08	-11.18	27.21	0.211	0.105
		Woman	Woman	1,628	0.028	73	0.025	2.64	-0.48	27.47	0.260	0.151
		Total		41,675	0.728	2,101	0.711	-5.10	-10.97	29.21	0.176	0.089
Minority	White	Man	Man	5,552	0.097	303	0.102	9.88	10.09	36.71	0.125	0.063
		Man	Woman	611	0.011	31	0.010	18.69	20.68	34.24	0.129	0.000
		Woman	Man	3,224	0.056	166	0.056	11.82	9.57	38.48	0.084	0.036
		Woman	Woman	641	0.011	26	0.009	15.79	9.62	34.13	0.269	0.192
		Total		10,028	0.175	526	0.178	11.30	11.21	36.99	0.120	0.057
White	Minority	Man	Man	1,420	0.025	79	0.027	18.12	22.42	28.60	0.127	0.076
		Man	Woman	594	0.010	32	0.011	21.48	31.53	29.15	0.250	0.094
		Woman	Man	769	0.013	39	0.013	14.06	22.66	32.64	0.128	0.103
		Woman	Woman	309	0.005	13	0.004	19.68	19.05	18.00	0.154	0.000
		Total		3,092	0.054	163	0.055	17.93	22.42	28.93	0.153	0.080
Minority	Minority	Man	Man	1,084	0.019	81	0.027	36.39	42.40	31.33	0.111	0.074
		Man	Woman	407	0.007	26	0.009	39.79	38.28	23.39	0.115	0.115
		Woman	Man	661	0.012	45	0.015	39.64	45.38	28.70	0.089	0.067
		Woman	Woman	274	0.005	15	0.005	38.60	44.06	21.65	0.000	0.000
		Total		2,426	0.042	167	0.056	37.99	42.83	28.56	0.096	0.072
All		Man	Man	33,360	0.583	1,796	0.607	-0.42	-6.07	33.08	0.144	0.073
		Man	Woman	5,254	0.092	268	0.091	8.03	8.17	31.75	0.201	0.101
		Woman	Man	15,755	0.275	766	0.259	2.19	-6.34	32.61	0.172	0.087
		Woman	Woman	2,852	0.050	127	0.043	11.32	12.39	29.81	0.220	0.126
		Total		57,221	1.000	2,957	1.000	1.52	-3.45	32.85	0.160	0.082

Note: This table presents the composition of the 2006-2020 U.S. House ballots in our sample, organized by the demographics of their major party candidates. *Major party* ballots in our sample are defined as ballots where the two front-runners are a Democrat and a Republican, both of whom receive over 5% of their district's vote, and where candidate demographics and *Sesame Street* coverage is observed. Mean election outcomes for ballots of each type are also reported.

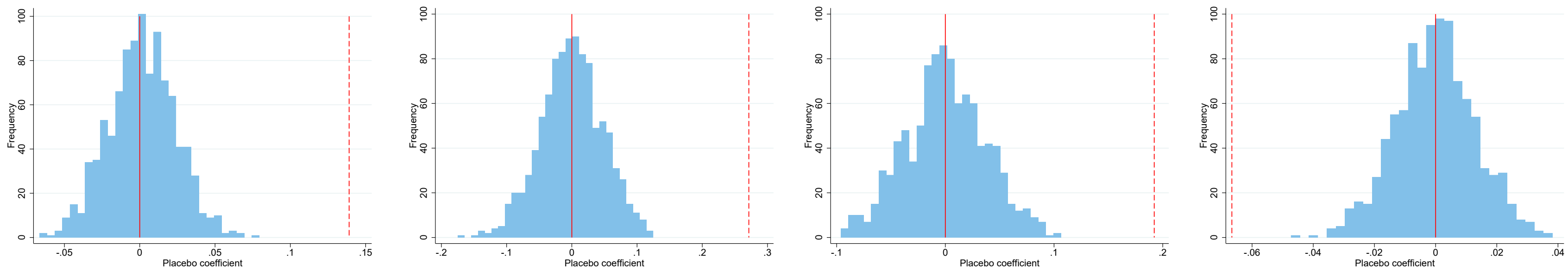
Table A3: Main results are robust to alternative specifications

	(1)	(2)	(3)	(4)
	Turnout	**Reports vote for a minority candidate**	**Reports vote for a woman candidate**	**White race IAT score**
Main results	0.139***	0.271***	0.192***	-0.067***
	(0.040)	(0.082)	(0.058)	(0.025)
	[51,723]	[11,811]	[19,034]	[261,189]
Panel a: Main specification with added controls				
Main specifications with the addition of controls for ...				
Respondent's education	0.131***	0.279***	0.181***	-0.067***
	(0.040)	(0.080)	(0.057)	(0.025)
	[51,713]	[11,809]	[19,031]	[261,189]
Respondent's family income	0.148***	0.278***	0.185***	.
	(0.041)	(0.082)	(0.064)	
	[46,596]	[10,574]	[17,151]	
County Black population share in 1970 $\times$ Cohort indicators	0.148***	0.262***	0.207***	-0.067***
	(0.040)	(0.084)	(0.062)	(0.025)
	[51,721]	[11,811]	[19,034]	[260,003]
County low income share in 1970 $\times$ Cohort indicators	0.114***	0.255***	0.177***	-0.067***
	(0.042)	(0.083)	(0.060)	(0.025)
	[51,721]	[11,811]	[19,034]	[260,003]
County Thurmond vote share in 1948 $\times$ Cohort indicators	0.138***	0.263***	0.194***	-0.070***
	(0.040)	(0.082)	(0.058)	(0.025)
	[51,239]	[11,628]	[18,846]	[257,687]
County Democratic vote share in 1968 $\times$ Cohort indicators	0.153***	0.279***	0.201***	-0.068***
	(0.040)	(0.082)	(0.060)	(0.025)
	[51,669]	[11,811]	[19,005]	[259,916]
School segregation $\times$ Cohort indicators	0.144***	0.260***	0.192***	-0.073***
	(0.041)	(0.084)	(0.060)	(0.026)
	[48,908]	[11,501]	[17,800]	[240,842]
Controls: Gender and race	Yes	Yes	Yes	Gender
FE: County x cong. district x year	Yes	Yes	Yes	No
FE: State x cohort x year	Yes	Yes	Yes	No
FE: County	.	.	.	Yes
FE: State x cohort	.	.	.	Yes
Panel b: Coefficients from a simple difference-in-differences specification with no fixed effects				
$preschool69_i \times SSCov_j$	0.103***	0.130**	0.125***	-0.057***
	(0.034)	(0.065)	(0.046)	(0.019)
$SSCov_j$	-0.049*	0.029	0.037	-0.070**
	(0.028)	(0.051)	(0.040)	(0.032)
$preschool69_i$	-0.087***	-0.085*	-0.076**	0.004
	(0.022)	(0.043)	(0.031)	(0.014)
N	[57,191]	[13,113]	[21,005]	[260,080]
Controls: Gender and race	Yes	Yes	Yes	Gender

