Sex Differences in Risk-Taking? Evidence from Female Representation in Legislatures[*]

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History's societies exhibit considerable variation, but appear relatively homogenous in how they distribute socioeconomic status across sexes. This homogeneity begs the question of whether forces that are *robust* to formal and informal institutions influence sex differences in status. We address this question by empirically evaluating evolutionary models of hierarchical attainment where sex-differences in risk-taking play an axiomatic role. These models imply that (1) winner-take-all games favor males, but (2) successful females maintain greater skill on average. We find support for these implications in how the sex composition of national legislatures differs across electoral mechanisms (i.e., majoritarian chambers employ a significantly smaller proportion of females) and how US Representatives' re-election prospects differ by sex (i.e., females enjoy significantly longer durations). These results cannot be easily dismissed as artifacts of endogeneity bias, and alternative models can (at best) rationalize our cross-sectional *or* time series results, but not both.

JEL: B52, D72, J16, J71, J78

Keywords: Winner-take-all games, electoral rules, dominant strategies, risk preferences, female representation, affirmative action

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1. Introduction

History's societies exhibit considerable variation. Their sex-disparities in socioeconomic status, nevertheless, appear relatively homogenous. Victor Fuchs (1989, p. 26) thus noted a "substantial gender inequality" in the US, even after decades of "major antidiscrimination legislation" and "massive social change."

This persistence begs the question of whether forces that are *robust* to formal and informal institutions might contribute to the tendency for sex-disparities to replicate themselves. We address this question by evaluating the extent to which evolutionary models of sex-differences in risk-taking can explain variation in hierarchical attainment. Our results are consistent with sex-differences playing an influential (but not exclusive) role.

These results build on an understanding from evolutionary biology that winner-take-all (WTA) games can favor risky strategies. To the extent that such strategies are heritable, and reproduction games in polygamous societies let males "take all," *nature may have selected human males that embody risky strategies*. Biologists and social scientists have leveraged this insight to rationalize the regularity with which mammalian males exhibit risk-loving tendencies (e.g., see Eddie Dekel and Suzanne Scotchmer 1999; Scotchmer 2005). This regularity's consequences, in turn, are potentially observable in how males fare versus females in WTA games. Here, risk-loving males should exhibit greater "fitness" than do their female counterparts in WTA games, but females who win such games should maintain greater ability.

Legislatures offer a relatively controlled setting in which to consider such implications. Winners take all, for example, in majoritarian elections, and legislative chambers vary considerably in whether prospective members gain entry via such a mechanism. Because success in elections is relatively transparent, then, this setting lets us see how male versus female outcomes relate to game-types. Here, we find evidence that sex differences in risk-appetites (i.e., differences that are institutionally robust) can rationalize a significant portion of associated differences in *legislatures*' composition and *legislators*' durations.

Dekel and Scotchmer (1999) help us rationalize this first difference. They showed that, if contemporary human males maintain risk-loving tendencies, then WTA games will adopt more males than they do females. We find empirical support for this implication from the manner in which female representation varies across 193 legislatures in 256 countries. Given that "winners take all" in majoritarian elections, Dekel and Scotchmer's (1999) model implies that female representation will be relatively low in legislatures whose members face such elections. Controlling for education, preferences for sex bias, and regional fixed effects, we find that female representation *is* relatively low, and significantly so, in national legislatures where members face majoritarian elections. In addition, for a subset of 26 countries where majoritarian elections govern promotions to only one legislative chamber, majoritarian chambers employ a significantly smaller percentage of females than do their non-majoritarian counterparts. This within-country across-chamber matching result corroborates other robustness checks to increase confidence that the sex composition of legislatures indeed changes with whether electoral mechanisms let winners take all.

But while this relationship appears robust, we remain concerned that something other than a tendency for WTA games to adopt risky strategies is generating our evidence. Richard Matland and Donley Studlar (1996), for example, developed a "contagion theory" whereby proportional representation systems employ increasingly large numbers of female legislators because small parties are more capable of gaining representation in those systems and, if one such party succeeds in electing a female candidate, others would have an incentive to mimic that innovator. So, while our identification strategy (e.g., matching estimate) offers confidence that the relationship between electoral mechanisms and legislatures' sex composition is not an artifact of endogeneity bias, it must offer less confidence that our research's *structural* motivation rationalizes this reduced form evidence.

Fortunately, the assumption that males embody risky strategies creates more than the observable implication that males perform relatively well in WTA games. Indeed, it can also imply that females who succeed in winner-take-all settings maintain greater *true* ability (on average) than do their successful male counterparts (Scotchmer 2005).[1] To see this implication, suppose that true ability identically distributes itself across sexes, but individuals win promotion only by sending ability-signals that exceed a sex-neutral standard. Suppose also that, if males play relatively risky strategies, then we can constructively model them as sending noisy signals and females as sending noiseless

[1] Torsten Persson and Guido Tabellini (2004, p. 80) define *ability* as "some mix of integrity, technical expertise or other intrinsic features valued by voters at large."

signals. Then, for a reasonably high standard, the number of males who send erroneously strong ability-signals will exceed that of males who send erroneously weak signals - i.e., promoted females' average ability will exceed that of promoted males. Sex-differences in risk appetites can thus push males and females onto different promotional trajectories, even if true-ability is identically distributed across sexes and promotion standards are sex-blind.

Here, a tendency for males to embody risky strategies creates observable implications for sex differences in *incumbent* players' ability. Consequently, if fundamental differences in risk-preferences are indeed driving our reduced form evidence on legislatures' sex composition, then we should *also* see a tendency for females who win WTA games to be "over-qualified."[2] We evaluate this implication by examining a panel of *re-elections* to the US House of Representatives. If majoritarian elections can be reduced to WTA games, then Scotchmer's (2005) model implies that "promoted" females (i.e., females who recognized an initial electoral success) enjoy a greater prospect for success in subsequent elections than do their male counterparts.

We find evidence for this prediction in the expected number of election successes for *incumbent* females significantly exceeding that for males. This result appears robust to controlling for forces that vary across time and states, such as those that might associate themselves with a preference for sex bias. It also appears robust to estimating a

[2] It is important to note that these "mistakes" need not reflect *individual* irrationality. Here, we treat the game as exogenous and individuals' strategies as having evolved to exploit related niches. Hence, *if individuals constitute units of selection*, then even "games" that aggregate information in an apparently irrational manner can persist. Rich literatures, including those on social choices and the political economy of macroeconomics highlight aggregate irrationality's capacity to emerge in this manner from individually rational agents.

likelihood that treats unobserved heterogeneity more flexibly, and to employing empirical methods that rest on different distributional assumptions (e.g., duration and fixed effects methods). And because potentially competing rationalizations of how electoral mechanisms influence the sex-composition of legislatures do not predict how incumbents' success varies with sex, this evidence increases confidence that deep-seated differences in risk-appetites have something to say about how persistent are sex-differences in hierarchical attainment.

Our research offers several contributions. A number of social scientists have argued that majoritarian chambers employ a relatively small number of females (e.g., see Robert Darcy, Susan Welch, and Janet Clark 1994),[3] and that human males exhibit relatively large risk appetites (e.g., see Nancy Jianakoplos and Alexandra Bernasek 1998; Annika Sunden and Brian Surette 1998; James Byrnes, David Miller, and William Schafer 1999; and Peggy Dwyer, James Gilkeson, and John List 2002). Evidence that we develop here extends that on majoritarian chambers favoring male-representation by exhibiting a considerable level of robustness to potential biases from unobserved covariates. In addition, our empirical investigation finds guidance from a theory that is grounded on fundamental forces that might have influenced preference formation, rather than hypotheses that lack a plausibly exogenous motivation. This feature of our research appears important by itself, and even more so because it encourages us to evaluate how

[3] Welch and Studlar (1990, p. 391) suggest that "political scientists have known for a long time that more women are elected in proportional representation (PR) systems than in plurality ones." These authors are critical of some such evidence, however, since non-PR systems may delineate districts as being either multi- or single-member (though they do not find strong evidence that such a delineation matters). In any event, this source of potential bias appears to create little difficulty for the present investigation since "multi-member non-PR elections are rare in national legislatures" (Welch and Studlar 1990, p. 394).

re-election success varies across sex, a relationship about which extant theories on cross-chamber differences say little (if anything), but for which our data speak strongly.[4] Finally, our research extends accumulated evidence that males are risk-loving to an important empirical setting (i.e., promotional patterns in hierarchical organizations), and does so via a design that carefully checks the potential for omitted variables and other methodological artifacts to introduce bias. As such, it may appeal to readers who find our substantive application interesting, as well as those who are more generally interested in endogenous preference formation.

