

From Lawyer to Judge: Advancement, Sex, and Name-Calling

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Working Paper

January 25, 2009

Abstract: This paper provides the first empirical test of the Portia Hypothesis: females with masculine monikers are more successful in legal careers. Utilizing South Carolina microdata, we look for correlation between an individual's advancement to a judgeship and his/her name's masculinity, which we construct from the joint empirical distribution of names and gender in the state's entire population of registered voters. We find robust evidence that nominally masculine females are favored over other females. Hence, our results support the Portia Hypothesis.

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1. Introduction

In Shakespeare's *Merchant of Venice*, a woman named Portia masquerades as a man in order to argue before the court as an attorney.¹ Indeed, for centuries the only way a woman could have practiced law was incognito because the courtroom was a domain reserved exclusively for men (Albisetti 2000). A notable exception on record is Miss Margaret Brent, circa 1640, who was permitted by Lord Baltimore to practice law as a woman; nonetheless, she was still addressed as "Gentleman Margaret Brent" during her several dozen appearances in the Maryland colonial court (Cup 2003).²

Most jurisdictions in the western world refused to admit women to the bar before World War I (Albisetti 2000). By the end of the nineteenth century, any woman attempting to practice law was labeled Portia, as was the first school established exclusively for the legal education of women.³ The first Portia to be admitted to the South Carolina bar was Miss James (Jim) Margrave Perry in 1918 (Cup 2003). Although women no longer needed a male disguise to practice law, a male persona or male moniker still might have helped.

¹ In Act IV, scene i, Portia's true love has failed to pay off his debts to a creditor who seeks retribution in the form of a pound of flesh. By coupling her famous plea for mercy with a slippery legal argument, Portia's persuasion wins over the judge.

² Miss Margaret Brent was not a licensed attorney, despite arguing cases before the court and her relationship to Lord Baltimore (she was both his kin and counselor).

³ The Portia Law School, established in Boston in 1908, is now the Northeastern School of Law.

Despite the fact that women made up half of the students graduating from law school in the past 15 years, the legal profession remains a male-dominated world (Harrington and Hsi 2007). Consequentially, one would suspect that having a male persona or male moniker might still be advantageous to a career in law. We dub this the Portia hypothesis: females with masculine names are more successful in legal careers than females with feminine names. The purpose of this paper is to conduct the first empirical test of the Portia hypothesis, using data from South Carolina.

We have good reason to expect to find the Portia hypothesis holding in our data. The first female lawyer in South Carolina had a masculine name and today many female lawyers privately express their belief that their nominal masculinity matters. Anecdotally, the legal profession remains one of the last bastions of the “good old boy network”, particularly in South Carolina. Even in Massachusetts – a state that is often viewed as less traditional than South Carolina – females comprise a small minority of all partners in law firms (Harrington and Hsi 2007). Just as precedent-bound law changes slowly, the legal profession is notoriously slow to embrace change.⁴ On the other hand, females are a protected class under the Equal Protection Clause of the US Constitution’s 14th Amendment, and no one should understand (and, arguably, respect) that better than lawyers and judges.

⁴ Witness the anachronistic attire of barristers throughout the Commonwealth, sluggish technological adoption, or the rarity of structuring law firm ownership in a form other than partnerships.

Yet judicial positions turn over rarely, some even being held for life, so that the equal status for women may not yet have propagated into the upper echelons of the legal profession.

Several different mechanisms could be at work to make the Portia hypothesis hold in the data. A lawyer's gender could explicitly matter for advancement to some decision makers; for example, some judicial positions are determined by popular election, and the electorate (or sufficiently large subset of it) could categorically prefer men to women. If nothing else were known about an individual besides that individual's name, the name itself could contain information on the gender of the individual, just as a name contains information on the race of an individual (Fryer and Levitt, 2004). Just as with the racial discrimination on call-backs for resumes submitted in job applications, individuals may be more likely to get into the pool of candidates receiving serious consideration for the sorts of positions that lead to potential judgeships, i.e. getting their "foot in the door", when they have a male moniker. Alternatively, nominal masculinity might matter when opinions are formed about a lawyer's work, not face-to-face, but through the written word, such as through briefs or publications in law journals. If there is some gender bias in the citation process – that is, if authors are generally more likely to cite a writer with a masculine name than with a feminine name – then we might observe female lawyers with masculine names receiving more citations than female lawyers with feminine

names, *ceteris paribus*, and having relatively fewer citations could affect career outcomes.⁵ The mechanism could be even subtler yet. There could be a subconscious preference for male names, even when the gender is known; jurists, clients, superiors, professors, legislators, etc., might just feel more comfortable with a woman called “George” than one called “Barbara”; in the context of the good old boy network, a woman with a male moniker might just feel more like “one of the boys”.⁶ Finally, it could just be that the parents who successfully nurture a girl’s ability are the same people who believe that bestowing a child with a masculine name would be advantageous in her future career path.

In this paper, we use the frequency of names and genders of all registered voters in South Carolina to construct a measure of nominal masculinity and assign this measure to each member of the South Carolina bar. Examining the correlation between a lawyer’s advancement to a judgeship and his/her name’s masculinity, we find that nominally masculine names appear to be favored over nominally feminine names. This could be due to the Portia Hypothesis. Alternatively, the correlation between attaining judgeship and masculine names could also arise from the fact that most judges are males, who tend to have more masculine names

⁵ Compared to advancing to a judgeship, citations probably do not matter nearly as much in advancing to partnership within a law firm.