Note: In panel a, each coefficient is the result of a separate regression. In columns 1, 2, and 3, CCES data is used to estimate $\hat{\beta}_1$ from equation 1 with the addition of the indicated controls in each row. These specifications also control for respondents' race and gender, *(county × congressional district × year)* and *(state × cohort × year)* fixed effects. Column 4 estimates $\hat{\beta}_1^{KL}$ on the race IAT data with the addition of the indicated controls in each row. This specification also controls for respondents' gender and *(county)* and *(state × cohort)* fixed effects. Panel b reports the coefficients from four simple difference-in-differences specifications regressing the main outcomes on $preschool69_i \times SSCov_j$; $SSCov_j$; $preschool69_i$; and controls for race and gender, but no fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Survey weights are used in columns 1-3. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.



(a) Placebo cohort estimates on turnout **(b) Placebo cohort estimates on reported voting for minorities** **(c) Placebo cohort estimates on reported voting for women** **(d) Placebo cohort estimates on white IAT scores**



(e) Placebo coverage estimates on turnout **(f) Placebo coverage estimates on reported for minorities** **(g) Placebo coverage estimates on reported voting for women** **(h) Placebo coverage estimates on white IAT scores**

Figure A1. Plots of 1000 placebo estimates randomizing cohort or predicted county coverage rates on turnout, reported voting for minority and women candidates, and white IAT scores

Notes: The figures plot the distribution of 1000 placebo estimates where either the observation was randomly assigned to treated or control cohorts top row) or the observation was randomly assigned a predicted *Sesame Street* coverage level from the predicted coverage distribution (bottom row). The dashed lines highlight the observed effects corresponding to the true cohort and predicted coverage rates.

Table A4: No evidence that exposure to *Sesame Street* coverage impacted moving between states, counties, cities, or towards counties with more or less *Sesame Street* coverage in 1969

	(1)	(2)	(3)	(4)	(5)
Panel a: Tests for selection in the CCES data					
	Lived in current city since under age 6	**Faces a minority-white ballot**	**Faces a woman-man ballot**	**White**	**Female**
Estimates of $\hat{\beta}_1^{KLadult}$					
$preschool69_i \times SSCov_j^{adult}$	0.046 (0.031)	0.007 (0.024)	0.011 (0.032)	-0.038 (0.037)	0.056 (0.039)
N	[34,909]	[56,850]	[56,850]	[56,850]	[56,850]
Controls: Gender and race	Yes	Yes	Yes	No	No
FE: Adult county	Yes	Yes	Yes	Yes	Yes
FE: Adult state x cohort	Yes	Yes	Yes	Yes	Yes
Coefficients from simplified specification					
$preschool69_i \times SSCov_j^{adult}$	0.037 (0.027)	0.003 (0.027)	-0.011 (0.032)	-0.021 (0.035)	0.032 (0.035)
$SSCov_j^{adult}$	0.005 (0.029)	0.090* (0.047)	0.113** (0.051)	-0.120*** (0.042)	-0.062** (0.025)
$preschool69_i$	-0.019 (0.018)	-0.007 (0.018)	0.002 (0.020)	-0.008 (0.023)	-0.020 (0.024)
N	[35,342]	[57,191]	[57,191]	[57,191]	[57,191]
Controls: Gender and race	Yes	Yes	Yes	No	No
FE: Adult county	No	No	No	No	No
FE: Adult state x cohort	No	No	No	No	No
Dependent variable mean	0.101	0.229	0.367	0.801	0.530
Panel b: Tests for selection in the PSID data					
	Attrits from PSID panel	**Lives in childhood state**	**Lives in childhood county**	**$SSCov_j^{adult}$ of adult county**	
Estimates of $\hat{\beta}_1^{KLchild}$					
$preschool69_i \times SSCov_j^{child}$	-0.024 (0.095)	0.073 (0.382)	-0.082 (0.293)	-0.024 (0.097)	
N	[4,190]	[1,052]	[1,052]	[1,044]	
Controls: Gender and race	Yes	Yes	Yes	Yes	
FE: Childhood county	Yes	Yes	Yes	Yes	
FE: Childhood state x cohort	Yes	Yes	Yes	Yes	
Dependent variable mean	0.721	0.746	0.473	0.630	

Note: Panel a uses CCES data to estimate differences between younger and older cohorts residing in high coverage counties on outcomes that could signal selection effects. $SScov_j^{adult}$ is the coverage rate in respondents' county of residence in the CCES when surveyed as adults. The top portion of panel a presents estimates of $\hat{\beta}_1^{KLadult}$ with $(adult\ county)$ and $(adult\ state \times cohort)$ fixed effects. Race and gender controls are used in columns 1-3. Each coefficient is the result of a separate regression. The bottom portion of panel a reports the coefficients from five simple difference-in-differences specifications regressing these outcomes on $preschool69_i \times SSCov_j^{adult}$; $SSCov_j^{adult}$; $preschool69_i$; and controls for race and gender in columns 1-3; with no fixed effects. Panel b tests for impacts of *Sesame Street* exposure on migration using data from the PSID where we observe both $SScov_j^{child}$ and $SScov_j^{adult}$. Each coefficient is the result of a separate regression estimating $\hat{\beta}_1^{KLchild}$ with $(child\ county)$ and $(child\ state \times cohort)$ fixed effects, and race and gender controls. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing outcome and control variables are omitted. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A5: Coverage rates of sending and destination counties are correlated for inter-county migrants

