Gender and quality at top economics journals*

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Abstract

We show articles published in "top-five" economics journals authored by men are cited less than articles those same journals publish by women. Additionally, men's citations rise when they co-author with women whereas women's citations fall while they co-author with men, conditional on acceptance. Under strong—but we believe reasonable—assumptions, our findings imply top economics journals hold female-authored papers to higher standards and, as a result, do not publish the highest quality research. They also suggest that authors will be less willing to collaborate with women, all else equal.

Keywords: Gender, Discrimination, Quality, Citations, Research, Productivity, Collboration; *JEL*: A11, J16, J24.

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1 Introduction

Discrimination hurts its victims and, sometimes, its perpetrators (Becker 1957). For example, if an academic journal only publishes papers authored by men, its quality should decline relative to one that is gender-blind; if a male economist refuses to co-author with women, his papers ought to publish less well than the men who don't.

For these reasons, discrimination is sometimes considered incompatible with competitive forces (see, *e.g.*, Summers 2005). When markets are complete, the argument is roughly as follows: sufficient competition between unprejudiced journals should ensure female-authored papers are accepted at rates just equal to their marginal quality; sufficient competition between prejudiced and unprejudiced journals should ensure each is ranked according to the quality of the articles it publishes.[1] As long as a journal's ranking prices its articles' quality, the quality of the papers it publishes should not vary by author gender nor should returns to co-authoring depend on a co-author's sex.[2]

As we show in this paper, however, articles published in top economics journals authored by men are cited less than articles those same journals publish by women. Moreover, men's citations rise when they co-author with women whereas women's citations fall while they co-author with men, conditional on acceptance.

These results are based on an analysis of gender differences in citations in over 11,000 full-length articles published in a "top-five" economics journal—*i.e.*, *American Economic Review* (*AER*), *Econometrica* (*ECA*), *Journal of Political Economy* (*JPE*), *Quarterly Journal of Economics* (*QJE*) and *Review of Economic Studies* (*REStud*). We follow previous research and use the inverse hyperbolic sine (asinh) of citations as our dependent variable (Card and DellaVigna 2020; Card *et al.* 2020). We stress, however, that results and conclusions are similar if raw citations or alternative transformations are used instead (see Appendices B and C).

Our first analysis suggests female-authored papers receive 22–24 log points more citations than male-authored papers. Differences in citation patterns across field coupled with higher female concentration in certain areas potentially explain, at most, half of the gap. Assuming quality positively associates with citations and the latter are not biased *in favour* of women (conditional on the former), these results suggest female-authored papers published in top economics journals are higher quality than male-authored papers.

Higher quality female-authored papers could be consistent with gender-neutral acceptance standards if women's papers are accepted more often or the variance in their quality is greater (see Theorem 3.1). Neither appears to be the case. Variance in quality is persistently lower in female-authored papers; recent evidence from a set of journals that partially overlaps with our own suggests men's and women's manuscripts are accepted at roughly equivalent rates (Card *et al.* 2020).

We also consider whether consistent same- or opposite-sex co-author complementarities meaningfully contribute to gender differences in quality, conditional on acceptance. For example, men and women may produce higher quality work when collaborating with one another; conversely, everybody could work better with members of their own sex.

Our evidence does not support either hypothesis. After accounting for author-specific fixed effects, we find men's citations increase 10–12 log points when they co-author with an equal share of women

[1] When markets are incomplete, however, even taste-based discrimination can persist in competitive equilibria or generate equilibria in which non-realised discrimination nevertheless results in an inefficient allocation of resources—*e.g.*, because some women flee to less discriminatory fields although their interests and talents would have been better matched to a career in economics. See, *e.g.*, Diamond (1971), Borjas and Bronars (1989) and Black (1995).

[2] If journal rankings perfectly price article quality, then every article published in the same journal must be exactly the same quality. If markets are complete but journal rankings *do not* precisely price article quality, then there must exist some other mechanism that does.

(compared to papers those same men co-author entirely with other men). The returns to women of co-authoring with an equal share of men, however, are -12 to -15 log points. These findings suggest women disproportionately contribute to the quality of co-authored work, conditional on acceptance, and are consistent with women facing tougher standards in peer review.

Moreover, the coefficients on co-author gender do not meaningfully change after accounting for field. This points to a relationship between citations and gender that is independent of field, conditional on author. Considered alongside the fall in coefficient value from our first analysis, it may also reveal an underlying association between field and author-specific unobservables that will partially bias estimates of gender differences in citations when controlling for the former but not the latter.

As a final exercise, we restrict our sample to senior male economists with at least two top-five papers co-authored with a single junior author of each sex. In addition to accounting for potential unobserved contributions from same-sex co-authors, these sample restrictions create a treatment group—senior male authors co-authoring with exactly one junior woman—that very closely resembles the counterfactual group—those very same seniors co-authoring with exactly one junior man.

Again, we find a senior man's work is higher quality when it is co-authored with a woman. Citations increase 60–70 log points—and 140 log points after accounting for field—when senior male economists co-author with junior women as opposed to junior men.

Combined, our evidence suggests journals subject female authors to higher standards and, as a result, their articles are better quality, conditional on acceptance. We emphasise, however, that these conclusions rely on the following (strong) assumptions: (i) citations are not biased in women's favour, conditional on quality; (ii) author-specific heterogeneity is fixed over time (author-level analyses, only); (iii) quality is normally (but not necessarily identically) distributed in male- and female-authored submissions; and (iv) conditional on controls, male- and female-authored papers are accepted at similar rates.

Assuming (i)–(iv) are satisfied, non-top-five economics journals could have published higher quality content than top-five economics journals simply by accepting more female-authored papers. The fact that journal rankings have not adjusted to reflect this suggests the market for academic research remains incomplete.

Journals function as price mechanisms—*i.e.*, the journals in which articles are published serve as nominal currency for their value. If women could hedge (without friction) against every possible publication outcome in every possible state of the world, then biased acceptance decisions at one journal could simply be "undone" by a costless change in one's submission and publication strategy the previous date—*e.g.*, women could publish their higher quality papers in currently lower-tiered journals, confident that their actions would lead to an appropriate relative change in journal rankings the very next period.

When competition isn't perfect, however, discrimination interacts with one or more market frictions to prevent those who discriminate from fully internalising its costs. Consequently, its victims will have to *partially* bear them. For example, imperfect information about journal rankings may mean tenure and promotion committees' expectations are slow to adjust to the lower quality of journals that reject too many women.[3] As a result, women (and the men they co-author with) are tenured and promoted at lower rates than they otherwise would be if markets were complete. To the extent that grant committees similarly rely on applicants' past publication histories to choose between projects, women will also have a harder time funding future work.

Moreover, discrimination undoubtedly distorts authors' decisions in ways that can further misallocate available resources. Indeed, our own evidence implies male and female economists are better off col-

[3] See Heckman and Moktan (2019) for evidence that tenure expectations are indeed sticky. Moreover, *AER*, *ECA*, *JPA*, *QJE* and *REStud* have remained at the top of the economics publishing hierarchy for the past 30–40 years (Ellison 2002).

laborating with men, all else equal. This incentivises authors of both sexes to forgo higher quality co-authoring opportunities with women in order to partner with men (see also Knobloch-Westerwick *et al.* 2013).

This paper makes four contributions. First, we join a large literature investigating gender differences in citations in economics (see, *e.g.*, Ferber 1988; Laband 1987; Smart and Waldfogel 1996; Ginther and Kahn 2004; Hamermesh 2018; Grossbard *et al.* 2018; Card *et al.* 2020). We add to this research by studying the comprehensive set of journals at the very top of the publishing hierarchy and examining gender-specific contributions to quality in mixed-sex co-authoring relationships, conditional on acceptance.

Second, we also contribute to a substantial body of research suggesting women are, in many situations, subjected to tougher standards and/or evaluated differently than men (see, *e.g.*, Foschi 1996; Moss-Racusin *et al.* 2012; Reuben *et al.* 2014; Krawczyk and Smyk 2016; Sarsons *et al.* 2019). Most relevant to our work, Hengel (2019) finds female-authored papers in top-four economics journals are held to higher writing standards in academic peer review. As a result, their manuscripts are subject to greater scrutiny, spend longer under review and women, in turn, respond by conforming to those standards. Card *et al.* (2020) finds female-authored papers are higher quality conditional on referee recommendations using submissions data from a partially overlapping set of journals (*Journal of the European Economic Association*, *Review of Economics and Statistics*, *QJE* and *REStud*).

Third, although market mechanisms undoubtedly alleviate discrimination's downstream effect, our results highlight that they will not, in general, fully absorb them. For example, the fall in women's citations when they co-author with men suggests adding male co-authors can mitigate higher acceptance standards; men also experience a rise in citations when they co-author with women, however, so co-authorship alone probably cannot eliminate them. And when combined with recent evidence of the "Matilda effect" in tenure decisions—as Sarsons *et al.* (2019) shows, tenure committees discount women's contributions to mixed-sex co-authored work—co-authoring with men may bring other consequences, as well.

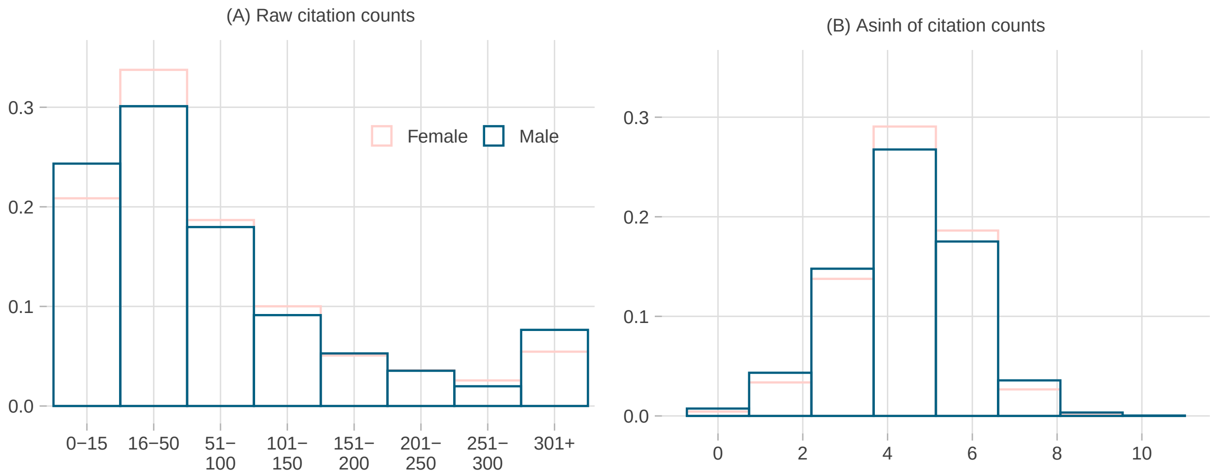
Finally, this paper builds on a broader literature studying editorial patterns (Card and DellaVigna 2013; Casnici *et al.* 2017; Clain and Leppel 2018; Ellison 2002), bias in editorial decisions (Abrevaya and Hamermesh 2012; Bransch and Kvansnicka 2017; Card and DellaVigna 2020; Hospido and Sanz 2019) and female academics' lagging productivity and under-representation (Auriol *et al.* 2019; Bayer and Rouse 2016; Chari and Goldsmith-Pinkham 2017; Ductor *et al.* 2018; Ginther and Kahn 2004; Heckman and Moktan 2019; Lundberg and Stearns 2019; Teele and Thelen 2017). We also join a wider debate about whether women are considered equal partners in research and given enough credit for their contributions (see, *e.g.*, Ferber 1986; Dion *et al.* 2018; Sarsons *et al.* 2019).

Our paper proceeds in the following order. Section 2 describes the data. Section 3 assesses gender differences in citations and quality. Section 4 investigates the returns to co-authoring with the opposite sex. Section 5 concludes.

2 Data and empirical setting

Our data include 11,336 full-length, English-language articles published between 1950–2015 in the *AER*, *ECA*, *JPE*, *QJE* and *REStud*. We define "full-length" as any article published with an abstract. Articles from *Papers & Proceedings* issues of *AER*, errata and corrigenda are excluded.

Each of the 7,574 unique authors in our dataset was manually assigned a gender based on (i) obviously gendered given names (*e.g.*, "James" or "Brenda"); (ii) photographs on personal or faculty websites; (iii) personal pronouns used in text written about the individual; and (iv) by contacting the author himself or people and institutions connected to him.



