Casting roles, casting votes:

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

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Abstract

Sesame Street's representation of minority characters, egalitarian minority-white interactions and portrayal of working women was distinctive in the mass media landscape of 1969, when it started airing. By exploiting both age variation and technological variation in broadcast reception, this paper contributes to the media and contact theory literatures by showing that positive representations of minorities via mass media can reduce long-run prejudice and impact voting, an important societal outcome. We find that for preschool-age children, a 20 percentage point (1 standard deviation) increase in *Sesame Street* coverage reduced adult measures of racial biases for white respondents and increased reported voting for minority and women candidates to the U.S. House of Representatives by 13% and 9.7% respectively. Voter turnout also increased by 4.4% in all elections. Voting for Democratic candidates increased because of the increase in voting for diverse candidates. When the sample is restricted to ballots featuring white men, turnout gains are split between parties.

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1 Introduction

Can child mass media change prejudices and racial biases in adulthood? Can it impact who we elect as our representatives as adults? Women and minority groups in the U.S. have historically 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 persistence and consequences of these patterns. As a result, much research on mass media has focused on it's 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, present great opportunity, with the potential to increase a population's exposure to, knowledge of, and respect for social groups. This is particularly true for young children. For many young kids, a blockbuster show like *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. In this paper, we examine whether early childhood exposure to *Sesame Street*, and its portrayal of an inclusive, egalitarian, and diverse America, impacted adult voting patterns and racial biases decades later.

To identify the causal impact of *Sesame Street* on adult biases and voting patterns we use two dimensions in exposure variation: age when the show started airing, and technologically induced variation in geographic coverage rates. This identification strategy was first used by Kearney and Levine (2019) in their work investigating *Sesame Street*'s impact on educational outcomes. Using data from a large election survey and an online survey of racial biases we apply the Kearney and Levine (2019) strategy using respondent's age and county of residence to identify potential exposure and explore our novel and socially important question. We find that exposure to *Sesame Street* increased political engagement, turnout, and changed voter preferences in adulthood. Exposed survey respondents were more likely to report voting for minority and women candidates to the U.S. House of Representatives. As a result, exposed respondents were more likely to report voting for Democratic candidates, because Democratic candidates are more likely to be women and minorities. In elections featuring two white men, turnout gains were evenly split between parties and we find little evidence that exposure to *Sesame Street* changed policy views or political identities in ways that favored one party over another. Instead, we find that *Sesame Street* reduced measures of racial bias against Blacks, suggesting that a long-run reduction in racial biases and prejudice explains these important differences in voting patterns.

These results provide causal evidence that child mass media programming can *positively* impact *long-run* patterns of prejudice and racial bias. We show that representation in mass media is indeed critical, not just as role models for underrepresented groups, but also for other audience members who's biases and prejudices are shaped by such representation. Finally, we show that these changes impact consequential decisions, namely one's participation in the electoral process and choice candidate to the U.S. House of Representatives.

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)). The show was strikingly different from contemporaneous children's programming.[1] 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 characters, including women, held jobs.[2] 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

[1] *Mr. Rodger's neighborhood*, which began airing in 1968, had pedagogical content, but focused on socio-emotional skills. Regular cast members in early seasons were European-American, with the exception of Officer Clemmons, an African-American police officer, first introduced in two appearances in August 1968 and then appearing in 15% of episodes between 1969 and 1972. *Captain Kangaroo* which aired on CBS from 1955 to 1984 was a popular children's show which had some educational content. It added its first Black cast member in 1968.

[2] Main human characters in the first season included Susan, an African-American nurse and her spouse Gordon, an African-American veteran and history teacher; Bob, a European-American music teacher and Mr. Hooper, a European-American shopkeeper; and a diverse and rotating cast of children. The cast expanded in season 2 adding David, an African-American law student; and Linda, a deaf European-American librarian. Season 3 saw the addition of Luis, the Mexican-American owner of the Fix-It shop; and Maria, a Puerto Rican librarian.

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 egalitarian interactions between cast members of different races and between adults, children, and Muppets.[3] 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.[4] As such, *Sesame Street* offers an opportunity to investigate the long-run effects of increased media exposure to non-stereotyped minority role models in childhood, an age when beliefs are thought to be especially malleable (Rutland and Killen (2015)).

In this paper, we show that exposure to *Sesame Street*, and its portrayal of an inclusive, egalitarian and diverse community of role models, reduced racial biases and prejudice in the long-run, impacting adult voting and candidate preferences. These results make important contribution to several strands of literature in economics and psychology including the literatures on mass media, prejudice interventions, contact theory, and role models.

We know that mass media can have important effects on a wide variety of economic behaviors such as savings and consumption, violent crime, health, divorce, and fertility (see the review by DellaVigna and La Ferrara (2015)) and the impacts of mass media on voting behavior and political preferences has been extensively studied (Gentzkow (2006), Gentzkow and Sinkinson (2011), DellaVigna and Kaplan (2007), Durante et al. (2019), Falck et al. (2014), Campante et al. (2017)). Though extensive, this literature has focused on adult content media to find effects on short-run voting outcomes. Similarly, a growing literature is exploring the media's impact on racial attitudes and in-group biases, and its role in exacerbating racial and ethnic tensions (Müller and Schwarz. (2023), Petrova et al. (2019), Adena et al. (2015), Yanagizawa-Drott (2014), Adena et al.

[3]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 being reinstated (Greene (2019)).

[4]The memorial comments on Twitter upon the death of Emilio Delgado, a long time cast member often expressed that he was, for many young watchers, 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?"*

(2015), Wang (2021), Ang (2023)). Few papers however have explored the mass media's potential to improve these outcomes. Paluck (2009) and Blouin and Mukand (2019), have shown how radio soap operas in Rwanda reduced inter-group prejudice and a recent working paper examines how exposure to the radio broadcast of *The Adventures of Superman* increased racial tolerance in the US (Armand et al. (2023)). Overall, little is known about how media exposure in childhood impacts later life voting and preferences over candidates' demographics.

Child media has long been a contentious topic with concern expressed by policy makers and parents about media content, and how race and gender topics are presented to children.[5] Yet evidence of how exposure to child media affects long-run outcomes is very limited and has focused primarily on education and human capital (Gentzkow and Shapiro (2008), Huang and Lee (2010), Hernæs et al. (2019), Fiorini and Keane (2014), Riley (2019)). Particularly relevant is Kearney and Levine (2019), which investigates *Sesame Street*'s impacts on academic and labor market outcomes. Using national census data and estimates of county level coverage rates, they find that *Sesame Street* 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.[6] Evidence of *Sesame Street*'s impacts on attitudes towards race and gender is limited and exclusively focused on short-run impacts, finding some positive effects.[7] The long-run impact of *Sesame Street* on race and gender attitudes has not been evaluated.

Our findings contribute to the literature on the contact theory hypothesis, which suggests that interpersonal contact is an effective way to reduce prejudice when interacting and cooperating with equal status and common goals (Allport et al. (1954)). Though *Sesame Street* did not increase contact between children of different backgrounds, it increased children's exposure to minority role

[5]Adukia et al. (2023) have recently documented patterns in race and gender representation in children's books and its correlation with local politics

[6]*Sesame Street* was also the subject of several early studies including impact evaluations commissioned by governments and the Children's Television Workshop (CTW) which produces it, as well as academic studies in psychology and education. These typically focused on *Sesame Street*'s educational impacts, generally concluding that it was effective at teaching preschool aged kids its targeted educational curriculum (Fisch et al. (1999), Fisch and Truglio (2014) and Murphy (1991)), with positive effects on pro-social behaviors (Paulson (1974), Leifer (1975), Bankart and Anderson (1979), Zielinska and Chambers (1995)).

[7]Outside of the US contexts, Cole et al. (2003) look at the impact of two CTW designed *Sesame Street* inspired shows broadcast in Israel and Palestine. After 4 months of exposure, they found an improvement in attitudes of children towards the other ethnic group for Israeli-Jewish preschoolers. Mares and Pan (2013) conducted an international meta-analysis of *Sesame Street* using CTW data finding positive short-run impacts on attitudes towards social differences.

models, and exemplified equal and cooperative interactions between in and out-group members. In economics and psychology, a number of papers have used field-experiments or exploited policy changes to test the contact theory hypothesis, most finding evidence in support of it (Paluck et al. (2019), Sacerdote (2001), Boisjoly et al. (2006), Carrell et al. (2019), Finseraas et al. (2019), Rao (2019)). Much of this literature evaluates short to medium-run effects, often looking at contact in college years. A set of recent papers has looked at long-run effects on school age children finding impacts on political preferences (Billings et al. (2021), Brown et al. (2021)) and neighborhood choice (Merlino et al. 2022).

The question of whether prejudice can be impacted by experimental interventions has been the focus of a substantial literature in the fields of psychology and economics (see the review by Bertrand and Duflo (2017) and meta-analysis by Paluck et al. (2021)). Yet few interventions examining prejudice reduction through entertainment and mass media have been evaluated, and the examination of long-term outcomes is exceedingly rare (Browne Graves (1999), Paluck et al. (2021)). This is surprising. The impact of mass media content, particularly for children's media, is frequently discussed in the popular press and existing work on media interventions finds promising results (Vittrup and Holden 2011).[8] Furthermore, mass media, by its technological nature, offers important advantages when it comes to implementation at scale as compared to the other types of interventions studied.