	(1)	(2)	(3)	(4)
	Destination county coverage rate			
	All observations	**All movers**	**Movers to neighboring counties**	**Movers to non-neighboring counties**
Sending county coverage rate (PSID)	0.547***	0.242***	0.789***	0.101*
	(0.040)	(0.050)	(0.036)	(0.054)
	[1,189]	[622]	[186]	[436]
Sending county coverage rate (Census)		0.347***	0.818***	0.124***
		(0.002)	(0.005)	(0.002)
		[236,880]	[13,964]	[222,916]

Note: For correlations using PSID data in row 1, each observation represents a respondent with observations used to estimate the correlation reported in brackets. PSID regressions are weighted by survey weights. For correlations using census data in row 2, each observation consists of migration flows observed between county pair combinations by the census between 2016 and 2020. Census regressions are weighted by the number of migrants between that sending and destination county. Numbers in brackets report the county pair observations used in each estimation. Standard errors are reported in parenthesis, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A6: Effects are generally larger for respondents who have resided in their city since before age 6

	(1)	(2)	*p-value of difference*	(3)
	$\hat{\beta}_1^{KL}$	**Respondent**		
Dependent indicator variable		**Lived in current city since under age 6**		**Moved to current city at age 6 or later**
Validated turnout	0.119**	0.234	—0.427—	0.102**
	(0.048)	(0.166)		(0.051)
	[34,909]	[3,101]		[31,308]
Active voter registration	0.113**	0.303*	—0.207—	0.089*
	(0.050)	(0.171)		(0.051)
	[30,372]	[2,602]		[27,285]
Reports voting for a minority candidate	0.263**	0.848***	—0.004—	0.216**
	(0.108)	(0.213)		(0.104)
	[7,489]	[574]		[6,674]
Reports voting for a female candidate	0.144**	-0.217	—0.216—	0.148**
	(0.069)	(0.316)		(0.074)
	[12,063]	[823]		[10,871]
Controls: Gender and race	Yes	Yes		Yes
FE: County	Yes	Yes		Yes
FE: State x cohort	Yes	Yes		Yes

Note: Each coefficient is the result of a separate regression. Given the small sub-samples, we estimate $\hat{\beta}_1^{KL}$ with $(county)$ and $(state \times cohort)$ fixed effects. All regressions control for the race and gender of respondents. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Column 1 uses all 35,372 observations where the duration of residence is observed. Column 2 limits the sample to respondents who report living in their city since before age 6. Column 3 limits the sample to respondents who report living in their city since age 6 or later. Values between columns 2-3 give the p-value on the interaction term of the equivalent fully interacted specification testing for heterogeneity between the two groups. All estimates employ survey weights. Standard errors are reported in parenthesis, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A7: No evidence that *Sesame Street* coverage impacted selection into taking the race IAT but it did impact selection into taking the gender-career IAT

	(1)	(2)	(3)
	Share of county's test takers in treated cohorts		
	Race IAT All test takers	**Race IAT White test takers**	**Gender-career IAT All test takers**
Predicted *Sesame Street* coverage rate ($SScov_j$	0.0080 (0.0078)	-0.0018 (0.0092)	0.0340** (0.0151)
Constant	0.55*** (0.01)	0.55*** (0.01)	0.54*** (0.01)
County observations	2,856	2,810	2,372
Total test takers	350,080	261,412	84,479

Note: Each observation represents a county. Coefficients measure the correlation between the the share of test takers in the county coming from treated cohorts with $SScov_j$, the county's predicted coverage rate, weighted by the total number of test takers in the 1959-1968 cohorts. Standard errors are reported in parentheses with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A8: Differences between younger and older cohort in high *Sesame Street* coverage counties in the reported reasons for not voting

	(1)	(2)	(3)
Dependent indicator variable	**Dependent indicator mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$
All reasons	0.151	-0.128***	-0.128***
		(0.033)	(0.035)
		[47,592]	[52,017]
...Did not like the candidates	0.024	-0.025**	-0.028**
		(0.012)	(0.012)
		[47,592]	[52,017]
...I am not registered	0.024	-0.030*	-0.024
		(0.016)	(0.016)
		[47,592]	[52,017]
...I am not interested	0.016	-0.014	-0.011
		(0.012)	(0.012)
		[47,592]	[52,017]
...Sick or disabled	0.017	-0.009	0.006
		(0.013)	(0.014)
		[47,592]	[52,017]
...I did not feel that I knew enough about the choices	0.014	-0.001	-0.019
		(0.011)	(0.016)
		[47,592]	[52,017]
...All other reasons listed	0.057	-0.048**	-0.052**
		(0.020)	(0.022)
		[47,592]	[52,017]
Controls: Gender and race		Yes	Yes
FE: County		.	Yes
FE: State x cohort		.	Yes
FE: County x cong. district x year		Yes	No
FE: State x cohort x year		Yes	No

Note: Each coefficient is the result of a separate regression. Outcome variables are coded as 0 if the respondent reports voting, or not-voting for a different reason, and 1 if the respondent gives the listed reason for non-turnout. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Other reasons for non-turnout include bad weather, not knowing why, lack of identification, lack of knowledge about polling locations, forgetting to vote, fear of covid exposure, non-receipt of absentee ballots, being out of town, long lines at polling stations, dismissal at the polling station, lack of transportation, being too busy, and other reasons. Reason for non-turnout was not asked in the 2006 survey, and is only asked to respondents who report not voting. All estimates employ survey weights. Standard errors are reported in parenthesis, clustered at the county level, with the following significance indicators: * $p<0.1$, ** $p<0.05$ and *** $p<0.01$.

