Equal Job, Unequal Pay? Evidence from 4 Million Regulatory Careers

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Keywords: Gender Pay Gaps, Labor Mobility, Economics of Regulation, Administrative Data

JEL Classifications: G3, J3, J6, K31

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Introduction

The existence of a gender pay gap is widely recognized, yet its origins remain deeply contested. In recent years, both public- and private-sector organizations have implemented programs aimed at improving outcomes for underrepresented groups, including women. More recently, however, such initiatives are being rapidly scaled back, often based on the assertion that they are unnecessary or ineffective. Despite the prominence of these policy debates, rigorous and large-scale empirical research on gender pay gaps is limited. Existing studies typically rely on small samples or coarse records, limiting the ability to observe full career trajectories and key determinants of pay. As a result, the magnitude, persistence, and drivers of gender pay gaps remain unclear. We leverage a unique setting to address these challenges: the U.S. Federal government. As the nation's largest continuous employer, it offers unusually rich, longitudinal data on employee pay, assignments, promotions, and job transitions. This setting is also especially timely. Shortly after taking office in January 2025, President Trump issued an executive order dissolving the White House Gender Policy Council, and the newly created Department of Government Efficiency (DOGE) began eliminating nearly all Diversity, Equity, and Inclusion (DEI) programs across federal agencies. Our analysis speaks directly to the ongoing debate regarding the role of such programs.

We assemble comprehensive records on nearly all civilian federal employees over five decades, excluding military agencies. We focus on employees whose gender can be reasonably identified based on their first name. Our final sample includes 36.5 million employee-level observations between 1973 and 2023, across 334 agencies in virtually every city in the United States. Women constitute approximately half of our sample, and their representation has increased over time, from 40.2% in 1973 to 54.4% in 2023. The granularity and scope of our dataset enables precise measurement of gender differences in pay across careers, agencies, and occupations.

In the first part of the paper, we document persistent and widespread gender pay gaps in the federal workforce. We regress employee pay on gender and gradually add fixed effects to control for observable determinants of pay. Without controls, women earn $18,089 less than men, expressed in 2023 USD. This equates to a pay gap of 19.8%. To identify the portion of this gap that is driven by women being paid less for the same job versus women holding lower-level jobs, we compare employees within the same local office (for example, the SEC office in Philadelphia), conditional on their tenure, and who have the same occupation and rank. We find that women earn $1,248 less than men, expressed in 2023 USD. This translates to a pay gap of 0.8%.

We find no clear time trend in the estimated gender pay gap. In fact, the pay gap as of 2023 is only marginally lower than that in 1973, nearly half a century ago. It has fluctuated over time, but there is no evidence of a consistent decrease. Moreover, the within-rank pay gap has remained essentially flat for the past fifty years, indicating that the modest decline in the raw gap reflects changes in job sorting rather than convergence in pay for equal work. In the cross section, gender pay gaps exist in all U.S. states, in large and small local offices, and in all major agencies. Given legislation requiring the U.S. Federal Government to operate under policies that ensure equal pay for equal work, the existence of such pay gaps is striking.

Over a longer horizon, the cumulative effects of this pay disparity are striking. To estimate this, we rely on the unusually long horizon of our database, which spans five decades, and thus tracks the entire careers of many federal employees. We focus on over 3 million unique employees who started their career in 1974 or later and ended their career in 2022 or earlier. We find that women receive $1,074 (or 0.6%) lower starting salary, $1,133 (or 0.6%) lower early-career salaries (years 2 to 5), $1,189 (or 0.7%) lower mid-career salaries (years 6 to 37), and $1,214 (or 0.7%) lower late-career salaries (final three years), compared to their male counterparts. To calculate the cumulative effect of pay disparity, we assume an annual rate of return of 7%, starting government employment age of 25 years, and an average retirement after 40 years of service. With those parameters, we estimate a cumulative loss of $250,689 upon retirement. For reference, according to the Federal Reserve's 2022 Survey of Consumer Finances, the mean and median retirement account savings for families at retirement age of

65-74 is \$609,230 and \$200,000, respectively.[1]

We examine three factors that potentially contribute to the observed pay gap. We begin with sorting. A long literature suggests that women may self-select into lower-paying parts of the labor market due to differences in risk tolerance or preferences for flexibility (e.g., Goldin, 2014; Niederle and Vesterlund, 2011; Bowles, Babcock, and Lai, 2007; Sapienza, Zingales, and Maestripieri, 2009). To quantify how much of the raw gender pay gap reflects differences in the types of jobs men and women hold, we compare regressions with increasingly rich sets of fixed effects. Controlling only for basic observables leaves the raw gap largely intact, but adding agency and occupation fixed effects eliminates roughly 72% of the unconditional pay gap. In parallel, the model's explanatory power increases sharply: the R^2 rises by 77 percentage points.

We next examine sorting more directly by testing whether women are disproportionately located in lower-paying sectors of the government. We define "sectors" at increasingly fine levels: year × agency, year × agency × county, year × agency × county × occupation, and ultimately year × agency × county × occupation × rank. For each definition, we compute the average male pay in that sector and regress it on the sector's share of female employees. This approach asks whether women work in sectors that pay less even for men, thereby avoiding mechanical correlations between gender composition and pay. We find strong evidence of sorting at every level of aggregation. At the most granular level – e.g., GS-12 attorneys in Philadelphia working at the SEC versus the EPA – work units with a one–standard-deviation higher share of female employees have average *male* pay that is 6.9% lower. Because the result are for male pay, it cannot be attributed to greater discrimination within female-heavy units. Rather, it reflects women clustering into organizational niches that, on average, offer lower compensation. In sum, a substantial share of the gender pay gap reflects systematic differences in where men and women work, consistent with earlier evidence that job segregation and industry choice explain much of the aggregate gender pay

[1] See https://www.federalreserve.gov/econres/scf/dataviz/scf/chart/ (last accessed: 11/14/2025).

disparity (Bertrand and Hallock, 2001).

The second factor we examine is the probability of promotion. Promotions represent a key opportunity for pay raises in the federal system, and unequal promotion process can generate persistent pay differences even when initial salaries are comparable. We find that, on average, women are less likely to be promoted than otherwise similar men, with effects larger at higher ranks, where promotions rely more heavily on subjective assessments and negotiation. Our estimates suggest that at upper quartile ranks, women face an 0.3% to 0.4% lower probability of promotion.

The third contributing factor is pay progression within rank. Even when men and women remain in the same GS grade, their salaries may evolve differently depending on how quickly they progress toward the top of the pay range. We find that this also plays a significant role. In our most stringent specification, women are 3.4 percentage points farther from the top of their rank's pay band. a gap equal to 4% of the sample mean. In sum, even when holding job level constant, women experience slower earnings growth over time, contributing to persistent lifetime disparities.

Having established that sorting, promotion likelihood, and pay progression within rank all contribute to the gender pay gap, we turn to the economic mechanisms that plausibly underlie these consistently worse outcomes for women. We consider three mechanisms: education and experience, managerial style, and labor mobility. We begin with education and experience, where the expected effects are ambiguous ex ante. On the one hand, higher levels of achievement provide a clearer signal of productivity and may strengthen women's bargaining power (Becker, 1957, 1962). On the other hand, higher-level jobs often involve more intangible and subjective tasks, creating greater scope for bias and for differences in negotiation behavior to influence pay (e.g., Dittrich, Knabe, and Leipold, 2015; Eckel and Grossman, 2008; Save-Soderbergh, 2019). To examine these forces empirically, we exploit the structure of the General Schedule (GS) pay scale, which assigns most federal employees to ranks from GS-1 to GS-15. Higher ranks reflect greater education, experience, and responsibility. We find that the gender pay gap is concentrated at the upper levels of the hierarchy, namely, GS-9 and above. Similarly, the gap widens with education, reaching 3% among employees with a Ph.D. These results are particularly striking given that women who remain in the most senior positions are a positively selected group (Blau and Kahn (2017)).

The second mechanism is managerial style, where we examine whether the presence of women in top leadership positions affects gender gaps within their organizations. Theory offers competing predictions. Homophily – the tendency of people to prefer others who are similar to them (Becker, 1957) – suggests that female leaders may foster more equitable outcomes for female employees. In contrast, evidence on implicit bias (e.g., Bertrand and Mullainathan, 2004; Reuben et al., 2014) implies that both male and female managers may associate certain demographics with certain roles, producing similar patterns of discrimination. A third view, the "Queen Bee" hypothesis (Staines, Tavris, and Jayaratne, 1973), predicts that female executives may distance themselves from other women, potentially widening gaps under female leadership. The federal setting allows us to identify leaders cleanly by focusing on political appointees in Washington, DC. Empirically, we find evidence most consistent with the role model channel: the gender pay gap is roughly 30% smaller in agencies with at least one female executive.

The third and final mechanism we examine is labor mobility. Across the population, women are less likely than men to be the highest-paid member of the household (Bertrand, Kamenica, and Pan, 2015), which may limit their ability to relocate for better employment opportunities. We both estimate regressions with finely defined fixed effects and examine exogenous shocks, office closures, that force employees to relocate. Across both specifications, we find that women are no less likely to move than their male counterparts. Thus, although labor mobility is frequently cited as a contributor to gender pay disparities, our results – drawn from a broad sample of more than 25 million employee-year observations – cast doubt on this explanation.

In sum, we document persistent and pervasive gender pay gaps in the federal workforce, which is driven by women sorting into lower-paid units, having a lower promotion likelihood, and experiencing slower pay progression within rank. Both managerial style and levels of experience and education contribute to the gaps, while constrained labor mobility plays a negligible role. Understanding these mechanisms helps clarify the structural forces that perpetuate the gender disparities. Our analysis benefits from three key advantages: first, we examine an exceptionally long time period of half a century; second, we control for a rich set of determinants of pay; and third, we reject a strong null hypothesis that a large, rule-bound public employer should not exhibit gender gaps.

We contribute to several streams of literature. First, we contribute to literature on the pay gap. Bertrand and Hallock (2001) find an unconditional pay gap of 45% among corporate executives, which reduces to 5% after controlling for observable firm and person characteristics. In a study of 5,000 US workers, Blau and Kahn (2017) find that the pay gap has decreased over time, reaching an unconditional 20% as of 2010 and 8% after controlling for observable factors. More recently, Maloney and Neumark (2025) examine the pay gap using state-level data. Our estimates of the unconditional gaps are similar to those of Blau and Kahn (2017), but our estimates after controlling for observable factors are smaller than these prior studies. This is at least partially due to our ability to better control for many determinants of pay.[2] There have also been several government analyses, for example by the General Accounting Office (GAO) and Office of Personnel Management (OPM), which find gaps in the 6–7% range as of 2022. Our analysis provides more detail on the drivers of this gap, as well as heterogeneity across agencies, ranks, and education levels. In sum, one of the main takeaways of our paper is the existence of a gender pay gap within the Federal government, across workers that are equal in so many precisely measured dimensions, and which exists in spite of decades of federal legislation advancing equal pay.