Our contributions are thus largely positive. In particular, they say nothing explicit about the normative properties of leaving unchecked deep forces that potentially influence sex-differences in socioeconomic status. Pushed in a normative direction, however, our analysis suggests that "affirmative action" (e.g., a policy that sets different promotion standards for different sexes) might enhance welfare, even if males and females have identically distributed ability and otherwise face identical promotion standards. To our knowledge, the present investigation is the first to develop an *empirical* platform from which future research can move in this important policy-direction.[5] *Rather than apologizing for sex differences in promotional attainment, our results thus highlight fundamental obstacles that policy makers might confront if they attempt to mitigate such disparities*.

[4] To be sure, Jeffrey Milyo and Samantha Schosberg (2000) also find evidence that incumbent female Representatives are more able than are their male counterparts. But where Milyo and Samantha attribute this evidence to sex-based prejudices, we find evidence that females' ability-advantage is robust to forces associated with prejudices.

[5] Scotchmer (2005) offers a theoretical treatment of this issue.

We develop these results more fully in the paper's remainder. In the following section, we outline the extant theory that sex differences in risk-appetites can induce robust differences in promotional patterns. In Section 3, we empirically evaluate this theory by examining its capacity to rationalize data on the sex composition of legislatures across countries and sex differences in the tenure of US Representatives. Finally, we conclude in Section 4 by considering how this research might inform associated policy debates, and identifying directions in which future research might accordingly move.

2. Theoretical Motivation

To motivate our empirical analysis, we draw on Dekel and Scotchmer's (1999) model of how risk preferences might have evolved differently across sexes, and Scotchmer's (2005) model of how such a difference can introduce sex biases to hierarchical *outcomes* (even if promotional standards are sex-blind). Exploring the potential for endogenous preference-formation, Dekel and Scotchmer (1999) found that winner-take-all games tend to adopt a certain type of risky strategy. If re-production games let male "winners" take all, then evolutionary forces may have thus endowed contemporary males with relatively large risk-appetites. Embodying risky strategies, these individuals would then exhibit a payoff advantage in other winner-take-all settings. For example, *if majoritarian elections are WTA games, then such elections should adopt more males, on average, than do their non-majoritarian counterparts*.

Scotchmer (2005) extends this formal insight to show how sex differences in risk-taking can induce associated differences in hierarchical promotions. If risk-taking introduces noise to signals of ability, and if males maintain relatively large risk-appetites, then while randomly selected males might exhibit an initial payoff advantage in WTA games, those who recognize a successful outcome will maintain lower ability (on average) than do their female counterparts. In short, sex differences in the distribution of *true* ability emerge from the large number of males that initially recognize success not as a function of ability, but rather from sending a highly variant signal, the particular realization of which happens to exceed a promotional standard. Consequently, if "winners take all" in majoritarian elections, then *females who realize an initial election success (i.e., incumbent females) should display greater ability than do their male counterparts*.

2.1 Observable Implication for Cross-Sectional Data

Dekel and Scotchmer's (1999) model lets us formally consider how cross-sectional data on sex representation in legislatures might be organized. These authors showed that, in large population WTA games, evolutionary forces favor a particular type of risk taking - i.e., "tail dominance." A lottery F_1 tail dominates a lottery F_2 if, evaluated at an element near the top of the lotteries' support, the probability of recognizing an inferior payoff under F_2 exceeds that of F_1. Since WTA games award strictly positive payoffs to only *one* player, they thus tend to favor tail dominant strategies. We summarize this implication here as Proposition 1.

> **Proposition 1** *Winner-take-all games favor risky strategies (Dekel and Scotchmer 1999)*.

Given this proposition, Dekel and Scotchmer (1999) argued that, if forces associated with WTA games influenced mammalian males' reproductive success, then surviving males tend to embody risky strategies. Here, they found considerable support from accumulated evidence that mammalian reproductive games let winners take all, and that mammalian males indeed tend to play risky strategies.[6] In light of mammals' shared ancestry, they concluded that such evolutionary forces could have very well encouraged risk taking in contemporary human males. We restate this conclusion here as Conjecture 1.

> **Conjecture 1** *Relative to their female counterparts, contemporary human males tend to embody risky strategies (Dekel and Scotchmer 1999).*

In addition to that offered by Dekel and Scotchmer (1999), support for Conjecture 1 comes from both evolutionary biologists and psychologists.[7] Male embodiment of risky strategies may also be evident in the relatively strong tendency for males to choose entrepreneurial employment (Joachim Wagner 2004) and behave in an overconfident manner (Claes Bengtsson, Mats Persson, and Peter Willenhag 2005). Moreover, Dwyer et al. (2002) and Robert Olsen and Constance Cox (2001) offered evidence from investment applications that males not only maintain relatively large risk appetites, but that this taste is especially intense for extreme risks.[8]

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[6] Douglas Futuyma (1998) offers a textbook treatment of this phenomenon.

[7] Richard Dawkins (1989) and Steven Pinker (2002), among others, reviewed relevant evidence.

[8] Scotchmer (2005) recounts additional evidence in this regard. Renate Schubert, Martin Brown, Matthias Gysler, and Hans Wolfgang Brachinger (1999), on the other hand, critically reviewed accumulated evidence for gender differences in risk preferences and offer experimental evidence against such differences being empirically important. Their characterization of risk taking (i.e., the magnitude of lotteries' certainty equivalents), however, departs from that which Dekel and Scotchmer (1999) formally identify as fitness-enhancing under winner-take-all mechanisms. Schubert et al.'s (1999) evidence thus does not speak as clearly to our conjecture's empirical relevance as does, say, that which Dwyer et al. (2002) or Olsen and Cox (2001) report. In addition, note that the tail dominance order characterizes preferences where recognizing a support's most preferred outcome produces significantly more utility than does recognizing its second most preferred outcome (Dekel and Scotchmer 1999). This definition of risk appears well suited for the present empirical application.

If Dekel and Scotchmer's (1999) model meaningfully characterizes our empirical reality, then WTA games should favor contemporary human males over their female counterparts (ceteris paribus). To evaluate this conjecture's empirical relevance, one can thus compare how males fare in games that award total payoffs to only one player to how they fare in games that distribute total payoffs across more than one player.

Majoritarian versus non-majoritarian elections offer a rich non-experimental setting in which to conduct this evaluation.[9] Majoritarian mechanisms formally award representation only to an election's winner. Alternative mechanisms (e.g., proportional representation systems), on the other hand, can award strictly positive payoffs (e.g., shared representation) to players that win less than a plurality of votes. *Hence, if evolutionary forces adopted human males that embody tail dominant strategies, and WTA games favor such strategies, then majoritarian elections should favor contemporary males*. We restate this implication as a corollary to Dekel and Scotchmer's (1999) proposition.

> **Corollary 1** *Majoritarian elections favor randomly selected males over randomly selected females*.

2.2 Observable Implication for Time Series Data

Corollary 1 predicts how sex composition, ceteris paribus, varies *across* election mechanisms. Its evolutionary grounding, however, also creates an observable implication for how elected individuals' ability distributes itself across sex, but *within* a

[9] Norman Schofield (2002) suggests that majoritarian systems can engender risky electoral strategies where risk defines itself in terms of how variant is a strategy's expected vote share.

majoritarian chamber. In short, it implies that female *incumbents* maintain (on average) greater true ability than do their male counterparts.

Scotchmer (2005) motivated this implication via the following proposition.

> **Proposition 2** *In winner-take-all games with equal promotion standards, the average ability of promoted females exceeds that of promoted males (Scotchmer 2005).*

To understand this motivation, suppose that playing risky strategies introduces noise to one's signal of ability – i.e., realizing a "successful outcome" conveys little information about the ability of players who place large bets. Under this condition, the *actual* ability of promoted risk-takers (e.g., those who win an initial election) tends to fall short of that which is necessary for promotion. Consequently, if females are more risk averse than are males, then incumbent females' average ability will exceed that of their male counterparts. The following figure illustrates Scotchmer's (2005) formal derivation of this implication.