Table A9: No clear heterogeneity in turnout and registration effects by respondent minority status and sex

	(1)	*p-value of difference*	(2)	(3)	*p-value of difference*	(4)
	Respondent is			**Respondent is**		
Dependent indicator variable	**White**		**Minority**	**Female**		**Male**
Panel a: Heterogeneity in turnout effects by respondent minority status and sex						
Verified general election turnout	0.132*** (0.044)	—0.602—	0.064 (0.130)	0.148** (0.061)	—0.912—	0.139** (0.055)
Self-reported election turnout	0.124*** (0.034)	—0.448—	0.043 (0.106)	0.136*** (0.048)	—0.399—	0.080* (0.045)
Inconsistent self-reported voting status with validation	-0.011 (0.036)	—0.507—	-0.091 (0.121)	-0.038 (0.052)	—0.842—	-0.052 (0.050)
N	[40,376]		[7,688]	[24,614]		[21,683]
Panel b: Heterogeneity in registration effects by respondent minority status and sex						
Verified active voter registration	0.074* (0.042)	—0.477—	-0.005 (0.108)	0.127** (0.058)	—0.577—	0.080 (0.060)
Self-reported voter registration	0.040 (0.025)	—0.528—	-0.019 (0.094)	0.053 (0.038)	—0.821—	0.043 (0.027)
Inconsistent self-reported registration status with validation	-0.048 (0.038)	—0.715—	-0.002 (0.125)	-0.080 (0.055)	—0.925—	-0.087 (0.056)
N	[37,330]		[7,049]	[22,911]		[19,918]
Controls: Gender and race	Gender only		Yes	Race only		Race only
FE: County x cong. dist. x year	Yes		Yes	Yes		Yes
FE: State x cohort x year	Yes		Yes	Yes		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Values between columns 1-2 and 3-4 give the p-value on the interaction term of the equivalent fully interacted specification testing for heterogeneity between the two groups. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A10: No evidence that turnout effects differ by ballot composition

Dependent variable and ballot sub-sample	(1) Dependent indicator mean	(2) $\hat{\beta}_1$
Verified general election turnout	0.634	0.139***
		(0.040)
		[51,723]
... turnout when ballot is white man vs. white man	0.618	0.178**
		(0.074)
		[22,143]
... turnout when ballot is white vs. white	0.632	0.163***
		(0.050)
		[37,139]
... turnout when ballot is minority vs. white	0.640	0.126
		(0.080)
		[11,811]
... turnout when ballot is man vs. man	0.620	0.158***
		(0.056)
		[29,573]
... turnout when ballot is woman vs. man	0.649	0.121*
		(0.064)
		[19,034]
... turnout when ballot is midterm elections	0.565	0.171***
		(0.054)
		[25,032]
Controls: Gender and race		Yes
FE: County x cong. district x year		Yes
FE: State x cohort x year		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A11: No evidence that younger cohorts in high *Sesame Street* coverage counties differ in their educational attainment or family income

	(1)	(2)	(3)	(4)	(5)
	CCES			**Race IAT**	
Dependent variable	**Mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	**Mean**	$\hat{\beta}_1^{KL}$
Educational attainment (years)	14	0.167 (0.244) [51,713]	0.283 (0.264) [56,839]	16	0.034 (0.059) [327,498]
Reported family income (in $1,000)	69	3.482 (3.831) [46,596]	1.895 (3.154) [51,729]		.
Controls: Gender and race		Yes	Yes		Yes
FE: County		.	Yes		Yes
FE: State x cohort		.	Yes		Yes
FE: County x cong. district x year		Yes	No		No
FE: State x cohort x year		Yes	No		No

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates are estimated using the same specification as Kearney and Levine (2019), with $(county)$ and $(state \times cohort)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Educational attainment (in years) and reported family income are continuous variables built from binned response options (6 and 12 bins respectively). Estimates using the CCES data employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A12: No evidence that younger cohorts in high *Sesame Street* coverage counties differ in reported indicators of civic engagement

Dependent variable	(1) **Dependent indicator mean**	(2) $\hat{\beta}_1$	(3) $\hat{\beta}_1^{KL}$
Donated blood	0.135	0.035	0.015
		(0.030)	(0.029)
		[43,637]	[47,405]
Was ever a union member	0.279	-0.021	-0.029
		(0.040)	(0.038)
		[47,646]	[52,114]
Was ever in the military	0.134	0.017	0.003
		(0.032)	(0.028)
		[51,723]	[56,850]
Controls: Gender and race		Yes	Yes
FE: County		.	Yes
FE: State x cohort		.	Yes
FE: County x cong. district x year		Yes	No
FE: State x cohort x year		Yes	No

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing outcome and control variables are omitted. Dependent variables are indicator variables. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A13: Impacts on general election participation are larger for non-voters in primary elections

	(1)	*p-value of difference*	(2)
	Primary turnout is		
	Verified		**Not-verified**
Panel a: Heterogeneity in turnout effects by verified primary election turnout			
Verified general election turnout	0.051** (0.024)	—0.067—	0.162*** (0.056)
Self-reported election turnout	0.012 (0.017)	—< 0.001—	0.183*** (0.048)
Inconsistent self-reported voting status with validation	-0.025 (0.021)	—0.369—	0.023 (0.049)
N	[13,937]		[29,421]
Panel b: Heterogeneity in registration effects by verified primary election turnout			
Verified active voter registration	0.015 (0.011)	—0.072—	0.116** (0.055)
Self-reported voter registration	0.003 (0.005)	—0.045—	0.074** (0.035)
Inconsistent self-reported registration status with validation	-0.017 (0.012)	—0.337—	-0.069 (0.053)
N	[13,937]		[29,421]
Controls: Gender and race	Yes		Yes
FE: County x cong. dist. x year	Yes		Yes
FE: State x cohort x year	Yes		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Values between columns 1 and 2 give the p-value on the interaction term of the equivalent fully interacted specification testing for heterogeneity between the two groups. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A14: Among major party voters, no evidence of heterogeneity in increased reported voting for minority, women, and Democratic candidates by respondents' turnout validation

	(1)	(2)	(3)	*p-value of difference*	(4)
		All major party	Major party voters whose turnout is		
	Share	voters	Validated		Not-Validated
Panel a: Reported voting of major party voters on white-minority ballots					
Minority	0.507	0.277*** (0.092)	0.250** (0.119)	—0.652—	0.143 (0.220)
N		[9,200]	[6,467]		[1,310]
Panel b: Reported voting of major party voters on woman-man ballots					
Woman	0.488	0.154** (0.066)	0.106 (0.082)	—0.772—	0.170 (0.218)
N		[14,994]	[10,851]		[2,092]
Panel c: Reported voting of major party voters by candidate party					
Democrat	0.495	0.082* (0.043)	0.075 (0.051)	—0.710—	0.019 (0.149)
N		[40,659]	[29,141]		[6,972]
Controls: Gender and race		Yes	Yes		Yes
FE: County x cong. dist. x year		Yes	Yes		Yes
FE: State x cohort x year		Yes	Yes		Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. The sample is limited to respondents who report a major party vote for the U.S. House on a white-minority ballot (panel a), a man-woman ballot (panel b), or all ballots (panel c). Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Values between columns 3 and 4 give the p-value on the interaction term of the equivalent fully interacted specification testing for heterogeneity between the two groups. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

A1 Implications for election outcomes

It is interesting to consider how these effects may have impacted election outcomes. Our estimates are based on a small share of the electorate born between 1959 and 1968, making projections onto electoral outcomes challenging for several reasons. First, broadcast availability of the show changed substantially over subsequent decades. More critically, although cohorts born after 1968 watched and continue to watch *Sesame Street* in large numbers, their counterfactual activities and programming options are quite different from those of the early 1970s making the relative impact of *Sesame Street* in later decades unclear. One could also argue that *Sesame Street* may have had "general equilibrium effects" as other shows adapted to compete with *Sesame Street*'s success, yet estimating such effects is beyond the scope of this paper.