Second, we contribute to work on the determinants of pay gaps. Unlike most literature in this area, our long time-series and ability to track millions of individual workers enables us

[2]The federal pay system is significantly more compressed than the private sector, which enables us to more precisely identify a person's level and as such causes estimates of pay gaps to be lower.

to precisely estimate differences in promotion likelihood and within-rank pay raises. In our analyses of the economic mechanisms driving these differential outcomes, we examine two channels that are frequently mentioned, but on which there is less conclusive evidence due in part to data limitations.[3] First, we examine whether women are less willing to relocate, in ways that facilitate career advancement and associated pay increases. We find no evidence that this plays a significant role. Second, we examine the influence of specific managers. In light of conflicting theory and empirical evidence (see, e.g., Staines, Tavris, and Jayaratne (1973) , Bertrand and Mullainathan (2004), Reuben et al. (2014), Tate and Yang (2015), Matsa and Miller (2013), and Srivastava and Sherman (2015)), our data again provides a unique advantage. We follow unique workers who are employed at different times for different managers, both within the same office and across different offices, and we find that women managers significantly mitigate the pay gap.

Finally, we add to the literature on regulatory incentives. Studies tend to focus on the level and composition of pay, such as Dal Bó, Finan, and Rossi (2013); Kalmenovitz (2021). Others examine organizational features such as fee schedules (Kisin and Manela, 2018), field offices (Gopalan, Kalda, and Manela, 2021), supervision (Hirtle, Kovner, and Plosser, 2020; Eisenbach, Lucca, and Townsend, 2016), and jurisdictional overlap (Kalmenovitz, Lowry, and Volkova, 2023). To our knowledge, ours is the first paper to directly study gender differences in pay within the U.S. federal government. While our findings may not apply equally across all sectors, at a minimum they contribute to our understanding of a large, influential, and yet understudied group of employees. Our work can also add to the policy discussion around regulatory pay. Gender pay gaps incurred at the onset of the regulator's career, including occupational and agency choice, have substantial downstream impact on their lifetime earnings. Limited vertical mobility inside the organization, and limited horizontal mobility

[3] Explanations that have been examined in prior work include preferences (Goldin (2014), Niederle and Vesterlund (2011)), discrimination (Arrow (1973), Becker (1957)), shorter work hours and more career discontinuities Goldin (2014), Bertrand, Goldin, and Katz (2010)), family-friendly policies that make it more costly to hire women (Blair and Posmanick (2023)), and human capital differences (Card, Cardoso, and Kline (2015), Lagaras et al. (2022)).

across geographic areas, exacerbate the persistence of gender pay gaps.

1 Data and Summary Statistics

We assemble employee-level payroll records of the entire civilian federal workforce. The information is derived from two separate datasets. The first dataset was released by BuzzFeed News following a Freedom of Information Act (FOIA) request, and it covers the years 1973-1995. The second dataset is a result of repeated FOIA requests submitted by one of the authors to various federal entities, and it covers the years 1996-2023 (inclusive). For ease of notation, we refer to the early dataset as HR-1 and to the latter as HR-2. Both datasets include employee-level information on agency, occupation, location (state, county, city), adjusted pay,[4] and rank. Additional information on education and age is included in HR-1 (but not in HR-2), and additional information on base pay, bonus, and hiring date is included in HR-2 (but not in HR-1). The bulk of the analysis relies on variables available in both datasets, while some subsequent tests utilize the unique variables in each one. To the best of our knowledge, the data include the universe of the civilian federal workforce from that period.

The raw dataset contains 39.9 million employee×year observations. To derive our final sample, we apply three main filters consecutively. We first remove employees whose gender cannot be reasonably identified. We follow methodologies similar to those of Niessen-Ruenzi and Ruenzi (2018), Kaplan and Yuan (2020) and Atanasov, Dana, and Teeselink (2024) to identify the employees' genders by matching their first name to Social Security frequency tables. We define $Female = 1$ if the first name has at least a 75% probability of belonging to a woman, and $Female = 0$ if the first name has 25% or less probability of belonging to a woman. We remove approximately 0.5 million observations in which the first name is either missing or ambiguous. This leaves a sample of 39.4 million employee×year observations.

[4]In the Federal government, adjusted pay is the employee's base pay plus a locality adjustment that varies across regions and over time.

The second filter removes individuals with duplicate names. Some of our tests require tracking individuals over time, for instance, to study promotions and relocations.[5] We therefore remove individuals with duplicate names within the same year, for instance, if there is more than one Jane Doe working in the Federal government in 2010. This filter removes approximately 1.7 million observations.

The third filter removes agencies with few observations. Since many of our tests are based on within-agency pay gaps, we remove agencies with fewer than 1,000 observations over the entire sample period or agencies that exist for less than five years. Those are typically ad-hoc commissions, formed around a concrete task and dissolved within several years.[6] This filter removes approximately 1.2 million observations.

The final sample includes 36.5 million employee×year observations. Table 1 provides descriptive statistics. Panel A summarizes the full sample, and Panels B and C focus on HR-1 and HR-2, respectively. Most variables are similar across the two subsamples, and we focus our discussion on the full sample. Approximately half (49.6%) are females, and the remainder are males. The average tenure is 11 years and the average pay is $50,275. Given that our sample spans five decades, we adjust salaries for inflation. The average pay in constant 2023 USD is $85,855. In a typical year, 1% of people are promoted, 6% move, and 64% receive a pay raise (in real terms).

2 Gender Pay Disparity

In this section, we examine the existence, prevalence, and magnitude of gender pay gaps in the federal government. Our workhorse specification is:

$$y_{i,t} = Female_i + Tenure_{i,t} + \lambda \tag{1}$$

[5] Due to different name conventions, it is difficult to track individual employees across the two datasets.

[6] Examples include the Commission on Federal Paperwork, the National Transportation Policy Study, the Commission on Review of National Policy to Gambling, and the Financial Crisis Inquiry Commission.

where $y_{i,t}$ is adjusted pay or the natural log of the adjusted pay, expressed in 2023 USD. The main independent variable of interest, $Female$, equals 1 if the employee is classified as female based on their first name, and 0 otherwise. Tenure is the number of years in government service. The baseline specification includes only the gender indicator, and we gradually add $Tenure$ and increasingly tight fixed effects to assess how pay disparities persist after accounting for various sources of heterogeneity. Standard errors are clustered by county and year.

2.1 Main result

Table 2 summarizes our main result. The dependent variable is pay and log of pay in Panels A and B, respectively. The estimates suggest that women earn significantly lower pay than their male counterparts. In the first specification, where we compare mean pay across genders without any controls, we find that females earn \$18,089 or 19.8% less than males. The inclusion of year fixed effects has little impact on these estimates, as shown in column (2). The addition of year×agency in column (3), and the addition of year×agency×county fixed effects in column (4) similarly has relatively small effects. Looking at column (4), estimated gender gaps equal \$16,490 or 17.2%. All fixed effects are interactive, meaning this specification compares employees within the same local office, for example: the SEC's office in Philadelphia in 2010.

In column (5), we add the interaction of occupation fixed effects, which accounts for a significant portion of the gender pay gap. The coefficient on $Female$ drops to \$6,181 or 5.5%. Finally, in column (6), we add the interaction with rank fixed effects, for instance, comparing SEC junior attorneys (say, level 9) in Philadelphia in 2010. Moving from across-ranks to within-ranks also explains a significant portion of the pay gap. As shown in column (6), in our tightest specification we find that females earn \$1,248 or 1% less than males.

It is helpful to summarize the contribution of each source to the gender pay gap. The unconditional gap is \$18,089. Including year×agency×county FE reduces this gap by 1.1% (to

$17,893), interacting with occupation FE further reduces the gap by 64.7% (to $6,181), and interacting with rank FE shrink the gap by additional 27.3% (to $1,248). The incremental increase in R^2 supports a similar interpretation, as it rises by over 40% once we account for occupation and for 17% when we account for rank. This suggests that sorting into different occupations is the primary source of gender gap in adjusted base pay.

Our findings relate to prior evidence regarding the influence of occupation on the gender pay gap. Blau and Kahn (2017) show that a substantive portion of the pay gap arises from differences in occupation choice, for example with women disproportionately entering traditionally low-paying jobs such as nursing and education, and men choosing fields in the STEM and finance domains which tend to pay more. Our findings suggest that such factors play significant roles, even when effects are limited to occupations covered by a single employer, in our setting the federal government. We will continue this line of investigation in Section 3.1.

For completeness, in Table A.1 we estimate the gap separately in HR-1 (1973-1995) and HR-2 (1996-2023). We find significant gender pay gaps in both datasets and time periods, limiting the concern that database-specific measurement issues contribute to the gap. In Table A.2, we remove the 1973 cohort from HR-1 since their tenure is measured imprecisely. The magnitude of the base gap, without any fixed effects, is over twice as large in the earlier dataset. In contrast, the magnitude of the gap after controlling for key determinants, including both occupation and rank, is similar across both datasets. Together, these findings suggest that women are gradually entering occupations with higher pay and advancing to higher ranks.

In sum, even after accounting for the richest set of controls, we document significant pay disparities within the government sector. The magnitude is even greater among certain subsamples and among certain demographics, as discussed in more detail in subsequent sections.

2.2 Aggregate patterns

In this section, we present additional stylized facts, which shed light on time trends and pervasiveness of the gender pay gaps. We start by investigating the evolution of the gaps over time. Revisiting our workhorse specification from Equation (1), we estimate each year separately. We first include no additional fixed effects, which yields a set of coefficients that represent the average gender gap in a given year. The results are in Figure 1, Panel A. We find that pay gaps have been generally declining over time, from approximately \$36,000 in 1973 to \$3,000 in 2023. We document similar trends when adding agency, agency×county, and agency×county×occupation fixed effects, as shown in Panels B, C, and D, respectively. The latter reflects a set of coefficients that represent the average gender gap within an agency, within a county, within an occupation, each year. This gap has decreased from approximately \$13,000 to \$2,000 over our sample period.