Finally, this paper also contributes to the literature on role models. There is a substantial literature that examines the impact of role models in economics, with numerous studies looking at the impact of exposure to under-represented teachers. The focus of this work has generally been on academic and aspirational outcomes, particularly for the minority group, be it female students in STEM fields or racial and ethnic minorities (Gershenson et al. (2022), Dee (2004), Dee (2005), Gershenson et al. (2016), Lindsay and Hart (2017), Fairlie et al. (2014), Lim and Meer (2020), Porter and Serra (2020), Canaan and Mouganie (2021), Lusher et al. (2018)). Less is known about how such role models affect prejudice, particularly for the majority group.[9] In their review chapter on

[8] Vittrup and Holden (2011) had children watch videos that featured positive relationships between a racially diverse cast. It should be noted that two of the five videos were excerpts from *Sesame Street*. They find that exposure to the videos had a positive effect on children's out-group attitudes.

[9] Beaman et al. (2009) observed reduced stereotyping of women by men in India when local village councils were

the impacts of minority leaders, Bertrand and Duflo (2017) conclude that more work investigating impacts on the attitudes of the majority is needed, a contribution of our paper.

The remainder of the paper is organized as follows. Section 2 presents the data. Section 3 details our empirical strategy. Section 4 shows that *Sesame Street* increased electoral participation and engagement. Section 5 shows how *Sesame Street* changed who respondents report voting for. Section 6 examines which candidate attributes drive voter behaviors. Section 7 shows *Sesame Street*'s impacts on racial biases. Finally, 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 (CCES) in election years from 2006 to 2020.[10] 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.[11] 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.[12] *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.[13] 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

randomly forced to elect female leaders.

[10]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.

[11]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.

[12]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* is associated with post-election survey completion (see appendix table A1).

[13]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).

these ballots.[14] Of the 5,914 candidacies in our data, 522 (8.8%) are election runs by Black candidates, 320 (5.4%) are by Hispanic candidates, 4,891 (82.7%) are by 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 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.

Race and gender biases: Data on race and gender biases is available from Harvard's *Project Implicit*, 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.[15] 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.[16] IATs have been used in psychology, and in economics, as a way of measuring implicit attitudes (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

[14]Basic race (Black, Hispanic, other, white) and gender (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 was collected by 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.

[15]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 7.

[16]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)), 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)).

IAT scores of US resident citizens born between 1959 and 1968, along with their survey responses.[17]

***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 and 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.[18] 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 is then applied to non-commercial stations 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 1 shows the map of broadcasting stations by signal type along with the Kearney and Levine (2019) estimates of simulated coverage levels in each county. 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. These estimated coverage rates are validated using Nielsen ratings for 28 metropolitan areas. They are strongly predictive of *Sesame Street* ratings with viewership in the past week increasing by 0.58 percent in the 2 to 5 year old demographic for each percentage point increase in 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.[19]

[17]Respondents that report having taken three or more IATs are dropped from the sample.

[18]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 received 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 (Kearney and Levine (2019)).

[19]More details on the construction of predicted coverage rates are available in Kearney and Levine (2019) and its appendix.

3 Empirical strategy

To identify the effects of early childhood exposure to *Sesame Street* on voting patterns 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.[20] We then 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 old's and 19% of 4 year old's attending preschool in the 1970 census. For 5 year old's, 57% attended kindergarten with 88% of these kindergartners enrolled in half day programs (Kearney and Levine (2019)). Not only were these 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.

In addition to the variation in cohort exposure, differences in broadcast coverage rates between counties generates a second dimension of exposure variation for identification. Treated kids are thus cohorts that were under the age of 6 when the show first aired *and* who lived in high broadcast coverage counties. Untreated kids are those who were 6 or older in 1969 as well as those who were younger but who lived in counties with a poor broadcast signal. The main empirical models used in estimations using the CCES data are thus specified below,

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

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

Y is the outcome variable of interest for individual i, in cohort c, residing in county j, and for the CCES data 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

[20]These respondents would have been between their late 30's and early 60's when they were surveyed in the election years between 2006 and 2020 or when they completed their IATs.

is interacted with the variable $SSCov_j$, the predicted level of *Sesame Street* coverage in county j. Figures also present disaggregated cohort level estimates, λ_{1c}, as specified in equation 2 where $Cohort_i$ 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, in addition to the γ and δ fixed effects. 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 in the same way (the presence of third party candidates, idiosyncratic shocks, up-ballot elections...).

For the IAT data, controlling for voting patterns in a particular election is not necessary. These results are estimated using the following specification which is identical to the specification used in Kearney and Levine (2019),

$$Y_{ijyc} = \beta_0 + \beta_1^{KL}(preschool69_i * SSCov_j) + \beta_2 X_i + \gamma_{sc}^{KL} + \delta_j^{KL} + \epsilon_i, \tag{3}$$

$$Y_{ijyc} = \lambda_0 + \sum_{c=59, c\neq 63}^{68} \lambda_{1c}^{KL}(Cohort_i * SSCov_j) + \lambda_2 X_i + \gamma_{sc}^{KL} + \delta_j^{KL} + \epsilon_i. \tag{4}$$

These specifications employ a (*state* $\times$ *cohort*) fixed effect, γ_{sc}^{KL}, and a *county* fixed effect, δ_j^{KL}. Estimates of β_1^{KL} using the CCES data will also be reported.

Finally, results for a specification labeled Edu are also reported. These are estimated using an augmented set of controls, X_i^{Edu}, that controls for a respondent's education and income.

It is important to note that our measure of exposure does not capture actual viewership of *Sesame Street* as national county level ratings data is unavailable. Rather, the treatment measures a county's exposure to a *Sesame Street* broadcast. Our estimates should thus be interpreted as "intent-to-treat" estimates, identifying the impact of making the show more available to children rather than the impact of watching the show on an individual child. As such, we will report the effects of a 20 percentage point (about one standard deviation) increase in the predicted coverage rate of a county. It is also important to highlight that treatment in our context is composed of both increased exposure to *Sesame Street*, 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.[21] While we frame our interpretation as the impacts of *Sesame Street* viewership, our estimates could also, for instance, 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

Migration is an important issue to address given our data limitations. Ideally we would observe respondents' counties of residence during their childhood years, but this is not available in the CCES and IAT data. Our measure of *Sesame Street* coverage exposure is thus assigned based on county of residence when surveyed, between 2006 and 2020, many decades after potential exposure. Substantial levels of out-of-county migration between childhood and adulthood would generate attrition bias, biasing our results towards zero. The possibility that exposure may selectively affect the likelihood of moving out-of-county is also a concern. We explore these issues below.

We find no evidence that *Sesame Street* impacted the probability of lifelong residence in one's

[21] 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.

childhood city, or birth state. Problematic selection bias could arise and confound our estimates if exposure to *Sesame Street* impacted respondents' inter-county mobility between the period of treatment and the survey dates. Though we do not observe county of residence in 1969, in many waves of the CCES data respondents were asked how long they have lived in their current *city* of residence. Using this, we identify city never-movers who have lived in their city since they were two or younger. To check for a relationship between exposure to *Sesame Street* and out-of-city mobility, we estimate equations 1 and 2 on the city never-mover indicator. Results are reported in table A3 and figure A1a. For all specifications, we fail to reject the null of no effect on mobility away from one's early childhood city (p-value= 0.23) , nor is there any visual evidence suggesting a discontinuity across treated and untreated cohorts in figure A1a. Patterns in the American Community Survey (ACS) are consistent with this finding. ACS respondents report their state of birth allowing the identification of inter-state movers. We build $SSCov_s$, a coarse state level measure of early childhood exposure to *Sesame Street* by calculating the population weighted mean coverage rate across a state's counties. We estimate specifications similar to equations 1 and 2 on the indicator for residing in the state of one's birth, replacing $SSCov_j$ with $SSCov_s$ and using coarser state and cohort fixed effects. Results are reported in the second row of table A3 and figure A1b. We fail to reject the null of no effect on mobility away from one's birth state (p-value=0.40), nor is there any visual evidence suggesting a discontinuity across treated and untreated cohorts.

Coverage rates between destination and sending counties are correlated for county-to-county migrants, counteracting attenuation bias. Only 9% of the CCES sample are city never-movers, raising the issue of attenuation bias. It should be noted however that the survey question only allows us to identify city never-movers. This will not include respondents who moved between cities within a county, and respondents who moved, but then returned, to their childhood city. Thus the share of respondents for whom our predicted county coverage rate is accurate lies between this 9% and the 61% of respondents observed in the ACS residing in their birth state. Attenuation bias will also be less severe if inter-county migration tends to occur between counties with similar coverage rates. 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 table A4 show a positive correlation of 0.347 (p-value$<$0.001) between the 1969 county coverage rates of destination and sending counties for inter-county movers. This is not surprising. 34 % of county-to-county moves in this data occur between neighboring counties that have highly correlated 1969 coverage rates: the mover weighted correlation between destination and sending county coverage rates for neighboring county movers is of 0.790 (p-value$<$0.001) as reported in column 2. Yet even for long distance moves, there is a positive correlation between destination and sending counties' coverage rates. Column 3 limits the pairs to non-neighboring different state counties and shows evidence of a statistically significant 0.066 (p-value$<$0.001) correlation. Overall, attenuation bias due to inter-county moving will attenuate our results though this effect is counteracted by the high level of correlation between destination and sending county coverage rates.

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* may affect the probability individuals select into the sample, biasing estimates.[22] 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 A5. Column 1 presents estimates using all the race IAT test takers, column 2 using only 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 *Sesame Street* exposure changed selection into the race IAT data. In contrast, *Sesame Street* increased the likelihood of taking

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

the (less widely used) gender-career IAT test.[23] Given this selection, our analysis in section 7 will focus on the race IAT data.

4 *Sesame Street* increased electoral participation and engagement

In this section, we examine the impacts of exposure to *Sesame Street* on voter turnout and registration. We then investigate mechanisms behind the observed increase in electoral engagement.