Note. Left-hand graph displays the fraction of authors with a top-five paper that was cited 0–15 times, 16–50 times, 51–100 times, *etc.* Right-hand graph plots the histogram of transformed citations (asinh).

Figure 1: Distribution of citations

As we highlight in Section 2.2, exclusively female-authored papers are only a very small proportion of articles published in top economics journals. We therefore define the gender of a paper by its fraction of female authors. (See Hengel (2019) for a theoretical justification of this indicator.) For robustness, we also replicate relevant analyses using a categorical variable to account for seven different gender categories and by comparing entirely male-authored papers to papers with: (a) a senior female author; (b) at least one female author; (c) a weak majority of female authors; and (d) a solo female author. Results are shown in Appendix D.

Citation data were obtained from Web of Science (2018), a comprehensive database of all social science research published since 1900. Counts correspond to the number of published papers in the Web of Science database that cite a given article and include self-citations to later work. Citations for *AER*, *ECA*, *JPE* and *QJE* were first collected in August 2017 and updated in January 2018; citations for *REStud* were collected in October 2018.

2.1 Issues when analysing gender citation gaps

Analysing citation gaps raises several issues: (i) younger articles have had less time to accrue citations; (ii) older articles are disproportionately male; (iii) mean citation counts may be distorted by a small number of superstar economists; (iv) superstar economists are *also* disproportionately male. (i) and (iii) skew the distribution of citations; (ii) and (iv) distort gender differences if they correlate with unobserved factors that generate citations—*e.g.*, winning a prestigious award—or produce self-reinforcing loops that create citations all on their own (Merton 1986; Merton 1988; Zuckerman 1977).[4]

Additionally, men disproportionately cluster at the very top and bottom of the distribution of citations but raw counts are censored below at zero and unbounded from above (Figure 1, Graph (A)). This generates a non-linear mapping from quality onto citations that depends on the former's variance. Thus even in the absence of (i)–(iv), average citations to male-authored papers likely place too much weight on high-citation papers and not enough weight on low-citation papers compared to the average for female-authored papers.[5]

[4] In Appendix F, we illustrate how (iii) and (iv) combine to distort gender differences at the mean in raw citations by constructing and controlling for a set of "superstar" and Nobel Prize fixed effects (Appendix F).

[5] For example, suppose the "quality" of female- and male-authored papers are each normally distributed with means zero and variances one and two, respectively, and the exponential function maps quality onto citations. Then male-authored papers are, on average, cited more than female-authored papers (2.7 versus 1.6, respectively) even though both groups' average quality is exactly the same.

We principally account for these issues by controlling for journal-year fixed effects and transforming raw counts with the inverse hyperbolic sine function (asinh). This reduces the impact of outlier observations and results in a far more symmetric distribution (Figure 1, Graph (B)). We emphasise, however, that the conclusions we draw do not depend on the transformation. We find similar results using a quantile regression model (Appendix B) and the log of 1 plus citations or raw counts themselves as the dependent variable (see Appendix C).

A related concern is that the citations a paper accumulates aren't fixed in time. As a result, they could be influenced by the future success or failure of a paper's authors—*i.e.*, even among non-superstar economists, a stronger publishing record later on probably drives citations to earlier work, all else equal (for evidence see, *e.g.*, Bjarnason and Sigfusdottir 2002). Female economists, however, have weaker publishing histories: men's mean number of top-five papers is 2.6; the corresponding figure for women is only 1.8.

To account for this, we additionally control for how prominent an author is at about the time citations were collected (max. T). Max. T is equal to the total number of top-five articles for the most prolific co-author as of the end of 2015. Because *ex post* success may itself be endogenous to citations, we also show results without max. T.

Finally, several studies show citation counts correlate with sub-field and author prestige at the time of publication, conditional on quality (see, *e.g.*, Bornmann *et al.* 2012; Wang 2014). Additionally, more co-authors can present a single paper more widely and self-cite it more frequently.[6] We therefore account for the following in most of our analyses: primary and tertiary *JEL* fixed effects,[7] number of co-authors (N) and author seniority at the time of publication (max. t)—*i.e.*, the most prolific co-author's total number of top-five articles at the time the paper was published.

2.2 The representation of women in top-five economics journals

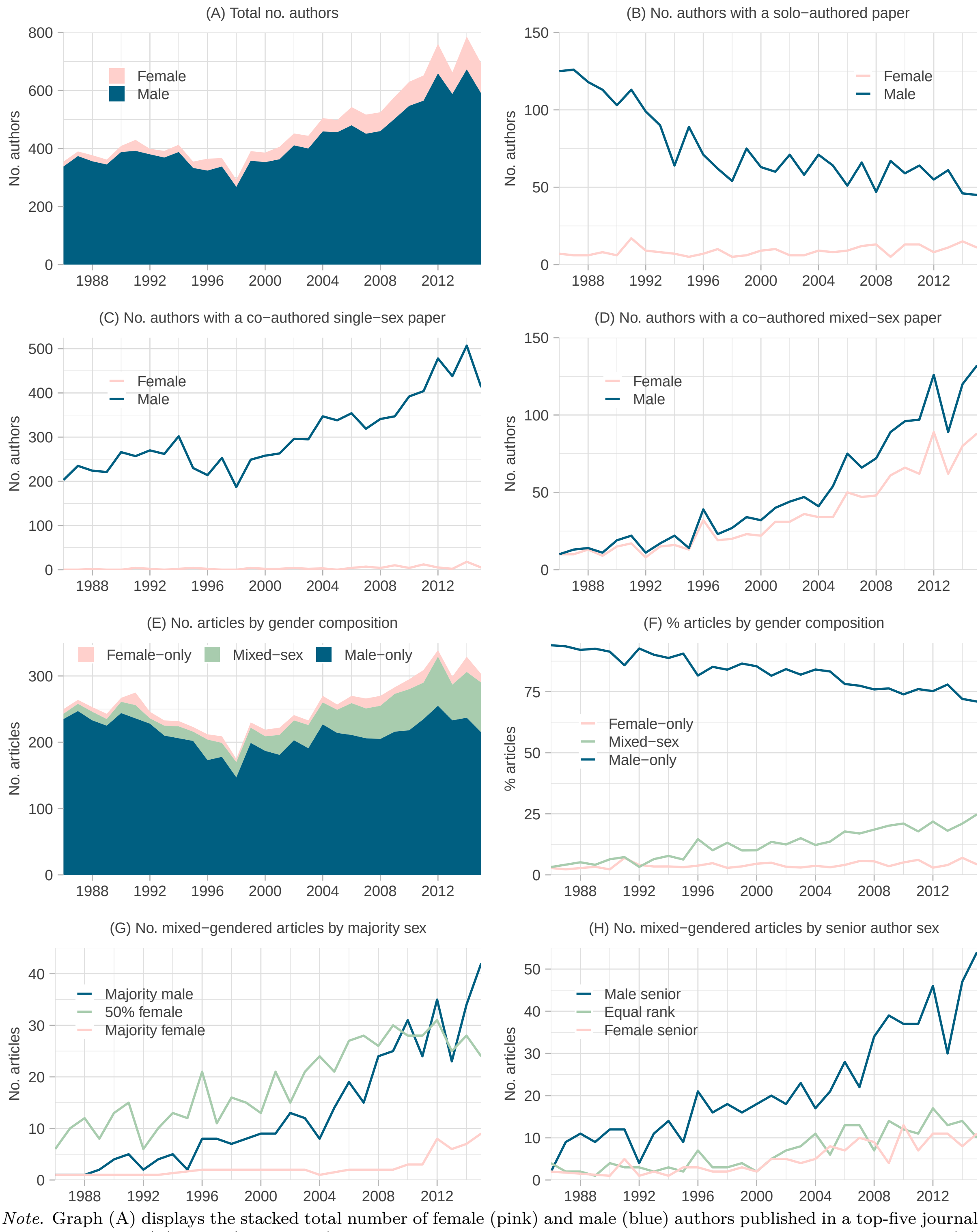
Women are under-represented in top-five economics journals. As Graph (A) in Figure 2 illustrates, the situation has improved little with time: women make up only 11 percent of all authors published since 1990, 12 percent since 2000 and 14 percent since 2010. Between 1986–2015, there has been zero growth in the number of exclusively female-authored papers; almost no growth in the number of majority female-authored papers; and no meaningful change in the number of mixed-gendered papers with a senior female co-author. The only tepid growth that *has* occurred, is largely—if not entirely—due to an increase in the number of articles by senior men co-authoring with a weak minority of junior women.

Top-five journals publish about as many solo female-authored papers today as they did in the late 1980s (Figure 2, graph (B)): seven in 1986, ten in 1997 and eleven in 2015. The number of solo male-authored papers, however, has declined: 125 were published in 1986, 62 in 1997 and 45 in 2015. As a result, the proportion of solo-authored papers by women has increased from five percent in 1986 to twenty percent in 2015.

But falling male solo-authored papers has been more than offset by rising male *co-authored* papers. As a result, the proportion of female authors on single-sex papers has remained stubbornly close to zero for the past 30 years (Figure 2, graph (C)). In 1987, top-five journals collectively published 105 articles co-authored by two men and zero articles co-authored by two women; in 2015, the corresponding figures were 104 and one. Meanwhile, journals have sharply increased the number of single-sex articles they publish by three or more men: 66 were published in 2015 versus 18 in 1986. As of 2015, however, only ***six*** had *ever* been published by women; no top-five journal had yet to publish a full-length paper

[6]See Tahamtan *et al.* (2016, p. 1208) for a review of the research on the relationship between author count and citations.

[7]*JEL* codes were significantly revised in 1990; comparable codes are not available for periods pre- and post-reform. We therefore only control for *JEL* fixed effects in articles published after the reform.



Note. Graph (A) displays the stacked total number of female (pink) and male (blue) authors published in a top-five journal each year. Graph (B) is the (non-unique) number of male and female economists with a solo-authored paper; Graphs (C) and (D) plot the corresponding number of authors with a co-authored single-sex paper and a co-authored mixed-gendered paper. Graphs (E) and (F) are the stacked total number and percentage, respectively, of exclusively female-authored, mixed-gendered (green) and exclusively male-authored papers. Graphs (G) and (H) plot the total number of mixed-gendered papers with a strict majority of male and female co-authors and a male and female senior author, respectively; papers with an equal number of each gender or two or more senior authors of the opposite gender shown in green.

Figure 2: Gender composition of top-five publications

exclusively authored by four or more women.[8]

Moreover, women *do not* make up a greater share of authors on mixed-gendered papers. Journals are publishing more articles with at least one female author, but the number of male authors on these papers has increased slightly faster than the number of female authors—meaning the share of women among authors on mixed-gendered papers has actually *declined*. Graph (D) in Figure 2 plots the number of authors with a co-authored mixed-sex top-five paper each year. In the late 1980s, men and women were about equally represented. Since then, however, mixed-gendered papers have tended to generate more publications for men than they do for women. Graphs (E) and (F) reinforce this conclusion. They plot the number and percentage of single- and mixed-gendered papers published in top-five journals, respectively: the latter has increased; the number and percentage of exclusively female-authored papers is completely flat.

Finally, majority- and senior-female-authored papers are almost as rare today as they were 30 years ago. Very few majority-female mixed-sex papers were published in top-five journals before 2000; since then, they publish about four a year (Figure 2, Graph (G)). Meanwhile, the number of mixed-gendered papers with a majority or equal share of *male* authors has risen. The result is little or no growth in majority-female papers. Similarly, mixed-gendered papers with male senior authors have steadily increased since the late 1980s (Graph (H)). But growth in papers with a senior female author or male and female co-authors of equal rank is basically zero.

3 Gender differences in quality

3.1 Empirical strategy

Equation (1) represents the statistical relationship between gender and quality for papers published in top economics journals:

$$\hat{q}_{kjt} = \alpha_0 + \alpha_1 \,\text{female}_k + \eta_{jt} + \varepsilon_{kjt}, \tag{1}$$

where $\hat{q}_{kjt}$ is a hypothetical indicator that perfectly captures the quality of paper k published in journal j at time t, female_k represents the gender composition of paper k, η_{jt} absorb journal- and year/volume-specific selection effects and ε_{kjt} is an error term.