A broad body of literature highlights that at least a portion of the gender pay gap is driven by the fact that women are less likely to have higher-level jobs. To examine trends within job level, we add rank fixed effects. Specifically, we now include agency× county × occupation ×rank fixed effects. Strikingly, there is no visible time trend in within-rank pay gaps, as shown in Figure 1, Panel E. The gap has remained quite constant over time. The stability of within-rank pay gaps adds nuance to governmentwide reports, such as those by the Office of Personnel Management, which document a declining overall pay gap.[7]

Second, we investigate whether the average gender pay gap is driven by a small subset of agencies or whether it is more pervasive across the government. As mentioned above, our sample includes only agencies with at least 1,000 observations and which have existed for at least 5 years. We estimate Equation (1) for each agency separately, with year, county, occupation, and rank fixed effects (interacted). Naturally, we omit agency fixed effects. We plot the resultant coefficients in Panel A of Figure 2. We find that the majority of federal

[7] Governmentwide Gender and Racial/Ethnic Pay Gap Analysis Summary, Office of Personnel Management, 2024. Link: `https://chcoc.gov/about-us/reports-publications/agency-reports/governmentwide-gender-and-racialethnic-pay-gap-analysis-summary/` (last visited: 11/13/2025)

agencies (76%) display statistically significant gender pay gaps. The highest statistically significant gap is estimated at the Office of Navajo and Hopi Indian Relocation (29% gap in favor of males) and the lowest statistically significant gap is at the Inter-American Foundation (1.6% gap in favor of females).

In Panel B, we replicate this analysis on the subset of agencies for which the gender pay gap is statistically significant, with a p-value of 0.05 or lower. Here, the takeaway is even more striking. Within every agency, the pay gap is negative. In other words, among agencies where there is a significant difference in the pay of women versus men (within the same office and the same rank), women always earn less than men in all but two agencies.

Third , we drill down even further to study local office-level pay gaps. We define the local office as the interaction of agency by county, for instance, the EPA office in Atlanta. We focus on local offices with at least 100 observations in each year to maintain sufficient statistical power. We then estimate Equation (1) for each office separately with year $\times$ occupation $\times$ rank fixed effects, naturally omitting agency and county fixed effects. Given the small number of observations in many of these specifications, it is not surprising that estimated effects are statistically insignificant in many cases. We focus on the 1,301 local offices in which the estimated effect is statistically significant. Results are shown in Figure 3. The average office has a 1.1% pay gap, and in the majority of offices women get paid less than men. The distribution has a long left tail, implying that while some offices show substantial negative pay gaps against women, large positive gaps in favor of women are rarer.

Fourth, we examine patterns in the gender pay gap across varying levels of tenure within the government. We distinguish between four career stages: first year; early-career stage, defined as years 2 through 5; mid-career stage, defined as years 6 till 3 years prior to retirement; and late-career stage, defined as the last 3 years of service.[8] We exclude individuals from the HR-1 sample who show up in 1973 for the first time because we cannot accurately

[8] In the event an individual can be categorized into more than one career stage, we assign the individual to the latest career stage. For example, an individual who works only two years at a federal agency and then leaves the government will be categorized as late-career stage.

measure their tenure for this exercise. We further exclude observations in 1993–1995 and 2021–2023 for individuals whose last observed year is 1995 and 2023, respectively, because we cannot determine whether they have exited government service. We then estimate Equation (1) separately for each stage, using our tightest specification with year, agency, county, occupation, and rank fixed effects (interactive).

The results are shown in Table 3. We find that pay gaps persist throughout all career stages, but they are significantly higher at later stages. The coefficients increase linearly across stages, from an average pay gap of $1,074 in starting salaries, to an average pay gap of $1,214 in late-career salaries. Those differences reflect, in part, the fact that late-career employees have a higher average pay, leaving more room for potential pay gaps. However, the fact that there is an upward trend in percentage gaps (as shown in Panel B), indicates that this is not the only factor. Importantly, pay differentials compound almost mechanically over time, as the employee continues to receive pay raises. In general, with an annual pay raise of p, an initial pay gap of $1 will reach $(1+p)^{1+t}$ after t years of service.

In sum, the aggregate patterns yield three main takeaways. First, while the unconditional gender pay gap has narrowed markedly over the past five decades, the gap within finely defined agency $\times$ county $\times$ occupation $\times$ rank cells is relatively flat. Most of the apparent progress therefore reflects changes in sorting across jobs and ranks rather than convergence in pay for equal work. Second, gender pay gaps are pervasive: in the vast majority of agencies and local offices the coefficient on *Female* is negative, and among the units with a statistically significant gap, men almost always earn more than women. Third, although the within-cell pay gap is modest in any given year, it is present at every career stage and grows in percentage terms over time, mechanically compounding into substantial lifetime income losses. We expand on this point in the next section.

2.3 Cumulative income loss

In this subsection, we employ the estimates from Table 3 to estimate the cumulative effect of pay disparities, over the employee's entire government career. We use the estimates of the gap in the first year, early-career stage, mid-career stage, and late-career stage to compute an illustrative example of potential lifetime losses due to unexplained gender pay gaps. We construct a representative female agent, who joins the federal government at the age of 25 and retires after 40 years at the age of 65. We assume the agent has an investment portfolio with a 7% annual rate of return, and we compute the future value of pay gaps in various career stages at the time of retirement.

We separately calculate the cumulative effects of the pay gap, by for each career stage. Table 4 summarizes our calculations. First, our estimated \$1,074 pay gap in hiring salaries will accumulate to \$16,083 by the time of retirement ($\$1*(1+0.07)^{40}$). Second, the \$1,133 pay gap per year in the early-career stage will accumulate to \$57,468 upon retirement ($\Sigma_{t=1}^{t=4}\$1*(1+0.07)^{40-t}$). Third, the \$1,189 pay gap per year in the mid-career stage will accumulate to \$173,236 upon retirement ($\Sigma_{t=5}^{t=37}\$1*(1+0.07)^{40-t}$). Fourth, the \$1,214 pay gap per year in the late-career stage will accumulate to \$3,903 upon retirement ($\Sigma_{t=38}^{t=40}\$1*(1+0.07)^{40-t}$). Combined, our representative female agent has a \$250,689 deficit upon retirement, due to persistent pay gaps over her career in the federal government.

Taken together, our analysis illustrates that even modest annual pay disparities – when persistent and compounded over a full career – translate into economically meaningful differences in wealth. It is important to note that the \$250,689 figure likely represents a lower bound. First, it does not account for differences in pension payments which are generally a function of salary. Second, it does not account for potential differences in promotion, which we examine directly in a subsequent section. Third, these calculations abstract from behavioral responses that may amplify lifetime gaps. If systematic pay penalties alter women's career trajectories – for example, by discouraging mobility into higher-paying units – the true cumulative loss could be substantially larger. Finally, our exercise assumes a representative career path with no employment interruptions. Any career breaks or transitions into lower-paid roles, which disproportionately affect women in many settings, would further widen the lifetime gap.

3 Contributing Factors

Findings in Section 2 illustrate that gender pay gaps within the federal government are pervasive, even after accounting for a rich set of pay determinants. In this section, we utilize the breadth of our dataset to expand on our main finding and investigate the institutional mechanisms that contribute to the gap. In the next section, we turn to the economic mechanisms driving this gap.

3.1 Sorting

The first mechanism we examine is sorting, whereby women tend to work in agencies or in offices with lower pay. Given that we have a single employer (the US Government), we are able to more precisely estimate sorting effects across various dimensions. In particular, we focus on the factors that potentially cause the average pay gap to be 0.8% when measured within extremely narrowly defined categories, but up to 19.8% when defined across occupations, ranks, and agencies (based on Table 2).

To evaluate sorting, we aggregate the data at several alternative levels. First, we aggregate to the year×rank level, for example, 2010×GS-10 and 2010×GS-11 (where GS-10 and GS-11 are government ranks). Within each bucket, we compute the average pay for males ($AvgPayMale$) and the percent of female employees ($Female^{\%}$). We multiply the latter by one hundred for ease of interpretation and omit work units with 0% or 100% female representation. We then estimate the following univariate regression:

$$AvgPayMale_{r,t} = Female^{\%}_{r,t} + \lambda_r \tag{2}$$

This specification compares ranks over time: we examine whether years in which a given rank has relatively more female also represent years in which that rank's pay level is relatively lower. Results are shown in column (1) of Table 5. A one–standard-deviation increase in female representation is associated with a $549 decrease in average *male* pay, or 15.4% of the sample mean. We do not interpret this pattern as causal evidence, since the fraction of female employees is highly endogenous. Instead, it provides suggestive evidence that women tend to sit in lower-paying parts of the federal pay structure.

In column (2), we aggregate to the year×rank×county level. Here, we effectively shut down the option to relocate, thus focusing on the selection of occupation and agency to work in, within a county and rank. Findings again suggest that women are significantly more likely to work in agencies and occupations that pay less even for males. In economic terms, a one-standard-deviation increase in female representation is associated with $281 decrease in male pay, which is 9.9% over the sample mean.

Finally, in column (3), we aggregate to the year×rank×county×occupation level. In this specification, we additionally hold constant a person's occupation, for example whether they are a scientist or an attorney. To provide an example, we are examining whether female attorneys of rank GS-12 in Philadelphia are more likely to work for agencies that tend to pay less, e.g., the SEC, versus the EPA, versus the Federal Reserve, etc. Findings indicate that a one-standard-deviation increase in female representation is associated with $216 decrease in male pay, which is 6.9% over the sample mean.

In sum, these patterns indicate that women are disproportionately located in work units that offer lower pay – even for male employees. This holds true within the same geographic location and within the same occupation, suggesting that sorting operates through more than just differences in local labor markets or job types. The findings raise a natural question: what drives this sorting? One possibility, following Goldin (2014), is that women may value nonpecuniary aspects of jobs, such as flexibility or predictability, and may accept lower compensation in exchange. However, the fact that all our analyses examine a single employer (U.S. government) with standardized employment contracts and benefits makes this explanation less compelling. A second possibility is that women are more likely to self-select into certain domains such as labor, environment, or public health rather than finance, enforcement, or technology, and that agencies in these domains tend to pay less on average.

3.2 Promotions

The next contributing factor we investigate is likelihood of promotion. The tendency of women to sort into lower-paying work units, as shown in Table 5, is likely driven at least in part by the voluntary choices of men compared to women. In contrast, any potential differences in promotion rates are less likely to be a choice variable. We estimate Equation (1) with the new outcome variable $Promotion$, which equals 1 if the individual's rank at time t is higher than their rank at time $t-1$. We multiply $Promotion$ by 100 to facilitate interpretation. We focus on individuals who are in the government in both time t and $t-1$, meaning that we exclude new hires at time t. We further exclude employees who switch pay plans, since the hierarchical order is ambiguous.[9] We add year $\times$ agency $\times$ county $\times$ occupation $\times$ rank fixed effects.