***Sesame Street* increased turnout in general elections.** Treated cohorts exposed to *Sesame Street* in preschool are more likely to vote in general elections. 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 for a 20 percentage point (1 sd) increase in the coverage rate, the probability of treated cohort members turning out to vote increases by 2.8 percentage points, a 4.4% increase. This effect is highly statistically significant, and robust to alternative specifications (columns 3 and 4). Figure 2a plots the cohort level $\hat{\lambda}_{1y}$ estimates, confirming a discontinuity in validated turnout between treated and untreated 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.

The similarity between estimates using verified and self-reported turnout is important. 85% of respondents report voting but verified turnout rates are substantially lower, and can only be validated for 63% of CCES respondents. 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 are reported the third row of table 1, panel a. We see no impact of treatment on the probability of having inconsistent verified and self-reported turnout

[23] The positive and statistically significant estimate of 0.034 implies that for a 20 percentage point increase in the *Sesame Street* coverage rate, the share of test takers from treated cohorts in a county increases by 0.68 percentage points.

records. There is no evidence that exposure to *Sesame Street* affects the likelihood respondents misrepresent their behavior to surveyors.

To better understand why *Sesame Street* impacts turnout, we examine respondents' reasons for not voting.[24] Results are presented in appendix table A6. Exposure to *Sesame Street* reduced the likelihood that respondents report not voting because they dislike the candidates and because they are not registered. These effects are robust across specifications and explain about 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.

***Sesame Street* increased voter registration.** Treated cohorts who were exposed to *Sesame Street* in preschool are more likely to be registered to vote. 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 20 percentage point (1 sd) increase in the coverage rate, the probability of having a verified active registration record is 1.8 percentage points higher, a 2.4% increase. This effect is statistically significant, and robust to alternative specifications (columns 3 and 4). Figure 2b plots the cohort level $\hat{\lambda}_{1y}$ estimates and confirms a discontinuity in validated registrations between treated and untreated cohorts. While still present, the estimate of $\hat{\beta}_1$ on self-reported voter registration is smaller in magnitude by about half with little evidence of a discontinuity between treated and untreated cohorts.

The difference between these two indicators is interesting. 94% of respondents self-report being registered to vote while only 74% have a verified active registration. A substantial share of respondents misreport their registration status, and 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 treated cohorts 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 treated cohorts are more politically

[24] 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.

informed, suggesting that *Sesame Street* could have reduced misinformation about respondents' voter registration status.

Heterogeneity in this increase in electoral participation is examined in the appendix. Table A7 evaluates differences by respondent demographics. We cannot reject the null of no heterogeneity in estimated effects. Turnout increases are also observed across all ballots, regardless of candidate demographics, with $\hat{\beta}_1 \in [0.121, 0.178]$, as reported in appendix table A8.

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

We find no evidence of a change 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 table 2. Like Kearney and Levine (2019) we cannot reject the null of no effect on educational attainment (in years) and log family income. Though the coefficients are positive at (0.167 (p-value= 0.49) and 0.016 (p-value= 0.78) respectively, they are not statistically significant. The same estimation applied to the IAT data also fails to reject the null of no effect on educational attainment at 0.034 (p-value= 0.56) . Furthermore, the $\hat{\beta}_1^{Edu}$ estimates on turnout and registration outcomes reported in column 4 of table 1 include controls for educational attainment and family income and do not substantially differ from the main result. Increased educational attainment and family income are not the main mechanisms behind the participation results.

***Sesame Street* increased political knowledge and interest.** Estimates in table 3, panel a, show that exposure to *Sesame Street* increased political knowledge in adulthood. Treated cohorts in higher coverage counties 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.

***Sesame Street* increased weak political identities, but not active political engagement or strong political identities.** With respect to political identification, a more complex picture emerges in table 3, panel b. Treated 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.[25] 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 active expressions 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 $\in [0.19, 0.77]$).[26] Consistent with this, table A10 shows that effects on electoral participation are larger for marginally engaged respondents, as proxied by their verified primary election turnout status.

Overall, these results suggest that *Sesame Street*'s impacts on registration and turnout stem from an increase in the political engagement and knowledge of marginal voters.

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

[26]We also find no evidence of a change 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 A9.

5 *Sesame Street* changed who respondents report voting for

In this section, we examine how exposure to *Sesame Street* 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 amongst the racially integrated cast increased reported voting for minority and women candidates.

5.1 *Sesame Street* increased reported voting 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 white candidate. 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 4, panel a. Estimates of β_1 using each indicator as an outcome are reported in column 2 of table 4, panel a.

Treated cohorts exposed to *Sesame Street* are substantially more likely to report voting for a minority candidate. The $\hat{\beta}_1$ estimate of 0.271 (p-value < 0.001) reported in the first row of table 4, panel a, implies that for a 20 percentage point increase in the coverage rate, the probability a voter reports a vote for a minority candidate is 5.4 percentage points higher, a 13% increase. This effect is highly statistically significant, and robust to alternative specifications (columns 3 and 4). 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 4, panel a. 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, with substantial gains for minority candidates. 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 A11. Potential impacts on election outcomes are discussed in appendix A1.

In columns 5 and 6, we split the sample to estimate the effect separately for white and minority voters. 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 *Sesame Street* increased reported voting for women candidates

Using respondents voting on ballots that feature a woman running against a man, 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 4, panel b.

Treated cohorts exposed to *Sesame Street* 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 4, panel b, implies that for a 20 percentage point increase in the coverage rate, the probability of reporting a vote for a woman candidate is 3.8 percentage points higher, a 9.7% increase. This effect is highly statistically significant, and robust to alternate specifications (columns 3 and 4). 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* had little effect on reported voting for men candidates as demonstrated 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. When the sample is limited to validated voters reporting a major party vote, in appendix table A11, we cannot reject the null of no effect, though the direction of the coefficient suggests a small positive effect on women candidates' vote share. Possible implications on election outcomes are discussed in appendix A1.

In columns 5 and 6, 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.

5.3 *Sesame Street* increased reported voting 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 4, panel c.

Cohorts exposed to *Sesame Street* 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 20 percentage point increase in the coverage rate, the probability a voter reports a Democratic vote is 2.0 percentage points higher, a 5.0% increase. Estimates using alternative specifications (columns 3 and 4) are of a slightly smaller magnitude, but still statistically significant at conventional levels (p-values $\in [0.05, 0.07]$). 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 and not generally statistically significant. 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 A11, we cannot reject the null of no effect, though the coefficient suggests a small positive effect favoring Democratic candidates.

It is worth highlighting that *Sesame Street*'s impact 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 A11 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* increases the likelihood respondents report voting for minority candidates, women candidates and Democratic candidates. What drives these patterns is unclear. Many of these 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% as opposed to 19.4% of white candidacies. In the following section 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 exposure to *Sesame Street* changed biases, the most likely explanation for the observed impacts on reported voting.

6.1 Diverse candidates drive the increase in reported voting for Democrats

Sesame Street increased reported voting for Democratic candidates. Is this effect due to changed political views or a response to candidate characteristics that correlate with candidate party?

Table 5, 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 *Sesame Street* had on the electorate.

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 Reported voting for all minority candidates increases

Sesame Street increased reported voting for minority candidates. In table 6, panel a, we consider whether candidate characteristics that correlate with a candidate's minority identity, namely political party and gender identity, explain this pattern.[27] Columns 1 and 2 of table 6, 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 6, panel a, we estimate β_1 separately on ballots where the minority candidate is a woman (column 3) and a man (column 4). 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) .

[27] 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.

6.3 Reported voting for Democratic women increases, with larger effects for minority women

Sesame Street increased reported voting for women candidates. In table 6, panel b, we consider if this a response to candidate characteristics that correlate with a candidate's gender, namely political party and minority identity. [28] Columns 1 and 2 of table 6, 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 given the small number of 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 6, 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 appears to be doubly affected by exposure to *Sesame Street* coverage.

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.

***Sesame Street* did not consistently change policy preferences in a pattern consistent with party platforms.** Exposure to *Sesame Street* may have changed treated respondents' policy preferences towards policies that are more likely to be espoused by minority and female candidates. To investi-

[28] 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. Overall, 26.2% of candidacies by women are also minority candidacies as opposed to 20.5% of candidacies by men.

gate this mechanism, we examine CCES respondents' views on specific policies. Table 7 presents aggregated indices of responses to broad policy topics.[29] Overall, we cannot rule out that exposure to *Sesame Street* has no impact on respondents' support for environmental policies or abortion rights. 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.[30] 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.[31] Overall, impacts on policy views show contrasting effects that do not clearly align with any particular political platform.

***Sesame Street* increased political and party identification without clearly favoring one end of the political spectrum.** Table 8, panel a, examines treatment effects on respondents' reported political ideology in both the CCES and IAT data. Treated respondents are more likely to report identifying with a liberal or a conservative ideology as opposed to being moderate or neutral, as reported in column 1 for the CCES data and column 3 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 2 and 4, 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).

[29] Administered questions vary across surveys. Selected questions used to construct indices were administered in multiple survey years. Disaggregated estimates are reported in table A12.

[30] Estimates for all questions estimated separately are presented in appendix table A13.

[31] 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.