⁶ During a focus group session, a female judge in South Carolina pointed out that when the good old boy network gathers for bonding activities that blur the line between professional and social interactions, wives are often not around. Husbands may prefer that the females involved in such activities have male monikers so that their wives are not suspicious of extra-marital affairs occurring on these trips. Going on a trip that combines business with some pleasure (e.g. fishing, hunting, golfing, football, or gambling) may just be easier with a “Jim” than an “Emily”.

(by definition). Because we do not observe the gender of South Carolina bar members, we are unable to control for male domination of the judiciary with that data source.

To separate these two possible causes of correlation between nominal masculinity and judgeship, we combine data on the names and genders of the entire population of registered voters in South Carolina with the publicly available names and genders of judges. Controlling for gender, we find a significant correlation between nominal masculinity and judgeship, supporting the Portia Hypothesis. A series of robustness checks confirm the Portia Hypothesis.

The remainder of the paper proceeds as follows. Section 2 gives background information and reviews the existing literature on the topic of race and gender bias in labor markets. Section 3 presents our data. Section 4 details our econometric model. Section 5 presents our results and checks their robustness. Section 6 concludes.

2. Background

The question of whether gender or race affects career success has been addressed in many scenarios. Since Becker (1957), most literature on the subject has focused on testing for and explaining the wage gap between, for example, equally productive male and female workers or equally qualified black and white workers. Of course, empirical determination of whether two individuals are equally productive potentially suffers from measurement error and omitted

variables. A more recently developed branch of labor literature circumvents the fact that employers have more information about employee characteristics than researchers by examining the outcomes of employment applications from different racial groups. Employment applications, even if they do not specifically state race or gender, contain the names of the applicants and therefore potentially transmit information about the applicants' race or gender. We focus our literature review on papers centered on names and career success.

Whether race or gender affects labor market opportunities remains an uncertain empirical question. Bertrand and Mullainathan (2004) report that, in correspondence testing, resumes with "black" names had a significantly lower callback rate than resumes with "white" names. Conversely, the aforementioned Levitt and Fryer (2004) study finds no significant difference between black and white names when controlling for circumstances at birth. Arai and Thoursie (2007) examine whether immigrants in Sweden who change their names to Swedish-sounding or neutral names are paid more than immigrants who retain non-Swedish sounding names; their results indicate that there is a statistically significant wage gap between the name-changers and the name-retainers after the name change but not before.

Tregenza (2002) and Budden et al. (2007) have examined a potential gender bias in the refereeing process for academic publications. Tregenza (2002) compared publication success rates for papers with male first authors and papers

with female first authors at five different ecology and evolution research journals. Editors of these journals noted gender of submitters as well as whether the paper was accepted.⁷ This study found no significant difference in overall acceptance rates between male and female submitters, although one of the five journals did, when analyzed individually, show some signs of a bias against female submitters. Budden et al. (2007) exploit the changes in policy at a primary research journal in 2001 to switch from a single-blind to double-blind review process. They find that switching to the double-blind process, which withholds the authors' names and therefore any information regarding gender from the referee, increases female authors' acceptance rates. As in Tregenza (2002), the gender of the submitting author was determined by journal editors' interpretation of the author's first name.

The method used in each of these papers of determining what information is contained in a name differs. The majority (Tregenza (2002), Budden et al. (2007), Arai and Thoursie (2007)) rely on either the author's discretion or a journal editor's discretion to decide what information (e.g., race or gender) a name conveys. Both Fryer and Levitt (2004) and Bertrand and Mullainathan (2004) rely on frequencies of names and race according to birth certificates registered in California and Massachusetts, respectively. Fryer and Levitt construct a "black name index," which is essentially the ratio of black babies born

⁷ The gender of the submitter was decided upon by the editor of each journal based on the first name of the first author. If the editor could not readily decide the gender based on first name, the gender was classified as unknown.

with a given name to the total number of babies with that name, multiplied by 100. Similarly, Bertrand and Mullainathan use the ratio of black babies with a given name to the total number of babies with that name to construct lists of names that are distinctly black and distinctly white.

We follow a method similar to Fryer and Levitt (2004) and Bertrand and Mullainathan (2004) in that we allow our data to dictate the information about gender that is conveyed by an individual's name. For example, there are 8,640 registered voters in South Carolina with the name "Carolyn." 8,615 are female, and the remaining 25 are male. We conclude, based on the data, that Carolyn is a rather feminine name; precisely, we calculate the masculinity of the name Carolyn to be 25 out of 8640, or 0.0029. A name that has only males registered (e.g., there are 118 voters named "Al," and they are all male) has a nominal masculinity of 1, and a name that has only females (e.g., all 246 voters named "Deidre" are female) has a nominal masculinity of 0. There are many names that convey relatively little information about gender, such as the name "Kerry," which has a maleness of 0.502. Tables 1 and 2 list the most masculine female names and the least masculine male names, according to frequency and gender, in our dataset.⁸

⁸ At the request of an anonymous referee, we perform a check on our nominal masculinity measure by examining the score for Bacon Magazine's "Top 10 Stripper Names". In theory, female exotic dancers choose hyper-feminized stage names. Only two of those names, Candy and Porsche, had a nominal masculinity of 0. Three other names had nominal masculinity names below the mean female voter. Two other names on the list actually scored quite high in nominal masculinity; Angel had a nominal masculinity of 0.15 (due to its popularity among Spanish speakers as a boy's name) and Houston had a nominal masculinity of 0.98. These findings suggest the potential for further research, which is beyond the scope of this paper.