The results are shown in the first three columns of Table 6. We find that females are 0.045% less likely to be promoted, compared to males. This is equivalent to 4.3% difference relative to the mean rate of promotion, which is 1.1%. As shown in columns (2) and (3), the gender gap in promotions exists in both HR-1 and HR-2. However, the magnitude is substantially larger in HR-1, even relative to the respective sample means (which are higher in HR-1 then HR-2, 1.5% compared to 0.8%). Relative to the sample mean, females were 7.7% less likely to be promoted during the earlier subsample, compared to 2.4% in the later subsample. The narrowing of promotion gaps over time is likely driven at least partially by females' higher levels of education in recent decades. Over the years of the HR1 sample, 1973 - 1995 period, 21.5% of males age 25 years an older had a college degree, compared

[9] There are several instances within our sample in which an agency moved from one pay plan to another. For example, in 2002, the SEC moved from the GS rank-based pay plan to the SK/SO pay plan.

to only 14.4% of females, a difference of 6.7%. In contrast, during 1996 - 2022 period, the difference had narrowed to only 1.3% (30.8% of males vs 29.5% of females).[10]

In columns (4) - (6) of Table 6, we examine whether promotion gaps vary by ranks. A large body of literature shows that women are less likely to negotiate aggressively (see, e.g., Bertrand (2011) and Mazei (2015) , among others). Moreover, these differences in negotiating power contribute to worse career outcomes for women (Dittrich, Knabe, and Leipold (2015), Eckel and Grossman (2008), and Save-Soderbergh (2019)). Because higher-level jobs involve more intangible tasks, where promotions are more influenced by soft skills, negotiating power arguably plays a bigger role at these levels. This suggests that differences in promotion will be greater at the higher ranks. In contrast, Becker (1957) and Becker (1962) argue that pay gaps should be lower when workers have higher levels of education and experience, as they should have more bargaining power. To test the competing predictions, we focus on the General Schedule, where the hierarchy is easily represented by the rank's number (for example, GS-2 has seniority over GS-1). We then estimate:

$$Promotion = \beta_1 \cdot Female + \beta_2 \cdot Rank \cdot Female + Rank, \tag{3}$$

where *Rank* is an integer from 1 through 15, representing the employee's post-promotion rank on the GS pay plan. We focus on the coefficient β_2, which captures how the promotion gender gap evolves as employees move up the ranks. Naturally, we omit rank FE from this specification. In the full sample, the estimated female coefficient is $\beta_1 = 0.452$ and the interaction with rank is $\beta_2 = -0.056$. Combined, this implies that women start with a hypothetical 0.452% promotion advantage, which erodes by 0.056 percentage points at each successive grade and eventually vanishes at higher levels. Dividing one coefficient by another, we estimate that the female advantage flips at GS-8. We estimate the same specification separately for HR-1 and HR-2 and find qualitatively similar results, identifying GS-6 and

[10] These statistics are based on the Current Population Survey (CPS) Historical Time Series Tables, published by the U.S. Census Bureau (`https://www.census.gov/data/tables/time-series/demo/educational-attainment/cps-historical-time-series.html`; last accessed: 11/13/2025).

GS-10 as the pivotal points, respectively.

In sum, our promotion results reinforce the earlier evidence: women advance more slowly than men, and these disparities tend to be larger at higher ranks. Rather than shrinking with experience – as suggested by Becker (1957, 1962) – gender gaps in advancement appear to widen in parts of the hierarchy where discretion and negotiation matter more. These patterns underscore the importance of upward mobility, as an important factor contributing to the long-run gender pay gaps in federal employment.

3.3 Pay raises within rank

In this subsection, we focus on factors that contribute to pay gaps within narrowly defined categories, for example year $\times$ agency $\times$ county $\times$ occupation $\times$ rank, for which we estimate a 0.8% pay gap in Table 2. We investigate women's ability to negotiate pay raises within their pay rank, conditional on not receiving a promotion to a higher rank. We again focus on the General Schedule (GS). In the GS, each rank is further divided into 10 steps, where the first (last) step represents the lower (upper) bound for this particular rank. There are therefore a total of 150 steps, 10 for each of the 15 ranks. The steps, especially the upper and lower bound for each rank, are updated annually (typically upwards). We submit a separate FOIA request and obtain the full data on all 150 steps since 1980. We then define a new outcome variable, $RankPosition$:

$$RankPosition_{i,r,t} = \frac{Pay_{i,r,t} - MaxPay_{r,t}}{MaxPay_{r,t} - MinPay_{r,t}}, \tag{4}$$

where $Pay_{i,r,t}$ is the adjusted pay for employee i in rank r at time t. $MaxPay_{r,t}$ and $MinPay_{r,t}$ represent the upper and lower bound of rank r at time t, respectively. Thus, the numerator is the employee's distance from the rank's upper bound, and the denominator is the rank's width. If $RankPosition = 0\%$, the employee is positioned at the upper end of his/her rank; if $RankPosition = -100\%$, the employee is at the lowest end of his/her rank.

We then estimate Equation (1) with *RankPosition* as the outcome, with gradually tighter fixed effects as in Table 2. The results are in Table 7.

Consistent with intuition, individuals with greater tenure are more likely to be at the top of their respective rank, as evidenced by the positive coefficient on Tenure. Turning to our coefficient of interest, across all specifications, the coefficient on Female is significantly negative: females are substantially less likely to be at the top of their respective rank, compared to males. That is, they are thus much closer to (farther from) the minimum (maximum) pay available for their GS rank. In our tightest specification, with year $\times$ agency $\times$ county $\times$ occupation $\times$ rank fixed effects, females are 3.4 percentage points farther from the upper bound of their rank than males, which represents 4% relative to the sample mean.

In sum, the findings in this section underscore systematic gender gaps at multiple stages of the career ladder. First, females are more likely to sort into offices and agencies that pay less. Second, they are less likely to be promoted. This gap is especially pronounced at higher ranks, where discretion and negotiation play a larger role. Third, conditional on remaining within the same rank, women are significantly less likely to be positioned near the top of the pay range. Taken together, these patterns reveal that gender pay gaps arise not from a single margin but from the cumulative effect of sorting, slower advancement, and smaller raises within rank.

4 Economic Mechanisms

In Section 2, we document persistent and substantial gender pay gaps, and Section 3 provides insight into the factors such as sorting and promotion likelihood that contribute to these gaps. In this section, we study three economic mechanisms that potentially underlie these channels, and we evaluate their relative importance.

4.1 Experience and education

The first mechanism we consider is employee and job type. In particular, we examine whether the gender pay gap varies systematically with the employee's career level (rank) and educational level. On the one hand, higher levels of achievement may provide a clearer signal of productivity and strengthen the employee's bargaining power (Becker, 1957, 1962). This should, in principle, narrow gender pay gaps. For example, federal pay scales explicitly tie many jobs to degree requirements, implying that men and women with equivalent qualifications should receive comparable compensation. On the other hand, advanced degrees and greater experience may exacerbate disparities if women face barriers in translating credentials into higher earnings. This could be due to differential access to high-paying assignments, differences in negotiation behavior, or persistent stereotypes about leadership and competence at the top of the pay scale (e.g., Dittrich, Knabe, and Leipold, 2015; Eckel and Grossman, 2008; Save-Soderbergh, 2019). These frictions are likely to be more salient in higher-level roles, where performance is less easily measured. For instance, jobs emphasizing regulatory quality or strategic oversight involve subjective assessments, whereas lower-level roles often rely on easily quantifiable outputs such as the number of cases processed. Analogously, our promotion results (Table 6) suggest that differences in negotiating style are more influential at higher ranks.

We begin by examining variation in pay gaps across ranks. In addition to the considerations mentioned above, the government-wide cap on pay creates a natural pay compression at the higher ranks (see, for example, Chen, Hajda, and Kalmenovitz (2024)), which could further mitigate any gender-based differences in earnings. To test these possibilities we again focus on the General Schedule (GS) and estimate Equation (1) with year $\times$ agency $\times$ county $\times$ occupation fixed effects, for each rank separately.[11] We naturally omit rank fixed effects. The results in Table 8 reveal substantial difference across ranks. Among lower-ranked em-

[11] Incidences of ranks from GS-16 and above that show up in the sample ended by 1991, and we therefore include them in GS-15.

ployees, GS-8 and lower, the gender pay gap is insignificant or slightly positive, implying that women in these positions face fewer earnings disparities relative to their male counterparts. Note that sample size is substantial in all ranks (30,000 observations at least), so the insignificant pay gap cannot be attributed to lack of statistical power. The gender gap starts to evolve only among higher-ranked employees (GS-9 and above). The effect does not seem to vary substantially from GS-9 and above, suggesting that gender pay gaps are persistent and similar across the government's higher-ranked employees.

Next we examine the influence of education. We focus on the early sample period (1973-1995), where we have information on the employee's education. We group employees into four categories: (i) those with associate degrees, high school education, or lower (including non-academic occupational training), (ii) those with bachelor degrees, (iii) those with master's degrees, and (iv) those with doctoral degrees. We estimate our tightest specification for each of those groups with year, agency, county, occupation, and rank fixed effects (interacted). Note that there is some overlap between education – our object of interest – and rank, since some ranks require minimum education level.[12] We circumvent this problem by adding rank fixed effects, thereby focusing on within-rank pay gaps as a function of education. The results in Table 9 show that gender gaps exist across all education levels. Focusing on the economic magnitude, we find that the gender pay gap increases with the level of education, which echoes prior results showing wider gaps at higher ranks. Among employees with high-school education, gender pay gaps are 1.1% on average, compared to 1.2% among those with a Bacherlor's, 1.3% among those with a Master's, and a striking 3.0% among those with a PhD. Our finding that the gender pay gap is over twice as large among individuals with the highest levels of education is particularly notable given our stringent set of fixed effects.

In sum, the gender pay gap is more pronounced at higher ranks and among those with higher levels of education. These patterns reinforce the broader set of results in the paper. Earlier, we showed that women are more likely to sort into lower-paying offices and agencies;

[12] For instance, according to current OPM guidelines (as of July 2025), to qualify for a GS-7 job one must have one year of specialized experience, but GS-9 job requires a master's degree.

are promoted at lower rates, especially in the more senior parts of the hierarchy; and are less likely to reach the upper end of the pay range within a given rank. The present findings add another layer: gender pay gaps widen as employees move into higher-skilled and higher-ranked positions, where subjective evaluation and negotiation play a larger role.