-----Insert Figure 1 Here-----

Suppose that ability distributes itself across individuals as illustrated and that, if a promotion mechanism could accurately observe ability, it would promote individuals whose ability exceeds a standard x. Now consider two sets of individuals where members of the first maintain an ability-level a_1, members of the second maintain an ability-level a_2, and $|a_1 - x| = |a_2 - x|$ (i.e., the true ability of group-1 and group-2 individuals is equidistant from the promotion standard). Then, if the promotion mechanism observes a noisy signal (rather than true ability), *the number of erroneously included individuals can exceed that of erroneously excluded individuals*.

Here, for example, signals for each type might be uniformly distributed over supports of equal length (i.e., $|s_{1L} - s_{1H}| = |s_{2L} - s_{2H}|$). But because the frequency of group-1 individuals (i.e., individuals whose true ability falls short of the promotion standard) exceeds that of group-2 individuals (i.e., individuals whose true ability surpasses the promotion standard), the number of "truly unqualified" individuals who appear qualified (represented here as the volume of the "Erroneously Included" rectangle) exceeds that of "truly qualified" individuals who appear unqualified (represented here as the volume of the "Erroneously Excluded" rectangle).

Having established that promoted female's average ability can exceed that of promoted males, Scotchmer (2005) showed that these females exhibit (at least initially) an increased probability of recognizing future promotions. This result follows from repeated play's capacity to increase the signal-to-noise ratio of players' strategies and thus more accurately reveal players' true abilities to the promotion mechanism. Its implication is thus potentially observable in how the electoral success of incumbent legislators differs across sexes. We refer to this implication as Corollary 2.

> **Corollary 2** *Female Representatives recognize more favorable re-election prospects than do their male counterparts*.

If majoritarian elections are winner-take-all games, then so too are elections to the US House of Representatives. Consequently, given Proposition 2's implication about how ability distributes itself across *promoted* males and females, we can expect previously elected females to maintain greater true ability than do their male counterparts. To the extent, then, that ability influences candidates' success (and given repeated play's capacity to increase signal-to-noise ratios), *incumbent females should exhibit a relatively strong persistence as majoritarian-elected officials*.

To be sure, it is important to distinguish this insight from its more direct relatives in the literature. Gautam Gowrisankaran, Matthew Mitchell, and Andrea Moro (2003), for example, developed evidence that popularly reported incumbency advantages are largely attributable to selection effects - i.e., the cumulative distribution function (cdf) of incumbent-quality is everywhere (weakly) less than that of non-incumbents. We depart from such treatments by examining how sex differences affect election probabilities *after* candidates have been selected for incumbency. In other words, we're not interested in whether incumbents maintain greater ability than do non-incumbents, but rather in whether female-incumbents maintain greater ability than do their male counterparts.

Milyo and Schosberg (2000) also pursue this interest, but do so via a model where a preference for sex discrimination induces a less than proportional representation of females in the pool of qualified candidates. Our inference, on the other hand, rests on preferences for risk-taking influencing the selection of incumbents and, in turn, these preferences' varying with sex. Here, we're interested not in discrimination's potential to contribute to sex differences in attainment per se, but rather in controlling for this potential so that we can carefully evaluate how differences in risk taking might *independently* influence variation in hierarchical attainment.

2.3 Summary

Before turning to our empirical investigation, we summarize this theoretical motivation via the following figure.

-----Insert Figure 2 Here-----

Both of our observable implications (i.e., Corollaries 1 and 2) ultimately rest on the argument that contemporary human males embody risky strategies. If these males are indeed risk-loving, then WTA games should promote "too many" of them, and thus leave promoted males with an inferior level of true ability (relative to promoted females' ability). The first of these observable implications (i.e., WTA games favor males) rationalizes evidence developed below that legislatures employ a significantly greater percentage of males, ceteris paribus, when majoritarian mechanisms (i.e., WTA games) govern their elections. The second (i.e., successful females maintain relatively high average ability) rationalizes evidence, also developed below, that female Representatives recognize more favorable *re*-election prospects than do their male counterparts. *The hypothesis that males and females differ fundamentally in their preferences for risk thus works, in a parsimonious manner, toward rationalizing what can otherwise appear to be unrelated sex-differences in hierarchical attainment*.

3. Empirical Results

To empirically evaluate our corollaries, we examine how female representation in legislatures varies across majoritarian and non-majoritarian chambers, and how re-election prospects to the US House of Representatives varies across sex. Legislatures offer a relatively controlled (though non-experimental) setting in which to evaluate models where sex differences in risk taking play a foundational role. For example, they offer a more transparent measure of promotional success than do perhaps more immediately obvious settings (e.g., labor markets) in which the phenomena of present interest might be investigated. In addition, forces that might produce observationally equivalent outcomes to those that our theoretical motivation implies (e.g., those associated with child-bearing decisions) should exert a relatively small influence on "promotions" in legislatures (e.g., the average age at which our sampled individuals *entered* the US House is almost 50 years). Finally, legislative data are relatively easy to access, and thus let us control for confounding forces that vary systematically across space and time - e.g., those associated with preferences for sex discrimination. In sum, our ability to check inferential biases here, including those that emerge from measurement error and omitted variables, appears relatively strong.

Results from this examination tend to support our corollaries. First, we find evidence that female representation is significantly lower in chambers where a majoritarian system governs elections. Perhaps the most persuasive evidence in this regard comes from a subset of bicameral legislatures where a majoritarian system governs only one chamber's election. Here, we can make a relatively strong case that potentially confounding, but otherwise unobservable, forces do not spuriously influence our evidence that majoritarian systems favor males. Second, we find evidence that elected females maintain greater ability, on average, than do their male counterparts. In particular, following their first election to the US House of Representatives, female members exhibit significantly longer expected durations as legislators, ceteris paribus, than do their male counterparts. This evidence appears robust to controlling for forces that vary across both time and space (e.g., preferences for sex discrimination), and to specifying duration distributions in a manner that addresses the problem of omitted variables more generally.

3.1 Do Majoritarian Elections Adopt More Males?

To evaluate our first corollary, we examine how the percentage of females in legislatures varies across countries with whether prospective legislators face majoritarian elections. Stated more formally, we estimate parameters from the following model

$$(1) \qquad Percent\ Female_i = \alpha_0 + \alpha_1 Majoritarian_i + \sum_{j=2}^{J} \alpha_j Controls_{i,\ j-1} + \varepsilon_i$$

where *Percent Female*$_i$ equals the percentage of chamber i's legislators who are female, *Majoritarian*$_i$ indicates whether a majoritarian system governs those legislators' election, each element of *Controls*$_i$ proxies for forces that might simultaneously vary with both a chamber's female representation and election mechanism, and ε_i measures residual variation. We can, potentially, observe these variables for 256 chambers from 193 countries. Sample sizes for our empirical results vary, however, since not all measures are accessible for all chambers. Our data appendix defines these variables, identifies their sources, summarizes their distributions, and details each observation.

3.1.1 OLS Results

If majoritarian elections can be formalized as winner-take-all games, and evolutionary forces adopted males whose strategies facilitate high average payoffs in such games, then our variables *Majoritarian* and *Percent Female* should exhibit a *negative* relationship. Table 1 reports parameter estimates that evidence this relationship.

-----Insert Table 1 Here-----

Holding regional effects constant, specification 1 produces evidence that female representation is significantly lower in legislatures whose members face majoritarian elections. Here, *Percent Female* is almost one-half of a standard deviation less in majoritarian chambers than it is in non-majoritarian chambers.

We employ a finer set of controls in specifications 2 and 3 to consider whether preferences for sex bias might spuriously create a negative relationship between *Majoritarian* and *Percent Female*. If a preference for bias shifts the promotion standard for females rightward, and variation in this preference coincides with that in electoral mechanisms, then a relatively large share of males can gain promotion even if sex differences in risk appetites are absent. Recognizing that such a shift might also check females' incentives to invest in requisite human capital, we thus control for education's level (*Education*) and sex-distribution (*Education Difference*). These controls, in other words, address the potential for our benchmark result to reflect (behavioral) bias either directly or through a less than proportional representation of women in majoritarian candidate pools.[10] The resulting evidence, nevertheless remains consistent with our corollary that WTA elections tend to adopt risk-loving males. Here, *Percent Female* continues to be almost one-half of a standard deviation less in majoritarian chambers than it is in non-majoritarian chambers.