Our method fails to account for names with closely related phonemes but distinct spellings. For instance, “Jean” and “Gene” are homophones but the differences in their spelling communicate a strong signal about gender. Other researchers have found that masculine names tend to use distinct sounds from feminine names—see Cutler, McQueen, and Robinson (1990) or Cassidy, Kelly, and Sharoni (1999). In general, male names tend to have stronger initial syllables and are less likely to be polysyllabic. If the primary causal mechanism works on the appearance of a name in text, then ignoring homophones is appropriate. In contrast, if the mechanism works on the sounding out a name in speech, then homophones should not be ignored. Had we chosen not to ignore homophones, a close examination of the names of the female judges in our data suggests that our findings would be even stronger.⁹

3. Data

The data come from several sources. First, we secured South Carolina’s voter registration dataset, which contains the first name and gender of every registered voter in the state. In instances where a voter registered with an initial for his/her first name and a non-abbreviated middle name, we assumed the middle name to be the primary name used. This resulted in a dataset of 2,243,980 registered voters. From these data, we constructed a list of 86,642 different names that are registered in South Carolina. Each unique name was associated with the count of

⁹ Indeed, a woman named “Jean” is the Chief Justice of the South Carolina Supreme Court.

the number of registered males with that name, the count of the number of registered females with that name, and the total count of registered voters with that name. These data were used to construct the subjective probability of gender conditional upon name, which we term nominal masculinity. Functionally, as described in Section 2, nominal masculinity is the ratio of the number of males with a given name over the total number of individuals with the same name. Summary statistics of the registered voter data are given in Table 3. The names and nominal masculinity scores of female judges in South Carolina are given in Table 4, as well as the means for the judges and voters as reference points.

Second, data on South Carolina bar membership were gathered from their website, <http://www.sctbar.org>. The gender of each bar member was unavailable. Summary statistics of bar membership data are given in Table 5.

Third, we gathered data on South Carolina judges from the state and federal judiciary's websites, <http://www.judicial.state.sc.us/> and <http://www.scd.uscourts.gov/> respectively. The gender of federal district court judges was available at <http://www.fjc.gov/>. For other judges, we determined the gender of each judge by examining each judge's digital photograph or interviewing judicial clerks.¹⁰ As of November 2007, there were 156 male judges and 52 female judges in South Carolina. Summary statistics of judicial membership data are given in Table 3 alongside the registered voter data

¹⁰ None of the judges' digital photographs appeared androgynous to us.

statistics. Part of this paper includes analyses of the distribution of the nominal masculinity of female voters and the distribution of nominal masculinity of female judges, which we thus describe further here. The nominal masculinity at the 15th percentile for all registered female voters is 0, while the nominal masculinity of the female judges at the 15th percentile is quite close to zero at 0.001. At the 85th percentile, the nominal masculinity of female voters is 0.007, and that of the female judges is 0.105.

Figure 1 displays a kernel density for nominal masculinity using the data whose summary statistics are displayed in Table 3. It is clear that most males have very masculine names and most females have very feminine names. In this sense, the distributions look very similar for both voters and judges. Indeed, with the shape of these densities, it is difficult to discern significant difference simply by inspection. The voter densities peak at their respective poles (0 for females, 1 for males), quickly and smoothly declining away from that pole. The judge data is similar except for two features. One, the densities peak just short of their respective poles. Two, there is bumpy decline away from the poles, due to individual observations standing out as blips in a small sample.

4. Model

We begin by modeling the data generating process. At birth, each individual receives a random draw for gender and ability. Then, depending upon their

gender, an individual draws his/her name.¹¹ We assume that the only aspect of a name that matters, for the purpose of our analysis, is its nominal masculinity. We define nominal masculinity to be the rational belief over gender given name:

$$\text{Nominal Masculinity} = \Pr(\text{♂}|\text{Name})$$

♂ is an indicator variable equal to 1 when male and 0 when female. With enough names for nominal masculinity to be sufficiently dense in its unit interval domain, an individual's name serves as his/her draw from a continuous distribution of nominal masculinity. Following the arguments of nonparametrics, the best way to construct this subjective probability of gender conditional upon name is to use the empirical distribution:

$$\Pr(\text{♂}|\text{Name}) = [\sum 1\{\text{♂}, \text{Name}\}] / [\sum 1\{\text{Name}\}]$$

Naturally, we need a very large data set with name and gender jointly observed; we found voter registration data to be the largest data set available.¹² To use voter registration data, we make an important assumption: within a given state (South Carolina, in our data), the distribution of names conditional on gender (hence nominal masculinity) is the same for registered voters, lawyers, and judges.

¹¹ This implicitly rules out the name depending upon the ability draw. See the results section for a discussion of the possibility that parents know (or have informative *a priori* beliefs about) their child's ability and name it accordingly. We aren't as concerned with amniocentesis as we are with serial persistence of (hereditary) ability across generations within a dynasty, where the high types would signal their class with their choice of name.