4.2 Female leadership

The next channel we examine is the influence of the manager, in particular whether the presence of females at top positions affects gender gaps throughout their organizations. Two competing hypotheses guide this analysis. On one hand, the "role model" hypothesis predicts that having at least one woman in a senior executive role signals a more inclusive culture and mitigates bias, thereby narrowing pay and promotion gaps Cohen and Huffman (2007); Tate and Yang (2015). Relatedly, to the extent that taste-based discrimination contributes to the gender gap (Becker (1957)), such bias would be eliminated under female leaders. On the other hand, the "queen bee" hypothesis suggests that female executives may less vigorously seek to narrow the gender gaps experienced by other females, yielding either larger or unchanged gender disparities in their organization (Derks et al. (2011); Srivastava and Sherman (2015).

To test the competing predictions, we must identify the head of each office. Here, we take advantage of the structure of government agencies, wherein political appointees generally exercise greater formal authority. Within our data, these individuals are designated by the separate pay plan "EX." We limit the analysis to the DC office of each agency, where the majority of the political appointees are located and thus can have the greatest influence if they choose to exert it. We define two alternative measures of the female presence within the political ranks, within each agency office. First, we define an indicator which equals 1 if agency a has at least one female executive at time t. Second, we define a continuous variable which represents the fraction of females among the agency's political leadership. We then estimate a version of Equation (1):

$$y_{i,a,t} = \beta_1 \cdot Female_i + \beta_2 \cdot TopFemales_{a,t} \cdot Female_i + Tenure + \lambda, \tag{5}$$

where the outcome is the employee's log of adjusted base pay (expressed in 2023 USD) and $TopFemales$ is either the indicator or the continuous measure of female leadership representation. All regressions control for tenure and include year, agency, county, occupation and rank fixed effects (interacted). Note that $TopFemales_{a,t}$ is subsumed under year×agency fixed effects, as it measures female leadership presence at the agency level. Naturally, we exclude the EX employees themselves from this analysis. As before, we cluster standard errors at the county and year level.

The results are shown in Table 10. In column (1), we estimate our workhorse specification (Equation (1)) on this sample, which consists only of DC offices. We find that women earn roughly 1.5% less than comparable men. This is slightly larger than the average gender gap in the full sample (Table 2). In column (2), we add the interaction of female with an indicator for at least one female executive. We find that the gender gap is approximately 30% lower when at least one female executive is present, relative to agencies without any female executive (β_2/β_1). In column (3), we replace the indicator with a continuous measure for the proportion of female executives out of all executives. We find that each 10 percentage-point increase in female executives reduces the gap by about 3.1%.[13]

In sum, we find that the presence of female executives mitigates gender pay gaps among lower-ranked employees, consistent with a positive "role model" effect rather than a "queen bee" pattern. These effects, however, are partial: female leadership narrows but does not eliminate pay disparities. Taken together with our earlier findings, this suggests a coherent pattern. Women face disadvantages along several margins: sorting into lower-paying units, slower promotion rates, smaller raises within rank, and wider gaps at higher levels of

[13] The variable, *Top Females*$^{(\%)}$, is measured in decimal form, so a 10 percentage point change is represented by 0.1 * 0.005, which equals 0.0005. Relative to the average effect of being Female of -0.016, this represents a 3.1% reduction.

education and seniority. Female leadership appears to attenuate one component of these disparities, particularly at the base of the hierarchy where executives may have greater visibility into day-to-day personnel practices. Yet the persistence of gaps even under female leadership indicates that structural features of pay setting continue to generate gender differences that leadership alone cannot fully offset.

4.3 Labor mobility

Finally, we investigate the role of labor mobility. Here, the question is whether females have the possibility to move to a different location with higher pay opportunities. We proceed in two steps. First, we ask if females are more constrained in their ability to move. Second, focusing on cases where males and females alike are forced to move, we examine whether females move to offices with smaller gender pay gaps.

To test the first question, we estimate Equation (1) with a different outcome variable, $Moved_{t+1}$, which equals one if the employee moved to a different office within the federal government at time $t+1$. We use the most stringent set of fixed effects and control for tenure, thus comparing employees with the opposite gender who work at the same office, occupation, and rank. The results are in Table 11, Panel A. We find virtually no difference between males and females in their tendency to move, as the coefficient on female indicator is zero and statistically insignificant. In columns (2)-(3) we estimate it separately on the early and the late time period (HR-1 and HR-2), and again find no gender gaps in their tendency to move.

Turning to the second question, we study the career trajectories of males and females whose office has been closed. For example, the Department of Justice closed four of its Antitrust Division regional field offices in 2011. In this scenario, all the employees are forced to move from their current locations. We examine whether females are able to capitalize on this opportunity, by moving to an office with smaller gender pay gap. We focus on offices which have closed and estimate the following regression:

$$y_{i,o,t} = \beta_1 \cdot Female_i + \beta_2 \cdot OfficeClosure_{o,t} \qquad (6)$$
$$+ \beta_3 \cdot OfficeClosure_{o,t} \cdot Female_i + Tenure_{i,t} + \lambda$$

where $y_{i,o,t}$ is the outcome for employee i working in office o, and $OfficeClosure_{o,t} = 1$ if the regional office has been closed between time $t-1$ and time t. The focus is on β_3, representing the differences in outcome for females after their office has been closed. We use the tightest set of fixed effects – year, agency, county, occupation, and rank (interacted) and continue to control for tenure.[14] The results are in Table 11, Panel B.

In column (1), the outcome is exit, which equals one if the employee was in the government at time $t-1$ but is no longer there at time t. On average throughout the sample period, females are 0.8% less likely than males to exit the federal government (β_1), suggesting that females view their outside option less favorably than males. Moreover, on average across males and females, office closure is associated with a 14.2% greater likelihood of leaving the government (β_2), confirming that office closure is a material shock to federal careers. However, following the shock that forces a career change of some form, we fail to find strong evidence that females are more likely than males to exit; the coefficient on β_3 is negative, but insignificant at conventional levels.

In column (2), the outcome is the employee's log pay at time t. We find that females earn less than males (β_1), consistent with the central result throughout the paper. We additionally find that office closure is associated with a pay decrease of 0.6% (β_2). However, more central to our main question, we find no evidence that females take advantage of this forced move to join an office where they are paid more equally to their male counterparts, as reflected by the insignificant coefficient on β_3.

In sum, we find no evidence that limited mobility contributes meaningfully to the gender

[14] Note that $OfficeClosure_{o,t}$ is not fully subsumed under year-agency-county FE, since the employee has moved to a new office and his/her new colleagues have not experienced an office closure in their past.

pay gap. First, men and women are equally likely to transfer across offices in normal times. Second, when office closures force employees to relocate or exit, we again see no systematic gender differences in separation rates. Third, among those who stay in the government and relocate, women do not appear to use the disruption as an opportunity to move into offices with more equitable pay structures. Taken together with earlier findings on sorting, promotions, and within-rank raises, these results suggest that gender disparities in federal pay are not driven by differential access to higher-paying locations or mobility constraints. Instead, the persistence of the gap may arise from the mechanisms documented elsewhere in the paper – sorting into lower-paying organizational units, slower advancement at higher ranks, and lower progression within rank – rather than from differences in geographic opportunity or relocation behavior.

5 Conclusions

We provide the first comprehensive, career-level analysis of gender pay gaps in the U.S. Federal workforce, leveraging a uniquely rich dataset that tracks nearly all civilian federal employees over five decades. We document substantial and persistent differences in earnings between male and female employees, which accumulate to at least $250,000 by retirement age for a representative female employee. Importantly, the gaps we identify are not driven by isolated agencies or specific subgroups: they appear across occupations, agencies, and cohorts, indicating a broad and persistent pattern in federal labor markets. Using detailed information on employee pay, rank, location, occupation, tenure, and promotions, we decompose these gaps into components associated with sorting across units, promotions, and pay growth. We show that female employees are disproportionately concentrated in lower-paying organizational units, face slower promotion rates, and receive lower pay raises. Taken together, the evidence points to structural forces – rather than short-term shocks or idiosyncratic practices – as the primary contributors to the gender pay gap in the federal workforce.

Beyond describing the aggregate patterns, our results also carry direct policy implications. The federal government has long positioned itself as a model employer with standardized pay schedules, formal promotion procedures, and workplace protections designed to limit discretion in pay-setting. Yet our findings indicate that large gender disparities survive even in this highly regulated environment. The recent dismantling of Diversity, Equity, and Inclusion (DEI) initiatives further elevates the importance of understanding how pay gaps arise and how they evolve over employees' careers. Our analysis provides timely evidence on the scope of these gaps and the structural mechanisms through which they accumulate.

Several avenues remain open for future work. First, while our study documents the magnitude and sources of gender pay gaps, it does not take a stand on welfare: future research could quantify how these disparities affect, for instance, the overall quality of government services. Second, our results uncover substantial heterogeneity across agencies and occupations, suggesting fertile ground for research on organizational practices that amplify or mitigate pay disparities. Third, linking our federal dataset to private-sector labor market outcomes would allow us to study how pay disparities inside the government translate into outside job opportunities. Finally, the recent elimination of federal DEI programs creates an important opportunity for future research to assess how major institutional changes affect the evolution of gender pay gaps in the public sector.

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Figure 1: **Time Series of the Gender Pay Gap**

This figure plots annual estimates of the gender pay gap in the U.S. federal workforce from 1973–2023, based on yearly regressions of pay on a female indicator. The panels show estimates under increasingly rich fixed effects: unconditional (top left), adding agency FE (top middle), agency×county FE (top left), agency×county×occupation FE (bottom left), and agency×county×occupation×rank FE (bottom right). Solid blue lines show point estimates and dashed red lines show 95% confidence intervals. The key takeaway is that while aggregate pay gaps decline, the within-job-cell gap remains persistent. See Section 2.2.