Because of these confounds, we estimate impacts on election outcomes very conservatively, assuming that only individuals born between 1964 and 1968 are treated. Using this assumption, and the full sample of all CCES respondents from all cohorts, we calculate a measure of electorate exposure for each election in each congressional district as

$$ElectorateExposure_{dy} = \frac{\sum_1^N SScov_j * \mathbb{1}(preschool69_i = 1)}{N_{dy}}.$$

An election outcome is considered impacted if $WinMargin_{dy} - ElectorateExposure_{dy} * \hat{\beta}_1 < 0$, where $WinMargin_{dy}$ is the difference in the received vote share of a winning minority or woman candidate and their major party opponent, and $\hat{\beta}_1$ is the coefficient in the relevant first row of either panel a or b of column 2, table 3. Using this conservative approach we find that the show may have contributed to the electoral victories of 6 minority candidates (1.75% of minority wins) and 7 women candidates (1.57% of woman wins).

Table A15: ***Sesame Street*** **coverage and cohort differences in policy preferences**

	(1)	(2)	(3)	(4)
Dependent variable	**Mean**	**All**		
Panel a: Policy questions in the CCES survey			**Respondent is**	
Index of support for environmental policies (Scale from 0 to 1)	0.596	0.019 (0.025) [56,680]		
Supportive of allowing gays and lesbians to marry legally	0.558	0.122** (0.050) [34,849]		
			Male	***Female***
Index of support for abortion rights (Scale from 0 to 1)	0.689	0.009 (0.029) [56,630]	-0.010 (0.044) [26,334]	0.006 (0.043) [29,782]
			Non-hispanic	***Hispanic***
Index for support of less restrictive immigration policies (Scale from 0 to 1)	0.448	-0.078** (0.033) [38,609]	-0.078** (0.032) [36,713]	-0.028 (0.170) [1,614]
Panel b: Race relations and race policy questions in the CCES survey				
			White	***Black***
Belief in structural racism index Respondent agrees with statements indicating that Blacks face barriers to socio-economic advancement compared to whites (scale 1 - 5)	2.7	-0.083 (0.118) [47,378]	0.070 (0.114) [37,679]	-0.722** (0.310) [4,741]
Supports minority affirmative action programs in employment and college admissions (scale 1 - 4)	2.1	-0.070 (0.097) [24,726]	0.033 (0.105) [19,538]	-0.433 (0.515) [2,403]
Panel c: Race relations and race policy questions in the IAT survey				
			White	***Black***
Supports affirmative action index Indicates support on questions regarding affirmative action in employment and college admissions (scale 0-1)	0.229	0.064** (0.032) [23,022]	0.057 (0.038) [16,525]	0.037 (0.132) [3,312]
Justifies racial profiling index Indicates racial profiling can be justified in certain situations (scale 0-1)	0.091	0.036 (0.023) [23,041]	0.032 (0.027) [16,568]	0.077 (0.091) [3,257]
Controls: Gender and race		Yes	Yes	Yes
FE: County		Yes	Yes	Yes
FE: State x cohort		Yes	Yes	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Indices are built as the average response to questions pertaining to that topic that were administered in the respondent's survey year. Estimates on individual questions are reported in appendix tables A16 and A17. Estimates using the CCES data employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A16: Question estimates of how younger cohorts in high *Sesame Street* coverage counties differ in policy views

	Dependent variable	(1) Mean	(2) All	(3) Respondent is	(4)
Environmental policy					
A	In a trade-off between environmental protection versus jobs and the economy: priority is the environment (1) or the economy (5). (Scale from 1 to 5)	2.9	-0.140		
			(0.129)		
			[22,604]		
B	Supports strengthening the Environmental Protection Agency enforcement of the Clean Air Act and Clean Water Act even if it costs U.S. jobs	0.524	0.026		
			(0.050)		
			[33,435]		
C	Supports giving the Environmental Protection Agency power to regulate carbon dioxide emissions	0.631	0.029		
			(0.049)		
			[33,413]		
D	Supports requiring a minimum amount of renewable fuels (wind, solar, and hydroelectric) in the generation of electricity even if electricity prices increase somewhat	0.585	0.038		
			(0.046)		
			[33,442]		
Abortion policy				***Male***	***Female***
E	Supports making all abortions illegal	0.127	-0.017	0.007	-0.009
			(0.028)	(0.041)	(0.039)
			[49,724]	[23,018]	[26,196]
F	Supports a woman always being able to obtain an abortion as a matter of personal choice	0.538	0.018	0.009	0.015
			(0.041)	(0.063)	(0.058)
			[56,619]	[26,331]	[29,773]
Immigration policy				***Non-hispanic***	***Hispanic***
G	Supports fining U.S. businesses that hire illegal immigrants	0.675	-0.025	0.019	-0.964**
			(0.066)	(0.067)	(0.378)
			[15,328]	[14,675]	[469]
H	Supports granting legal status to all immigrants who have held jobs and paid taxes for at least 5 years, and not been convicted of any felony crimes.	0.481	-0.077*	-0.066	-0.002
			(0.046)	(0.046)	(0.281)
			[38,601]	[36,705]	[1,614]
I	Supports increasing the number of border patrols on the U.S.-Mexican border	0.621	0.073	0.085*	-0.157
			(0.047)	(0.045)	(0.233)
			[38,595]	[36,699]	[1,614]
J	Supports allowing police to question anyone they think may be in the country illegally	0.425	0.111*	0.141**	0.455
			(0.058)	(0.059)	(0.388)
			[20,173]	[19,155]	[773]
	Controls: Gender and race		Yes	Yes	Yes
	FE: County		Yes	Yes	Yes
	FE: State x cohort		Yes	Yes	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1^{KL}$ estimates are reported using $(county)$ and $(state \times cohort)$ fixed effects. All outcome variables other than question A are binary indicators set to 1 if the respondent supports the stated policy. Question A is a scale from 1 (prioritize the environment) to 5 (prioritize the economy). Not all statements are administered in all survey years. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