¹² The sample size of this dataset is several orders of magnitude larger than the data we use for the remainder of the analysis. Hence, even if one thought that it was more appropriate to consider these beliefs as estimates with variance, the standard errors would be negligible. See Imbens and Lancaster (1994).

Success in a legal career depends certainly upon ability and possibly upon luck; it may also depend upon gender or nominal masculinity. If success depends upon gender and gender were known with certainty, then we would expect a correlation between success and gender but no (partial) correlation between success and nominal masculinity having controlled for gender. If gender were unknown and success favored men, then we would expect a positive correlation between nominal masculinity and success; if gender did not matter, then we would expect no correlation between nominal masculinity and success. This testable implication forms the basis of our first regression, discussed in section 5 and reported in Table 6, using the data from the bar membership directory (which lacks gender):

$$\Pr(\text{♂} | \text{Name}_i) = \beta_0 + \beta_1 \mathbb{1}\{\text{Judge}_i\} + \varepsilon_i$$

β_0 is the mean nominal masculinity for all members of the bar and β_1 is the change in that mean when we condition upon a judicial membership. A significantly positive β_1 implies that nominal masculinity matters for success, defined as attaining a judgeship. We could interpret a positive β_1 as nominally masculine monikers improving a female attorney's chances of success. However, we would find the same correlation if gender was known by decision-makers (but unobserved to the econometrician), males were favored in success, and males tended to have more nominally masculine names.

To construct an unbiased estimate of the effect of nominal masculinity on the success of female attorneys, we need to observe and control for gender. To accomplish this, we broaden the scope of analysis from members of the bar to all registered voters and from judicial bar members to all judges.¹³ Simple regressions on this larger data set produced estimates of the parameters in variants of the following equation:

$$\Pr(\hat{O} | \text{Name}_i) = \beta_0 + \beta_1 1\{\text{Judge}_i\} + \beta_2 1\{\hat{O}_i\} + \beta_3 1\{\hat{O}_i\} \times 1\{\text{Judge}_i\} + \varepsilon_i$$

Table 7 presents the results from estimating this equation as it appears above, as well as a regression where β_3 is constrained to be 0. These results are discussed in section 5.

Because the number of female judges in the data is small, the asymptotic justification for conducting standard hypothesis tests is suspect. If the disturbance term were distributed normally, then conducting a t-test of the difference in means between female judges and female voters is still valid. However, the kernel density of nominal masculinity, shown in Figure 1, clearly reveals strong non-normality, implying that the disturbance term is also non-normal.

To address concerns over inference validity when using conventional t-tests, we perform a couple of non-parametric tests of the equality of the distribution of nominal masculinity conditioning on gender and judgeship. To compare the

¹³ Judges for some lower courts, like the Probate and Magistrate Courts, are not required to be members of the bar.

difference in the medians between the two distributions, we employ the Kruskal-Wallis test; we also examine a Kolmogorov-Smirnov test for the (maximum) difference in the CDFs across the two distributions. The results of these nonparametric tests are reported in Table 8. We also bootstrap the distribution of the estimated mean nominal masculinity for female judges, with the results displayed in Figure 2. All of these robustness checks are described in further detail in section 5.

5. Results

Using the data described in Section 3, we conducted a test of the hypothesis that females with relatively masculine names are more likely to succeed in the legal profession than females with relatively feminine names. Nominal masculinity is defined as the number of registered male voters with a given name divided by the total number of registered voters with that name. We constructed this measure using a database of all registered South Carolina voters. To determine whether nominal masculinity plays a role in a lawyer rising to the judgeship level, we ran an ordinary least squares regression of nominal masculinity on judgeship for all SC Bar members, where judgeship equals unity if the bar member is a judge and equals zero otherwise. The results, as reported in Table 6, indicate that the mean nominal masculinity for an attorney is 0.71 and the mean nominal masculinity for a judge is higher at 0.76. That difference is significant at the 6% level. This difference could be driven by the Portia

Hypothesis, or it may just be due to male judges, who have relatively masculine names, outnumbering female judges three to one. To disentangle these two causes, we need to observe and control for gender.

Table 7 reports the results for the regression analysis when we control for gender using the voter registration data. Note that the inclusion of gender produced a sizeable increase in the goodness of fit, implying that the gender informational content signaled by a name's masculinity is quite high relative to the noise. When the effect of conditioning on a judgeship is held to be the same across genders (the first regression presented in Table 5), judges do not appear to be significantly more nominally masculine; this is due to the much greater number of male judges swamping the effect for the females.

Column 2 of Table 7 shows that the average nominal masculinity for a female without a judgeship is slightly above 0.02, while female judges have a mean nominal masculinity of just above 0.08.¹⁴ The difference between the nominal masculinity of female voters and female judges is significant at the 1% level, providing further support of the Portia hypothesis. In contrast, the average nominal masculinity for males, whether they are judges or not, hovers just above 0.94. This possibly indicates a sort of diminishing return to nominal masculinity: a marginal increase in nominal masculinity above 0.94 yields little additional information about gender.