We consider other potentially important covariates in specifications (4) - (6). In specification (4), we control for whether electoral rules formally require a minimal level of female representation. Interestingly, our proxy (*Quota*) maintains negligible

[10] Milyo and Schosberg (2000) call particular attention to this latter possibility.

explanatory power, perhaps because quotas can *result from* (rather than *influence*) a chamber's gender composition. In specification (5), we control for the intensity of party competition (*Cohesion*), but find that female representation continues to decrease by almost one-half of a standard deviation when moving to a majoritarian chamber.

Finally, in specification (6), we partition chambers more finely by adding the covariate *Threshold*. For chambers where a proportional representation mechanism governs elections, *Threshold* equals the average (1975-1997) minimum vote share that a party must earn to gain at least one seat. As *Threshold* increases, then, the corresponding chamber's election begins to look like one that is governed by a majoritarian mechanism. Distinguishing chambers in this more "continuous" manner, we find additional evidence for our first corollary - i.e., *female representation not only decreases when moving discretely to majoritarian chambers, it does so also when moving smoothly away from non-majoritarian chambers*. In this latter case, a standard deviation increase in *Threshold* is associated with about a $1/5$ standard deviation decrease in *Percent Female*.

3.1.2 Matching Results

So far, our results appear robust to particular assumptions about regression specification. To *more generally* address our concern that the negative relationship between *Percent Female* and *Majoritarian* is artifactual, we take advantage of a special feature of our data (i.e., a subset of bicameral legislatures maintain both majoritarian and non-majoritarian chambers) to develop a non-parametric matching estimator. Here, unobservables maintain relatively little in the way of channels through which to influence both electoral mechanisms and the sex composition of legislatures. Results from this method are consistent with omitted variables having contributed little to the explanatory power of our OLS estimates.

Among the countries that we observe, 26 maintain bicameral legislatures where prospective members face majoritarian elections to only one chamber. For *20 of these 26* legislatures, the majoritarian chamber employs a smaller percentage of females than does its non-majoritarian counterpart. Table 2 reports these differences.

-----Insert Table 2 Here-----

Majoritarian chambers employ, on average, 5.38% fewer females than do their matched non-majoritarian chambers. This difference is statistically significant at any reasonable level of confidence (i.e., $p = 0.0085$), and exhibits the same magnitude as do our OLS estimates. In addition, variation in other potentially influential forces does not coincide with that in whether a majoritarian system governs a chamber's admission. For example, in unreported regressions, a coarse control for whether the majoritarian-chamber also tends to be the upper chamber does not alter the inference that our matching estimator makes available. Here, holding differences in the number of seats across chambers constant, we continue to find that majoritarian chambers employ significantly fewer females.[11]

Exceptional observations to our corollary provide additional confidence in this regard. For example, Grenada’s majoritarian chamber employs almost 19% *more* females than does its non-majoritarian chamber. However, its non-majoritarian chamber is also its upper house. If forces peculiar to upper houses check female representation in

[11] Our data source - i.e., the Inter-Parliamentary Union's (IPU) PARLINE Database - does not explicitly distinguish upper from lower chambers for all bicameral legislatures.

legislatures (e.g., if preferences for sex discrimination exert an especially strong influence here), then this particular inconsistency with our corollary might not create significant concern. A similar rationalization is available for the UK's inconsistency, while Argentina's and France's might emanate from gender quotas for party lists. On the other hand, similar confounding forces do not systematically influence observations that *are* consistent with our corollary. For example, Belize's non-majoritarian chamber employs over 30% more females than does its majoritarian chamber, even though the non-majoritarian chamber is the upper house.

3.1.3 Alternative Explanations

As reported thus far, our findings remain exposed to at least two potential difficulties. First, our corollary assumes that female- and male-ability are drawn from the same distribution. However, our regression analysis cannot distinguish observations of previously elected legislators from those of never-elected legislators. Consequently, if our second corollary is insightful, then observations of previously majoritarian-elected females are drawn from a distribution whose average ability exceeds that of previously elected males, as well as that of never-elected individuals.[12] But to the extent that any such distributional differences characterize our data, they work *against* finding evidence for Corollary 1.

[12] Recall Corollary 2's implication that, for previously elected individuals, the distribution of female-ability stochastically dominates that of male-ability. On the other hand, a preference for gender discrimination could cause the distribution of male-ability to stochastically dominate that of female-ability. If gender discrimination checks labor market opportunities, for example, then female investment in human capital could be discouraged. We attempt to address this latter possibility in specifications 2 and 3 (see Table 1).

Second, females' electoral advantage in non-majoritarian systems might emanate from sex differences in policy preferences (as opposed to risk preferences).[13] If such preferences constrain prospective legislators' platforms, then female candidates may be "stuck" in a policy space's neighborhood where relatively few electoral members reside.[14] Here, females might exhibit a disadvantage under majoritarian promotion mechanisms that is observationally equivalent to our first corollary's implication.

This alternative rationalization encounters difficulties, however, where our corollaries do not. First, if sex differences in educational attainment are associated with sex differences in policy preferences, then sex disparities in education should *exacerbate* females' disadvantage in majoritarian elections. The coefficient estimate on *Education Difference* reported in Table 1, however, is positive (though insignificant).[15] Second, while sex differences in policy preferences might produce different patterns of female representation *across* majoritarian and non-majoritarian chambers, they would not explain sex differences in electoral outcomes *within* a majoritarian chamber. Our working assumption that sex differences in risk taking are fundamental, on the other hand, generates predictions for *both* cases. Finally, a "sex difference in policy preferences" story encounters difficulty even when attempting to rationalize patterns of female representation across legislatures. The extreme case of a one-district legislature is illustrative. Suppose, for example, that a female list / candidate receives 40% of an

[13] The "gender differences in policy preferences" rationalization enjoys strong support from advocates of PR systems - e.g., see Cynthia Terrell (2000).

[14] This sort of friction can emerge from legislators' inability to make binding promises to electorates. Such a friction can produce net electoral advantages, however, by enhancing associated candidates' commitment capacity (e.g., see David Lee, Enrico Moretti, and Matthew Butler 2002).

[15] In unreported regressions, the interaction between *Education Difference* and *Majoritarian* is similarly uninformative.

electorate's support while each of two male lists / candidates receives 30%. Then *Percent Female* would equal 40 in a proportional representation system, but 100 in a majoritarian system.

3.2 Do Majoritarian Elections Adopt More Able Females?

Our attention turns now to evaluating Corollary 2 - i.e., majoritarian promotional mechanisms adopt more able (and thus more persistent) females. Here, we employ methods of duration analysis to examine whether females, once elected to the US House of Representatives, enjoy a better chance of winning subsequent elections than do their male counterparts. We thus define "spells" as the number of Congresses over which an individual finds continuous employment as a Representative following his or her first election.[16] Spells end upon an individual's first unsuccessful re-election bid, and are censored for members of our last observed Congress (i.e., the 107th Congress) and those who left Congress for reasons other than losing a re-election bid.

Stated more formally, we choose parameters that maximize the following log likelihood function

$$\log L = \sum_{uncensored} \log f(t \mid x) + \sum_{censored} \log S(t \mid x) \tag{2}$$

[16] By defining spells as starting with an *election*, we attempt to mitigate the potential for bias from observations where legislators first gained membership by replacing a deceased spouse or parent (such replacements characterize 14 of our 150 female observations, but none of our male-observations). Controlling for this phenomenon by either adding a "replacement-dummy" to the set of regressors or omitting "replacement-observations" from the sample also produces evidence that "successful" females enjoy longer durations than do their male counterparts, but this evidence exhibits less robustness than does that reported here.

where $f(\cdot)$ denotes the probability that an uncensored observation associated with covariates x survives *exactly* t periods, and $S(\cdot)$ denotes the probability that a censored observation associated with covariates x survives *at least* t periods.

To produce our benchmark result, we postulate that durations follow a Weibull duration. Such a modeling decision appears reasonable when the probability of recognizing a hazard changes monotonically over time. The present application would satisfy this condition, for example, if the probability of recognizing an election-loss diminishes monotonically over an individual's service period.

We evaluate this benchmark's robustness by first controlling for unobservables that might trend over time or vary systematically across states (e.g., preferences for gender bias). To more generally address the potential for endogeneity bias, we also estimate parameters from the likelihood function (2) by postulating that durations follow a *mixed* distribution. Here, we follow Scott Atkinson and John Tschirhart (1986) by employing the Burr-12 density to correct for omitted variables bias.[17] This method flexibly models unobservables as following a gamma distribution, combines that distribution with its less flexible Weibull counterpart for observables, and finally integrates out the unobserved heterogeneity component (from the resulting mixed distribution) to produce an estimable survivor function.[18]

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[17] Atkinson and Tschirhart (1986) argue that this increase in accuracy is especially large when one is considering durations of high-risk / high-stress careers, where career-switching (rather than retirement) can often end an employment spell.