Although it does not identify a particular causal mechanism, a non-zero α_1 means papers published by one gender are higher quality than papers published by the opposite gender.[9] It may also imply that quality is not the overriding objective determining what is and is not published in top economics journals. To see this, assume papers are only accepted if their quality $q \sim \Phi_g$ exceeds the gender-specific threshold θ_g where Φ_g is a cumulative normal distribution function and $g \in \{M, F\}$. Let $\sigma_g^2(\theta_g)$ denote group g's variance in quality, conditional on acceptance and define α_1 as in Equation (1). If male- and female-authored papers are accepted at similar rates, conditional on submission ($\Phi_M(\theta_M) = \Phi_F(\theta_F)$), $\alpha_1 > 0$ but $\sigma_F^2(\theta_F) \leq \sigma_M^2(\theta_M)$, then female-authored papers have a lower probability of acceptance, conditional on their quality—*i.e.*, $\theta_M < \theta_F$ (Theorem 3.1).

Theorem 3.1. *Suppose Φ_g is a cumulative normal distribution function specific to group $g \in \{M, F\}$ and assume papers are only accepted if their quality $q \sim \Phi_g$ exceeds the threshold θ_g. If $\Phi_M(\theta_M) = \Phi_F(\theta_F)$, $\alpha_1 > 0$ and $\sigma_F^2(\theta_F) \leq \sigma_M^2(\theta_M)$, then $\theta_M < \theta_F$.*

Bringing Theorem 3.1 to the data poses two problems. First, the articles we evaluate have already been published, so we cannot directly test for average gender differences in acceptances. Other studies, however, suggest male- and female-authored papers are accepted at roughly similar rates. According to

[8] Since then, *AER* has published two and *REStud* one (current as of May 2019).

[9] *E.g.*, if female economists co-author with higher quality research teams or have access to better data, $\alpha_1 > 0$.

Blank (1991), 12.7 and 10.6 percent of male- and female-authored papers are accepted at *AER*. A recent study of four journals that semi-overlap with our own suggests exclusively male- and female-authored manuscripts receive a revise and resubmit decision 8 and 6 percent of the time, respectively (Card *et al.* 2020); the weighted average across all female-led papers is 7.4 percent (Card *et al.* 2018, Table 1, p. 39). For a review of other relevant studies, see Ceci *et al.* (2014, p. 7) and Hengel (2019, pp. 28–29).

Our second problem is that we do not know $\hat{q}_{kjt}$. We use citations to proxy. Conditional on $\hat{q}_{kjt}$, citations are assumed to depend on the additional factors represented in Equation (2):

$$\text{citations}_{kjt} = \beta_0 + \beta_1\,\text{female}_k + \gamma\,\hat{q}_{kjt} + \theta\,\mathbf{X}_k + \lambda_k + u_{kjt}, \tag{2}$$

where $\mathbf{X}_k$ is a vector of observable controls that affect citations conditional on quality (see Section 2). λ_k captures unobservable confounders specific to citations. β_1 picks up gender differences in citations conditional on $\hat{q}_{kjt}$ (and $\mathbf{X}_k$). u_{kjt} is the error term.

Let $\widetilde{\beta}_0 = \gamma\,\alpha_0 + \beta_0$, $\widetilde{\alpha}_1 = \gamma\,\alpha_1 + \beta_1$, $\widetilde{\varepsilon}_{kjt} = \gamma\,\varepsilon_{kjt} + u_{kjt}$ and combine Equations (1) and (2) as follows:

$$\text{citations}_{kjt} = \widetilde{\beta}_0 + \widetilde{\alpha}_1\,\text{female}_k + \theta\,\mathbf{X}_k + \eta_{jt} + \lambda_k + \widetilde{\varepsilon}_{kjt}. \tag{3}$$

If the gender composition of article k does not partially correlate with unobserved factors picked up by λ_k, OLS generates an unbiased estimate of $\widetilde{\alpha}_1$.[10] As long as citations are not biased in women's favour conditional on quality ($\beta_1 \leq 0$) and positively correlate with $\hat{q}_{kjt}$ ($\gamma > 0$), OLS *also* conservatively estimates α_1 in Equation (1) conditional on $\mathbf{X}_k$.

Prior research suggests both assumptions hold. Citations have been consistently found to positively correlate with alternative measures of research quality (see, *e.g.*, Stremersch *et al.* 2007; Buela-Casal and Zych 2010). Economists believe female-authored papers are cited less, holding quality constant (Card *et al.* 2020). Men are disproportionately more likely to cite their own (King *et al.* 2017) and other male-authored work (Dion *et al.* 2018; Dworkin *et al.* 2020; Ferber 1986).

3.2 Results

Table 1 displays results from OLS estimation of Equation (3). Column (1) suggests female-authored papers receive 16 log points more citations than male-authored papers conditional on author seniority at the time of publication (max. t), number of co-authors (N) and journal-year fixed effects. Adjusting for co-author prominence at the time citations were collected (max. T) increases the coefficient on female to 22 log points (column(2)).

Columns (3)–(5) estimate Equation (3) in the sub-sample of articles published after 1990 and controlling for primary and tertiary *JEL* categories. Column (3) omits *JEL* fixed effects; the coefficient on female is roughly equivalent to the estimate shown in column (2). It falls by 9–12 log points once field controls are added (columns (4) and (5)). We also re-estimate Equation (3) without $\mathbf{X}_k$ on the sample of articles published between 2000–2015.[11] Results are shown in columns (6) and (7). The coefficient on female is similar to the corresponding coefficients shown in columns (2) and (3).

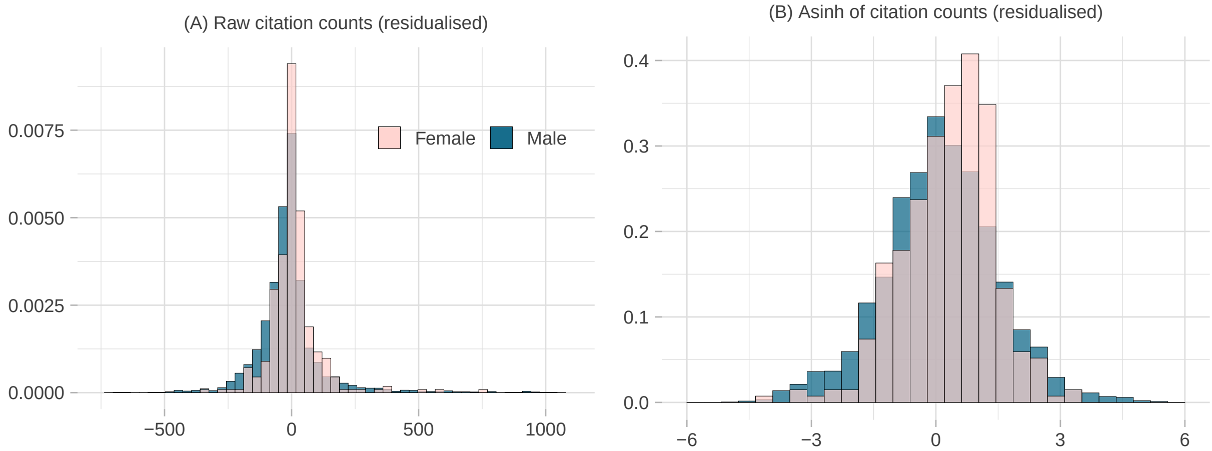
[10] λ_k almost certainly captures unobservables associated with authors and articles located in the right tail of the distribution (see Section 2.1 and Appendix F). Because they probably positively associate with citations and are located on a segment of the distribution with few female authors, they bias our estimate of $\widetilde{\alpha}_1$ *downward*. Thus, a positive OLS estimate of $\widetilde{\alpha}_1$ underrates the magnitude—but not the direction—of the coefficient's true value. For robustness, we also control for other observable factors that correlate with gender and might disproportionately impact right-tail citations, conditional on quality. Results are shown in Appendix F.

[11] Older male-authored papers likely drive the bulk of right-tail bias (see Section 2.1) and are the primary reason we control for $\mathbf{X}_k$ (and, especially, max. T_k). Their impact should, however, attenuate the closer an article is to its date of publication—citations to younger articles are less skewed (Bornmann and Leydesdorff 2017); in our own data, gender differences at the mean appear less influenced by superstar authors (see Appendix F).

Table 1: Gender differences in the quality of research output

	1950–2015		1990–2015			2000–2015	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
female	0.162***	0.218***	0.235***	0.148***	0.119**	0.194***	0.252***
	(0.051)	(0.050)	(0.051)	(0.051)	(0.055)	(0.060)	(0.061)
N	0.208***	0.192***	0.173***	0.164***	0.165***		
	(0.016)	(0.015)	(0.016)	(0.016)	(0.017)		
max. t	0.021***	−0.046***	−0.041***	−0.039***	−0.039***		
	(0.002)	(0.004)	(0.004)	(0.004)	(0.005)		
max. T		0.052***	0.047***	0.047***	0.047***		
		(0.002)	(0.003)	(0.003)	(0.003)		
Year×Journal	✓	✓	✓	✓	✓	✓	
JEL (primary)				✓			
JEL (tertiary)					✓		
Year							✓
$\sigma^2_M(\theta_M)$	1.723	1.640	1.134	1.081	0.883	1.028	1.110
$\sigma^2_F(\theta_F)$	0.935	0.910	0.867	0.832	0.422	0.855	0.982
Ratio p-value	0.000	0.000	0.000	0.000	0.000	0.001	0.022
Obs.	11,335	11,335	6,475	6,475	6,475	4,165	4,165
R^2	0.288	0.322	0.401	0.426	0.523	0.395	0.347

Note. Figures correspond to coefficients from OLS estimation of Equation (3). The dependent variable is citation counts (asinh). Independent variables are the ratio of female authors (female), number of co-authors (N), author seniority at the time of publication (max. t) and author prominence at the time citations were collected (max. T). Columns (1)–(2) are estimated on the entire sample. Columns (3)–(5) and (6)–(7) restrict the sample to articles published between 1990–2015 and 2000–2015, respectively. $\sigma^2_M(\theta_M)$ and $\sigma^2_F(\theta_F)$ are residual variances from estimating Equation (3) in the samples of entirely male-authored papers and papers with at least one female author, respectively (see Footnote 12). They are followed by p-values from testing the null hypothesis $\sigma^2_M(\theta_M)/\sigma^2_F(\theta_F) = 1$. Robust standard errors in parentheses. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.



Note. Graphs display the histograms of raw (left) and asinh transformed (right) citations for solo-authored papers (4,548 male-authored papers and 326 female-authored papers). Citations have been residualised with respect to the following controls: max. t, max. T and journal-year fixed effects.

Figure 3: Distribution of residualised citations, solo-authored papers

Figure 3 plots the distribution of citations—asinh (right) and raw counts (left)—for solo-authored papers, controlling for max. t, max. T and journal-year fixed effects. Consistent with results in Table 1, mean citations to women's papers are higher than they are to men's papers. Female-authored papers are also relatively absent from the left- (and to a lesser extent right-) hand tail of both distributions. This suggests the variance in women's citations is lower than men's.

$\sigma^2_M(\theta_M)$ and $\sigma^2_F(\theta_F)$ in Table 1 confirm this hypothesis. They represent residual variances from estimating Equation (3) in the sub-samples of entirely male-authored papers and papers with at least one female author, respectively;[12] they are followed by p-values from testing the null hypothesis $\sigma^2_M(\theta_M)/\sigma^2_F(\theta_F) = 1$. In line with Figure 3, $\sigma^2_F(\theta_F)$ is consistently lower than $\sigma^2_M(\theta_M)$. Assuming quality is normally distributed and male- and female-authored papers are accepted at roughly equivalent rates (conditional on controls) then women's higher mean citations but lower variance suggest they are subject to higher standards in peer review (Theorem 3.1).

Figure 3 and our estimates of $\sigma^2_M(\theta_M)$ and $\sigma^2_F(\theta_F)$ may also signal that higher standards operate in ways that affect more than just gender differences at the mean. In particular, women's sparser representation in the left-hand tail of the distribution (relative to the right-hand tail) could indicate that referees and editors are less willing to gamble on their riskiest work.[13]

Coefficients on remaining co-variates coincide with findings from previous research. Papers with more co-authors and by more prominent authors are cited more. The experience of an author at the time of publication is positive and significant when max. T is *not* controlled for (column (1)) but negative and significant when it *is* (columns (2)–(5))—*i.e.*, authors' publish their most cited work earlier in their careers, conditional on reputation.

[12] We do not control for the ratio of female authors in the sample of papers with at least one female author, but results are very similar if we do. See Appendix D for variance estimates comparing exclusively male-authored papers to papers: by a senior female author, with at least 50 percent female authors or only female authors.