¹⁴ Although the absolute difference in these means may seem small, the percentage difference is

The significant relationship that we have found may not readily lend itself to a familiar interpretation. In order to ease interpretation, we use Bayes' rule to construct an estimate of the effect of bestowing a girl with a more masculine name on the probability of attaining a judgeship:

$$\Pr(\text{Judge}|\text{Name}, \varphi) = \Pr(\text{Judge}|\varphi) \text{pdf}(\text{Name}|\text{Judge}, \varphi) / \text{pdf}(\text{Name}|\varphi)$$

Where φ is an indicator variable equaling 1 if female and 0 otherwise.

Nonparametrically estimating these probabilities is straightforward. The $\Pr(\text{Judge}|\varphi)$ term can be estimated with the count of female judges divided by the count of females in the voting population and the remainder of the right hand side is simply the ratio of the kernel densities for female judges and voters presented in Figure 1.¹⁵ The results are quite informative. The probability of any woman becoming a judge, regardless of her name, is virtually the same as the probability that a woman with an extremely feminine name becomes a judge. Changing a girl's name from something fairly feminine, like "Sue" (which is less masculine than the mean female voter's name), to a more gender-ambiguous name, like "Kelly", increases her probability of becoming a judge by roughly 5%.¹⁶ This effect may appear small, but it is highly nonlinear in nominal masculinity; changing a girl's name from "Sue" to a predominantly male name, like

¹⁵ Figure 1 selects a different bandwidth for each density estimated. For the sake of making a fairer comparison, we used the same (larger) bandwidth in performing this calculation.

¹⁶ To put this finding in context, consider the findings of Pelham, Mirenberg, and Jones (2002). They find that if a subject's name starts with the same phoneme as the name of a profession (eg "Denny" and "Dentist"), then the subject's probability of choosing that profession increases by 15%.

“Cameron” (75% of those named “Cameron” in South Carolina’s voting population are male), increases her probability of becoming a judge by a factor of 3 (roughly). Moreover, changing a girl’s name from “Sue” to “Bruce” (99% of those named “Bruce” in South Carolina’s voting population are males) increases her probability of becoming a judge by a factor of 5 (roughly).

The robustness checks in the Kruskal-Wallis and Kolmogorov-Smirnov tests, shown in Table 8, support the findings of the differences we found in the conditional means using regression analysis. The median nominal masculinity of judges is significantly higher than voters; comparing their CDFs reveals a significant difference in those populations. The significant discrepancy in distributions disappears when we condition upon the male gender, implying that the significant difference we found before is at least partially due to the overrepresentation of males in the judiciary. Conditioning upon the female gender yields a difference in medians (significant at the 2% level) that is qualitatively similar to the difference in means. Likewise, the Kolmogorov-Smirnov test finds that the greatest difference in the CDFs is statistically significant (at the 5% level).

To further address concerns over this small sample, we bootstrap the distribution of estimated mean nominal masculinity for females. In theory, bootstrapping can provide more accurate confidence intervals when a medium-

sized sample is drawn from an extremely non-normal distribution.¹⁷ The results of 10,000 bootstrap replications are presented as a kernel density in Figure 2. Note that there is little overlap in the densities for female voters and female judges. The 99% confidence interval on the mean nominal masculinity of female voters spans about 1% of the probability mass for female judges. About 2% of that probability mass lies below the confidence interval and 97% lies above it. Hence, we are 97% confident that the mean nominal masculinity for female judges is greater than the mean nominal masculinity of female voters.

One might be concerned with changing in the popularity of names over time. To score the nominal masculinity of our judges we have used voters of all ages. For 21 of the 52 female judges, age is publically available and ranges from 34 to 69 with a mean of 51.5 and a standard deviation of 9.25. This could present a problem if those under age 30 or over 70 have a systematically different association of names and genders (e.g., if the name “Bruce” is given to just a few females between ages 30 and 70 but no males and it is given predominantly to males for those under 30 or over 70, then we would be improperly scoring it as masculine). To gauge this potential problem, we compute a weighted nominal

¹⁷ The bootstrap’s small-sample properties are an area of ongoing research with some promising preliminary findings. To our knowledge, only Bayesian estimation is completely robust to small sample sizes. We performed a Bayesian estimation assuming that the female judges’ nominal masculinity measures were draws from a Beta distribution. With flat priors on the mean parameter (as well as on the spread parameter), we find even stronger support of the Portia Hypothesis. However, such results are susceptible to misspecification arguments and have been omitted here (they are available from the authors upon request).

masculinity score with voters whose ages are closest to the female judges receiving the highest weight.¹⁸ This age-weighted nominal masculinity measure would only provide a mild strengthening of our results because the age-weighted nominal masculinity measure is virtually indistinguishable from its unweighted counterpart, with a 0.998 correlation, as can be seen in Figure 3.¹⁹

We have discussed many different causal mechanisms for our finding that female judges are more nominally masculine than voters. The explanation favored by many who have reviewed our work is that there is a common cause: wealthier families give their female children “stronger” (i.e. gender-neutral) names and a daughter of a wealthier family is more likely to become a judge.²⁰ Unfortunately, we do not observe wealth in our data sets. However, we can still perform some exploratory analysis by linking the voting registration data to income data from the Census Bureau’s 2005-2007 American Community Survey via geography. We link at the county level because that is the finest level of geographic resolution in the data available to us. If wealth is the common cause, then we might expect to find a positive correlation between a county’s median income and the mean nominal masculinity of its female voters. Figure 4 shows a scatterplot of that data, with an insignificant negative correlation (p-value is 0.46). If wealth is a common

¹⁸ We used the normal density as the weighting function with mean and standard deviation equal to the empirical density of the sample moments of the observed ages for female judges.