***Sesame Street* did not change voting for incumbents.** As exposure to *Sesame Street* increased political knowledge it may have reduced the propensity to default towards voting for incumbent candidates which we explore in appendix table A14. There is no evidence that exposure changed relative reported voting for incumbent candidates. Turnout effects are split. Respondents report increased voting for both incumbent and non-incumbent candidates.[32]

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 *Sesame Street's* impacts on measures of racial biases

In this section we present evidence that the observed changes in voter preferences are the result of a change 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. We focus our analysis on the race IAT given the evidence of selection into the gender-career IAT discussed in section 3.2. It should be noted however that the selection pattern is itself suggestive that exposure to *Sesame Street* also increased interest in issues surrounding gender-career biases, though we cannot make conclusive statements on how it impacted gender-career IAT scores.[33]

***Sesame Street* reduced measures of white racial bias towards Blacks in the long run.** Column 2 of table 9 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 white test takers. The β_1^{KL} estimate of -0.067 sds (p-value = 0.007) reported in row 1 of column 2 implies that for a 20

[32]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) .

[33]If less biased test takers select into taking the test, estimates will be biased towards zero. Estimates of β_1^{KL} on the gender-career IAT scores are, unsurprisingly, small and not statistically significant (appendix table A15).

percentage point (1 sd) increase in the coverage rate, standardized IAT scores decrease by -0.01 standard deviations.[34] 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.[35] 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 administers two race thermology questions which we examine in table 10. Column 1 pools results for white and Black test takers, results for white test takers are presented in column 2, and for Black test takers in column 3. 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).[36] In this context, it is not surprising that we detect no discernible effect of *Sesame Street* on the explicit preferences expressed by white test takers (p-value $\in [0.17, 0.51]$). Estimates of β_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, the magnitude of the effects become larger, and a small effect signalling reduced in-group preferences on the warmth questions becomes statistically distinguishable from zero (p-value$= 0.09$) .[37] 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

[34]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.

[35]The estimate for test takers whose racial identifier does not fit in either of these categories is smaller in magnitude and not statistically significant.

[36]Expressing racial preferences, particularly for white Americans, is stigmatized in contemporary US 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.

[37]The -0.192 point estimate suggests that for a 20 percentage point (1sd) increase in the coverage rate, the difference in expressed warmth for European versus African Americans for these white non-zero test takers shrinks by -0.04 points, a 0.017 sd decrease.

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.

Non-incumbent minority candidates gain most from their electorate's exposure to *Sesame Street*. 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 11 presents heterogeneity results by minority candidate incumbency status. 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).[38] 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. For incumbent minority candidates, while still positive, the estimate is substantially smaller at 0.070 (p-value= 0.62) and not statistically significant with the difference between these two groups approaching statistically significance at the 10% level (p-value= 0.11) when comparing impacts in columns 2 and 3.

8 Conclusion

Can child mass media change 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 topics of gender and racial identity are approached in child media, has also been the subject of recent politicized debate. Yet despite this

[38]Estimates in column 1 are imprecise given the small number of white-minority ballots that have no incumbent candidates.

flood of attention, there is little causal evidence on how positive non-stereotyped representation affects social and economic outcomes.

This paper help inform these debates. In this paper, we show that preschool-age exposure to *Sesame Street*, 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 were 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 were 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 reported voter preferences over candidate demographics, consistent with observed reductions in the racial bias measures of white test takers observed for these cohorts.

Our results suggest multiple new directions for future research. Research in economics on the effects of child 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 had 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 behavioral lever for efforts in prejudice reduction that has been understudied in academic research.

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9 Figures

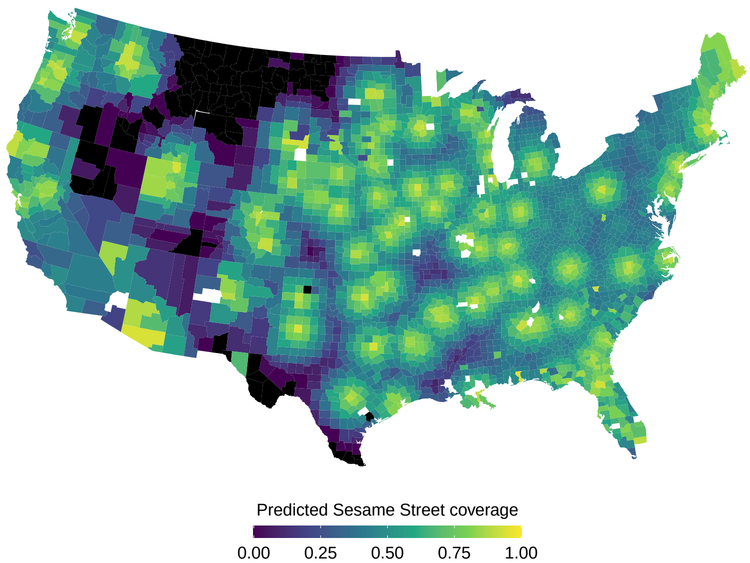
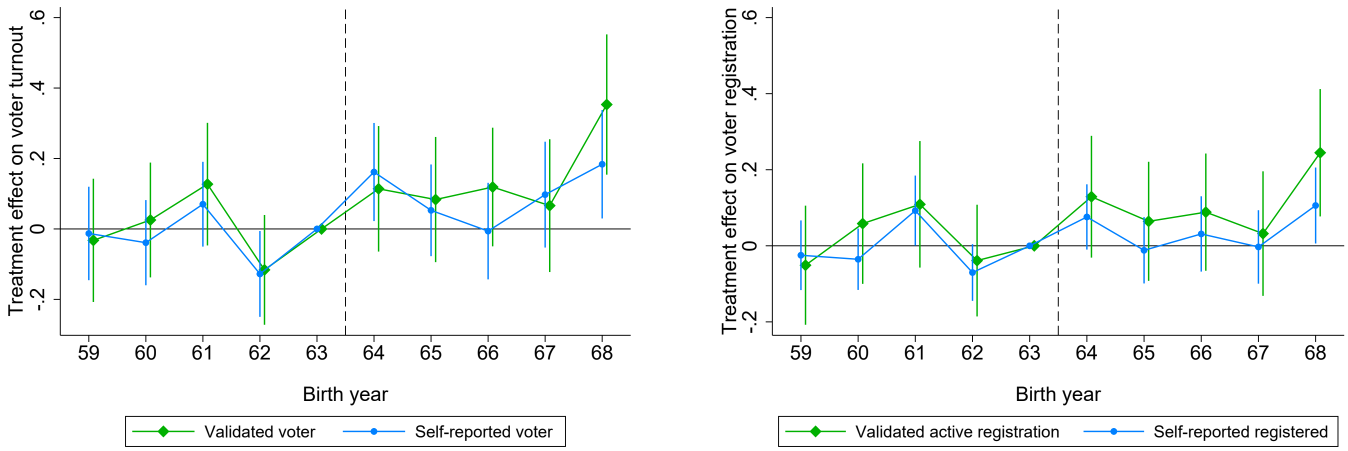


Figure 1. Broadcast signal towers and estimated county coverage rates of *Sesame Street* in 1969

Notes: Color gradient represents predicted share of households that could watch *Sesame Street* in that county in 1969 as calculated in Kearney and Levine (2019). Estimated coverage is simulated using the average relationship between the broadcast technology of signal towers and signal receipt.

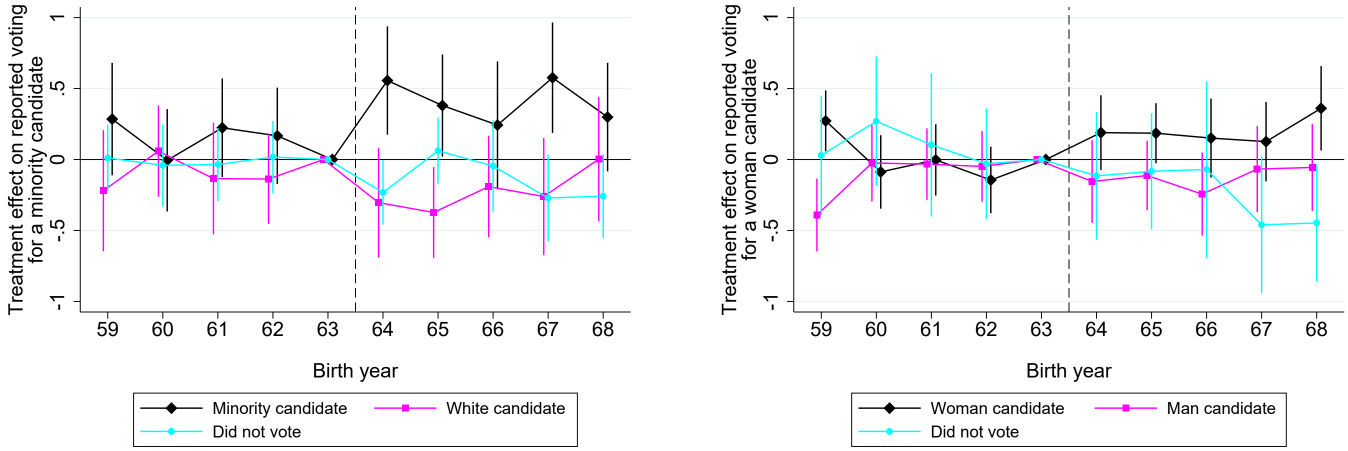


(a) Self-reported and validated turnout

(b) Self-reported and validated registration

Figure 2. ***Sesame Street*** **increased voter turnout and registration**

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*×*cohort*×*year*) and (*county*×*congressional district*×*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. ***Sesame Street*** **increased reported voting for minority and women candidates**

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.