[13] In any case, the "greater male variability" hypothesis is unlikely to be consistent with Figure 3. Gender differences in variability are equivalent to gender differences in (conditional) averages. Presumably, all academic economists publishing in top journals are drawn from the top half of the distribution of "quality". Thus, greater variability in men generally implies that average male quality is higher than average female quality, conditional on being an academic economist published in a top-five journal. (See also Ball *et al.* (2020) for similar arguments and evidence using citation data from fundamental physics.)

4 Returns to co-authoring with the opposite sex

4.1 Empirical specification

Equation (4) models the value added to author i from co-authoring his tth top-five paper with members of the opposite sex:

$$\text{citations}_{it} = \widetilde{\alpha}_1^{-i}\, \text{sex}_{it}^{-i} + \theta\, \mathbf{X}_{it} + \widetilde{\lambda}_i + \widetilde{\varepsilon}_{it}, \tag{4}$$

where sex_{it}^{-i} is the proportional contribution opposite sex co-authors ($-i \in \{M, F\}$) made to i's tth paper, $\mathbf{X}_{it}$ is a vector of time-varying controls that impact citations and correlate with sex_{it}^{-i}, $\widetilde{\lambda}_i$ are author-specific effects and $\widetilde{\varepsilon}_{it}$ is an error term.

As we show in Appendix A.2, if citations and quality positively correlate, the female composition of a paper is weakly negatively associated with citations conditional on quality and $\widetilde{\lambda}_i$ is independent of time, then a non-zero $\widetilde{\alpha}_1^{-i}$ from a fixed effects regression of Equation (4) indicates sex_{it}^{-i} determines the quality of i's top-five papers. Considered in isolation, however, $\widetilde{\alpha}_1^{-i} \neq 0$ is vague about why. For example, $\widetilde{\alpha}_1^{F} > 0$ might mean men work harder when they co-author with women or women work harder when they co-author with men (opposite-sex complementarities); $\widetilde{\alpha}_1^{M} < 0$ could suggest women are more productive when they collaborate with other women (same-sex complementaries). Both effects are also consistent with male authors contributing less to the quality of a paper, conditional on acceptance (sex-specific contributions).

$\widetilde{\alpha}_1^{M}$ and $\widetilde{\alpha}_1^{F}$ considered together provide more information. With opposite-sex complementaries, $\widetilde{\alpha}_1^{M}$ and $\widetilde{\alpha}_1^{F}$ will be positive; with same-sex complementaries, both should be negative. A positive value for one gender and negative value for the other suggests sex-specific contributions conditional on acceptance—*e.g.*, because papers by a particular gender are held to higher standards in peer review.

4.2 Results

To estimate Equation (4), we duplicate each article N_k times and assign observation k_n article k's nth$\in \{1, \ldots, N_k\}$ co-author. This generates a panel dataset following author i over the $t \in \{1, \ldots, T_i\}$ papers he publishes in top-five journals.

Table 2 displays results from fixed effects regressions on the resulting sub-samples of male (first panel) and female (second panel) authors, respectively. Assuming conditions outlined in Section 4.1 apply, coefficients on sex_{it}^{-i} represent the average change in quality authors would experience if they switched from co-authoring only with members of their own sex ($\text{sex}_{it}^{-i} = 0$) to co-authoring entirely with members of the opposite sex ($\text{sex}_{it}^{-i} = 1$), conditional on acceptance. Given i's own contribution to a paper implies $\text{sex}_{it}^{-i} < 1$, however, we always interpret and discuss the coefficient at $\text{sex}_{it}^{-i} = 0.5$.[14]

According to Table 2, the quality of men's papers *rises* when they co-author with the opposite sex whereas the quality of women's papers *falls*. Compared to the baseline of $\text{sex}_{it}^{F} = 0$, men's papers improve 10–12 log points when co-authored with an equal share of women (Table 2, first panel, columns (1)–(2)). The returns to women of co-authoring with an equal share of men, however, are -12 to -15 log points (second panel, columns (1)–(2)).

Moreover, results for both genders do not meaningfully change when articles published before 1990 are excluded or if *JEL* fixed effects are added (columns (3)–(6)).[15] This latter result may also suggest that the relationship between citations and gender does not depend on field after accounting for author fixed effects. Combined with results in Table 1 (columns (3)–(5)), it could also reveal an underlying association

[14] For the range of citation counts above 2–3—about 98 percent of our data—the impact of co-authoring with x percent sex_{it}^{-i} is approximately $\widetilde{\alpha}_1 \times x$ (Bellemare and Wichman 2020).

[15] The sample of female authors with two or more publications is too small to control for tertiary *JEL* categories.

Table 2: Returns to co-authoring with the opposite sex

	1950–2015		1990–2015			
	(1)	(2)	(3)	(4)	(5)	(6)
Returns to men from co-authoring with women						
sex^F	0.200**	0.239***	0.197**	0.227**	0.168*	0.211**
	(0.088)	(0.088)	(0.090)	(0.091)	(0.091)	(0.096)
N	0.165***	0.148***	0.161***	0.151***	0.145***	0.160***
	(0.016)	(0.016)	(0.016)	(0.016)	(0.015)	(0.017)
max. t	0.004	−0.012***	0.003	−0.019***	−0.018***	−0.019***
	(0.003)	(0.004)	(0.003)	(0.004)	(0.004)	(0.004)
max. T		0.021***		0.024***	0.025***	0.026***
		(0.003)		(0.003)	(0.004)	(0.003)
Returns to women from co-authoring with men						
sex^M	−0.238	−0.296**	−0.285*	−0.334**	−0.372**	
	(0.153)	(0.146)	(0.152)	(0.149)	(0.159)	
N	0.223***	0.224***	0.208***	0.211***	0.209***	
	(0.051)	(0.052)	(0.050)	(0.052)	(0.052)	
max. t	0.028***	−0.024	0.023***	−0.034*	−0.042**	
	(0.009)	(0.020)	(0.009)	(0.019)	(0.018)	
max. T		0.048***		0.052***	0.058***	
		(0.018)		(0.017)	(0.016)	
Year×Journal	✓	✓	✓	✓	✓	✓
JEL (primary)					✓	
JEL (tertiary)						✓

Note. Figures correspond to coefficients from fixed effects estimation of Equation (4). The dependent variable is citations (asinh). Independent variables are the ratio of opposite-sex co-authors (sex^{-i}), number of co-authors (N), author seniority at the time of publication (max. t) and author prominence at the time citations were collected (max. T). Panel one is estimated on the sample of male authors only. Columns (1)–(2) are estimated on all men (6,701 unique authors; 2,965 with two or more top-five publications); columns (3)–(6) restrict the sample to articles published between 1990–2015 (4,447 unique authors; 2,276 with two or more top-five publications). Panel two is similarly estimated on the sample of female authors only (873 unique authors and 287 authors with two or more top-five publications in columns (1)–(2); 751 unique authors and 270 authors with two or more top-five publications in columns (3)–(5)). Standard errors clustered at the author-level in parentheses. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table 3: Returns to senior men from co-authoring with junior women

	1950–2015		1990–2015	
	(1)	(2)	(3)	(4)
sex^F	0.302	0.599*	0.668*	1.352***
	(0.328)	(0.322)	(0.344)	(0.267)
max. $t = t$		−0.145***	−0.172***	−0.270**
		(0.048)	(0.049)	(0.074)
Year×Journal	✓	✓	✓	✓
JEL (primary)				✓
Obs.	242	242	179	179

Note. Figures correspond to coefficients from fixed effects estimation of Equation (4) on the sub-sample of authors and papers that satisfy the following criteria: senior male authors ($t_{it} = \text{max. } t_{it}$) with at least two top-five papers (max. $t_{it} > 1$) co-authored with exactly one ($N_{it} = 2$) economist of each sex ($\text{sex}_{it}^F = 0.5$ and $\text{sex}_{it'}^F = 0$, $t \neq t'$) who has no previous top-five publications (min. $t_{it} = 1$). Sample includes 73 senior men, 154 papers co-authored with exactly one junior man and 88 papers co-authored with exactly one junior woman. Standard errors clustered at the author level in parentheses. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

between field and author-specific unobservables that biases estimates of gender differences in citations that only control for the former.

As argued in Section 4.1, $\widetilde{\alpha}_1^F > 0$ and $\widetilde{\alpha}_1^M < 0$ imply female-specific contributions drive gender differences in quality, conditional on acceptance, and are consistent with Section 3's conclusion that women are held to higher standards in top-five peer review. Our estimates may, however, be biased by contributions from unobserved co-authors—*e.g.*, male economists may be more likely to collaborate with high-quality men on projects with at least one female author.

To account for this, we additionally limit our sample to papers and authors that satisfy the following criteria: senior male authors ($t_{it} = \text{max. } t_{it}$) with at least two top-five papers (max. $t_{it} > 1$) co-authored with exactly one ($N_{it} = 2$) junior (min. $t_{it} = 1$) economist of each sex ($\text{sex}_{it}^F = 0.5$ and $\text{sex}_{it'}^F = 0$, $t \neq t'$).[16] The subsequent sub-sample yields one treatment group—senior male economists co-authoring with exactly one junior woman—and one control group—those same men co-authoring with exactly one junior man. Appendix E.2 lists the names of all 73 senior men.

Coefficients from fixed effects regressions are shown in Table 3. Column (1) does not control for max. $t_{it} = t_{it}$: the coefficient on sex^F is similar in magnitude to estimates in Table 2, but its standard error is noticeably larger. As shown in Appendix E.1, t_{it} is somewhat imbalanced between treatment and control groups; after adjusting for it, $\widetilde{\alpha}_1^F$ doubles (column (2)) and becomes weakly significant. Excluding articles published before 1990 does not noticeably impact $\widetilde{\alpha}_1^F$ (column (3)). Adding *JEL* fixed effects does: $\widetilde{\alpha}_1^F$ doubles again and is highly significant (column (4)).

5 Conclusion

Male-authored papers published in top economics journals are cited less than female-authored papers. Moreover, citations to men's papers rise with they co-author with women whereas citations to women's papers fall when they co-author with men, conditional on acceptance. Under strong, but reasonable assumptions, we argue these findings suggest female authors face tougher standards in peer review.

Ideally, publication in a top journal would send a weaker signal about the quality of male-authored papers than it would about the quality of female-authored papers. In the real world, however, expectations are

[16] The number of senior female economists satisfying these conditions is too small to similarly analyse senior women.

slow to adjust (see, *e.g.*, Heckman and Moktan 2019). As a result, higher standards in peer review create higher standards for tenure and promotion. They also incentivise both genders to collaborate with men.

In economics, we tend to favour policies targeted at individual market imperfections. But when the space of information asymmetries and transaction costs is large and poorly understood, active policy solutions—including formal and informal quotas—may be sensible alternatives (Lundberg 1991; Lundberg and Startz 1983). They are not only objective and non-punitive but may also create positive externalities that could not have been achieved using markets alone (see, *e.g.*, Niederle *et al.* 2013; Besley *et al.* 2017). For example, clearly signalling a determination to publish more female authors will likely decrease the relative price of co-authoring with women and encourage more fruitful collaborations.

But active policy interventions are only Pareto improving when based on an adequate understanding of the context. More research is certainly needed. We hope journals are challenged to address the tougher standards they likely impose on women, willing to support the access and research needed to better understand them and open to whatever policy options most effectively check them.