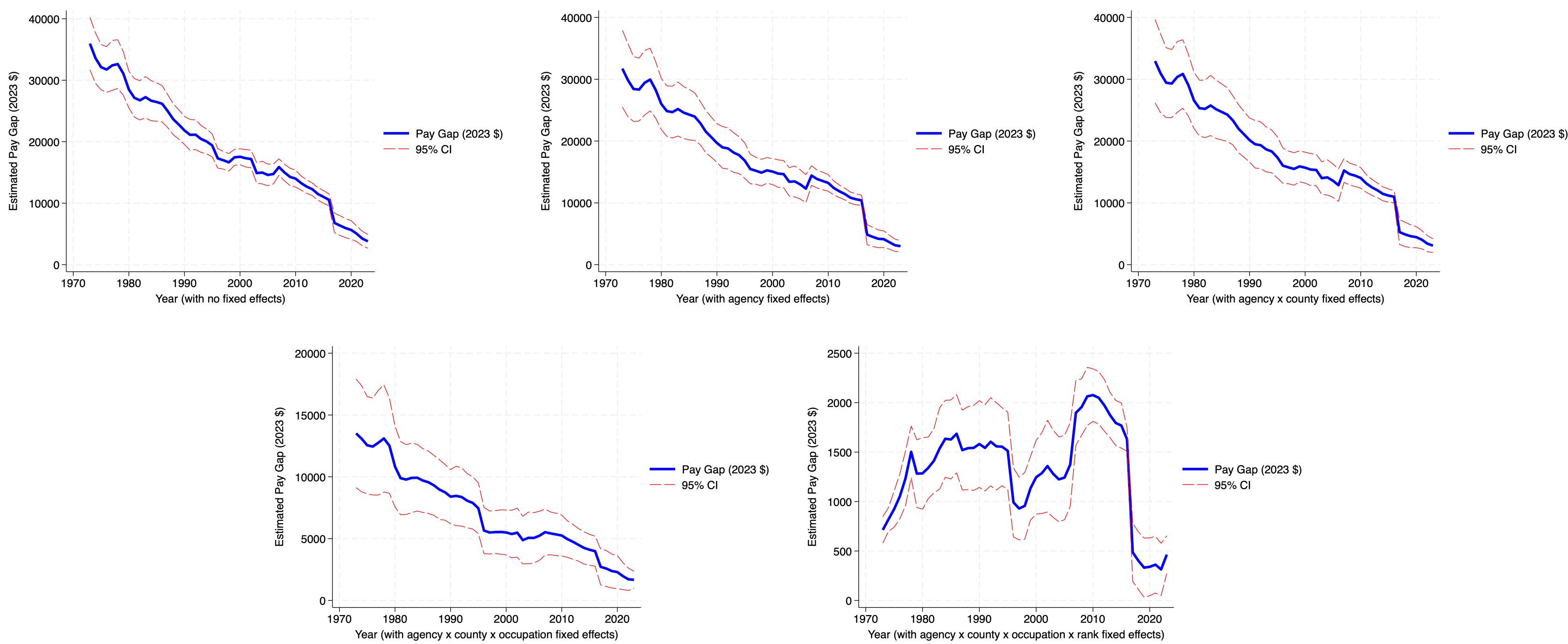

Figure 2: **Distribution of Gender Pay Gap across Agencies**

This figure shows the distribution of agency-level gender pay gaps based on separate regressions estimated for each federal agency using *LogPay* (2023 USD) as the outcome. All specifications include year, county, occupation, and rank fixed effects (interacted), and coefficients represent within-agency gender gaps after accounting for these controls. Standard errors are clustered at the county and year level. Panel A plots the distribution across all agencies, while Panel B restricts to agencies with statistically significant coefficients (p-value of 0.05 or lower). The key takeaway is that gender pay disparities are widespread across the federal government, with most agencies exhibiting significant gaps in favor of men. See Section 2.2.

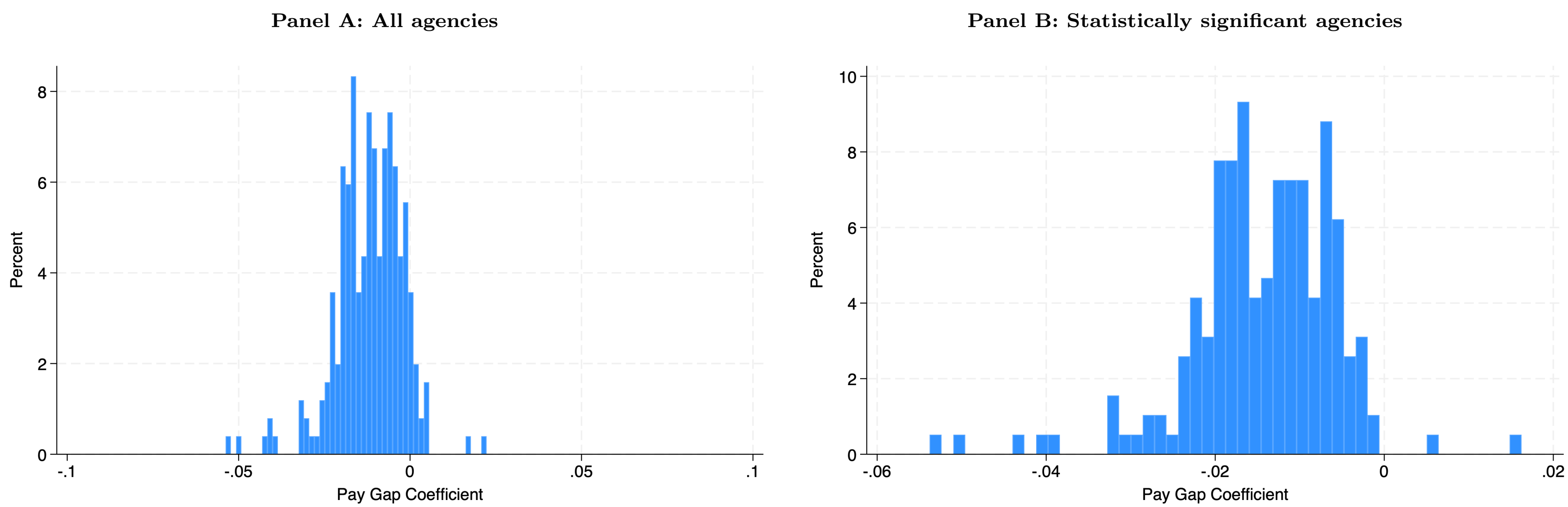

Figure 3: **Geographic Distribution of Gender Pay Gaps**

This figure shows the distribution of office-level gender pay gaps based on separate regressions estimated for each local office, defined as an agency-by-county unit, using $LogPay$ (2023 USD) as the outcome. All specifications include year, occupation, and rank fixed effects (interacted), and coefficients represent within-office gender gaps after accounting for these controls. Standard errors are clustered at the county and year level. We plot the distribution for the subset of 1,301 local offices in which the estimated coefficient is statistically significant (p-value of 0.05 or lower). Gender pay disparities persist even within narrowly defined local offices, and some offices show substantial negative pay gaps against women, while large positive gaps in favor of women are rarer. See Section 2.2.

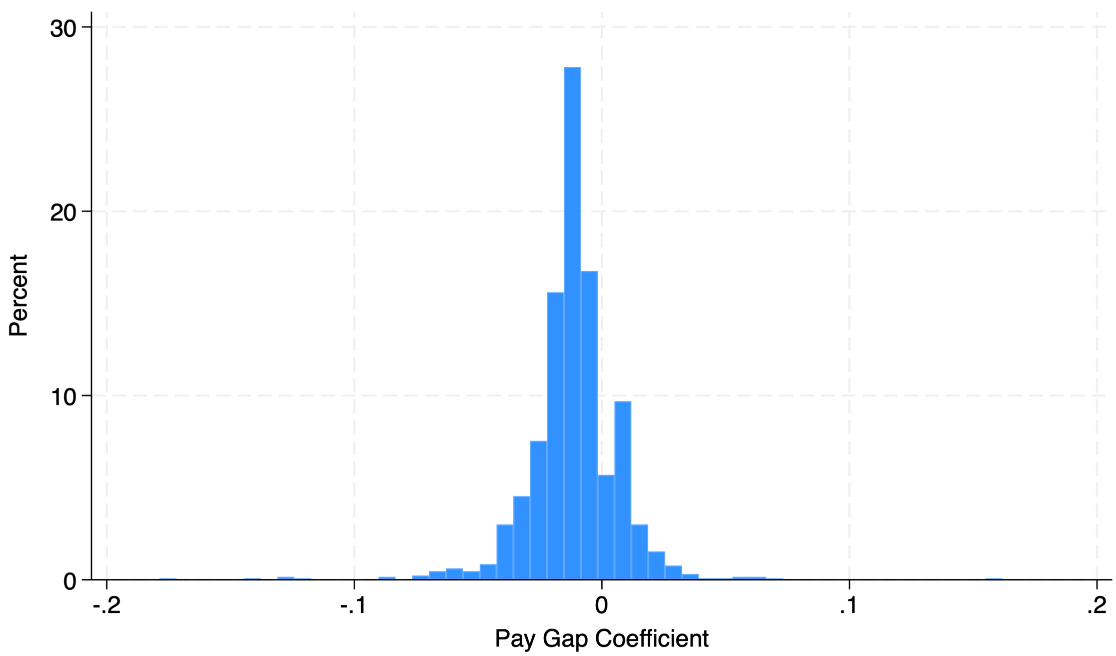

Table 1: **Summary Statistics**

This table reports summary statistics of employee-year level variables. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's number of years in federal service. *Pay* is the employee's adjusted base pay (base pay plus locality adjustment) for years prior to 1995 and the employee's base pay for years after 1995. *Pay (2023USD)* is the employee's pay expressed in 2023 USD. *Promoted*, *Moved*, and *Pay Raise* are indicators equal to one if, respectively, the employee is promoted within the same pay scale, relocates to a different office (agency × county), or receives a higher real salary than in the prior year. For presentation, all indicators are multiplied by 100. Panel A includes the full sample from 1973–2023, Panel B covers 1973–1995, and Panel C covers 1996–2023. See Section 1.

Panel A: Full sample (1973–2023)

	count	mean	sd	min	p50	max
Female (%)	36,474,721	49.62	50.00	0.00	0.00	100.00
Tenure	36,474,721	10.84	9.60	0.00	8.00	76.00
Pay	36,474,721	50,275.13	34,930.38	1.00	41,339.00	492,643.00
Pay (2023USD)	36,474,721	85,855.45	43,280.29	2.33	76,616.55	686,254.44
Promoted (%)	30,436,313	1.10	10.44	0.00	0.00	100.00
Moved (%)	31,138,363	6.00	23.75	0.00	0.00	100.00
Pay Raise (%)	31,138,363	63.76	48.07	0.00	100.00	100.00

Panel B: HR-1 sample (1973–1995)

	count	mean	sd	min	p50	max
Female (%)	14,094,381	45.83	49.83	0.00	0.00	100.00
Tenure	14,094,381	6.52	5.77	0.00	5.00	22.00
Pay	14,094,381	26,037.40	16,018.24	1.00	21,724.00	184,008.00
Pay (2023USD)	14,094,381	79,470.44	39,908.99	2.33	68,532.36	686,254.44
Promoted (%)	11,473,671	1.52	12.24	0.00	0.00	100.00
Moved (%)	11,816,651	7.47	26.29	0.00	0.00	100.00
Pay Raise (%)	11,816,651	60.20	48.95	0.00	100.00	100.00

Panel C: HR-2 sample (1996–2023)

	count	mean	sd	min	p50	max
Female (%)	22,380,340	52.01	49.96	0.00	100.00	100.00
Tenure	22,380,340	13.57	10.50	0.00	12.00	76.00
Pay	22,380,340	65,539.23	34,985.44	3.00	58,808.00	492,643.00
Pay (2023USD)	22,380,340	89,876.51	44,810.29	5.61	82,733.55	565,426.50
Promoted (%)	18,962,642	0.85	9.17	0.00	0.00	100.00
Moved (%)	19,321,712	5.10	22.00	0.00	0.00	100.00
Pay Raise (%)	19,321,712	65.93	47.39	0.00	100.00	100.00

Table 2: **Existence and Magnitue of the Gender Pay Gap**

This table reports estimates of the gender pay gap using the full employee-year sample from 1973–2023. *Pay (2023 USD)* is the employee's pay expressed in 2023 USD, and *Log Pay (2023 USD)* is its natural logarithm. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's years of federal service. Across columns, we progressively introduce interacted fixed effects for year, agency, county, occupation, and rank, allowing us to compare employees within increasingly narrow job cells. Standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. See Section 2.1.