[18] See Atkinson and Tschirhar's (1986, p. 560) equations (16) and (17).

3.2.1 Weibull Results

The covariate of present interest is the variable *Female*, which equals one for observations of female legislators and zero otherwise. Table 3 summarizes the distributions of this and other variables of interest.[19]

-----Insert Table 3 Here-----

The following Table 4 reports estimated parameters from our likelihood function (4). For specification (1), we distinguish observations only by whether corresponding Congress members are female. Here, the estimated coefficient on *Female* is positive and significant, implying that female legislators enjoy an expected duration that is almost *50%* longer than that of their male counterparts. In specification (2), we begin to evaluate the potential for omitted variables to bias available inference by controlling for party affiliation, the year in which members first faced re-election, and the age at which members first faced re-election. These controls rationalize a significant proportion of variation in individual spells. In particular, older entrants exhibit relatively short spells while more recent incumbents exhibit relatively long spells.[20] Nevertheless, we find marginally significant evidence (i.e., p = 0.17) that females' expected durations (evaluated at *Democrat* = 1 and the means of *Start Year* and *Start Age*) are almost *30%* longer than are those of corresponding males.

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[19] Our sample period covers all Congressional election years for which US women were formally eligible to vote (i.e., 1920 through 2000, so $T = 41$). In the cross-sectional dimension, our sample includes 150 female and 520 male observations (i.e., $N = 670$). Resources permitted us to observe all of the female Representatives during our sample period, but only 5% of corresponding males. Here, we selected for observation every twentieth male that the "Biographical Directory" (United States Congress) reports as having served during our sample period.

[20] This latter relationship appears consistent with Gary Jacobson's (1987) evidence that the frequency of first-term House members losing elections appears to have decreased since the 1950s.

-----Insert Table 4 Here-----

We push harder on our check for omitted variables bias in specification (3). Here, we add state dummies to the set of regressors, and continue to find evidence that females recognize considerably longer expected durations than do their male counterparts. For example, evaluated at the other regressors' means (and letting the indicators for *Democrat* and *New York* equal one), the expected duration for females is again about *50%* longer than is that for males.

3.2.2 Other Results

Our evidence from estimating duration models appears robust across several specifications, including those that control for unspecified forces that might vary systematically across time and states (e.g., preferences for gender bias). To control more generally for unobserved heterogeneity, we postulate that durations follow a Burr-12 distribution and report corresponding parameter estimates in specification (4). While this assumption imposes relatively little structure on how unobserved heterogeneity might affect the distribution of durations,[21] we continue to find evidence that female incumbents enjoy significantly longer spells than do their male counterparts.

In unreported results, we also consider the potential for our results to be an artifact of model uncertainty. For example, employing a state / year fixed effects model, we examine a panel of re-election outcomes to evaluate whether previously elected members' probability of wining differs across sexes. We also examine a cross section of electoral success measures to evaluate whether the percentage of re-elections won, or the

[21] See Atkinson and Tschirhart (1986).

incidence of never winning a re-election, differs across sexes. Evidence that emerges from either examination is independent of the distributional assumptions upon which our duration analysis rests, but remains largely consistent with female incumbents enjoying more favorable re-election prospects.

3.2.3 Alternative Interpretations

Just as Corollary 1's structural rationalization of our reduced form evidence appears robust to received alternative interpretations, so does Corollary 2's rationalization. One such alternative is that forces associated with child-bearing cause females to enter Congress with less experience than do comparably aged males. Here, holding our covariates constant, female incumbents might accumulate human capital at a faster rate than do their male counterparts, and thus exhibit longer durations.[22] In addition, females might enjoy longer durations than do comparably aged males because females enjoy longer expected lifetimes.[23] In either case, durations should not vary with *Female* per se, but rather with the interaction of a Representative's sex and his or her starting age. In the following Table 5's first specification, we thus let the duration-effect of Representatives' starting age vary with sex. In this and unreported regressions (e.g., those including fixed effects), however, we find no evidence that the starting age effect on durations varies significantly with sex.[24]

-----Insert Table 5 Here-----

[22] We thank Myles Watts for highlighting this possibility.

[23] We thank Rob Fleck for highlighting this possibility.

[24] In related work, Ghazala Azmat, Maia Guell, and Alan Manning (2003) found evidence that cross-country gender gaps in accumulated labor market experience share an insignificant relationship with associated differences in unemployment rates. In addition, they offered evidence that the US male-female unemployment gap is negligible for the second half of our study period (i.e., 1960-2000, Azmat et al. did not report data for the period prior to 1960).

A third alternative interpretation is that preferences (electoral or otherwise) for gender bias influenced our reduced form evidence. Here, discrimination might increase the standard to which electorate's hold females, and thus promote only "overqualified" females. This implication is observationally equivalent to that of Scotchmer's (2005) model where sex differences in risk taking are primitive.[25]

In the gender bias case, durations might not vary with *Female* per se, but rather with the interaction of a Representative's sex with his or her starting year. For example, if a preference for sex discrimination significantly influenced our data's generation, and if this preference weakened over our sample period (e.g., see Claudia Goldin 2002), then female-observations from relatively early periods in our data should exhibit, ceteris paribus, relatively long durations. We thus interact the variables *Female* and *Start Year* in specification (2).

Recall from Table 4 that estimated coefficients on *Start Year* are positive and significant (see specifications (2) - (4)). These estimates are consistent with *all* legislators enjoying longer durations over time. By interacting *Start Year* and *Female*, we can thus evaluate the extent to which this relationship differs for female observations. We find no evidence of this difference in our data, however. Indeed, the estimated coefficient on *Female* × *Start Year* takes the wrong sign for a gender bias story, and does so also in unreported specifications (e.g., those including state fixed effects).[26]

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[25] Milyo and Schosberg (2000) employed a model of discrimination to rationalize evidence of female candidates' superior quality.

[26] In one such unreported specification, we split the sample into pre- and post-World War II samples. Estimates from both samples evidence longer expected durations for females than for males. In addition, we find evidence that this difference is more robust for the post-war sample. This robustness is also

4. Conclusion

Treating sex differences in risk taking as exogenous, we highlight the potential for promotions in hierarchies to differ systematically across males and females. We argue in particular that winner-take-all games can adopt more males than they do females. We also argue, however, that this tendency can leave promoted-females that (on average) maintain a higher ability-level than do their male counterparts. Finally, we find empirical support for these implications from a cross-section of female representation in legislatures and a panel of re-elections to the US House.

This evidence appears robust not only to our modeling assumptions, but also to numerous alternative rationalizations. While these alternatives might rationalize either our cross-sectional *or* time series results, we find little support for them in our data. Moreover, our corollaries rest on a common theoretical foundation and can rationalize both sets of results. This parsimony strengthens the candidacy of sex-differences in risk appetites as explaining at least some of the associated variation in hierarchical attainment.

In addition to helping us understand how evolutionary forces might influence social status, then, our research may also have important policy implications. Here, a model in which (initial) ability and promotion standards do not vary with sex offers a contending, and certainly parsimonious, rationalization of sex differences in hierarchical attainment. While largely positive, though, this finding suggests that affirmative action policies (e.g., those that would implement lower promotion standards for females) might improve

interesting in that early women in office tended to be "placeholders" for deceased husbands or fathers (we thank Chris Fastnow for bringing this stylized fact to our attention), further suggesting that such females did not spuriously create our duration estimates.

measures of *general* welfare, even if popular normative motivations for such a policy (e.g., gender biases) are absent. Future research could pursue this suggestion by distinguishing how such policies affect welfare in general, rather than the costs and benefits to any particular constituency.