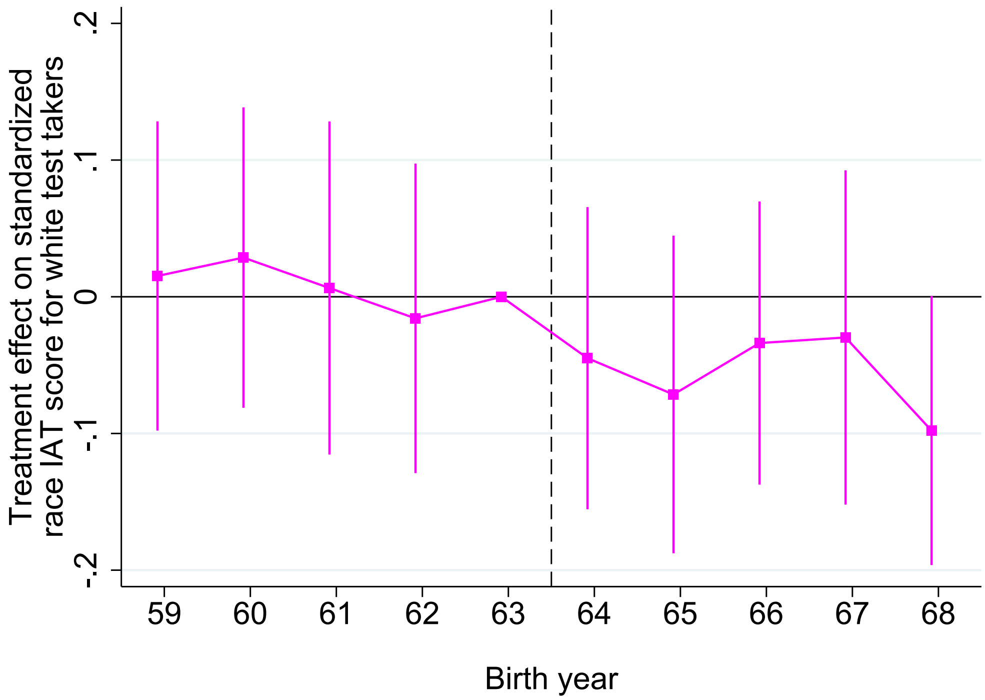


Figure 4. ***Sesame Street*** **reduced measures of implicit bias against Blacks for white test takers**

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 as specified in equation 4. 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 a stronger Black-bad and white-good associations. 95% confidence intervals are depicted using standard errors clustered at the county level.

10 Tables

Table 1: ***Sesame Street*** **increased voter turnout and registration**

Dependent indicator variable	(1) Dependent indicator mean	(2) $\hat{\beta}_1$	(3) $\hat{\beta}_1^{KL}$	(4) $\hat{\beta}_1^{Edu}$
Panel a: ***Sesame Street*** **increased self-reported and validated voter turnout**				
Verified general election turnout	0.634	0.139*** (0.040) [51,723]	0.109*** (0.040) [56,850]	0.146*** (0.040) [46,587]
Self-reported election turnout	0.851	0.120*** (0.032) [51,723]	0.118*** (0.033) [56,850]	0.113*** (0.032) [46,587]
Inconsistent self-reported voting status with validation	0.229	-0.034 (0.034) [51,723]	-0.018 (0.034) [56,850]	-0.047 (0.037) [46,587]
Panel b: ***Sesame Street*** **increased voter registration and knowledge of registration status**				
Verified active voter registration	0.744	0.091** (0.038) [47,604]	0.081** (0.038) [52,030]	0.089** (0.039) [42,981]
Self-reported voter registration	0.944	0.044* (0.024) [47,353]	0.049* (0.028) [51,777]	0.039 (0.024) [42,753]
Inconsistent self-reported registration status with validation	0.210	-0.066* (0.037) [47,353]	-0.062* (0.033) [51,777]	-0.072* (0.037) [42,753]
Controls: Gender and race		Yes	Yes	Yes
FE: County		.	Yes	.
FE: State x cohort		.	Yes	.
FE: County x cong. district x year		Yes	No	Yes
FE: State x cohort x year		Yes	No	Yes
Controls: Educ., family income		No	No	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using our preferred specification, with the addition of controls for education and family income. Numbers in brackets report the observations used in each estimation, once omitted singletons and observations with missing control variables are omitted. Registration data is not available for the 2006 survey. 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: No evidence that *Sesame Street* changed educational attainment or family income

	(1)	(2)	(3)
	CCES		Race IAT
Dependent variable	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	$\hat{\beta}_1^{KL}$
Educational attainment (years)	0.167 (0.244) [51,713] ⟨13.973⟩	0.283 (0.264) [56,839] ⟨13.973⟩	0.034 (0.059) [327,498] ⟨16.272⟩
Log reported family income (in $1,000)	0.016 (0.057) [46,596] ⟨4.061⟩	-0.011 (0.048) [51,729] ⟨4.061⟩	.
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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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 built 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 outcome and control variables are omitted. Numbers in carrots report the mean of the dependent variable. 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 3: *Sesame Street* increased political knowledge and moderate political identification, but not active engagement

	(1)	(2)	(3)	(4)	(5)	(6)
		CCES			Race IAT	
Dependent variable	**Mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	$\hat{\beta}_1^{Edu}$	**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]	0.162** (0.075) [45,784]		
Interest in government and public affairs (scale 1-4)	3.4	0.155** (0.079) [46,835]	0.165** (0.078) [51,248]	0.169** (0.076) [42,342]		
Accessed no news media in the past 24 hours	0.056	-0.029 (0.020) [41,890]	-0.018 (0.021) [45,616]	-0.033 (0.021) [37,829]		
Panel b: Political identification						
Has a political ideology	0.615	0.024 (0.039) [51,575]	0.015 (0.038) [56,707]	0.025 (0.041) [46,457]	0.804	0.037*** (0.009) [319,563]
..... Weakly identifies with a political ideology	0.396	0.017 (0.041) [51,575]	0.034 (0.038) [56,707]	0.024 (0.042) [46,457]	0.599	0.027** (0.012) [319,563]
..... Strongly identifies with a political ideology	0.220	0.007 (0.032) [51,575]	-0.018 (0.028) [56,707]	0.001 (0.035) [46,457]	0.205	0.010 (0.010) [319,563]
Identifies with a major political party	0.839	0.055* (0.031) [51,527]	0.069** (0.033) [56,659]	0.051 (0.033) [46,407]		
..... Weakly identifies with a major political party	0.418	0.073* (0.043) [51,527]	0.091** (0.040) [56,659]	0.060 (0.045) [46,407]		
..... Strongly identifies with a major political party	0.421	-0.018 (0.040) [51,527]	-0.022 (0.039) [56,659]	-0.009 (0.042) [46,407]		
Panel c: Active political engagement						
Verified congressional primary turnout	0.347	0.026 (0.042) [47,604]	0.014 (0.041) [52,030]	-0.000 (0.041) [42,981]		
Donated money to a political campaign or organization	0.241	0.011 (0.038) [43,637]	0.016 (0.033) [47,405]	0.008 (0.037) [39,206]		
Put up a political sign	0.198	0.033 (0.038) [43,637]	0.045 (0.035) [47,405]	0.025 (0.037) [39,206]		
Attended a political meeting	0.134	-0.041 (0.031) [43,637]	-0.006 (0.028) [47,405]	-0.025 (0.032) [39,206]		
Worked for a candidate or campaign	0.061	-0.020 (0.021) [43,637]	-0.008 (0.018) [47,405]	-0.021 (0.022) [39,206]		
Reports having run for office	0.028	0.012 (0.013) [43,491]	0.014 (0.013) [47,267]	0.021 (0.013) [39,067]		
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	No	Yes		No
FE: State x cohort x year		Yes	No	Yes		No
Controls: Educ., family income		No	No	Yes		No

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using our preferred specification, with the addition of controls for education and family income. 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 4: *Sesame Street* increased reported voting for minority, women, and Democratic candidates

	(1)	(2)	(3)	(4)	(5)	*p-value of difference*	(6)
	Share	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	$\hat{\beta}_1^{Edu}$			
Panel a: Reported voting on minority-white ballots					**Respondent is**		
					White		**Minority**
Minority	0.409	0.271***	0.195***	0.293***	0.312***	—0.790—	0.383
		(0.082)	(0.074)	(0.081)	(0.095)		(0.260)
White	0.398	-0.144*	-0.083	-0.109	-0.266***	—0.289—	0.022
		(0.085)	(0.070)	(0.082)	(0.092)		(0.268)
Third party	0.028	0.004	-0.006	0.012	-0.009	—0.918—	-0.016
		(0.021)	(0.017)	(0.020)	(0.028)		(0.063)
Not voting	0.165	-0.132*	-0.106*	-0.195***	-0.037	—0.050—	-0.389**
		(0.069)	(0.063)	(0.064)	(0.077)		(0.169)
N		[11,811]	[12,819]	[10,573]	[8,055]		[2,469]
Panel b: Reported voting on woman-man ballots							
					Female		**Male**
Woman	0.396	0.192***	0.192***	0.187***	0.214**	—0.794—	0.170
		(0.058)	(0.050)	(0.064)	(0.106)		(0.129)
Man	0.417	-0.031	-0.004	-0.029	-0.164*	—0.127—	0.055
		(0.062)	(0.059)	(0.067)	(0.094)		(0.109)
Third party	0.025	-0.041**	-0.056***	-0.040**	-0.032	—0.874—	-0.039
		(0.017)	(0.020)	(0.016)	(0.021)		(0.037)
Not voting	0.162	-0.121***	-0.132***	-0.118***	-0.017	—0.164—	-0.186*
		(0.046)	(0.047)	(0.046)	(0.070)		(0.100)
N		[19,034]	[20,602]	[17,148]	[8,952]		[7,565]
Panel c: Reported voting by candidate party							
Democrat	0.398	0.100***	0.065*	0.077**			
		(0.037)	(0.036)	(0.039)			
Republican	0.407	0.033	0.059*	0.052			
		(0.037)	(0.036)	(0.039)			
Third party	0.027	-0.021*	-0.021*	-0.019*			
		(0.011)	(0.011)	(0.012)			
Not voting	0.168	-0.113***	-0.103***	-0.111***			
		(0.033)	(0.034)	(0.033)			
N		[51,723]	[56,850]	[46,587]			
Controls: Gender and race		Yes	Yes	Yes	Yes		Yes
FE: County		.	Yes	.	.		.
FE: State x cohort		.	Yes	.	.		.
FE: County x cong. dist. x year		Yes	No	Yes	Yes		Yes
FE: State x cohort x year		Yes	No	Yes	Yes		Yes
Controls: Educ., family inc.		No	No	Yes	No		No