Panel A: Pay as outcome

	Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-18089***	-17893***	-15867***	-16490***	-6181***	-1248***
	(1509.8)	(1643.4)	(1754.6)	(1934.2)	(1125.8)	(166.1)
Tenure		1481***	1321***	1274***	1267***	602***
		(57.5)	(52.4)	(47.1)	(42.5)	(16.7)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	85855	85855	85855	85916	86530	86699
Effect (%)	-21.07	-20.84	-18.48	-19.19	-7.14	-1.44
Change in Adjusted R^2		0.113	0.150	0.038	0.428	0.173
Adjusted R^2	0.044	0.157	0.306	0.344	0.773	0.946
Observations	36,474,721	36,474,721	36,474,693	36,126,747	33,756,528	29,964,231

Panel B: Log pay as outcome

	Log Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.198***	-0.195***	-0.167***	-0.172***	-0.055***	-0.008***
	(0.017)	(0.018)	(0.019)	(0.020)	(0.009)	(0.001)
Tenure		0.019***	0.017***	0.016***	0.015***	0.007***
		(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Change in Adjusted R^2		0.142	0.153	0.043	0.434	0.154
Adjusted R^2	0.042	0.184	0.337	0.380	0.814	0.968
Observations	36,474,721	36,474,721	36,474,693	36,126,747	33,756,528	29,964,231

Table 3: **Gender Pay Gap by Career Stage**

Pay (2023 USD) is the employee's pay in 2023 USD. *Log Pay (2023 USD)* is the natural logarithm of the employee's pay in 2023 USD. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. To estimate career stages, this table excludes the following from the sample: (1) employees in HR-1 sample whose first observed year is 1973, (2) observations in 1993–1995 for employees whose last observed year is 1995, and (3) observations in 2021–2023 for employees whose last observed year is 2023. Late-career stage employees are defined as individuals in their last three years of working in the federal government. Mid-career stage employees are defined as individuals working in at least their sixth year of employment in the federal government, conditional on not meeting the definition of late-career employees. Early-career employees are defined as individuals working in their second through sixth year of employment, conditional on not meeting the definition of late-career or mid-career employees. First-year employees are defined as individuals working in their first of employment, conditional on not meeting the definition of late-career, mid-career, or early-career employees. Robust standard errors are adjusted for clustering at the county and year level and are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively. See Section 2.3.

Panel A: Pay as outcome

	Pay (2023 USD)			
	First Year (1)	Early Career (2)	Mid Career (3)	Late Career (4)
Female	-1074***	-1133***	-1189***	-1214***
	(200.4)	(186.2)	(158.0)	(149.3)
YxAxCtxOxR FE	Yes	Yes	Yes	Yes
Sample Mean	60883	72552	96480	78332
Effect (%)	-1.76	-1.56	-1.23	-1.55
Adjusted R^2	0.916	0.929	0.940	0.935
Observations	843,640	3,208,776	11,211,925	3,973,832

Panel B: Log pay as outcome

	Log Pay (2023 USD)			
	First Year (1)	Early Career (2)	Mid Career (3)	Late Career (4)
Female	-0.006***	-0.006***	-0.007***	-0.007***
	(0.001)	(0.001)	(0.001)	(0.001)
YxAxCtxOxR FE	Yes	Yes	Yes	Yes
Adjusted R^2	0.961	0.964	0.962	0.963
Observations	843,640	3,208,776	11,211,925	3,973,832

Table 4: **Cumulative Income Loss Due to Gender Pay Gaps**

This table reports estimates of the gender pay gap at different career stages and its cumulative value at retirement. We divide federal careers into four segments: first year of service, early career (years 2–5), mid-career (years 6 until three years prior to retirement), and late career (final three years). For each stage, we re-estimate the baseline specification and obtain the implied annual gender pay gap (results in Table 3). To assess long-run implications, we assume a 7% annual rate of return and compute the future value of stage-specific pay gaps at the point of retirement, thereby translating differences in annual earnings into long-term wealth consequences. See Section 2.3.

Career level	Pay gap at this stage	Years at this stage	Future value of losses at this stage
Hiring	\$1,074	1	\$16,083
Early-Career	\$1,133	4	\$57,468
Mid-Career	\$1,189	32	\$173,236
Late-Career	\$1,214	3	\$3,903
Total			\$250,689

Table 5: **Contribution of Sorting to Observed Gender Pay Gaps**

This table examines whether sorting across organizational units contributes to the observed genders pay disparities. We aggregate the employee-level data into progressively finer levels of aggregation: year×rank (column (1)), year×rank×county (column (2)), and year×rank×county×occupation (column (3)). In each column, the outcome is the mean pay of male employees in 2023 USD, and *Female Employee (%)* is the share of employees identified as female in this organizational unit, multiplied by one hundred for clarity. Standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. For example, in column (3), we compare GS-10 attorneys in Philadelphia in 2010 across different agencies, and ask whether units with a higher share of women also differ systematically in male pay. We find that females tend to work in organizational units where pay is lower, even for males. See Section 3.1.

	Average Male Pay (USD 2023)		
	Year × Rank (1)	Year × Rank × County (2)	Year × Rank × County × Occupation (3)
Female Employee (%)	-549***		
	(56.4)		
Female Employee (%)		-281***	
		(18.3)	
Female Employee (%)			-216***
			(19.0)
Sample Mean	100967	81336	84943
Effect (%)	-15.41	-9.87	-6.87
Adjusted R^2	0.070	0.036	0.022
Observations	24,365	1,619,002	5,552,063

Table 6: **Gender Differences in Promotion Rates**

This table examines whether gender differences in promotion rates contribute to the observed pay disparities. *Promoted* = 1 if the employee's current pay grade is higher than in the prior year (conditional on remaining within the same pay scale), multiplied by one hundred for clarity. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's number of years in government service. *Rank* is the employee's numerical GS pay grade. Columns (4)–(6) include only employees on the GS pay schedule. Standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. We find that women are systematically less likely to be promoted (columns (1)–(3)), but those gaps only begin at relatively higher ranks (columns (4)–(6)). See Section 3.2.

	Promoted (%)					
	Full Sample (1)	HR1 (2)	HR2 (3)	Full Sample (GS) (4)	HR1 (GS) (5)	HR2 (GS) (6)
Female (β_1)	-0.045***	-0.115***	-0.019***	0.452***	0.616***	0.486***
	(0.011)	(0.033)	(0.006)	(0.041)	(0.084)	(0.051)
Tenure	0.010***	0.071***	0.004***	0.015***	0.067***	0.008***
	(0.003)	(0.024)	(0.001)	(0.003)	(0.012)	(0.002)
Female × Rank (β_2)				-0.056***	-0.103***	-0.050***
				(0.005)	(0.011)	(0.006)
Rank				-0.392***	-0.463***	-0.344***
				(0.028)	(0.051)	(0.027)
YxAxCtxOxR FE	Yes	Yes	Yes	No	No	No
YxAxCtxO FE	No	No	No	Yes	Yes	Yes
Sample Mean	1.07	1.53	0.79	-	-	-
Effect (%)	-4.26	-7.53	-2.41	-	-	-
$\|\beta_1/\beta_2\|$	-	-	-	8.03	5.99	9.63
Adjusted R^2	0.299	0.393	0.192	0.042	0.051	0.035
Observations	24,710,939	9,085,625	15,625,314	19,881,853	7,481,708	12,400,145

Table 7: **Pay Raises Absent Promotions**

This table examines gender gaps in how employees progress within the same General Schedule pay grade ("rank"). *Distance to Max Band* is (current pay minus the minimum pay for the GS rank) divided by (maximum minus minimum pay for that rank). *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's number of years in federal service. The sample is restricted to employees on the GS pay schedule. Robust standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The key takeaway is that women are systematically farther from the top pay of their GS pay grade, indicating slower within-grade progression that contributes to the persistent gender pay gaps. See Section 3.3.

	Distance to Max Band = (Pay - Max Pay)/(Max Pay - Min Pay)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.030***	-0.023***	-0.029***	-0.029***	-0.030***	-0.034***
	(0.005)	(0.006)	(0.005)	(0.005)	(0.005)	(0.005)
Tenure		0.013***	0.013***	0.013***	0.014***	0.016***
		(0.000)	(0.000)	(0.000)	(0.000)	(0.001)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	-0.84	-0.84	-0.84	-0.84	-0.84	-0.84
Effect (%)	3.57	2.74	3.45	3.45	3.57	4.05
Adjusted R^2	0.002	0.193	0.212	0.231	0.292	0.337
Observations	20,423,943	20,423,943	20,423,890	20,164,200	18,555,925	16,333,357

Table 8: **Gender Pay Gaps Along the General Schedule Grade Ladder**

This table estimates rank-by-rank gender gaps in federal pay. *Log Pay (2023 USD)* is the natural logarithm of an employee's pay expressed in constant 2023 dollars. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's number of years in federal service. The sample includes only employees on the GS pay schedule. Robust standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The estimates show that gender pay gaps are concentrated entirely at higher GS grades, with little to no gap among lower-ranked employees. See Section 4.1.