Table A17: Question estimates of how younger cohorts in high *Sesame Street* coverage counties differ in responses to race policy and race relations questions

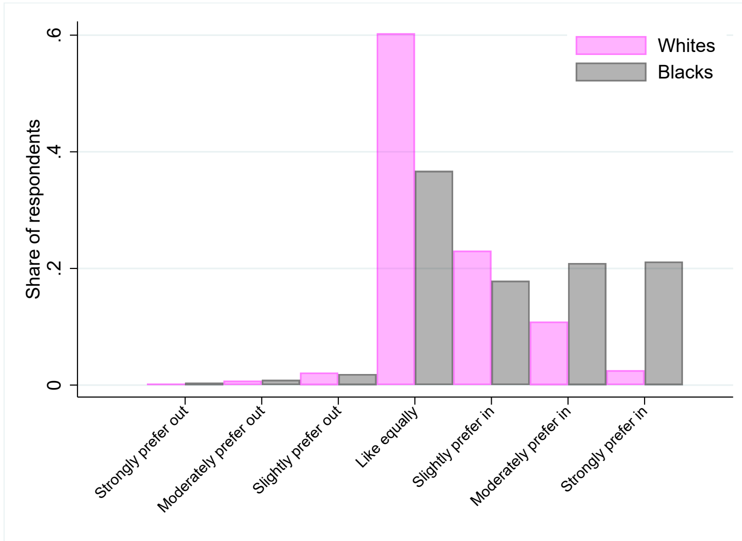
	Dependent variable	(1) Mean	(2) All	(3) Respondent is White	(4) Respondent is Black
Panel a: CCES survey					
Structural disadvantage questions					
A	Irish, Italians, Jewish and many other minorities overcame prejudice and worked their way up. Blacks should do the same without any special favors.	3.6	-0.100 (0.128) [38,769]	-0.219* (0.121) [30,766]	1.376*** (0.492) [3,898]
B	Generations of slavery and discrimination have created conditions that make it difficult for blacks to work their way out of the lower class.	2.7	-0.132 (0.139) [38,780]	0.056 (0.142) [30,772]	-0.625 (0.516) [3,910]
C	White people in the U.S. have certain advantages because of the color of their skin.	3.1	0.035 (0.165) [26,632]	0.076 (0.185) [21,459]	-0.322 (0.556) [2,297]
D	It's really a matter of some people not trying hard enough, if blacks would only try harder they could be just as well off as whites.	3.0	-0.354 (0.252) [8,112]	-0.283 (0.262) [6,626]	4.186** (1.658) [490]
E	Over the past few years, blacks have gotten less than they deserve.	2.6	-0.083 (0.237) [8,120]	0.122 (0.267) [6,633]	-1.680 (1.233) [490]
Other race relations questions					
F	Racial problems in the U.S. are rare, isolated situations.	2.4	-0.039 (0.148) [26,137]	-0.098 (0.166) [21,046]	0.169 (0.605) [2,245]
G	I am angry that racism exists.	4.3	0.145 (0.191) [7,945]	-0.027 (0.231) [6,290]	-0.143 (0.691) [560]
H	I often find myself fearful of people of other races.	2.1	0.079 (0.248) [7,937]	-0.043 (0.251) [6,281]	-0.373 (1.489) [562]
Panel b: Race IAT survey					
Affirmative action questions					
A	A college admissions officer considers applications from African American and European American applicants with similar credentials and cannot accept all. Should the admissions officer more often accept African American than European American applicants?	0.196	0.045 (0.037) [17,934]	0.010 (0.045) [12,953]	0.407** (0.180) [2,416]
B	A corporate personnel officer is evaluating an African American and a European American job applicant who are identically qualified except the European American has more prior experience in related work. Is there a reasonable justification for this personnel officer hiring the African American applicant rather than the European American?	0.237	0.045 (0.041) [17,985]	0.025 (0.053) [12,901]	-0.145 (0.170) [2,483]
Racial profiling questions					
C	Air passengers arriving in the United States must pass through a checkpoint where customs officers may examine contents of baggage in search of contraband such as illegal drugs. Should customs officers be more ready to examine contents of baggage for an African American passenger than a European American passenger?	0.032	-0.017 (0.019) [18,091]	-0.022 (0.020) [13,023]	-0.006 (0.065) [2,473]
D	Do cab drivers in big cities who occasionally choose to pass by an African American person seeking a cab ride, then pick up a nearby European American person, have a reasonable justification for doing this?	0.150	0.092** (0.036) [17,905]	0.085* (0.046) [12,914]	0.177 (0.140) [2,405]
	Controls: Gender and race		Yes	Yes	Yes
	FE: County		Yes	Yes	Yes
	FE: State x cohort		Yes	Yes	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. For CCES questions, the outcome variables represents the degree to which the respondent agrees with the listed statement with responses homogenized across years to fit a 1-5 point scale with: *1-Strongly disagree; 2-Somewhat disagree; 3- Neither agree nor disagree; 4-Somewhat agree; 5-Strongly agree*. Not all statements are administered in all survey years. For race IAT questions, the outcome variable is a binary indicating agreement with the question. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Estimates using the CCES data employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.