¹⁹ Weighting by age actually raises the mean nominal masculinity of judges to 0.093.

²⁰ Most lawyers aware of our research have actually favored a more direct explanation of gender discrimination.

cause, then county-level linking on income is simply too noisy and coarse to reveal such a relationship. Future research should delve into this potential causal mechanism with more appropriate data.

6. Conclusions

The Portia Hypothesis tested in this paper is that females with masculine names fare better in legal careers than females with feminine names. Using data containing the first name and gender of all registered voters in South Carolina, we create a measure of nominal masculinity. We test whether nominal masculinity is a significant determinant of achieving judicial status by combining the nominal masculinity measure with data on all SC Bar Association members. Each Bar member's name is associated with the nominal masculinity of that name as determined by the registered voter data. The SC Bar data indicate that nominal masculinity is significantly higher in the judiciary; however, these data cannot pin down whether the higher nominal masculinity in members with judicial status is due to the Portia Hypothesis or an overrepresentation of males. Expanding the scope to voters allows us to controlling for gender, and our results support the Portia hypothesis. This result seems to be robust across different tests constructed from different assumptions.

Our work represents the first empirical test of the Portia Hypothesis. Although we find robust support of the Portia Hypothesis, our data do not allow us to isolate the source of this outcome asymmetry across genders. In future work,

we hope to collect additional data on women in the legal profession to expand our understanding of this phenomenon. Ultimately, this line of research could answer the question of exactly why and how males dominate leadership positions in the legal profession. Possible explanations abound, ranging from gender bias to lifestyle choices. This paper shows that gender bias might indeed play a role in the advancement of women to judicial positions, but we caution that this evidence should not be taken as conclusive.

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Table 1: Top 25 Most Masculine Female Names

(given to at least 100 Females)

Name	Females Observed	Males Observed	Nominal Masculinity
JAMES	113	51320	0.998
JOHN	102	38730	0.997
MICHAEL	130	28066	0.995
BOBBY	104	4455	0.977
JERRY	140	5967	0.977
EDDIE	114	1819	0.941
FRANCIS	148	1664	0.918
CHRIS	113	1131	0.909
FREDDIE	122	973	0.889
CARROLL	102	697	0.872
BENNIE	120	675	0.849
JIMMIE	234	942	0.801
SHAWN	358	1392	0.795
CHRISTIAN	118	433	0.786
TERRY	1138	3900	0.774
DALE	453	1286	0.740
ADRIAN	193	527	0.732
LEE	601	1571	0.723
JORDAN	107	249	0.699
MARION	926	2068	0.691
TOMMIE	173	303	0.637
CAREY	146	229	0.611
JOHNNIE	781	1118	0.589
JODY	302	386	0.561
KERRY	341	344	0.502

Table 2: Top 25 Least Masculine Male Names

(given to at least 100 Males)

Name	Females Observed	Males Observed	Nominal Masculinity
CAROL	6906	144	0.020
ASHLEY	4405	307	0.065
LAURIE	1260	100	0.074
HAZEL	1683	145	0.079
ROBIN	3782	396	0.095
KELLY	3906	442	0.102
COURTNEY	1281	149	0.104
LYNN	2395	298	0.111
STACEY	1549	204	0.116
KIM	1563	213	0.120
SHANNON	3077	461	0.130
BILLIE	910	146	0.138
DANA	2025	325	0.138
ANGEL	629	108	0.147
STACY	1477	287	0.163
TRACY	3040	691	0.185
JAN	697	169	0.195
LESLIE	2494	678	0.214
SANDY	485	132	0.214
JAIME	339	101	0.230
OLLIE	320	125	0.281
JAMIE	1635	819	0.334
JESSIE	1269	683	0.350
MORGAN	251	161	0.391
CASEY	470	309	0.397

Table 3: Summary Statistics of maleness for data taken from the SC Judiciary

Database and Registered Voters Database

Group	Obs.	Mean (Nominal Masculinity)	Std Dev (Nominal Masculinity)	Min (Nominal Masculinity)	Max (Nominal Masculinity)
Female Judges	52	0.0840887	0.2354271	0	0.993311
Female Voters	1246881	0.0256224	0.1112169	0	0.999262
Male Judges	156	0.9651312	0.1181546	0.0021987	1
Male Voters	999519	0.9680387	0.121208	0.0004812	1

Table 4: Nominal Masculinity for the Female Judges of South Carolina

	Maleness		Maleness
BRUCE	0.993311	GEORGIA	0.002634
BARNEY	0.97619	DIANE	0.002542
CAMERON	0.762533	MARGARET (x2)	0.002541
DALE	0.739356	JANE	0.002272
LESLIE	0.213745	ANNE	0.002226
JAN	0.19515	DEBORAH	0.002184
RUDELL	0.166667	JUDY	0.002104
KELLY	0.101656	AMY	0.002076
MEAN FEMALE JUDGE	0.084089	HARRIETT	0.002045
CARMEN	0.05421	DONNA (x2)	0.002
ROCHELLE	0.028037	SANDRA (x2)	0.001977
MEAN FEMALE VOTER	0.025622	MARTHA	0.001882
JEAN	0.018872	PAMELA	0.001837
KAY	0.015668	CATHERINE	0.001788
JOCELYN	0.00995	KATHY	0.00162
ALISON	0.008141	PAULA	0.001463
DEBORA	0.007246	BRENDA	0.001325
KAYE (x2)	0.006494	SHEILA	0.000992
TIFFANY	0.004341	APHRODITE	0
FRANCES	0.003953	BRANNA	0
SUE	0.00342	DEADRA	0
LOIS	0.003178	DEIRDRE	0
DOROTHY	0.003064	PONDA	0
NANCY (x2)	0.00305	PANDORA	0
MICHELLE	0.002811		

Note: (x2) indicates that there are two judges with that name.