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Data Appendix
Variable Definitions and Sources

Variable	Definition	Source
Percent Female	Percentage of legislature's seats held by females as of August 12, 2002	IPU
Majoritarian	Indicates chambers for which a majoritarian system governed elections as of August 12, 2002	IPU
Education	Average years of schooling (1960-2000) for over-25 population	Barro and Lee
Education Difference	Percentage difference between male and female education	Barro and Lee
Quota	Indicates electoral rules that mandate a minimum female representation	IPU
Cohesion	Index of political cohesion (0 - 3 with 0 indicating the most cohesive government - e.g., one party majority government)	Beck et al.
Threshold	For proportional representation systems, equals the average (1975-1997) minimum vote share that a party must earn to gain at least one seat	Beck et al.
Africa	Indicates chambers residing in Africa	IPU
Americas	Indicates chambers residing in the Americas	IPU
Asia	Indicates chambers residing in Asia	IPU
Europe	Indicates chambers residing in Europe	IPU
Oceana	Indicates chambers residing in Oceana	IPU
Advanced Economy	Indicates advanced economies	Barro and Lee

Data Appendix (cont'd)
Distribution of Variables

	Percent Female	Majoritarian	Education	Education Difference	Quota	Cohesion	Threshold	Africa	Americas	Asia	Europe	Oceana	Advanced Economy
Mean	15.29	0.37	4.98	0.31	0.25	0.80	1.35	0.22	0.31	0.18	0.26	0.02	0.28
Median	12.50	0.00	4.69	0.14	0.00	0.67	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Maximum	42.69	1.00	10.86	1.23	1.00	2.91	17.00	1.00	1.00	1.00	1.00	1.00	1.00
Minimum	0.00	0.00	0.41	-0.13	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
Std. Dev.	9.74	0.48	2.64	0.33	0.43	0.66	2.53	0.42	0.46	0.39	0.44	0.15	0.45
Observations	130	130	130	130	130	118	108	130	130	130	130	130	130

Data Appendix (cont'd)
Correlation of Variables

	Percent Female	Majoritarian	Education	Education Difference	Quota	Cohesion	Threshold	Africa	Americas	Asia	Europe	Oceana	Advanced Economy
% - Female	1.00												
Majoritarian	-0.31	1.00											
Education	0.49	-0.17	1.00										
Ed. Diff.	-0.36	0.25	-0.72	1.00									
Quota	-0.04	-0.03	-0.11	-0.05	1.00								
Cohesion	0.24	-0.16	0.39	-0.16	-0.13	1.00							
Threshold	0.06	-0.26	0.15	-0.08	0.05	0.10	1.00						
Africa	-0.18	0.16	-0.57	0.57	-0.18	-0.37	-0.14	1.00					
Americas	0.04	0.01	0.04	-0.40	0.32	-0.31	-0.14	-0.35	1.00				
Asia	-0.30	0.13	-0.11	0.23	0.05	0.14	-0.16	-0.25	-0.32	1.00			
Europe	0.36	-0.31	0.53	-0.30	-0.18	0.53	0.44	-0.32	-0.40	-0.29	1.00		
Oceana	0.07	0.09	0.22	-0.09	-0.09	0.02	-0.07	-0.07	-0.09	-0.07	-0.08	1.00	
Advanced	0.44	-0.22	0.67	-0.36	-0.19	0.49	0.15	-0.34	-0.24	-0.30	0.77	0.22	1.00

Data Appendix (cont'd)
The Data
Observations = 130
Majoritarian-Observations = 48

Country	% Female	Maj	Ed	Ed Diff	Quota	Cohes'n	Thresh	Afr	Amr	Asa	Eur	Oca	Adv Econ
Algeria	6.17	0	2.10	0.63	0	0.22	0.00	1	0	0	0	0	0
Algeria	5.56	1	2.10	0.63	0	0.22	0.00	1	0	0	0	0	0
Argentina	30.74	0	6.63	0.01	1	0.70	3.00	0	1	0	0	0	0
Argentina	33.33	1	6.63	0.01	1	0.70	3.00	0	1	0	0	0	0
Australia	25.33	1	9.97	0.08	0	0.87	0.00	0	0	0	0	1	1
Australia	28.95	0	9.97	0.08	0	0.87	0.00	0	0	0	0	1	1
Austria	26.78	0	7.73	0.33	0	0.61	4.00	0	0	0	1	0	1
Austria	20.31	0	7.73	0.33	0	0.61	4.00	0	0	0	1	0	1
Bangladesh	2.00	1	1.57	1.05	0	1.26	0.00	0	0	1	0	0	0
Barbados	10.71	1	7.65	-0.01	0			0	1	0	0	0	0
Barbados	33.33	0	7.65	-0.01	0			0	1	0	0	0	0
Belgium	23.33	0	8.12	0.08	1	2.26		0	0	0	1	0	1
Belgium	28.17	0	8.12	0.08	1	2.26		0	0	0	1	0	1
Benin	6.02	0	1.13	0.99	0	0.22	0.00	1	0	0	0	0	0
Bolivia	18.46	0	4.36	0.43	1	1.00		0	1	0	0	0	0
Bolivia	14.81	1	4.36	0.43	1	1.00		0	1	0	0	0	0
Botswana	17.02	1	2.81	0.07	0	0.00	0.00	1	0	0	0	0	0
Brazil	6.82	0	3.33	0.09	1	0.74	4.35	0	1	0	0	0	0
Brazil	6.25	1	3.33	0.09	1	0.74	4.35	0	1	0	0	0	0
Burundi	19.53	0	1.23	0.90	0	0.65	5.00	1	0	0	0	0	0
Canada	20.60	1	9.84	0.02	0	0.13	0.00	0	1	0	0	0	1
Canada	35.48	0	9.84	0.02	0	0.13	0.00	0	1	0	0	0	1
CAR	7.34	1	1.04	1.07	0	0.83	0.00	1	0	0	0	0	0
Chile	12.50	1	6.15	0.04	0	1.00	0.00	0	1	0	0	0	0
Chile	4.17	1	6.15	0.04	0	1.00	0.00	0	1	0	0	0	0
China	21.78	1	4.60	0.57	0	0.00		0	0	1	0	0	0
Columbia	12.05	0	3.85	0.03	0	0.30	0.00	0	1	0	0	0	0
Columbia	8.82	0	3.85	0.03	0	0.30	0.00	0	1	0	0	0	0
Costa Rica	31.58	0	4.79	0.01	1	0.48		0	1	0	0	0	0
Cyprus	10.71	0	6.71	0.24	0			0	0	0	1	0	0
Denmark	37.99	0	9.35	0.15	0	2.91	0.00	0	0	0	1	0	1
Dominican	17.33	0	3.57	0.09	1	0.48	0.00	0	1	0	0	0	0
Dominican	6.25	1	3.57	0.09	1	0.48	0.00	0	1	0	0	0	0
Ecuador	14.63	0	4.75	0.15	1	1.00	0.00	0	1	0	0	0	0
Egypt	2.42	1	3.22	0.71	0	0.17	0.00	1	0	0	0	0	0
El Salvador	9.52	0	3.02	0.21	0	0.57	0.00	0	1	0	0	0	0
Fiji Islands	5.71	1	6.20	0.19	0			0	0	0	0	1	0
Finland	36.50	0	7.85	0.03	0	2.00	0.00	0	0	0	1	0	1
France	12.31	1	6.84	0.06	1	1.61	5.00	0	0	0	1	0	1
France	10.90	0	6.84	0.06	1	1.61	5.00	0	0	0	1	0	1