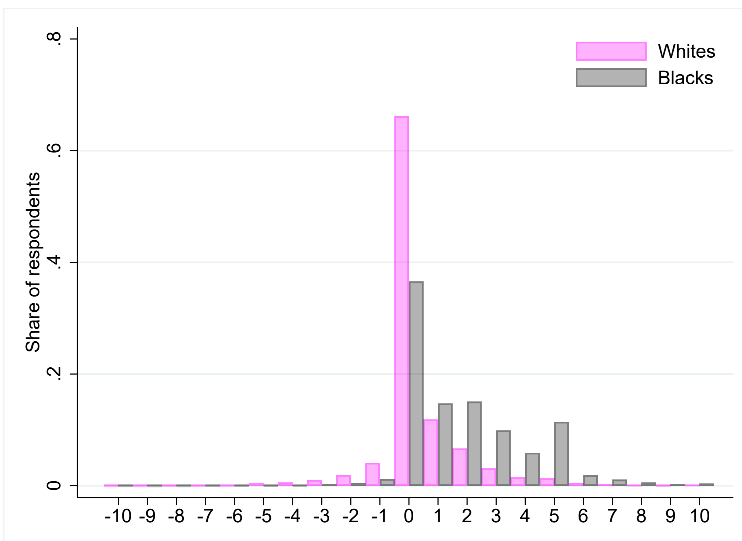
Table A18: No evidence that younger cohorts in high *Sesame Street* coverage counties differ in their reported voting for incumbents

	(1)	(2)	(3)	(4)	(5)
				Major party voters	
	Share	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	**All**	**Validated**
Incumbent	0.495	0.087* (0.044)	0.087** (0.043)	0.008 (0.050)	-0.044 (0.057)
Non-incumbent	0.313	0.051 (0.040)	0.041 (0.039)		
Third party	0.025	-0.029** (0.012)	-0.023* (0.013)		
Not voting	0.167	-0.108*** (0.034)	-0.105*** (0.036)		
N		[43,968]	[48,252]	[34,623]	[24,615]
Controls: Gender and race		Yes	Yes	Yes	Yes
FE: County		.	Yes	.	.
FE: State x cohort		.	Yes	.	.
FE: County x cong. district x year		Yes	No	Yes	Yes
FE: State x cohort x year		Yes	No	Yes	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ is estimated using equation 1 which controls for $(county \times congressional\ district \times year)$ and $(state \times cohort \times year)$ fixed effects. $\hat{\beta}_1^{KL}$ estimates use $(county)$ and $(state \times cohort)$ fixed effects. The sample is limited to respondents voting in U.S. House elections that feature a major party incumbent and non-incumbent candidate. Each observation in the sample has one of the mutually exclusive voting behaviors listed set to 1 and all others set to 0. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.



(a) Preference for in-group and out-group Americans



(b) Difference in warmth towards in-group and out-group Americans

Figure A2. Distribution of explicit racial preferences between European and African Americans

Note: Figure a presents responses to a seven option question on preferences for European and African Americans asked in the race IAT survey. Responses are flipped for African Americans for comparability. Figure b plots the net difference between race IAT test takers' rated warmth (0-coldest to 10-warmest) on two questions, one for their in-group and the other their out-group. Negative values indicates greater expressed explicit warmth towards their out-group while positive values indicates greater expressed explicit warmth towards their in-group.

A2 Heterogeneity by county characteristics

How did the impacts of *Sesame Street* differ based on the racial context in which it aired? A priori it is not clear whether media "contact" with minority role models would have a larger effect in areas where opportunities for other forms of contact were high or low. Where contact opportunities were few, the show may have filled a vacuum allowing for a large effect. Alternatively, effects may have been larger in areas where preschool age exposure to the show could later be reinforced through other interactions. This question is of academic interest as it would shed light on contact theory mechanisms.

To explore this, using the sub-sample of white respondents, we estimate modified versions of our specifications that include the relevant interactions with a county characteristics ($CountyChar_j$),

$$\begin{aligned} VotesMinority_{icjdy} &= \alpha_0 + \alpha_1(preschool69_i * SSCov_j) + \alpha_2(preschool69_i * SSCov_j * CountyChar_j) \\ &+ \alpha_3(preschool69_i * CountyChar_j) + \alpha_4 X_i + \gamma_{scy} + \delta_{jdy} + \epsilon_i, \end{aligned}$$

$$\begin{aligned} IATscore_{icj} &= \alpha_0^{KL} + \alpha_1^{KL}(preschool69_i * SSCov_j) + \alpha_2^{KL}(preschool69_i * SSCov_j * CountyChar_j) \\ &+ \alpha_3^{KL}(preschool69_i * CountyChar_j) + \alpha_4^{KL} X_i + \gamma_{sc}^{KL} + \delta_j^{KL} + \epsilon_i. \end{aligned}$$

Estimates in column 2 of table A19 report α_1 and α_2 on reporting a vote for a minority candidate using several different county characteristics that help characterize the racial context of counties. Estimates of α_1^{KL} and α_2^{KL} on race IAT scores are reported in column 4. Subsequent odd numbered columns report $\hat{\beta}_1$ and $\hat{\beta}_1^{KL}$ using the same sample of observations, as not all county characteristics are available for all counties.

From a policy perspective, α_2 and α_2^{KL} inform us on whether increasing coverage rates and show access might affect certain types of counties more than others. However, it must be highlighted that using α_2 and α_2^{KL} to understand how seeing the show might differently impact individuals in differing county environments is challenging and subject to ambiguity. Heterogeneity in response to predicted coverage rates could reflect a heterogeneous response to watching the show, with the environment amplifying or subduing the effects of the show. Alternatively, county characteristics could impact viewership rates and who selects into viewing the show. Given the lack of national county level data on *Sesame Street* viewership it is not possible for us to fully disentangle these competing mechanisms. Nevertheless, data on ratings for preschool age viewers is available for 28 metro media market areas. Using this, we estimate

$$Ratings_m = \pi_0 + \pi_1 MediaMarketSSCov_m + \pi_2 MediaMarketChar_m + \epsilon_m, \qquad (1)$$

where $Ratings_m$ is the number of viewers in the metropolitan area m, between the ages of 2 and 5 who watched *Sesame Street*, as reported in Haydon (1973), divided by the population count for this age group in the 1970 census. $MediaMarketSSCov_m$ and $MediaMarketChar_m$ are the metropolitan area's *Sesame Street* coverage rate and the area's characteristic, calculated as the population weighted mean of these attributes over the media market's composing counties. Estimates of π_1 and π_2 are reported in column 1 of table A19.

We examine heterogeneity by the standardized 1970 Black population share of counties, the standardized county vote share received by the staunchly segregationist presidential candidate Strom Thurmond in 1948, the standardized county vote share received by the democratic presidential candidate in 1968, the county's metro area's schools' standardized measures of white-Black segregation, as well as an indicator for being in a state that required school segregation prior to 1954.[1] Overall, we find little evidence of meaningful heterogeneity in our estimates by county characteristics. For all estimates of α_2 and α_2^{KL} we fail to reject the null of no heterogeneity by examined county characteristics. From a policy perspective, there is no evidence that increasing coverage rates and access to the show affected certain types of counties more than others.