Table 5: Summary Statistics for Data taken from the SC Bar Membership

Variable	Obs	Mean	Std. Dev.	Min	Max
Nominal Masculinity	8731	0.712735	0.438829	0	1
Judicial	8731	0.024281	0.15393	0	1

Table 6: OLS regression of nominal masculinity on judicial bar membership

	Nominal Masculinity
Judicial Membership to Bar	0.05668 (0.06321)*
Constant	0.71136 (0.00000)***
Observations	8731
R-squared	0.00040

Absolute value of t-statistics in parentheses

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 7: OLS regression of nominal masculinity on judicial status and gender,
using voter registration data

	(1) Maleness	(2) Maleness
Male	0.942 (6063.34)***	0.942 (6063.16)***
Judge	0.012 (1.55)	0.058 (3.64)***
MaleXJudge		-0.061 (3.31)***
Constant	0.026 (247.15)***	0.026 (247.13)***
Observations	2246608	2246608
R-squared	0.94	0.94

Absolute value of t statistics in parentheses

* significant at 10%; ** significant at 5%; *** significant at 1%

Table 8: Nonparametric Tests for Equality of the Conditional Distributions of
Nominal Masculinity

Groups Tested	Kruskal-Wallis Test* (Equality of Medians)		Kolmogorov-Smirnov Test (Equality of CDFs)	
	Test stat	p-value	Test stat	p-value
All Judges v. All Voters	67.093	0.00010	0.3183	0.00
Male Judges v. Male Voters	-0.166	1	0.0389	0.624
Female Judges v. Female Voters	5.164	0.0231	0.1695	0.05

* This test is alternatively referred to as the Mann-Whitney-Wilcoxon Test.

Figure 1: Kernel Density of Nominal Masculine Monikers

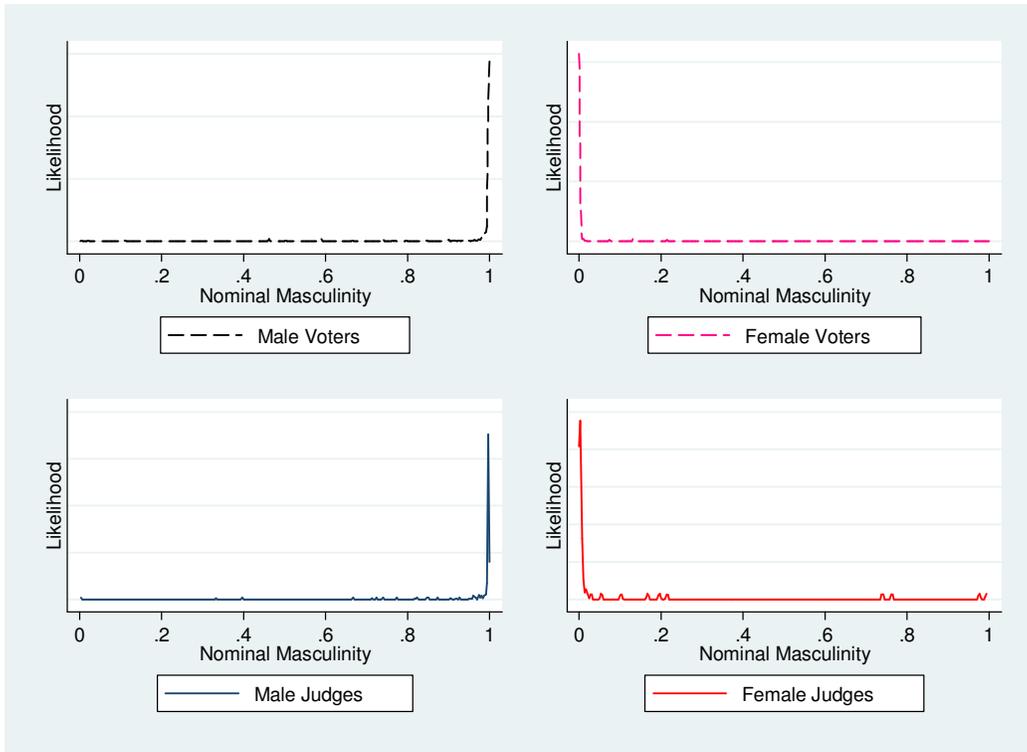
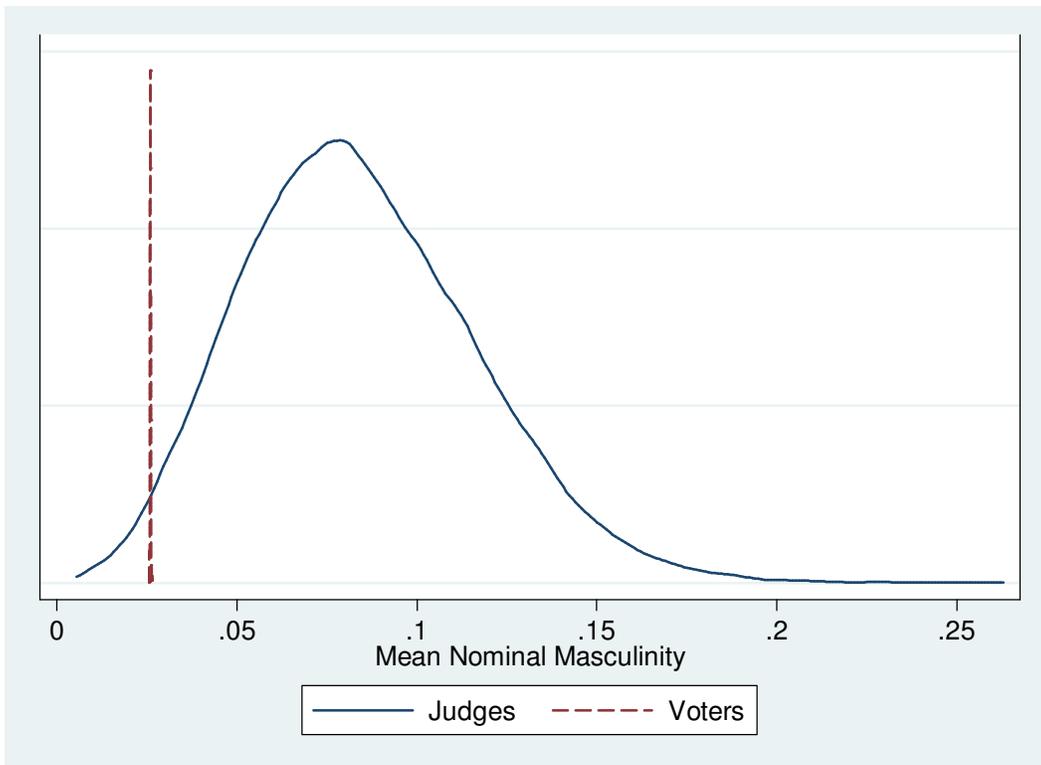