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Appendices

A Proofs and derivations

A.1 Proofs

Proof of Theorem 3.1. Conditional on acceptance, the mean quality of papers by group $g \in \{M, F\}$ is

$$\mathbb{E}_g[q|q \geq \theta_g] = \int_{\theta_g}^{\infty} \frac{q\,\Phi_g'(q)}{1-\Phi_g(\theta_g)}\,\mathrm{d}q = \int_{\theta_g}^{\infty} \frac{1-\Phi_g(q)}{1-\Phi_g(\theta_g)}\,\mathrm{d}q + \theta_g, \tag{A.1}$$

where the last equality is obtained using integration by parts (see for example Hajeck (2015), p. 19; the Remark following this proof provides a full derivation). Thus, $\alpha_1 > 0$ is equivalent to

$$\int_{\theta_M}^{\infty} \frac{1-\Phi_M(q)}{1-\Phi_M(\theta_M)}\,\mathrm{d}q < \int_{\theta_M}^{\infty} \frac{1-\Phi_F(q)}{1-\Phi_F(\theta_F)}\,\mathrm{d}q - \int_{\theta_F}^{\theta_M} \frac{\Phi_F(q)-\Phi_F(\theta_F)}{1-\Phi_F(\theta_F)}\,\mathrm{d}q. \tag{A.2}$$

By way of a contradiction, assume $\theta_F \leq \theta_M$. Thus, $\Phi_F(\theta_F) \leq \Phi_F(q)$ for all $q \in (\theta_F, \theta_M)$, so Equation (A.2) together with $\Phi_M(\theta_M) = \Phi_F(\theta_F)$ implies

$$\int_{\theta_M}^{\infty} (1-\Phi_M(q))\,\mathrm{d}q < \int_{\theta_M}^{\infty} (1-\Phi_F(q))\,\mathrm{d}q. \tag{A.3}$$

Note that

$$\lim_{x\to\infty} \int_{y}^{x} \Phi_g(q)\,\mathrm{d}q = \infty. \tag{A.4}$$

Since Φ_F and Φ_M are continuous distributions, however, there exists a sufficiently large $\bar{q}$ such that Equation (A.3) implies

$$\int_{\theta_M}^{\bar{q}} \Phi_F(q)\,\mathrm{d}q < \int_{\theta_M}^{\bar{q}} \Phi_M(q)\,\mathrm{d}q. \tag{A.5}$$

Suppose $\sigma_M^2 = \sigma_F^2$. If $\mu_F \leq \mu_M$, then $\Phi_M(q) \leq \Phi_F(q)$ for all $q \in \mathbb{R}$, contradicting the inequality in Equation (A.5). But if $\mu_M < \mu_F$, $\Phi_F(q) < \Phi_M(q)$ for all $q \in \mathbb{R}$; combined with $\theta_F \leq \theta_M$, this implies

$$\Phi_F(\theta_F) \leq \Phi_F(\theta_M) < \Phi_M(\theta_M),$$

contradicting our assumption that $\Phi_F(\theta_F) = \Phi_M(\theta_M)$. Thus, $\sigma_M^2 \neq \sigma_F^2$.

Normal distributions are ordered in dispersion according to their variances (Lewis and Thompson 1981, Section 6.3). That is, the distribution with the greater variance dominates the other in the dispersive order (denoted by $<_{\text{disp}}$). $\Phi_g <_{\text{disp}} \Phi_{g'}$ and $\sigma_g^2 \neq \sigma_{g'}^2$ imply Φ_g intersects $\Phi_{g'}$ exactly once and from below (Shaked 1982, Theorem 2.1). Thus, $\Phi_{g'}(q) \leq \Phi_g(q)$ for all $q \geq q^\star$ where $q^\star < \infty$ uniquely satisfies $\Phi_g(q^\star) = \Phi_{g'}(q^\star)$.

If $q^\star \leq \theta_M$, then Equation (A.5) implies that Φ_M lies above Φ_F for all $q \geq q^\star$. To see that the same is true when $\theta_M < q^\star$, rewrite Equation (A.5) as

$$\int_{q^\star}^{\bar{q}} \Phi_F(q)\,\mathrm{d}q + \int_{\theta_M}^{q^\star} \Phi_F(q)\,\mathrm{d}q < \int_{q^\star}^{\bar{q}} \Phi_M(q)\,\mathrm{d}q + \int_{\theta_M}^{q^\star} \Phi_M(q)\,\mathrm{d}q. \tag{A.6}$$

As $\bar{q} \to \infty$, the limits of the first terms on each side of the inequality in Equation (A.6) are infinite (Equation (A.4)) whereas the second terms are not. Thus, for a sufficiently large $\bar{q}'$, Equation (A.6) implies

$$\int_{q^\star}^{\bar{q}'} \Phi_F(q)\,\mathrm{d}q < \int_{q^\star}^{\bar{q}'} \Phi_M(q)\,\mathrm{d}q.$$

We therefore conclude that Φ_M lies above Φ_F for all $q \geq q^\star$. Thus, $\Phi_M <_{\text{disp}} \Phi_F$ and so $\sigma_M^2 < \sigma_F^2$ and also $\sigma_M^2(\theta_M) < \sigma_F^2(\theta_F)$ (without proof). This establishes the desired contradiction. $\square$

Remark (Derivation of Equation A.1). Recall from the first part of Equation (A.1) that

$$\begin{aligned}\mathbb{E}_g[q|q \geq \theta_g] &= \int_{\theta_g}^{\infty} \frac{q\,\Phi_g'(q)}{1-\Phi_g(\theta_g)}\,\mathrm{d}q \\ &= -\frac{1}{1-\Phi_g(\theta_g)} \int_{\theta_g}^{\infty} q\,\mathrm{d}(1-\Phi_g(q)). \end{aligned} \tag{A.7}$$

Using integration by parts on the last step of Equation (A.7), we get

$$\begin{aligned} &= -\frac{1}{1-\Phi_g(\theta_g)}\left(\lim_{q\to\infty}\{q(1-\Phi_g(q))\} - \theta_g\,(1-\Phi(\theta_g)) - \int_{\theta_g}^{\infty}(1-\Phi_g(q))\,\mathrm{d}q\right) \\ &= \int_{\theta_g}^{\infty} \frac{1-\Phi_g(q)}{1-\Phi_g(\theta_g)}\,\mathrm{d}q + \theta_g - \frac{1}{1-\Phi_g(\theta_g)} \lim_{q\to\infty} q(1-\Phi_g(q)). \end{aligned} \tag{A.8}$$

It remains to show that the limit in Equation (A.8) is zero. Note that

$$\lim_{q\to\infty} q(1-\Phi_g(q)) = \lim_{q\to\infty} \frac{1-\Phi_g(q)}{1/q}.$$

Applying L'Hôpital's rule, we have

$$\lim_{q\to\infty} \frac{1-\Phi_g(q)}{1/q} = \lim_{q\to\infty} \frac{\Phi_g'(q)}{1/q^2}. \tag{A.9}$$

Since Φ_g' is the density function for the normal distribution, Equation (A.9) is equivalent to

$$\begin{aligned}\lim_{q\to\infty} \frac{\Phi_g'(q)}{1/q^2} &= \lim_{q\to\infty} \frac{\frac{1}{\sqrt{2\pi\sigma_g^2}}\exp\left\{-\frac{(q-\mu_g)^2}{2\pi\sigma_g^2}\right\}}{1/q^2} \\ &= \frac{1}{\sqrt{2\pi\sigma_g^2}} \lim_{q\to\infty} \frac{q^2}{\exp\left\{\frac{(q-\mu_g)^2}{2\pi\sigma_g^2}\right\}} \\ &= 0. \end{aligned}$$

A.2 Section 4.1, full empirical strategy

Equation (A.10) models the value added to author i from co-authoring his tth top-five paper with members of the opposite sex:

$$\hat{q}_{it} = \alpha_1^{-i}\,\text{sex}_{it}^{-i} + \theta^{\hat{q}}\,\mathbf{X}_{it}^{\hat{q}} + \alpha_i + \varepsilon_{it}, \tag{A.10}$$

where sex_{it}^{-i} is the proportional contribution of opposite sex co-authors ($-i \in \{M, F\}$) to i's tth paper. As before, $\hat{q}_{it}$ is a hypothetical indicator that perfectly captures paper it's quality. $\mathbf{X}_{it}^{\hat{q}}$ is a vector of time-varying controls that impact quality and correlate with sex_{it}^{-i}. α_i are author-specific effects. ε_{it} is an error term.

Unfortunately, $\hat{q}_{it}$ is not known. Again, we use citations to proxy. As in Section 3, we assume $\hat{q}_{it}$ positively relates to citations as well as the additional factors shown in Equation (A.11):

$$\text{citations}_{it} = \beta_1^{-i}\,\text{sex}_{it}^{-i} + \gamma\,\hat{q}_{it} + \theta^c\,\mathbf{X}_{it}^c + \lambda_i + u_{it}, \tag{A.11}$$

where $\mathbf{X}_{it}^c$ is a vector of observable controls that affect citations conditional on $\hat{q}_{it}$—*e.g.*, the field in which a paper was published—and β_1^{-i} measures the impact sex_{it}^{-i} has on citations conditional on $\hat{q}_{it}$ and $\mathbf{X}_{it}^c$. λ_i captures residual factors specific to citations and author i. u_{it} is the error term.

Combine Equation (A.10) with Equation (A.11) and let $\widetilde{\alpha}_1^{-i} = \gamma\,\alpha_1^{-i} + \beta_1^{-i}$, $\widetilde{\lambda}_i = \gamma\,\alpha_i + \lambda_i$, $\widetilde{\varepsilon}_{it} = \gamma\,\varepsilon_{it} + u_{it}$ and $\theta\,\mathbf{X}_{it}$ represent $\theta^{\hat{q}}\,\mathbf{X}_{it}^{\hat{q}} + \theta^c\,\mathbf{X}_{it}^c$:

$$\text{citations}_{it} = \widetilde{\alpha}_1^{-i}\,\text{sex}_{it}^{-i} + \theta\,\mathbf{X}_{it} + \widetilde{\lambda}_i + \widetilde{\varepsilon}_{it}. \tag{A.12}$$

If citations and quality positively correlate ($\gamma > 0$), the female composition of a paper is weakly negatively associated with citations conditional on quality ($\beta_1^F \leq 0$ and $\beta_1^M \geq 0$) and $\widetilde{\lambda}_i$ is independent of time, conditional on $\mathbf{X}_{it}$, then $\widetilde{\alpha}_1^{-i}$ from a fixed effects regression of Equation (A.12) conservatively estimates α_1^{-i} in Equation (A.10).[1]

[1]See Section 3.1 for evidence justifying $\beta_1^F \leq 0$ and $\beta_1^M \geq 0$. Note also that our assumption of time-independence is stronger than required; a fixed effects estimate of $\widetilde{\alpha}_1^{-i}$ is consistent or conservative as long as the time-varying component of $\widetilde{\lambda}_i$ doesn't partially correlate with sex_{it}^{-i}.

B Quantile regression

Table B.1 re-estimates Table 1 using a quantile regression model and raw citation counts as the dependent variable. The first panel replicates Table 1, column (2) at the 25th, median and 75th percentiles of review times; the second panel similarly replicates column (4). The coefficient on female ratio is positive across all three percentiles.

Table B.1: Table 1, quantile regression

	1950–2015			1990–2015		
	25 pc.	50 pc.	75 pc.	25 pc.	50 pc.	75 pc.
	(1)	(2)	(3)	(4)	(5)	(6)
female	4.618***	6.583***	5.714*	4.396**	6.271**	1.569
	(1.591)	(2.107)	(3.180)	(1.992)	(2.528)	(3.716)
N	3.109***	4.812***	7.000***	3.276***	4.301***	5.916***
	(0.468)	(0.660)	(1.159)	(0.618)	(0.796)	(1.193)
max. t	−1.551***	−2.902***	−6.173***	−2.887***	−4.273***	−7.272***
	(0.164)	(0.211)	(0.680)	(0.238)	(0.408)	(0.922)
max. T	1.583***	3.159***	6.745***	2.850***	4.556***	7.666***
	(0.132)	(0.171)	(0.594)	(0.205)	(0.375)	(0.861)
Year×Journal	✓	✓	✓	✓	✓	✓
JEL (primary)				✓	✓	✓

Note. First panel replicates results shown in Table 1, column (2) across different percentiles of the distribution using quantile regressions and raw citation counts as the dependent variable; second panel similarly replicates results from column (4). Robust standard errors in parentheses. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

C Alternative dependent variables

The following sections replicate results presented in Tables 1, 2 and 3 using the log of 1 plus citations (Section C.1) and raw citation counts (Section C.2) as dependent variables.