Data Appendix (cont'd)
The Data

Country	% Female	Maj	Ed	Ed Diff	Quota	Cohes'n	Thresh	Afr	Amr	Asa	Eur	Oca	Adv Econ
Gambia	5.66	1	1.04	0.88	0	0.13	0.00	1	0	0	0	0	0
Germany	31.68	0	8.70	0.11	0	1.65	5.00	0	0	0	1	0	1
Germany	26.64	0	8.70	0.11	0	1.65	5.00	0	0	0	1	0	1
Ghana	9.00	1	2.44	0.69	0	0.70	0.00	1	0	0	0	0	0
Greece	8.67	0	6.46	0.30	0	0.61	3.00	0	0	0	1	0	1
Guatemala	8.85	0	2.14	0.29	0	0.65	4.00	0	1	0	0	0	0
Guyana	20.00	0	4.74	0.04	0			0	1	0	0	0	0
Haiti	3.61	1	1.63	0.61	0	0.52	0.00	0	1	0	0	0	0
Haiti	25.93	1	1.63	0.61	0			0	1	0	0	0	0
Honduras	5.47	0	2.74	0.10	1	0.30	0.00	0	1	0	0	0	0
Hungary	9.07	0	8.06	0.07	0	0.48	5.00	0	0	0	1	0	0
Iceland	34.92	0	7.12	0.08	0			0	0	0	1	0	1
India	8.84	1	2.85	0.85	0	0.74	0.00	0	0	1	0	0	0
India	9.09	0	2.85	0.85	0	0.74	0.00	0	0	1	0	0	0
Indonesia	8.00	0	2.89	0.51	1	0.00	0.00	0	0	1	0	0	0
Iran	4.14	1	2.29	0.65	0	1.00	0.00	0	0	1	0	0	0
Iraq	7.60	1	2.05	0.75	1	1.00	0.00	0	0	1	0	0	0
Ireland	13.25	0	7.54	-0.02	0	1.26	0.00	0	0	0	1	0	1
Ireland	18.33	0	7.54	-0.02	0	1.26	0.00	0	0	0	1	0	1
Israel	13.33	0	8.30	0.12	0	2.22	1.11	0	0	1	0	0	0
Italy	9.84	0	5.63	0.17	0	1.96	4.00	0	0	0	1	0	1
Italy	7.79	0	5.63	0.17	0	1.96	4.00	0	0	0	1	0	1
Jamaica	13.33	1	3.75	-0.13	0	0.00	0.00	0	1	0	0	0	0
Jamaica	23.81	0	3.75	-0.13	0	0.00	0.00	0	1	0	0	0	0
Japan	7.29	0	8.16	0.11	0	1.48		0	0	1	0	0	1
Japan	15.38	0	8.16	0.11	0	1.48		0	0	1	0	0	1
Jordan	1.25	1	3.82	0.57	1	1.00	0.00	0	0	1	0	0	0
Kenya	3.57	1	2.30	0.67	0	0.00	0.00	1	0	0	0	0	0
Kuwait	0.00	1	4.34	0.15	0	1.00	0.00	0	0	1	0	0	0
Liberia	7.81	0	1.39	0.98	0			1	0	0	0	0	0
Liberia	19.23	0	1.39	0.98	0			1	0	0	0	0	0
Malawi	9.33	1	2.22	0.77	0	0.13	0.00	1	0	0	0	0	0
Malaysia	10.36	1	4.69	0.54	0	0.96	0.00	0	0	1	0	0	0
Malaysia	26.09	0	4.69	0.54	0	0.96	0.00	0	0	1	0	0	0
Mali	12.24	1	0.41	1.03	0	0.26	0.00	1	0	0	0	0	0
Malta	9.23	0	6.15	0.13	0			0	0	0	1	0	0
Mauritius	5.71	1	4.18	0.37	0	1.61	0.00	1	0	0	0	0	0
Mexico	16.00	0	4.35	0.24	1	0.00	0.85	0	1	0	0	0	0
Mexico	15.63	1	4.35	0.24	1	0.00	0.85	0	1	0	0	0	0
Mozambique	30.00	0	0.63	0.78	0	0.09	5.00	1	0	0	0	0	0

Data Appendix (cont'd)
The Data

Country	% Female	Maj	Ed	Ed Diff	Quota	Cohes'n	Thresh	Afr	Amr	Asa	Eur	Oca	Adv Econ
Nepal	5.85	1	0.69	1.23	1	1.52	0.00	0	0	1	0	0	0
Netherlands	34.00	0	7.70	0.09	0	1.30	0.67	0	0	0	1	0	1
Netherlands	26.67	0	7.70	0.09	0	1.30	0.67	0	0	0	1	0	1
Nicaragua	20.65	0	3.08	0.02	0	0.43	0.00	0	1	0	0	0	0
Niger	1.20	0	0.42	1.11	0	0.83	0.00	1	0	0	0	0	0
Norway	35.76	0	8.73	0.09	0	2.30	0.00	0	0	0	1	0	1
Panama	9.86	0	5.87	0.00	1	0.74	5.00	0	1	0	0	0	0
Paraguay	2.50	0	4.56	0.14	1	0.17	0.00	0	1	0	0	0	0
Paraguay	17.78	0	4.56	0.14	1	0.17	0.00	0	1	0	0	0	0
Peru	18.33	0	5.03	0.30	1	0.35	1.18	0	1	0	0	0	0
Philippines	17.76	0	5.86	0.06	1	0.48	2.00	0	0	1	0	0	0
Philippines	12.50	1	5.86	0.06	1	0.48	2.00	0	0	1	0	0	0
Poland	20.22	0	8.44	0.09	0	1.00	5.00	0	0	0	1	0	0
Poland	23.00	1	8.44	0.09	0	1.00	5.00	0	0	0	1	0	0
Portugal	19.13	0	3.34	0.26	0	0.96	0.00	0	0	0	1	0	1
Korea	5.86	0	6.98	0.37	1	0.61	5.00	0	0	1	0	0	0
Rwanda	25.68	0	1.36	0.64	1	0.57	0.00	1	0	0	0	0	0
Senegal	19.17	0	1.80	0.63	0	0.00	0.00	1	0	0	0	0	0
Singapore	10.64	1	4.90	0.36	0	0.00	0.00	0	0	1	0	0	0
South Africa	29.82	0	5.31	0.07	0	0.00	0.65	1	0	0	0	0	0
South Africa	18.89	0	5.31	0.07	0	0.00	0.65	1	0	0	0	0	0
Spain	28.29	0	5.22	0.09	0	1.26	3.00	0	0	0	1	0	1
Spain	24.32	0	5.22	0.09	0	1.26	3.00	0	0	0	1	0	1
Sri Lanka	4.44	0	4.73	0.25	0	0.39	0.00	0	0	1	0	0	0
Sudan	9.72	0	0.87	0.97	1	0.65	0.00	1	0	0	0	0	0
Swaziland	3.08	1	3.53	-0.01	0			1	0	0	0	0	0
Swaziland	13.33	0	3.53	-0.01	0			1	0	0	0	0	0
Sweden	42.69	0	9.12	0.03	0	2.13	4.00	0	0	0	1	0	1
Switzerland	22.50	0	9.07	0.12	0	2.00	0.00	0	0	0	1	0	1
Switzerland	19.57	1	9.07	0.12	0	2.00	0.00	0	0	0	1	0	1
Syria	10.40	1	3.10	0.79	0	1.00	0.00	0	0	1	0	0	0
Thailand	9.20	0	4.38	0.26	0	2.00	0.00	0	0	1	0	0	0
Thailand	10.50	0	4.38	0.26	0	2.00	0.00	0	0	1	0	0	0
Togo	4.94	1	1.46	1.03	0	0.35	0.00	1	0	0	0	0	0
Trin. & Tob.	16.67	1	5.83	0.04	0	0.09	0.00	0	1	0	0	0	0
Trin. & Tob.	32.26	0	5.83	0.04	0	0.09	0.00	0	1	0	0	0	0
Tunisia	11.54	0	2.07	0.74	0	0.04	5.00	1	0	0	0	0	0
Turkey	4.18	0	3.11	0.66	0	1.22	10.00	0	0	0	1	0	1

Data Appendix (cont'd)
The Data

Country	% Female	Maj	Ed	Ed Diff	Quota	Cohes'n	Thresh	Afr	Amr	Asa	Eur	Oca	Adv Econ
Uganda	24.67	1	1.78	0.78	1	0.30		1	0	0	0	0	0
UAE	0.00	0	2.88	0.31	0	1.00		0	0	1	0	0	0
UK	17.91	1	8.25	0.01	0	0.30	0.00	0	0	0	1	0	1
UK	16.41	0	8.25	0.01	0	0.30	0.00	0	0	0	1	0	1
USA	14.02	1	10.86	0.01	0	0.65	0.00	0	1	0	0	0	1
USA	13.00	1	10.86	0.01	0	0.65	0.00	0	1	0	0	0	1
Uruguay	12.12	0	5.98	-0.05	0	1.00	0.00	0	1	0	0	0	0
Uruguay	9.68	0	5.98	-0.05	0	1.00	0.00	0	1	0	0	0	0
Venezuela	9.70	0	4.15	0.12	1	0.61	0.00	0	1	0	0	0	0
Yugoslavia	7.25	0	6.02	0.35	0	0.26	17.00	0	0	0	1	0	0
Zambia	12.03	1	3.28	0.56	0	0.00	0.00	1	0	0	0	0	0
Zimbabwe	10.00	1	2.92	0.49	0	0.26	0.00	1	0	0	0	0	0
Observations	130	130	130	130	130	118	108	130	130	130	130	130	130
Mean (Maj = 0)	17.63	0	5.33	0.24	0.26	0.91	1.87	0.17	0.30	0.15	0.37	0.01	0.35
Mean (Maj = 1)	11.30	1	4.39	0.42	0.23	0.63	0.49	0.31	0.31	0.25	0.08	0.04	0.15
p-value (diff)	0.00	NA	0.05	0.00	0.73	0.03	0.01	0.06	0.93	0.14	0.00	0.28	0.01