Heterogeneity in the effects of watching the show could be masked by how these characteristics correlate with viewership rates. Estimates of π_2 using our limited sample are imprecise but suggest a small positive correlation between metro area viewership and the area's Democratic vote share in 1968. Such differences in viewership could mask heterogeneous effects of the show on the children who actually watched the show. Given our data limitations, it is not possible for us to properly evaluate this question. We leave it to be examined in future research.

[1] Data on the 1970 Black population share of counties is sourced from the 1972 County and City Data Book, which is available as part of the replication package for Kearney and Levine (2019). County vote share data are available from Clubb et al. (1987). Metro area school measures of white-Black segregation is available for download from the American Communities Project (Logan 2017). The measure of Black-white segregation is the index of dissimilarity which gives the percentage of children in one group who would have to attend a different school to achieve racial balance across schools in the city.

Table A19: No evidence of clear heterogeneity in effects on white respondents by racial characteristics of counties

	(1)		(2)	(3)	(4)	(5)
	Media market preschool age viewership		**White reported voting for a minority candidate**		**White race IAT score**	
			With county characteristic	**$\hat{\beta}_1$ on same sample**	**With county characteristic**	**$\hat{\beta}_1^{KL}$ on same sample**
$SSCov_m$	0.595***	$SSCov_j \times preschool69_i$		0.312***		-0.067***
	(0.186)			(0.095)		(0.025)
	[28]			[8,055]		[261,189]
Standardized county Black population share in 1970						
$SSCov_m$	0.615***	$SSCov_j \times preschool69_i$	0.304***	0.312***	-0.067***	-0.067***
	(0.185)		(0.095)	(0.095)	(0.025)	(0.025)
$Std.\ 1970\ Black\ pop.\ share_m$	-0.051	$SSCov_j \times preschool69_i$	0.110		-0.001	
	(0.074)	$\times\ Std.\ 1970\ Black\ pop.\ share_j$	(0.140)		(0.032)	
N	[28]		[8,055]	[8,055]	[260,003]	[260,003]
Standardized county Thurmond vote share in 1948						
$SSCov_m$	0.596***	$SSCov_j \times preschool69_i$	0.303***	0.289***	-0.064**	-0.069***
	(0.192)		(0.097)	(0.096)	(0.026)	(0.025)
$Std.\ 1948\ Thurmond\ vote\ share_m$	-0.009	$SSCov_j \times preschool69_i$	0.098		0.022	
	(0.079)	$\times Std.\ 1948\ Thurmond\ vote\ share_j$	(0.162)		(0.039)	
N	[28]		[7,949]	[7,949]	[257,687]	[257,687]
Standardized county Democrat vote share in 1968						
$SSCov_m$	0.546***	$SSCov_j \times preschool69_i$	0.361***	0.312***	-0.070**	-0.067***
	(0.175)		(0.101)	(0.095)	(0.028)	(0.025)
$Std.\ 1968\ democratic\ vote\ share_m$	0.075*	$SSCov_j \times preschool69_i$	-0.094		0.006	
	(0.040)	$\times Std.\ 1968\ democratic\ vote\ share_j$	(0.104)		(0.024)	
N	[28]		[8,055]	[8,055]	[259,916]	[259,916]
Standardized metro area school white-Black segregation						
$SSCov_m$	0.530***	$SSCov_j \times preschool69_i$	0.346***	0.331***	-0.071***	-0.073***
	(0.170)		(0.102)	(0.098)	(0.026)	(0.025)
$Std.\ school\ segregation_m$	-0.077	$SSCov_j \times preschool69_i$	0.095		0.021	
	(0.046)	$\times\ Std.\ school\ segregation_j$	(0.080)		(0.023)	
N	[28]		[7,805]	[7,805]	[240,842]	[240,842]
Southern states requiring school segregation prior to 1954						
$SSCov_m$	0.592***	$SSCov_j \times preschool69_i$	0.245**	0.312***	-0.074**	-0.067***
	(0.197)		(0.124)	(0.095)	(0.030)	(0.025)
$Pop.\ share\ in\ state\ requiring\ sch.\ seg._m$	-0.029	$SSCov_j \times preschool69_i$	0.149		0.027	
	(0.062)	$\times\ State\ required\ sch.\ seg._j$	(0.191)		(0.052)	
N	[28]		[8,055]	[8,055]	[261,189]	[261,189]
Controls: Gender			Yes	Yes	Yes	Yes
Controls: Cohort x county characteristic			Yes	Yes	Yes	Yes
FE: County x cong. district x year			Yes	Yes	No	No
FE: State x cohort x year			Yes	Yes	No	No
FE: County			.	.	Yes	Yes
FE: State x cohort			.	.	Yes	Yes

Note: Each column of coefficients in each panel are estimated together. Column 1 reports the π_1 and π_2 coefficients described in appendix A2 from a regression of 28 media markets' preschool-age viewership rates on the media market's coverage rate $SScov_m$, controlling for a different metro area characteristic in each panel. Observations are weighted by the media markets' preschool aged population. Column 2 uses CCES data to estimate α_1 and α_2 as specified in appendix A2 on white respondents facing a minority-white ballots with the addition of $(preschool69_i * SSCov_j * CountyChar_j)$, reported in the table. These specifications control for $(preschool69_i * CountyChar_j)$ as well as respondents' gender, *(county × congressional district × year)* and *(state × cohort × year)* fixed effects. Column 3 reports $\hat{\beta}_1$ for the same sample used in column 2. Estimates for columns 2 and 3 employ survey weights. Column 4 estimates α_1^{KL} and α_2^{KL} as specified in appendix A2 on white race IAT test takers with the addition of $(preschool69_i * SSCov_j * CountyChar_j)$, reported in the table. This specification controls for $(preschool69_i * CountyChar_j)$ as well as respondents' gender, *(county)* and *(state × cohort)* fixed effects. Column 5, reports $\hat{\beta}_1^{KL}$ for the same sample used in column 4. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Standard errors are reported in parentheses, clustered at the county level, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.