C.1 Log citations

Table C.1: Table 1, log of 1 + citations as the dependent variable

	1950–2015		1990–2015			2000–2015	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
female	0.150***	0.204***	0.221***	0.139***	0.114**	0.180***	0.236***
	(0.048)	(0.047)	(0.049)	(0.048)	(0.052)	(0.057)	(0.058)
N	0.197***	0.182***	0.166***	0.158***	0.159***		
	(0.015)	(0.014)	(0.015)	(0.015)	(0.016)		
max. t	0.020***	−0.044***	−0.040***	−0.038***	−0.038***		
	(0.002)	(0.003)	(0.004)	(0.004)	(0.004)		
max. T		0.049***	0.046***	0.046***	0.046***		
		(0.002)	(0.003)	(0.003)	(0.003)		
Year×Journal	✓	✓	✓	✓	✓	✓	
JEL (primary)				✓			
JEL (tertiary)					✓		
Year							✓
$\sigma^2_M(\theta_M)$	1.530	1.454	1.033	0.985	0.804	0.914	0.989
$\sigma^2_F(\theta_F)$	0.842	0.818	0.783	0.751	0.379	0.761	0.876
Ratio p-value	0.000	0.000	0.000	0.000	0.000	0.001	0.023
Obs.	11,335	11,335	6,475	6,475	6,475	4,165	4,165
R^2	0.289	0.323	0.401	0.426	0.523	0.395	0.346

Note. Columns display estimates identical to those in Table 1, except that the dependent variable is the natural logarithm of 1 plus an article's raw citation count. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table C.2: Table 2, log of 1 + citations as the dependent variable

	1950–2015		1990–2015			
	(1)	(2)	(3)	(4)	(5)	(6)
Returns to men from co-authoring with women						
sex^F	0.189**	0.228***	0.187**	0.217**	0.161*	0.204**
	(0.084)	(0.084)	(0.087)	(0.087)	(0.087)	(0.092)
N	0.158***	0.141***	0.155***	0.145***	0.139***	0.153***
	(0.015)	(0.015)	(0.015)	(0.015)	(0.015)	(0.016)
max. t	0.004	−0.012***	0.003	−0.019***	−0.019***	−0.020***
	(0.003)	(0.004)	(0.003)	(0.004)	(0.004)	(0.004)
max. T		0.020***		0.024***	0.025***	0.026***
		(0.003)		(0.003)	(0.003)	(0.003)
Returns to women from co-authoring with men						
sex^M	−0.222	−0.279**	−0.273*	−0.321**	−0.358**	
	(0.147)	(0.141)	(0.147)	(0.144)	(0.152)	
N	0.217***	0.218***	0.202***	0.206***	0.203***	
	(0.049)	(0.050)	(0.048)	(0.050)	(0.050)	
max. t	0.026***	−0.025	0.022***	−0.034*	−0.042**	
	(0.008)	(0.019)	(0.008)	(0.018)	(0.017)	
max. T		0.047***		0.051***	0.058***	
		(0.017)		(0.016)	(0.015)	
Year×Journal	✓	✓	✓	✓	✓	✓
JEL (primary)					✓	
JEL (tertiary)						✓

Note. Columns display estimates identical to those in Table 2, except that the dependent variable is the natural logarithm of 1 plus an article's raw citation count. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table C.3: Table 3, log of 1 + citations as the dependent variable

	1950–2015		1990–2015	
	(1)	(2)	(3)	(4)
sex^F	0.270	0.565*	0.637*	1.329***
	(0.316)	(0.308)	(0.340)	(0.258)
max. $t = t$		−0.144***	−0.173***	−0.272**
		(0.047)	(0.049)	(0.071)
Year×Journal	✓	✓	✓	✓
JEL (primary)				✓
Obs.	242	242	179	179

Note. Columns display estimates identical to those in Table 3, except that the dependent variable is the natural logarithm of 1 plus an article's raw citation count. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

C.2 Raw citations

The following tables replicate Tables 1, 2 and 3 using raw citation counts as the dependent variable. Tables C.5 and C.6 are otherwise identical to their corresponding tables in Section 4. Because non-transformed citations are sensitive to a very small number of "superstar" authors, estimates shown in Table C.4 always control for N, max. t and max. T. Results that do not control for these variables are shown in Appendix F.

Table C.4: Table 1, raw citations as the dependent variable

	1950–2015	1990–2015			2000–2015		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
female	0.003	9.090	1.564	3.782	16.855**	11.319	9.920
	(6.731)	(7.581)	(7.732)	(8.894)	(6.925)	(7.118)	(7.344)
N	21.996***	18.939***	18.068***	19.206***	10.337***	10.157***	11.074***
	(4.302)	(3.539)	(3.485)	(3.854)	(2.008)	(1.986)	(2.357)
max. t	−6.576***	−6.797***	−6.562***	−7.017***	−5.624***	−5.665***	−4.322***
	(0.993)	(1.131)	(1.126)	(1.071)	(2.176)	(2.063)	(1.051)
max. T	6.502***	6.851***	6.697***	6.928***	5.510***	5.585***	4.326***
	(0.803)	(0.971)	(0.966)	(0.869)	(1.938)	(1.830)	(0.913)
Year×Journal	✓	✓	✓	✓	✓	✓	✓
JEL (primary)			✓			✓	
JEL (tertiary)				✓			✓
Obs.	11,335	6,475	6,475	6,475	4,165	4,165	4,165
R^2	0.072	0.139	0.150	0.252	0.211	0.227	0.405

Note. Columns display estimates similar to those in Table 1, except that the dependent variable is an article's raw citation count. (See Appendix F for results that do not control for N, max. t and max. T.) ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table C.5: Table 2, raw citations as the dependent variable

	1950–2015		1990–2015			
	(1)	(2)	(3)	(4)	(5)	(6)
Returns to men from co-authoring with women						
sex^F	21.462	28.503*	14.609	21.384	16.746	39.343*
	(16.675)	(16.664)	(18.418)	(18.129)	(17.964)	(23.678)
N	17.077***	13.884***	21.891***	19.504***	18.881***	19.605***
	(4.335)	(4.209)	(4.332)	(4.054)	(4.040)	(4.030)
max. t	0.011	−2.897**	−0.724	−5.713***	−5.535***	−4.349***
	(0.715)	(1.282)	(0.676)	(1.555)	(1.549)	(1.249)
max. T		3.733***		5.525***	5.410***	4.702***
		(0.979)		(1.373)	(1.374)	(1.064)
Returns to women from co-authoring with men						
sex^M	−8.889	−20.910	−21.175	−31.441	−30.281	
	(33.957)	(33.946)	(35.218)	(35.320)	(34.586)	
N	23.883***	24.224***	21.583***	22.300***	19.389**	
	(8.307)	(8.420)	(8.263)	(8.423)	(8.963)	
max. t	2.106	−8.589*	1.757	−10.108**	−11.823***	
	(1.431)	(4.373)	(1.413)	(4.589)	(4.366)	
max. T		9.922**		10.925***	12.340***	
		(4.048)		(4.135)	(3.749)	
Year×Journal	✓	✓	✓	✓	✓	✓
JEL (primary)					✓	
JEL (tertiary)						✓

Note. Columns display estimates identical to those in Table 2, except that the dependent variable is an article's raw citation count. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table C.6: Table 3, raw citations as the dependent variable

	1950–2015		1990–2015	
	(1)	(2)	(3)	(4)
sex^F	−17.516	50.345	74.417*	211.099***
	(47.429)	(39.971)	(42.038)	(22.120)
max. $t = t$		−33.214***	−32.036***	−56.197***
		(8.782)	(7.444)	(3.663)
Year×Journal	✓	✓	✓	✓
JEL (primary)				✓
Obs.	242	242	179	179

Note. Columns display estimates identical to those in Table 3, except that the dependent variable is an article's raw citation count. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

D Alternative proxies for article gender

The following tables replicate Table 1 using alternative definitions of female authorship. The first four tables compare entirely male-authored papers to papers with a senior female author (Table D.1), a weak majority of female authors (Table D.2), at least one female author (Table D.3) and a solo female author (Table D.4). Mixed-gendered papers not satisfying the relevant "female" criteria in Tables D.1, D.2 and D.3 and all co-authored papers in Table D.4 are dropped.

The final table uses a categorical variable to account for seven different gender categories: (i) female solo-authored, (ii) female co-authored, (iii) mixed sex co-authored with a senior female author, (iv) mixed sex co-authored with senior male and female authors of equal rank, (v) mixed sex co-authored with a senior male author, (iv) male solo-authored and (vii) male co-authored.

Table D.1: Table 1, senior female author

	1950–2015		1990–2015			2000–2015	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
female	0.099**	0.125***	0.139***	0.083**	0.066	0.080*	0.130***
	(0.043)	(0.042)	(0.042)	(0.042)	(0.045)	(0.047)	(0.048)
N	0.208***	0.192***	0.173***	0.164***	0.165***		
	(0.016)	(0.015)	(0.016)	(0.016)	(0.017)		
max. t	0.021***	−0.045***	−0.040***	−0.039***	−0.039***		
	(0.002)	(0.004)	(0.004)	(0.004)	(0.005)		
max. T		0.052***	0.047***	0.046***	0.047***		
		(0.002)	(0.003)	(0.003)	(0.003)		
Year×Journal	✓	✓	✓	✓	✓	✓	
JEL (primary)				✓			
JEL (tertiary)					✓		
Year							✓
$\sigma^2_M(\theta_M)$	1.686	1.607	1.126	1.076	0.887	1.027	1.108
$\sigma^2_F(\theta_F)$	0.763	0.736	0.699	0.659	0.196	0.703	0.922
Ratio p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.011
Obs.	11,335	11,335	6,475	6,475	6,475	4,165	4,165
R^2	0.288	0.321	0.400	0.426	0.523	0.394	0.345

Note. Columns display estimates identical to those in Table 1, except that the independent variable female is equal to a dummy variable equal to 1 if a female author had at least as many top-five papers as her co-authors at the time the paper was published (758 papers). (Mixed-gendered papers with a senior male co-author are excluded.) ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table D.2: Table 1, majority female-authored

	1950–2015		1990–2015			2000–2015	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
female	0.120***	0.158***	0.165***	0.110***	0.098**	0.131***	0.169***
	(0.037)	(0.037)	(0.038)	(0.038)	(0.041)	(0.043)	(0.044)
N	0.207***	0.189***	0.168***	0.164***	0.165***		
	(0.017)	(0.017)	(0.018)	(0.017)	(0.019)		
max. t	0.020***	−0.046***	−0.041***	−0.039***	−0.039***		
	(0.002)	(0.004)	(0.004)	(0.004)	(0.005)		
max. T		0.051***	0.046***	0.046***	0.047***		
		(0.002)	(0.003)	(0.003)	(0.003)		
Year×Journal	✓	✓	✓	✓	✓	✓	
JEL (primary)				✓			
JEL (tertiary)					✓		
Year							✓
$\sigma^2_M(\theta_M)$	1.723	1.640	1.134	1.081	0.883	1.028	1.110
$\sigma^2_F(\theta_F)$	0.891	0.874	0.826	0.791	0.301	0.810	0.962
Ratio p-value	0.000	0.000	0.000	0.000	0.000	0.000	0.025
Obs.	10,930	10,930	6,094	6,094	6,094	3,842	3,842
R^2	0.282	0.316	0.395	0.421	0.521	0.393	0.345

Note. Columns display estimates identical to those in Table 1, except that the independent variable female is a dummy variable equal to 1 if a weak majority (50% or more) of authors are female (1,046 papers). (Papers with a minority—but positive—number of female authors are excluded.) ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table D.3: Table 1, at least one female author

	1950–2015		1990–2015			2000–2015	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
female	0.101***	0.141***	0.149***	0.092***	0.080**	0.182***	0.221***
	(0.033)	(0.032)	(0.033)	(0.033)	(0.035)	(0.037)	(0.038)
N	0.201***	0.182***	0.160***	0.157***	0.159***		
	(0.016)	(0.015)	(0.016)	(0.016)	(0.017)		
max. t	0.020***	−0.046***	−0.041***	−0.039***	−0.039***		
	(0.002)	(0.004)	(0.004)	(0.004)	(0.005)		
max. T		0.052***	0.047***	0.047***	0.047***		
		(0.002)	(0.003)	(0.003)	(0.003)		
Year×Journal	✓	✓	✓	✓	✓	✓	
JEL (primary)				✓			
JEL (tertiary)					✓		
Year							✓
$\sigma^2_M(\theta_M)$	1.723	1.640	1.134	1.081	0.883	1.028	1.110
$\sigma^2_F(\theta_F)$	0.935	0.910	0.867	0.832	0.422	0.855	0.982
Ratio p-value	0.000	0.000	0.000	0.000	0.000	0.001	0.022
Obs.	11,335	11,335	6,475	6,475	6,475	4,165	4,165
R^2	0.288	0.322	0.401	0.426	0.523	0.397	0.349

Note. Columns display estimates identical to those in Table 1, except that the independent variable female is a dummy variable equal to 1 if at least one author on a paper is female (1,451 papers). ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table D.4: Table 1, solo-authored papers

	1950–2015		1990–2015			2000–2015	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
female	0.135*	0.196***	0.189**	0.095	0.087	0.195**	0.162*
	(0.071)	(0.071)	(0.075)	(0.073)	(0.085)	(0.093)	(0.090)
max. t	0.029***	−0.050***	−0.050***	−0.045***	−0.045***		
	(0.005)	(0.007)	(0.009)	(0.009)	(0.011)		
max. T		0.060***	0.056***	0.053***	0.052***		
		(0.004)	(0.007)	(0.007)	(0.008)		
Year×Journal	✓	✓	✓	✓	✓	✓	
JEL (primary)				✓			
JEL (tertiary)					✓		
Year							✓
$\sigma_M^2(\theta_M)$	1.990	1.890	1.155	1.096	0.727	0.974	1.098
$\sigma_F^2(\theta_F)$	0.644	0.635	0.613	0.491	—	0.613	0.943
Ratio p-value	0.000	0.000	0.000	0.000	—	0.001	0.245
Obs.	4,874	4,874	1,916	1,916	1,916	1,031	1,031
R^2	0.224	0.262	0.350	0.383	0.588	0.382	0.316