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using our preferred specification, with the addition of controls for education and family income. 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 (panel c); 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 5-6 give the p-value on the interaction term of a 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 5: No evidence that *Sesame Street* changed 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 our preferred specification presented in 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 6: *Sesame Street* increased reported voting for women Democrats and minority candidates of both parties

	(1)	*p-value of difference*	(2)	(3)	*p-value of difference*	(4)
Panel a: Reported voting for minority candidates when they 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 for women candidates when they 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 estimates are estimated using our preferred specification presented in 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 a 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 7: ***Sesame Street*****'s impacts on policy preferences**

	(1)	(2)	(3)
Dependent variable	**All**		
Panel a: Policy questions in the CCES survey		**Respondent is**	
Index of support for environmental policies (Scale from 0 to 1)	0.019 (0.025) [56,680] ⟨0.596⟩		
Supportive of allowing gays and lesbians to marry legally	0.122** (0.050) [34,849] ⟨0.558⟩		
		Male	***Female***
Index of support for abortion rights (Scale from 0 to 1)	0.009 (0.029) [56,630] ⟨0.689⟩	-0.010 (0.044) [26,334] ⟨0.661⟩	0.006 (0.043) [29,782] ⟨0.713⟩
		Non-hispanic	***Hispanic***
Index for support of less restrictive immigration policies (Scale from 0 to 1)	-0.078** (0.033) [38,609] ⟨0.448⟩	-0.078** (0.032) [36,713] ⟨0.441⟩	-0.028 (0.170) [1,614] ⟨0.582⟩
Panel b: Race relations and race policy 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)	-0.083 (0.118) [47,378] ⟨2.685⟩	0.070 (0.114) [37,679] ⟨2.529⟩	-0.722** (0.310) [4,741] ⟨3.837⟩
Supports minority affirmative action programs in employment and college admissions (scale 1 - 4)	-0.070 (0.097) [24,726] ⟨2.058⟩	0.033 (0.105) [19,538] ⟨1.868⟩	-0.433 (0.515) [2,403] ⟨3.327⟩
Panel c: Race relations and race policy 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)	0.064** (0.032) [23,022] ⟨0.229⟩	0.057 (0.038) [16,525] ⟨0.212⟩	0.037 (0.132) [3,312] ⟨0.293⟩
Justifies racial profiling index Indicates racial profiling can be justified in certain situations (scale 0-1)	0.036 (0.023) [23,041] ⟨0.091⟩	0.032 (0.027) [16,568] ⟨0.090⟩	0.077 (0.091) [3,257] ⟨0.088⟩
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, 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. Numbers in carrots report the mean of the dependent variable. 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 A12 and A13. 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 8: *Sesame Street* increased political and party and identification

	(1)	(2)	(3)	(4)
	CESS		**IAT**	
	$\hat{\beta}_1$	**Conditional on lib. or cons. dem. or rep.**	$\hat{\beta}_1^{KL}$	**Conditional on lib. or cons**
Panel a: Impacts on political ideology				
Liberal	0.012 (0.033) ⟨0.239⟩	0.084 (0.056) ⟨0.388⟩	0.025** (0.012) ⟨0.516⟩	-0.006 (0.013) ⟨0.359⟩
Moderate/Neutral	-0.030 (0.039) ⟨0.324⟩		-0.037*** (0.009) ⟨0.196⟩	
Conservative	0.012 (0.040) ⟨0.376⟩		0.012 (0.011) ⟨0.288⟩	
Not sure	0.006 (0.024) ⟨0.061⟩			
Observations	[51,575]	[29,549]	[319,563]	[256,923]
Panel b: Impacts on party identity				
Democrat	0.071* (0.038) ⟨0.429⟩	0.057 (0.046) ⟨0.512⟩		
Independent	-0.058** (0.028) ⟨0.139⟩			
Republican	-0.016 (0.044) ⟨0.410⟩			
Not Sure	0.003 (0.013) ⟨0.022⟩			
Observations	[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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. 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. Numbers in carrots report the mean of the dependent variable. 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 9: *Sesame Street* reduced measures of implicit bias against Blacks for white test takers

	(1)	(2)	(3)	(4)
Sample	**Mean score**	$\hat{\beta}_1^{KL}$		
White test takers	0.145	-0.067*** (0.025) [261,189]	-0.067*** (0.025) [261,189]	-0.051* (0.028) [252,098]
Black test takers	-0.800	0.065 (0.079) [43,397]	0.052 (0.078) [43,397]	0.097 (0.098) [38,601]
Hispanic/Other/Unreported	-0.063	0.041 (0.065) [44,356]	0.043 (0.064) [44,356]	-0.104 (0.074) [38,288]
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: Larger positive values indicate a stronger Black-bad and white-good associations. Each coefficient is the result of a separate regression using the indicated controls and fixed effects on standardized IAT scores. 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 10: *Sesame Street* had muted impacts on explicit racial preferences reported on the IAT survey

	(1)	(2)	(3)
		Test taker is	
Dependent variable and sub-sample	**White and Black**	**White**	**Black**
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)	-0.028 (0.024) [288,344] ⟨4.580⟩	-0.016 (0.025) [247,539] ⟨4.480⟩	-0.175* (0.105) [40,376] ⟨5.183⟩
..........Reported preference for non-equal respondents	-0.045 (0.039) [123,781] ⟨5.347⟩	-0.031 (0.042) [97,944] ⟨5.210⟩	-0.130 (0.136) [25,443] ⟨5.870⟩
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)	-0.061 (0.040) [297,737] ⟨0.584⟩	-0.056 (0.041) [255,188] ⟨0.381⟩	-0.090 (0.178) [42,109] ⟨1.803⟩
..........Difference in warmth for non-zero respondents	-0.231** (0.095) [112,997] ⟨1.534⟩	-0.192* (0.112) [85,991] ⟨1.126⟩	-0.378* (0.200) [26,614] ⟨2.844⟩
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, 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. Numbers in carrots report the mean of the dependent variable. Difference in warmth is only calculated for white and Black Americans and is the difference in the warmth reported between the respondent's in-group and respondent's out-group. Estimates on individual questions are reported in appendix table A13. 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 11: Non-incumbent minority candidates gain most from electorate exposure to *Sesame Street*

	(1)	*p-value of difference*	(2)	*p-value of difference*	(3)
			Minority candidate is		
	No incumbents		**Incumbent**		**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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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 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 a 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.

Appendix

Table A1: Exposure to *Sesame Street* is not associated with selection into post-election survey response

	(1)	(2)	(3)	(4)
Dependent indicator variable	**Dependent indicator mean**	$\hat{\boldsymbol{\beta}}_{\mathbf{1}}$	$\hat{\boldsymbol{\beta}}_{\mathbf{1}}^{\boldsymbol{KL}}$	$\hat{\boldsymbol{\beta}}_{\mathbf{1}}^{\boldsymbol{Edu}}$
Took the post-election survey	0.888	0.001 (0.024) [59,251]	-0.008 (0.024) [64,335]	-0.006 (0.024) [53,374]
Controls: Gender and race		Yes	Yes	Yes
FE: County		.	Yes	.
FE: State x cohort		.	Yes	.
FE: County x cong. district x year		Yes	No	Yes
FE: State x cohort x year		Yes	No	Yes
Controls: Educ., family income		No	No	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using our preferred specification, with the addition of controls for education and family income. 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

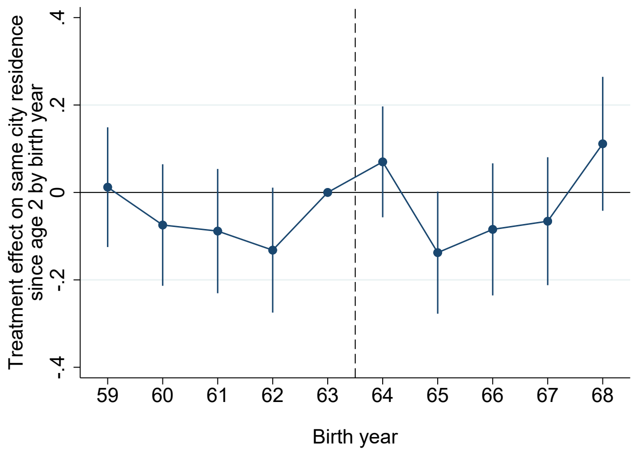
Race		Gender						Election outcomes				
								Democratic margin (ppts)			Share	
Democrat	Republican	Democrat	Republican	Respondents	Share	Ballots	Share	Mean	Median	Std. Dev.	Close (±10)	Very close (±5)
		Man	Man	25,304	0.442	1,333	0.451	-6.10	-12.35	30.05	0.152	0.075
		Man	Woman	3,642	0.064	179	0.061	-0.83	1.34	28.47	0.218	0.117
White	White	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
		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
Minority	White	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
		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
White	Minority	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
		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
Minority	Minority	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
		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
All		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.