Figure 2: Kernel Density of Mean Female Nominal Masculinity
Using 10,000 Bootstrap Realizations



Note: The narrowness of the confidence interval kernel density of voters gives it the appearance vertical line in this graph.

Figure 3: Scatterplot of Nominal Masculinity for Female Judges

vs.

Nominal Masculinity Weighted by Age for Female Judges

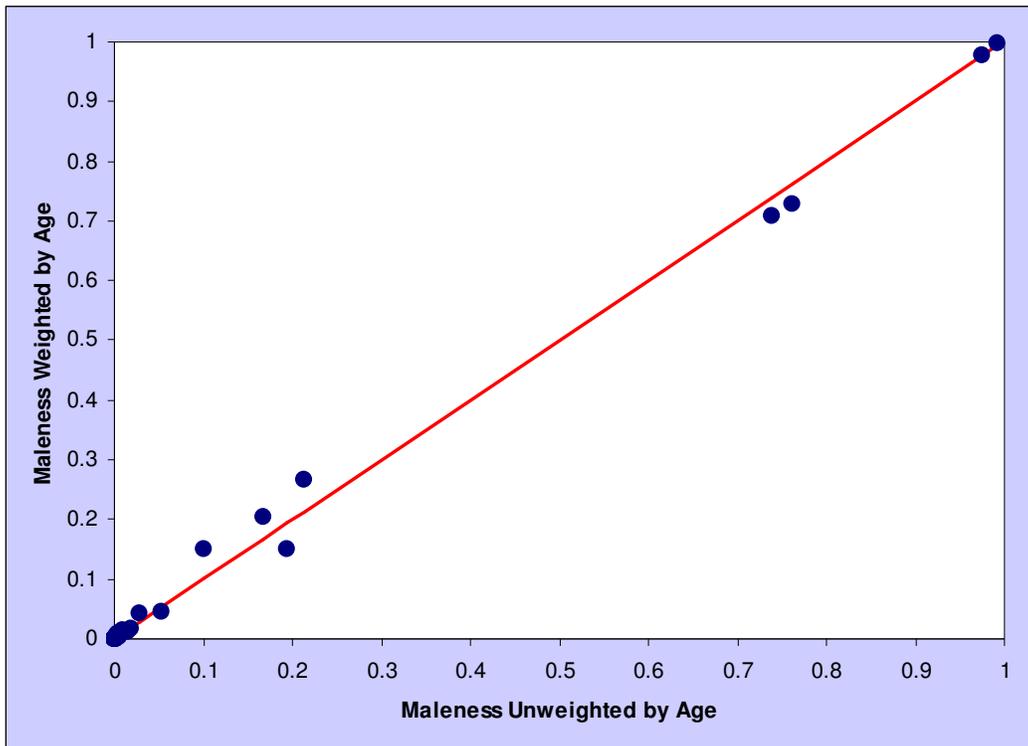


Figure 4: Scatterplot of County-level Mean Nominal Masculinity for Females

vs.

County-level Median Household Income