Note. Columns display estimates identical to those in Table 1, except that the independent variable female is a dummy variable equal to 1 if the paper was solo-authored by a woman (326 papers) and zero if it was solo-authored by a man (4,548 papers). (All co-authored papers are excluded.) Due to small sample sizes, we omit $\sigma_F^2(\theta_F)$ from column (5). ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table D.5: Table 1, multiple gender categories

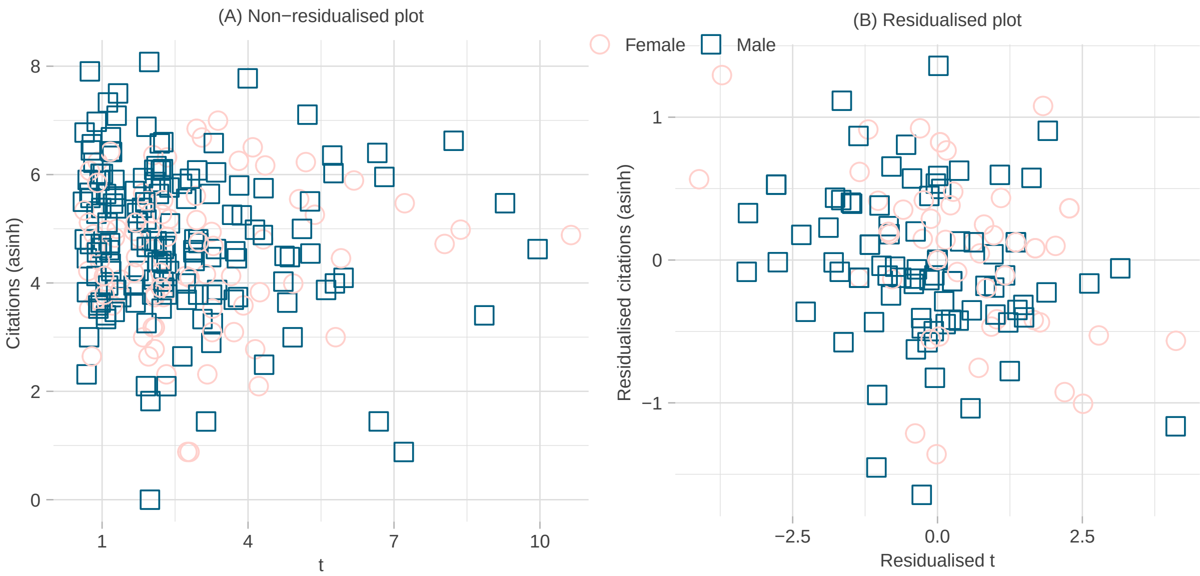
	1950–2015		1990–2015			2000–2015	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Female solo	−0.019	0.081	0.166**	0.096	0.044	−0.096	−0.099
	(0.074)	(0.073)	(0.077)	(0.076)	(0.081)	(0.086)	(0.086)
Female co-authored	0.271**	0.308***	0.397***	0.288**	0.318**	0.297***	0.447***
	(0.119)	(0.118)	(0.115)	(0.117)	(0.124)	(0.107)	(0.114)
Mixed sex, senior female	0.132	0.160**	0.181**	0.119	0.192**	0.184**	0.261***
	(0.082)	(0.081)	(0.082)	(0.082)	(0.093)	(0.090)	(0.092)
Mixed sex, equal rank	0.029	0.046	0.095	0.054	0.006	0.044	0.077
	(0.075)	(0.073)	(0.071)	(0.071)	(0.074)	(0.073)	(0.077)
Mixed sex, senior male	0.101**	0.148***	0.144***	0.094**	0.088*	0.195***	0.210***
	(0.046)	(0.045)	(0.047)	(0.046)	(0.049)	(0.051)	(0.052)
Male solo	−0.151***	−0.088**	0.034	0.029	0.028	−0.221***	−0.254***
	(0.042)	(0.041)	(0.046)	(0.045)	(0.048)	(0.041)	(0.042)
N	0.130***	0.142***	0.173***	0.167***	0.166***		
	(0.023)	(0.023)	(0.023)	(0.022)	(0.024)		
max. t	0.020***	−0.046***	−0.041***	−0.039***	−0.039***		
	(0.002)	(0.004)	(0.004)	(0.004)	(0.005)		
max. T		0.052***	0.047***	0.047***	0.047***		
		(0.002)	(0.003)	(0.003)	(0.003)		
Year×Journal	✓	✓	✓	✓	✓	✓	
JEL (primary)				✓			
JEL (tertiary)					✓		
Year							✓
$\sigma^2_M(\theta_M)$	1.723	1.640	1.134	1.081	0.883	1.028	1.110
$\sigma^2_F(\theta_F)$	0.638	0.633	0.608	0.513	—	0.614	0.909
Ratio p-value	0.000	0.000	0.000	0.000	—	0.000	0.075
Obs.	11,335	11,335	6,475	6,475	6,475	4,165	4,165
R^2	0.289	0.322	0.402	0.427	0.524	0.403	0.358

Note. Columns display estimates identical to those in Table 1, except that the independent variable female is replaced by a categorical variable that classifies papers as female solo-authored (326), female co-authored (50), mixed-sex co-authored with a senior female author (148), mixed-sex co-authored by men and women of equal rank (234), mixed-sex co-authored with a senior male author (693), male solo-authored (4,548) and the reference category, male co-authored (5,337 papers). Residual variances are estimated on the samples of exclusively male- and female-authored papers, respectively. Due to small sample sizes, we omit $\sigma^2_F(\theta_F)$ from column (5). ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

E Section 4, supplemental output

E.1 Table 3, covariate balance

By design, the sample of senior authors used to estimate Table 3 fixes N and max. T, conditional on author. For each author, however, t varies over time and appears somewhat imbalanced between treatment and control groups, particularly after accounting for author-specific fixed effects and journal-year interaction dummies (Figure E.1)—*i.e.*, conditional on author, year and journal, the senior men in our sample were slightly more experienced when they co-authored with junior women than they were when they co-authored with junior men. For that reason, we additionally control for max. t in columns (2)–(4) of Table 3.



Note. Graph (A) plots max. $t = t$ (x-axis) against asinh-transformed citations (y-axis) by co-author sex for the sample of senior male authors satisfying the conditions outlined in Section 4.2 and Table 3. Graph (B) plots the residuals of both variables after accounting for author-specific fixed effects and journal-year interaction dummies.

Figure E.1: Max. t balance among senior men

E.2 Table 3, list of senior men

Table E.1: Table 3, list of senior men

Acemoglu, Daron	Gastwirth, Joseph L.	Moulin, Hervé
Alesina, Alberto	Gertler, Mark	Muller, Ulrich K.
Andreoni, James	Hall, Robert E.	Palfrey, Thomas R.
Andrews, Donald W. K.	Hamilton, James D.	Pesendorfer, Martin
Ashenfelter, Orley	Helpman, Elhanan	Phillips, Peter C. B.
Barro, Robert J.	Hong, Yongmiao	Plott, Charles R.
Barsky, Robert B.	Hopenhayn, Hugo A.	Polemarchakis, Herakles
Bernheim, B. Douglas	Horowitz, Joel L.	Pollak, Robert A.
Boldrin, Michele	Jacoby, Hanan G.	Ray, Debraj
Borjas, George J.	Jovanovic, Boyan	Restuccia, Diego
Bronars, Stephen G.	Karni, Edi	Robin, Jean-Marc
Browning, Martin J.	Knight, Brian	Rodríguez-Clare, Andrés
Burdett, Kenneth	Kotlikoff, Laurence J.	Roth, Alvin E.
Card, David E.	Kremer, Michael	Rubinstein, Ariel
Chiappori, Pierre-André	Krishna, Pravin	Smith, Lones
Cochrane, John H.	Krueger, Alan B.	Smith, V. Kerry
Cogley, Timothy	Kuhn, Peter	Stockman, Alan C.
Crawford, Vincent P.	Libecap, Gary D.	Tollison, Robert D.
Deneckere, Raymond J.	Matthews, Steven A.	Waldfogel, Joel
Dow, Gregory K.	Mertens, Jean-François	Weibull, Jörgen W.
Duffy, John	Milgrom, Paul R.	Weinstein, David E.
Ehrenberg, Ronald G.	Miller, Robert A.	White, Halbert
Feldstein, Martin S.	Moen, Espen R.	Wright, Randall
Flinn, Christopher J.	Mookherjee, Dilip	
Frankel, Jeffrey A.	Morgan, John	

F Right-tail confounders

In order to illustrate how gender differences in raw citation counts at the mean may be distorted by a small number of extremely famous—and disproportionately male—economists, we control for "superstar" (Appendix F.1) and Nobel Prize winning authors (Appendix F.2).

F.1 Superstar authors

We define "superstars" as authors who satisfy one or more of the following criteria:

1. 17 or more top-five publications (1 percent of all authors);
2. 10 or more top-five publications, one of which is cited at least 2,500 times (0.2 percent of all authors);
3. 5 or more top-five publications, one of which is cited at least 5,000 times (0.1 percent of all authors).

The first criteria defines superstar according to quantity, alone. It is set as one plus the lifetime number of publications of the most prolific female economist as of December 2015 (Esther Duflo). Criteria two and three account for famous economists who are less prolific—*e.g.*, Paul Krugman—operate in fields with slower production functions—*e.g.*, industrial organisation—or publish extensively in other disciplines—*e.g.*, Daniel Kahneman. General results and conclusions do not change by making marginal adjustments to any criteria—including redefining condition (1) to include every male and female author with at least 10–15 publications.

1.2 percent of authors satisfy at least one condition. On average, each has published 21 times in a top-five journal; his highest cited paper is cited 1,844 times. Almost a third either won the Nobel Prize, the John Bates Clark medal or both. All are male. See Table F.1 for a list of their names.

F.1.1 Results

Tables F.2–F.4 illustrate the effect of superstardom on gender differences in raw citation counts using articles as the unit of analysis. Table F.2 is estimated using all observations. Column (1) controls only for journal-year fixed effects and the female composition of a paper. It suggests that male-authored papers are cited, on average, 15 more times than female-authored papers. The sign on the coefficient flips, however, after including the superstar dummy (column (2)). Column (3) adds fixed effects for each superstar author; results are similar to those in column (2). Columns (4)–(9) control for N, max. t and max. T. The coefficient on female generally hovers around zero, but jumps to 10 in the final column.

Older male-authored papers likely drive the bulk of superstar bias. Their impact, however, should attenuate the closer an article is to its date of publication. Tables F.3 and F.4 support this hypothesis. They reproduce results from Table F.2, but restrict the sample to papers published after 1990 and 2000, respectively. The coefficients on female in Table F.3 are universally larger than corresponding figures from Table F.2; the estimate in the final column suggests female-authored papers are cited 16 more times than male-authored papers. When data are restricted to articles published after 2000, female-authored papers are consistently cited more frequently than male-authored papers. Moreover, superstardom has little or no impact on the observed relationship between the female composition of a paper and its citations (Table F.4).