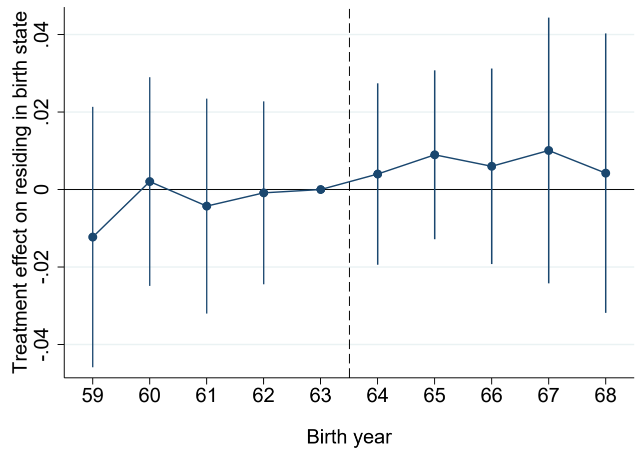
Table A3: No evidence that exposure to *Sesame Street* is associated with city or state moving

	(1)	(2)	(3)	(4)	(5)
Dependent indicator variable	**Dependent indicator mean**	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	$\hat{\beta}_1^{Edu}$	$\hat{\beta}_1^{ACS}$
Lived in current city since age 2	0.088	0.032 (0.027) [31,903]	0.039 (0.028) [34,909]	0.036 (0.029) [28,505]	
Living in birth state	0.613				0.0116 (0.0137) [6,124,373]
Controls: Gender and race		Yes	Yes	Yes	Yes
FE: Birth State		.	.	.	Yes
FE: Cohort		.	.	.	Yes
FE: County		.	Yes	.	.
FE: State x cohort		.	Yes	.	.
FE: County x cong. district x year		Yes	No	Yes	.
FE: State x cohort x year		Yes	No	Yes	.
Controls: Educ., family income		No	No	Yes	.

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using out preferred specification, with the addition of controls for education and family income. $\hat{\beta}_1^{ACS}$ is the coefficient on the interaction between the treated cohort indicator and the population weighted mean *Sesame Street* coverage rate in the state, estimated with *(birth state)* and *(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. All estimates employ survey weights. Standard errors are reported in parentheses, clustered at the county level for columns 2-4, and the state level for column 5, with the following significance indicators: * p<0.1, ** p<0.05 and *** p<0.01.



(a) Being a city non-mover



(b) Being a state non-mover

Figure A1. No evidence that exposure to *Sesame Street* is associated with city or state moving

Notes: Figure a uses CCES data and plots 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*×*cohort*×*year*) and (*county*×*congressional district*×*year*) fixed effects are included. The outcome variable is an indicator variable set to one if the respondent reports that they have lived in the same city since the age of 2. Figure b uses ACS data and plots the $\hat{\lambda}_{1c}$ estimates on the interaction between birth year indicators and the state's predicted *Sesame Street* coverage in 1969 calculated as the population weighted mean county coverage rates. Controls for race and gender as well a (*birth state*) and (*cohort*) fixed effects are included. The outcome variable is an indicator variable set to one if the respondent lives in the same state as their state of birth. The 1963 cohort is the omitted category. Estimates include survey weights. 95% confidence intervals are depicted using standard errors clustered at the county level in figure a, and at the state level in figure b.

Table A4: Coverage rates of sending and destination counties are correlated for inter-county migrants

	(1)	(2)	(3)
	Destination county coverage rate		
	All	**Neighboring**	**Out-of-state non-neighbor**
Sending county coverage rate	0.3469*** (0.0019) [236,880]	0.7899*** (0.0046) [17,796]	0.0656*** (0.0025) [151,735]

Note: Each observation consists of migration flows observed between county pair combinations by the census between 2016 and 2020. All regressions are weighted by the number of migrants between that sending and destination county. Numbers in brackets report the county pair observations with migrants 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 A5: No evidence that *Sesame Street* 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	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
Original observations	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 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 A6: *Sesame Street*'s impacts on reported reasons for not voting

Dependent indicator variable	(1) Dependent indicator mean	(2) $\hat{\beta}_1$	(3) $\hat{\beta}_1^{KL}$	(4) $\hat{\beta}_1^{Edu}$
All reasons	0.151	-0.128***	-0.128***	-0.113***
		(0.033)	(0.035)	(0.033)
		[47,592]	[52,017]	[42,972]
...Did not like the candidates	0.024	-0.025**	-0.028**	-0.023*
		(0.012)	(0.012)	(0.014)
		[47,592]	[52,017]	[42,972]
...I am not registered	0.024	-0.030*	-0.024	-0.024
		(0.016)	(0.016)	(0.016)
		[47,592]	[52,017]	[42,972]
...I am not interested	0.016	-0.014	-0.011	-0.008
		(0.012)	(0.012)	(0.013)
		[47,592]	[52,017]	[42,972]
...Sick or disabled	0.017	-0.009	0.006	-0.009
		(0.013)	(0.014)	(0.013)
		[47,592]	[52,017]	[42,972]
...I did not feel that I knew enough about the choices	0.014	-0.001	-0.019	-0.002
		(0.011)	(0.016)	(0.011)
		[47,592]	[52,017]	[42,972]
...All other reasons listed	0.057	-0.048**	-0.052**	-0.046**
		(0.020)	(0.022)	(0.021)
		[47,592]	[52,017]	[42,972]
Controls: Gender and race		Yes	Yes	Yes
FE: County		.	Yes	.
FE: State x cohort		.	Yes	.
FE: County x cong. district x year		Yes	No	Yes
FE: State x cohort x year		Yes	No	Yes
Controls: Educ., family income		No	No	Yes

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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using out preferred specification, with the addition of controls for education and family income. 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 A7: 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	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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. 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 a 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 A8: No evidence that turnout effects differ by ballot composition

Dependent indicator variable ... and ballot sub-sample	(1) Observations	(2) Dependent indicator mean	(3) $\hat{\beta}_1$
Verified general election turnout	[51,723]	0.634	0.139*** (0.040)
... white man vs. white man	[22,143]	0.618	0.178** (0.074)
... white vs. white	[37,139]	0.632	0.163*** (0.050)
... minority vs. white	[11,811]	0.640	0.126 (0.080)
... man vs. man	[29,573]	0.620	0.158*** (0.056)
... woman vs. man	[19,034]	0.649	0.121* (0.064)
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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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 A9: No evidence that *Sesame Street* impacted observed measures of civic engagement

	(1)	(2)	(3)	(4)
Dependent variable	**Dependent indicator mean**	$\hat{\boldsymbol{\beta}}_{\mathbf{1}}$	$\hat{\boldsymbol{\beta}}_{\mathbf{1}}^{\boldsymbol{KL}}$	$\hat{\boldsymbol{\beta}}_{\mathbf{1}}^{\boldsymbol{Edu}}$
Donated blood	0.135	0.035 (0.030) [43,637]	0.015 (0.029) [47,405]	0.041 (0.030) [39,206]
Was ever a union member	0.279	-0.021 (0.040) [47,646]	-0.029 (0.038) [52,114]	-0.041 (0.043) [42,708]
Was ever in the military	0.134	0.017 (0.032) [51,723]	0.003 (0.028) [56,850]	0.021 (0.034) [46,587]
Controls: Gender and race		Yes	Yes	Yes
FE: County		.	Yes	.
FE: State x cohort		.	Yes	.
FE: County x cong. district x year		Yes	No	Yes
FE: State x cohort x year		Yes	No	Yes
Controls: Educ., family income		No	No	Yes

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using our preferred specification, with the addition of controls for education and family income. 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. All estimates using 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: Heterogeneity in turnout and registration effects by verified primary election turnout

	(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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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 a 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 A11: *Sesame Street* increased reported voting for minority, women, and Democratic candidates amongst major party voters, with no evidence of heterogeneity by respondents' turnout validation

	(1)	(2)	(3)	*p-value of difference*	(4)
			Respondent is		
	Share	**All**	**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. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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 a 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, even though younger cohorts watched and continue to watch *Sesame Street* in large numbers, their counterfactual activities and programming options are quite different from those of the early 1970's 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 then considered impacted if $WinMargin_{dy} - ElectorateExposure_{dy} * \hat{\beta}_1 < 0$, where $WinMargin_{dy}$ in 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 coefficients in the relevant first row of either panel a or b of column 2, table 4 . 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 A12: Question estimates of *Sesame Street*'s impact on policy views

		(1)	(2)	(3)
	Dependent variable	**All**	**Respondent is**	
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)	-0.140 (0.129) [22,604] ⟨2.909⟩		
B	Supports strengthening the Environmental Protection Agency enforcement of the Clean Air Act and Clean Water Act even if it costs U.S. jobs	0.026 (0.050) [33,435] ⟨0.524⟩		
C	Supports giving the Environmental Protection Agency power to regulate Carbon Dioxide emissions	0.029 (0.049) [33,413] ⟨0.631⟩		
D	Supports requiring a minimum amount of renewable fuels (wind, solar, and hydroelectric) in the generation of electricity even if electricity prices increase somewhat	0.038 (0.046) [33,442] ⟨0.585⟩		
Abortion policy			***Male***	***Female***
E	Supports making all abortions illegal	-0.017 (0.028) [49,724] ⟨0.127⟩	0.007 (0.041) [23,018] ⟨0.123⟩	-0.009 (0.039) [26,196] ⟨0.131⟩
F	Supports a woman always being able to obtain an abortion as a matter of personal choice	0.018 (0.041) [56,619] ⟨0.538⟩	0.009 (0.063) [26,331] ⟨0.485⟩	0.015 (0.058) [29,773] ⟨0.585⟩
Immigration policy			***Non-hispanic***	***Hispanic***
G	Supports fining US businesses that hire illegal immigrants	-0.025 (0.066) [15,328] ⟨0.675⟩	0.019 (0.067) [14,675] ⟨0.682⟩	-0.964** (0.378) [469] ⟨0.514⟩
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.	-0.077* (0.046) [38,601] ⟨0.481⟩	-0.066 (0.046) [36,705] ⟨0.473⟩	-0.002 (0.281) [1,614] ⟨0.635⟩
I	Supports increasing the number of border patrols on the US-Mexican border	0.073 (0.047) [38,595] ⟨0.621⟩	0.085* (0.045) [36,699] ⟨0.626⟩	-0.157 (0.233) [1,614] ⟨0.513⟩
J	Supports allowing police to question anyone they think may be in the country illegally	0.111* (0.058) [20,173] ⟨0.425⟩	0.141** (0.059) [19,155] ⟨0.433⟩	0.455 (0.388) [773] ⟨0.270⟩
	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, estimated using the same specification as Kearney and Levine (2019), with $(county)$ 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. Numbers in carrots report the mean of the dependent variable. 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: Estimates of *Sesame Street*'s impact on race policy and race relations questions