Figure 1
Risk Taking Induces (Net) Erroneous Promotions

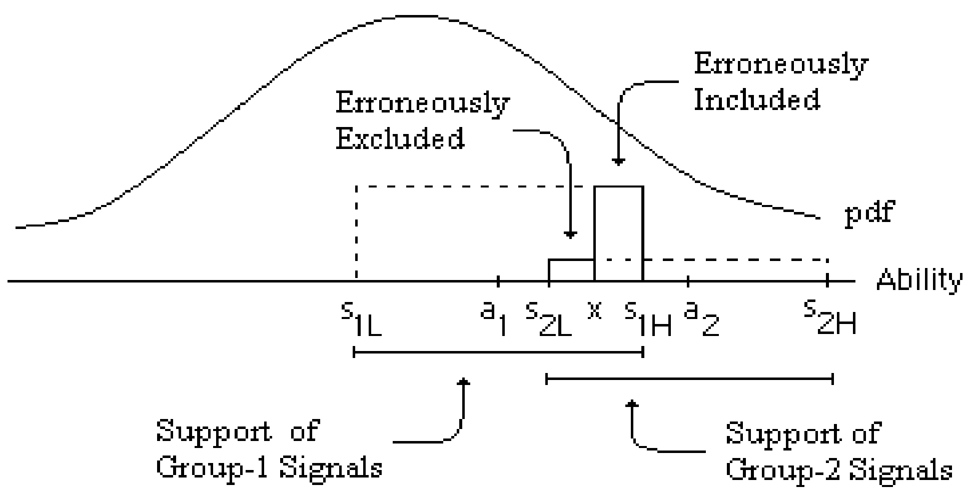

Figure 2
Development of Observable Implications

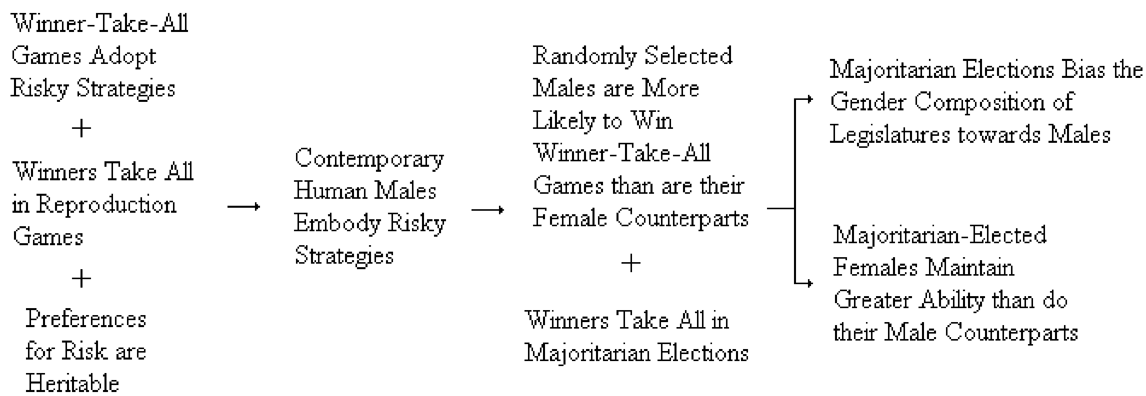

Table 1
Dependent Variable = *Percent Female*
Estimation Method = OLS
White-Corrected Standard Errors

	(1)		(2)		(3)		(4)		(5)		(6)	
Variable	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.	Coeff.	S.E.
Constant	17.68	3.83***	7.69	4.91	6.11	5.73	7.34	4.96	11.07	4.50***	11.01	4.52**
Majoritarian	-4.18	1.49***	-4.32	1.46***	-4.49	1.47***	-4.27	1.47***	-4.18	1.47***	-6.11	1.52***
Education			1.47	0.46***	1.66	0.58***	1.50	0.48***	1.74	0.45***	1.77	0.42***
Education Difference					2.11	3.32						
Quota							1.00	1.82				
Cohesion									0.40	1.52		
Threshold											-0.66	0.33**
N	130		130		130		130		118		108	
R^2	0.29		0.35		0.35		0.35		0.38		0.44	
Adj. R^2	0.25		0.31		0.31		0.31		0.34		0.40	
$\bar{y}$	15.29		15.29		15.29		15.29		15.20		14.89	
σ_y	9.74		9.74		9.74		9.74		9.69		9.69	

Notes: ***, **, and * indicate confidence at the 99%, 95%, and 90% levels, respectively. Each specification also includes regional dummies for Africa, Americas, Asia, and Europe (Oceana is omitted), and an indicator for advanced economies. Results reported here are robust to a finer regional partition – i.e., specifications that include dummies for East Asia and Pacific, Latin America and Caribbean, Middle East and North Africa, South Asia, and Sub-Saharan Africa.

Table 2
Within-Country / Between-Chamber
Differences in Female Representation

Country	% Female (Majoritarian) – % Female (Non-Majoritarian)
Algeria	-0.61
Antigua and Barbuda	-6.50
Argentina	2.59
Australia	-3.62
Barbados	-22.62
Belize	-30.60
Bolivia	-3.65
Brazil	-0.57
Canada	-14.88
Czech Republic	-4.65
Dominican Republic	-11.08
Ethiopia	-0.65
France	1.41
Grenada	18.98
India	-0.25
Jamaica	-10.48
Malaysia	-15.73
Mexico	-0.37
Morocco	0.25
Philippines	-5.26
Poland	2.78
Saint Lucia	-7.07
Swaziland	-10.25
Switzerland	-2.93
Trinidad and Tobago	-15.59
United Kingdom	1.50
Count	26
Average	-5.38
S.D.	9.60
Max	18.98
Min	-30.60

Table 3
Description of Variables
Re-Elections to the US House

Variable	Definition	Mean	S.D.	Max	Min	Source
Spell	Number of Congresses (following an individual's first successful election) during which an individual continuously served as a Representative	4.19	3.57	25	1	US Congress
Female	Indicates female observations	0.23	0.42	1	0	US Congress
Democrat	Indicates Democrat observations	0.55	0.50	1	0	US Congress
Start Year	Year in which individual first gained admission to the House	1957	25.8	2000	1920	US Congress
Start Age	Age at which individual first gained admission to the House	47.7	9.09	79	28	US Congress

Table 4
Dependent Variable = *Spell*
Method of Estimation = Maximum Likelihood
Observations = 670

Variable	(1) Coeff.	S.E.	(2) Coeff.	S.E.	(3) Coeff.	S.E.	(4) Coeff.	S.E.
Constant	2.24	0.06***	-19.01	5.50***	-15.40	5.37***	-20.51	2.34***
Female	0.39	0.17**	0.23	0.17	0.37	0.17**	0.29	0.01***
Democrat			-0.13	0.11	-0.35	0.12***	-0.28	0.01***
Start Year			0.01	0.00***	0.01	0.00***	0.01	0.00***
Start Age			-0.02	0.01***	-0.02	0.01***	-0.02	0.00***
Distribution	Weibull		Weibull		Weibull		Burr-12	
State Effects?	No		No		Yes		No	
Log *L*	-694		-674		-625		8,020	
Avg. Log *L*	-1.04		-1.01		-0.93		11.97	
Akaike	2.08		2.03		2.01		-23.92	

Notes: ***, **, and * indicate confidence at the 99%, 95%, and 90% levels, respectively.

Table 5
Dependent Variable = *Spell*
Method of Estimation = Maximum Likelihood
Observations = 670

	(1)			(2)		
Variable	Coeff.	S.E.		Coeff.	S.E.	
Constant	-18.7514	5.4952	***	-18.9554	5.5096	***
Female x Start Age	0.0049	0.0033				
Female x Start Year				0.0001	0.0001	
Democrat	-0.1332	0.1139		-0.1344	0.1140	
Start Year	0.0114	0.0028	***	0.0115	0.0028	***
Start Age	-0.0233	0.0061	***	-0.0227	0.0060	***
Distribution	Weibull			Weibull		
Log L	-674			-674		
Avg. Log L	-1.01			-1.01		
Akaike	2.03			2.03		
State Indicators	No			No		

Notes: ***, **, and * indicate confidence at the 99%, 95%, and 90% levels, respectively.