Table F.1: List of superstar men

Abel, Andrew B.	Fisher, Franklin M.	Pakes, Ariel
Acemoglu, Daron	Fudenberg, Drew	Palfrey, Thomas R.
Aghion, Philippe	Gale, Douglas	Persson, Torsten
Alesina, Alberto	Granger, Clive W. J.	Phillips, Peter C. B.
Andrews, Donald W. K.	Green, Jerry R.	Plott, Charles R.
Arellano, Manuel	Grossman, Gene M.	Postlewaite, Andrew
Banerjee, Abhijit V.	Grossman, Sanford J.	Ray, Debraj
Barro, Robert J.	Gruber, Jonathan	Robinson, Peter M.
Becker, Gary S.	Gul, Faruk	Romer, David H.
Bénabou, Roland	Hamilton, James D.	Rosen, Sherwin
Bernheim, B. Douglas	Hansen, Lars Peter	Rosenzweig, Mark R.
Besley, Timothy J.	Hart, Oliver D.	Roth, Alvin E.
Blackorby, Charles	Hausman, Jerry A.	Rubinstein, Ariel
Blanchard, Olivier J.	Heckman, James J.	Saez, Emmanuel
Blundell, Richard W.	Helpman, Elhanan	Samuelson, Larry
Bolton, Patrick	Jackson, Matthew O.	Sargent, Thomas J.
Browning, Martin J.	Jovanovic, Boyan	Scheinkman, José A.
Caballero, Ricardo J.	Kahneman, Daniel	Shleifer, Andrei
Campbell, John Y.	Kehoe, Patrick J.	Stein, Jeremy C.
Caplin, Andrew S.	Kremer, Michael	Stiglitz, Joseph E.
Card, David E.	Krugman, Paul R.	Tirole, Jean
Chiappori, Pierre-André	Laffont, Jean-Jacques	Tversky, Amos
Cooper, Russell	Laroque, Guy	Vishny, Robert W.
Crawford, Vincent P.	Levine, David K.	Weil, David N.
Deaton, Angus S.	Levitt, Steven D.	Weitzman, Martin L.
Diamond, Peter A.	List, John A.	White, Halbert
Dixit, Avinash K.	Mankiw, N. Gregory	Wolpin, Kenneth I.
Engle, Robert F.	Maskin, Eric S.	Wright, Randall
Epstein, Larry G.	Milgrom, Paul R.	Zame, William R.
Fehr, Ernst	Murphy, Kevin M.	
Feldstein, Martin S.	Newey, Whitney K.	

Table F.2: The impact of superstardom (1950–2015)

	Model 1			Model 2			Model 3		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
female	−15.080**	1.929	4.031	−7.131	−3.584	1.790	0.003	−1.119	10.217
	(6.834)	(6.803)	(6.493)	(6.819)	(6.733)	(6.517)	(6.731)	(6.834)	(6.423)
superstar		117.278***			151.754***			120.566***	
		(16.491)			(26.661)			(35.374)	
N				24.073***	21.531***	12.604***	21.996***	21.148***	8.204*
				(4.290)	(4.164)	(4.110)	(4.302)	(4.216)	(4.223)
max. t				1.768***	−4.278***	−2.205***	−6.576***	−6.774***	−6.481***
				(0.327)	(1.148)	(0.739)	(0.993)	(1.005)	(0.983)
max. T							6.502***	2.914**	8.271***
							(0.803)	(1.224)	(0.799)
Year×Journal	✓	✓	✓	✓	✓	✓	✓	✓	✓
Superstar authors			✓			✓			✓
Obs.	11,335	11,320	11,320	11,335	11,320	11,320	11,335	11,320	11,320
R^2	0.059	0.073	0.169	0.063	0.077	0.170	0.072	0.078	0.176

Note. Figures correspond to coefficients from an OLS regression of raw citation counts on the ratio of female authors and the indicated control variables. Superstar is a binary variable equal to 1 if at least one author on a paper satisfies the criteria defined in Appendix F.1. Superstar fixed effects account for each superstar author. Robust standard errors in parentheses. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table F.3: The impact of superstardom (1990–2015)

	Model 1			Model 2			Model 3		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
female	−3.629	5.158	8.748	4.324	4.940	7.894	9.930	9.632	15.625**
	(7.765)	(7.421)	(7.309)	(7.728)	(7.675)	(7.353)	(7.596)	(7.696)	(7.283)
superstar		62.409***			69.532***			28.885	
		(10.287)			(18.380)			(20.497)	
N				19.106***	18.551***	18.287***	18.693***	18.541***	15.745***
				(3.604)	(3.553)	(3.131)	(3.536)	(3.517)	(3.088)
max. t				1.575***	−1.074	−1.779***	−6.917***	−7.109***	−8.037***
				(0.284)	(0.750)	(0.656)	(1.151)	(1.166)	(1.057)
max. T							6.977***	6.232***	8.874***
							(0.991)	(1.107)	(1.025)
Year×Journal	✓	✓	✓	✓	✓	✓	✓	✓	✓
Superstar authors			✓			✓			✓
Obs.	6,440	6,434	6,434	6,440	6,434	6,434	6,440	6,434	6,434
R^2	0.114	0.125	0.205	0.123	0.130	0.209	0.140	0.141	0.222

Note. Columns display estimates identical to those in Table F.2 except that only articles published after 1990 are included. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table F.4: The impact of superstardom (2000–2015)

	Model 1			Model 2			Model 3		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
female	10.673	12.454*	14.145**	14.803**	14.780**	15.037**	16.742**	17.102**	17.701**
	(6.974)	(6.871)	(6.940)	(6.933)	(6.952)	(7.137)	(6.902)	(6.921)	(7.115)
superstar		15.388***			−2.705			−23.685***	
		(5.674)			(8.538)			(7.390)	
N				10.586***	10.572***	11.335***	10.264***	10.331***	10.151***
				(2.044)	(2.042)	(2.335)	(1.992)	(1.999)	(2.189)
max. t				0.726***	0.825**	−0.097	−5.626***	−5.366**	−5.311**
				(0.229)	(0.352)	(0.838)	(2.169)	(2.149)	(2.064)
max. T							5.512***	6.005***	6.082***
							(1.935)	(1.978)	(1.673)
Year×Journal	✓	✓	✓	✓	✓	✓	✓	✓	✓
Superstar authors			✓			✓			✓
Obs.	4,186	4,183	4,183	4,186	4,183	4,183	4,186	4,183	4,183
R^2	0.184	0.186	0.215	0.193	0.193	0.220	0.212	0.214	0.233

Note. Columns display estimates identical to those in Table F.2 except that only articles published after 2000 are included. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

F.2 Nobel Prize-winning authors

In this Appendix, we swap our *ad hoc* definition of "superstar" (Appendix F.1) with fixed effects (and a binary variable) for papers by Nobel Prize winning authors.[2]

About 0.9 percent of authors in our data are Nobel Prize winners. (See Table F.5 for a list of their names.) On average, each has published five papers in a top-five journal; their highest cited paper was cited 690 times.

F.2.1 Results

Results in Tables F.6, F.7 and F.8 closely mirror corresponding results from Appendix F.1. Controlling for Nobel Prize winners reduces the magnitude of the coefficient on female authorship (Table F.6) but the change is less pronounced when the sample is restricted to later years (Tables F.7 and F.8). Among articles published after 2000 (Table F.8), female-authored papers are cited 11–17 times more frequently and accounting for Nobel Prize winners does not observably impact this gap.

Table F.5: List of Nobel Prize winners

Akerlof, George A.	Koopmans, Tjalling C.	Samuelson, Paul A.
Allais, Maurice	Krugman, Paul R.	Sargent, Thomas J.
Arrow, Kenneth J.	Kydland, Finn E.	Scholes, Myron S.
Aumann, Robert J.	Lucas, Robert E. (Jr.)	Schultz, Theodore W.
Becker, Gary S.	Markowitz, Harry M.	Selten, Reinhard
Buchanan, James M.	Maskin, Eric S.	Sen, Amartya K.
Deaton, Angus S.	McFadden, Daniel L.	Shapley, Lloyd S.
Debreu, Gerard	Merton, Robert C.	Shiller, Robert J.
Diamond, Peter A.	Miller, Merton H.	Simon, Herbert A.
Engle, Robert F.	Mirrlees, James A.	Sims, Christopher A.
Fama, Eugene F.	Modigliani, Franco	Smith, Vernon L.
Friedman, Milton	Mortensen, Dale T.	Solow, Robert M.
Frisch, Ragnar	Mundell, Robert A.	Spence, A. Michael
Granger, Clive W. J.	Myerson, Roger B.	Stigler, George J.
Hansen, Lars Peter	Nordhaus, William D.	Stiglitz, Joseph E.
Harsanyi, John C.	North, Douglass C.	Stone, Richard
Hart, Oliver D.	Ostrom, Elinor	Thaler, Richard H.
Heckman, James J.	Phelps, Edmund S.	Tinbergen, Jan
Holmström, Bengt	Pissarides, Christopher A.	Tirole, Jean
Hurwicz, Leonid	Prescott, Edward C.	Tobin, James
Kahneman, Daniel	Romer, Paul M.	Williamson, Oliver E.
Klein, Lawrence R.	Roth, Alvin E.	

[2] We count only authors who had won the Nobel Prize when the data were last updated (October 2018; see Section 2); additionally controlling for the 2019 winners does not meaningfully change results.

Table F.6: The impact of the Nobel Prize (1950–2015)

	Model 1			Model 2			Model 3		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
female	−15.080**	−5.359	−1.225	−7.131	−8.091	1.088	0.003	−2.684	9.798
	(6.834)	(6.857)	(6.844)	(6.819)	(6.832)	(6.811)	(6.731)	(6.808)	(6.708)
Nobel		199.332***			204.649***			184.897***	
		(29.009)			(33.401)			(34.266)	
N				24.073***	18.740***	12.520***	21.996***	17.707***	9.395**
				(4.290)	(4.000)	(3.691)	(4.302)	(4.023)	(3.725)
max. t				1.768***	−1.673**	0.520	−6.576***	−7.558***	−6.990***
				(0.327)	(0.716)	(0.456)	(0.993)	(1.046)	(0.966)
max. T							6.502***	4.845***	6.965***
							(0.803)	(0.839)	(0.753)
Year×Journal	✓	✓	✓	✓	✓	✓	✓	✓	✓
Nobel authors			✓			✓			✓
Obs.	11,335	11,335	11,335	11,335	11,335	11,335	11,335	11,335	11,335
R^2	0.059	0.081	0.231	0.063	0.083	0.232	0.072	0.088	0.241

Note. Figures correspond to coefficients from an OLS regression of raw citation counts on the ratio of female authors and the indicated control variables. Nobel is a dummy variable equal to 1 if at least one author on a paper is a Nobel Prize winner; Nobel fixed effects account for each Nobel Prize winning author. Robust standard errors in parentheses. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table F.7: The impact of the Nobel Prize (1990–2015)

	Model 1			Model 2			Model 3		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
female	−3.629	−1.110	−0.726	4.324	3.597	4.085	9.930	9.205	10.630
	(7.765)	(7.789)	(7.672)	(7.728)	(7.700)	(7.550)	(7.596)	(7.578)	(7.416)
Nobel		68.656***			51.001***			52.658***	
		(12.656)			(15.192)			(14.964)	
N				19.106***	19.018***	19.226***	18.693***	18.601***	17.753***
				(3.604)	(3.615)	(3.651)	(3.536)	(3.547)	(3.562)
max. t				1.575***	0.825**	0.953**	−6.917***	−7.730***	−7.764***
				(0.284)	(0.356)	(0.384)	(1.151)	(1.130)	(1.137)
max. T							6.977***	7.009***	7.514***
							(0.991)	(0.991)	(1.059)
Year×Journal	✓	✓	✓	✓	✓	✓	✓	✓	✓
Nobel authors			✓			✓			✓
Obs.	6,440	6,440	6,440	6,440	6,440	6,440	6,440	6,440	6,440
R^2	0.114	0.119	0.138	0.123	0.125	0.145	0.140	0.143	0.164

Note. Columns display estimates identical to those in Table F.6 except that only articles published after 1990 are included. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

Table F.8: The impact of the Nobel Prize (2000–2015)

	Model 1			Model 2			Model 3		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
female	10.673	11.431	11.994*	14.803**	14.807**	15.086**	16.742**	16.813**	17.210**
	(6.974)	(6.983)	(7.064)	(6.933)	(6.935)	(7.013)	(6.902)	(6.905)	(6.979)
Nobel		15.985*			0.850			9.618	
		(8.174)			(8.657)			(8.064)	
N				10.586***	10.592***	11.545***	10.264***	10.331***	10.943***
				(2.044)	(2.029)	(2.069)	(1.992)	(1.986)	(2.008)
max. t				0.726***	0.716***	0.638**	−5.626***	−5.797***	−5.692***
				(0.229)	(0.240)	(0.248)	(2.169)	(2.177)	(2.177)
max. T							5.512***	5.564***	5.480***
							(1.935)	(1.937)	(1.962)
Year×Journal	✓	✓	✓	✓	✓	✓	✓	✓	✓
Nobel authors			✓			✓			✓
Obs.	4,186	4,186	4,186	4,186	4,186	4,186	4,186	4,186	4,186
R^2	0.184	0.184	0.191	0.193	0.193	0.199	0.212	0.212	0.218

Note. Columns display estimates identical to those in Table F.6 except that only articles published after 2000 are included. ***, ** and * statistically significant at 1%, 5% and 10%, respectively.

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