	Dependent variable	(1) All	(2) Respondent is White	(3) 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.	-0.100 (0.128) [38,769] ⟨3.635⟩	-0.219* (0.121) [30,766] ⟨3.768⟩	1.376*** (0.492) [3,898] ⟨2.610⟩
B	Generations of slavery and discrimination have created conditions that make it difficult for blacks to work their way out of the lower class.	-0.132 (0.139) [38,780] ⟨2.677⟩	0.056 (0.142) [30,772] ⟨2.512⟩	-0.625 (0.516) [3,910] ⟨3.894⟩
C	White people in the U.S. have certain advantages because of the color of their skin.	0.035 (0.165) [26,632] ⟨3.126⟩	0.076 (0.185) [21,459] ⟨2.951⟩	-0.322 (0.556) [2,297] ⟨4.468⟩
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.	-0.354 (0.252) [8,112] ⟨3.037⟩	-0.283 (0.262) [6,626] ⟨3.148⟩	4.186** (1.658) [490] ⟨2.052⟩
E	Over the past few years, blacks have gotten less than they deserve.	-0.083 (0.237) [8,120] ⟨2.634⟩	0.122 (0.267) [6,633] ⟨2.470⟩	-1.680 (1.233) [490] ⟨4.043⟩
Other race relations questions				
F	Racial problems in the U.S. are rare, isolated situations.	-0.039 (0.148) [26,137] ⟨2.366⟩	-0.098 (0.166) [21,046] ⟨2.445⟩	0.169 (0.605) [2,245] ⟨1.650⟩
G	I am angry that racism exists.	0.145 (0.191) [7,945] ⟨4.256⟩	-0.027 (0.231) [6,290] ⟨4.212⟩	-0.143 (0.691) [560] ⟨4.663⟩
H	I often find myself fearful of people of other races.	0.079 (0.248) [7,937] ⟨2.072⟩	-0.043 (0.251) [6,281] ⟨2.088⟩	-0.373 (1.489) [562] ⟨1.981⟩
Panel b: 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?	0.045 (0.037) [17,934] ⟨0.196⟩	0.010 (0.045) [12,953] ⟨0.171⟩	0.407** (0.180) [2,416] ⟨0.289⟩
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?	0.045 (0.041) [17,985] ⟨0.237⟩	0.025 (0.053) [12,901] ⟨0.229⟩	-0.145 (0.170) [2,483] ⟨0.278⟩
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?	-0.017 (0.019) [18,091] ⟨0.032⟩	-0.022 (0.020) [13,023] ⟨0.029⟩	-0.006 (0.065) [2,473] ⟨0.040⟩
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?	0.092** (0.036) [17,905] ⟨0.150⟩	0.085* (0.046) [12,914] ⟨0.151⟩	0.177 (0.140) [2,405] ⟨0.139⟩
	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, estimated using the same specification as Kearney and Levine (2019), with $(county)$ 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 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. Numbers in carrots report the mean of the dependent variable. 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 A14: No evidence that *Sesame Street* changed reported voting for incumbents

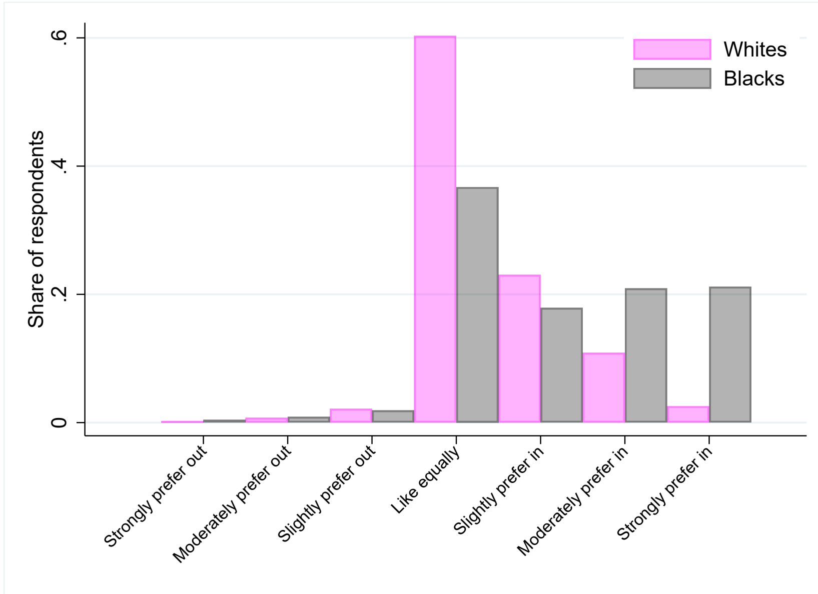
	(1)	(2)	(3)	(4)	(5)	(6)
					Major party voters	
	Share	$\hat{\beta}_1$	$\hat{\beta}_1^{KL}$	$\hat{\beta}_1^{Edu}$	**All**	**Validated**
Incumbent	0.495	0.087* (0.044)	0.087** (0.043)	0.086* (0.044)	0.008 (0.050)	-0.044 (0.057)
Non-incumbent	0.313	0.051 (0.040)	0.041 (0.039)	0.053 (0.041)		
Third party	0.025	-0.029** (0.012)	-0.023* (0.013)	-0.032*** (0.012)		
Not voting	0.167	-0.108*** (0.034)	-0.105*** (0.036)	-0.107*** (0.033)		
N		[43,968]	[48,252]	[39,613]	[34,623]	[24,615]
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	Yes	Yes
FE: State x cohort x year		Yes	No	Yes	Yes	Yes
Controls: Educ., family income		No	No	Yes	No	No

Note: Each coefficient is the result of a separate regression. All regressions control for the race and gender of respondents. $\hat{\beta}_1$ estimates are estimated using our preferred specification presented in 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. $\hat{\beta}_1^{Edu}$ estimates are estimated using out preferred specification, with the addition of controls for education and family income. 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.

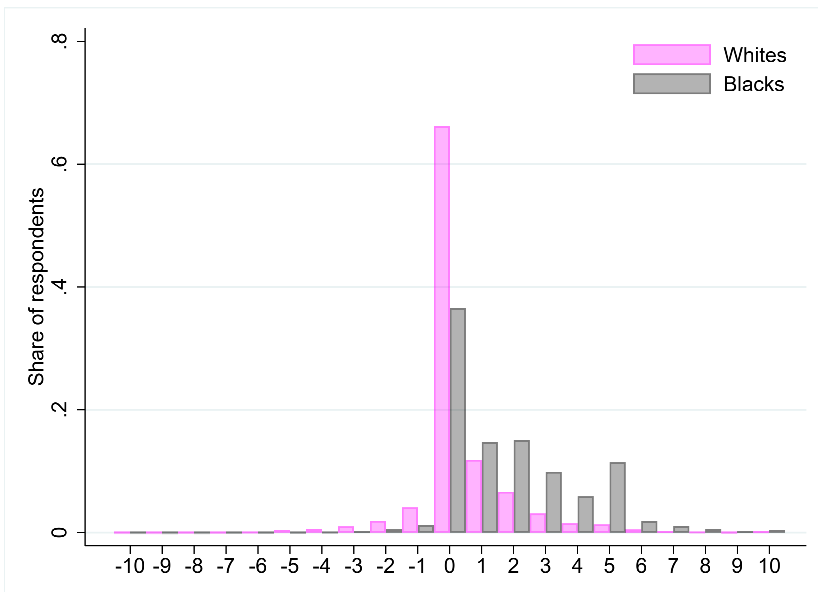
Table A15: Within the selected gender and career IAT data, there is no evidence that *Sesame Street* changed measures of gender and career biases

	(1)	(2)	(3)	(4)
Sample	**Mean score**	$\hat{\beta}_1^{KL}$		
All test takers	0.107	0.012 (0.047) [83,687]	0.016 (0.047) [83,687]	0.046 (0.060) [76,329]
Female test takers	0.210	-0.023 (0.059) [54,826]	-0.012 (0.059) [54,826]	0.022 (0.077) [47,638]
Male test takers	-0.089	0.091 (0.096) [28,280]	0.092 (0.096) [28,280]	-0.023 (0.142) [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 on standardized IAT scores. Larger positive values indicate a 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.



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



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

Figure A2. Explicit racial preferences between European and African Americans

Note: Figure a presents responses to questions on racial preferences asked in the IAT survey. Responses are flipped for African Americans for comparability. Figure b plots the net difference between IAT test takers' rated warmth (0-coldest to 10-warmest) towards their in-group and 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.