Panel A: GS 01-08

	Log Pay (2023 USD)							
	01	02	03	04	05	06	07	08
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Female	-0.000	0.001	0.001	0.001	0.002	0.006***	0.003***	0.003***
	(0.001)	(0.001)	(0.001)	(0.002)	(0.002)	(0.001)	(0.001)	(0.001)
Tenure	0.005***	0.008***	0.012***	0.011***	0.008***	0.007***	0.007***	0.006***
	(0.001)	(0.000)	(0.001)	(0.001)	(0.000)	(0.000)	(0.000)	(0.000)
YxAxCtxO FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.949	0.913	0.865	0.849	0.853	0.865	0.831	0.806
Observations	32,423	120,579	526,802	1,490,746	2,175,790	1,549,008	1,659,735	682,093

Panel B: GS 09-18

	Log Pay (2023 USD)						
	09	10	11	12	13	14	15-18
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Female	-0.003**	-0.017***	-0.010***	-0.015***	-0.016***	-0.014***	-0.013***
	(0.001)	(0.002)	(0.001)	(0.002)	(0.001)	(0.001)	(0.001)
Tenure	0.006***	0.007***	0.006***	0.005***	0.005***	0.005***	0.004***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
YxAxCtxO FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.825	0.771	0.804	0.760	0.733	0.725	0.720
Observations	1,665,344	454,129	2,420,950	2,867,381	2,536,795	1,491,971	825,798

Table 9: **Gender Pay Gaps Across Education Groups**

This table estimates gender pay gaps within education groups. *Log Pay (2023 USD)* is the natural logarithm of an employee's pay expressed in constant 2023 dollars. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's number of years in federal service. Column (1) includes employees with associate degrees, high school education, or lower (including terminal occupation programs). Column (2) includes employees with Bachelor's degrees. Column (3) includes employees with Master's degrees. Column (4) includes employees with PhDs. Robust standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The estimates show that gender pay gaps widen markedly with education, becoming largest among employees with advanced degrees. See Section 4.1.

	Log Pay (2023 USD)			
	High School	Bachelor's	Master's	PhD
	(1)	(2)	(3)	(4)
Female	-0.011***	-0.012***	-0.013***	-0.030***
	(0.001)	(0.001)	(0.002)	(0.004)
Tenure	0.011***	0.010***	0.009***	0.009***
	(0.001)	(0.001)	(0.001)	(0.001)
YxAxCtxOxR FE	Yes	Yes	Yes	Yes
Adjusted R^2	0.973	0.945	0.962	0.884
Observations	6,399,450	2,841,180	606,885	159,644

Table 10: **Gender Pay Gaps and Female Leadership**

This table examines whether the gender pay gap varies with the presence of females among the agency's leadership. We identify leadership using the "EX" pay plan, which designates senior political appointees, and restrict the sample to employees located in Washington, DC, where these appointees are concentrated. We construct two measures of female political presence at the agency level: *Top Females*$^{(0/1)}$ is an indicator equal to one if the agency has at least one female EX executive in year t, and *Top Females*$^{(\%)}$ is the proportion of EX executives who are female. We then estimate Equation (5), interacting each measure with an indicator for female employees. The dependent variable is *Log Pay (2023 USD)*. All regressions include fixed effects for year, agency, county, occupation, and rank (interacted). *Top Females* is absorbed by the year $\times$ agency $\times$ occupation $\times$ rank fixed effects. EX executives themselves are excluded from the analysis. Standard errors are clustered at the year level. We find that the presence of female executives mitigates gender pay gaps among lower-ranked employees. See Section 4.2.

	Log Pay (2023 USD)		
	(1)	(2)	(3)
Female	-0.015***	-0.018***	-0.016***
	(0.001)	(0.001)	(0.001)
Tenure	0.005***	0.005***	0.005***
	(0.000)	(0.000)	(0.000)
Female $\times$ Top Females$^{(0/1)}$		0.005***	
		(0.001)	
Female $\times$ Top Females$^{(\%)}$			0.005***
			(0.002)
YxAxOxR FE	Yes	Yes	Yes
Adjusted R^2	0.964	0.964	0.964
Observations	2,458,224	2,458,224	2,458,224

Table 11: **Mobility Constraints and Gender Pay Gaps**

Panel A. Mobility likelihood. This table estimates gender differences in the likelihood that employees change offices from one year to the next. *Moved (%)* is a variable that equals one hundred if the individual's office (agency $\times$ county) changes from that of the prior year, and zero otherwise. $Female$ equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's number of years in government service. Robust standard errors are adjusted for clustering at the county and year level and are reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% level, respectively. The results show that women are slightly more mobile than comparable men, but the differences are not statistically significant. See Section 4.3.

	Moved (%)		
	Full Sample	HR1	HR2
	(1)	(2)	(3)
Female	0.036	0.048	0.034
	(0.023)	(0.031)	(0.031)
Tenure	-0.137***	-0.181***	-0.133***
	(0.005)	(0.018)	(0.005)
YxAxCtxOxR FE	Yes	Yes	Yes
Sample Mean	5.19	6.68	4.32
Adjusted R^2	0.384	0.513	0.268
Observations	25,299,611	9,359,478	15,940,133

Panel B. Office closures. This panel tests whether gender differences in responses to forced relocation, triggered by regional office closures, contribute to gender pay gaps. *Office Closure* equals 1 if the employee worked in a regional office that closed between year t and year $t+1$. Offices in a given year with less than 10 employees are excluded from this analysis. $Female$ equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's years of federal service. The interaction term captures whether females respond differently to closure-induced relocation. All regressions include year, agency, county, occupation, and rank fixed effects (interacted). Standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. The estimates show no meaningful gender differences in closure-induced exits or pay changes, indicating that forced relocation does not contribute to the observed gender pay gaps. See Section 4.3.

	Exit Government Next Year (%)	Next Year Log Pay (2023 USD)
	(1)	(2)
Female (β_1)	-0.757***	-0.009***
	(0.121)	(0.001)
Office Closure (β_2)	14.216*	-0.006**
	(7.342)	(0.003)
Office Closure × Female (β_3)	-0.008	0.003
	(0.296)	(0.002)
Tenure	0.104***	0.005***
	(0.026)	(0.000)
YxCtxOxR FE	Yes	Yes
Sample Mean	10.85	11.30
Effect (%)	-0.08	0.29
Adjusted R^2	0.151	0.961
Observations	29,678,562	26,245,218

Internet Appendix

Table A.1: **Gender pay gap: robustness**

Panel A. This table is similar to Table 2, Panel A, except that we estimate pay gaps separately in HR-1 (1973–1995) and in HR-2 (1996–2023). *Pay (2023 USD)* is the employee's pay expressed in 2023 USD. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's years of federal service. Across columns, we progressively introduce interacted fixed effects for year, agency, county, occupation, and rank, allowing us to compare employees within increasingly narrow job cells. Standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. See Section 2.1.

Panel A: HR-1 sample (1973–1995)

	Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-28343***	-27084***	-24239***	-24873***	-10046***	-1359***
	(1955.5)	(1904.2)	(2355.5)	(2592.3)	(1555.0)	(174.3)
Tenure		2509***	2397***	2395***	2288***	763***
		(203.4)	(184.7)	(177.4)	(159.2)	(52.6)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	79470	79470	79470	79545	79958	79391
Effect (%)	-35.66	-34.08	-30.50	-31.27	-12.56	-1.71
Change in Adjusted R^2		0.096	0.118	0.035	0.354	0.233
Adjusted R^2	0.125	0.222	0.340	0.375	0.729	0.962
Observations	14,094,381	14,094,381	14,094,371	13,969,912	13,003,688	11,324,182

Panel B: HR-2 sample (1996–2023)

	Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-12769***	-11948***	-10545***	-11130***	-4193***	-1209***
	(1014.4)	(1109.8)	(1152.6)	(1302.2)	(827.8)	(185.4)
Tenure		1350***	1179***	1126***	1140***	586***
		(42.4)	(33.5)	(24.7)	(28.0)	(17.6)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	89877	89877	89876	89933	90648	91138
Effect (%)	-14.21	-13.29	-11.73	-12.38	-4.63	-1.33
Change in Adjusted R^2		0.106	0.166	0.041	0.463	0.141
Adjusted R^2	0.020	0.127	0.293	0.333	0.797	0.938
Observations	22,380,340	22,380,340	22,380,322	22,156,835	20,752,840	18,640,049

Panel B. This table is similar to Table 2, Panel B, except that we estimate pay gaps separately in HR-1 (1973–1995) and in HR-2 (1996–2023). *Log Pay (2023 USD)* is the natural logarithm of the employee's pay expressed in 2023 USD. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's years of federal service. Across columns, we progressively introduce interacted fixed effects for year, agency, county, occupation, and rank, allowing us to compare employees within increasingly narrow job cells. Standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. See Section 2.1.

Panel A: HR-1 sample (1973–1995)

	Log Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.334***	-0.317***	-0.280***	-0.284***	-0.099***	-0.012***
	(0.018)	(0.017)	(0.021)	(0.023)	(0.011)	(0.001)
Tenure		0.035***	0.033***	0.033***	0.029***	0.010***
		(0.003)	(0.003)	(0.003)	(0.002)	(0.001)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	11.17	11.17	11.17	11.17	11.17	11.17
Effect (%)	-28.41	-27.19	-24.42	-24.70	-9.45	-1.22
Change in Adjusted R^2		0.127	0.113	0.038	0.391	0.180
Adjusted R^2	0.122	0.249	0.362	0.400	0.792	0.972
Observations	14,094,381	14,094,381	14,094,371	13,969,912	13,003,688	11,324,182

Panel B: HR-2 sample (1996–2023)

	Log Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.126***	-0.116***	-0.095***	-0.100***	-0.032***	-0.007***
	(0.010)	(0.010)	(0.011)	(0.012)	(0.006)	(0.001)
Tenure		0.017***	0.015***	0.014***	0.013***	0.006***
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	11.29	11.29	11.29	11.29	11.30	11.30
Effect (%)	-11.84	-10.92	-9.09	-9.49	-3.19	-0.67
Change in Adjusted R^2		0.141	0.178	0.048	0.451	0.131
Adjusted R^2	0.017	0.158	0.336	0.384	0.835	0.965
Observations	22,380,340	22,380,340	22,380,322	22,156,835	20,752,840	18,640,049

Table A.2: **Gender pay gap excluding HR-1 1973 cohort**

This table is similar to Table 2, except that we exclude the cohort of 1973 from the HR-1 sample. *Pay (2023 USD)* is the employee's pay expressed in 2023 USD, and *Log Pay (2023 USD)* is its natural logarithm. *Female* equals 1 if the employee's first name has at least a 75% probability of belonging to a female, according to the Social Security frequency tables, and 0 if the probability is below 25%. *Tenure* is the employee's years of federal service. Across columns, we progressively introduce interacted fixed effects for year, agency, county, occupation, and rank, allowing us to compare employees within increasingly narrow job cells. Standard errors are clustered at the county and year level and reported in parentheses. *, **, and *** indicate significance at the 10%, 5%, and 1% levels, respectively. See Section 2.1.

Panel A: HR-1 sample (1973–1995) excluding HR-1 1973 cohort

	Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-13900***	-13641***	-12392***	-12868***	-4467***	-1125***
	(1007.0)	(1123.3)	(1231.8)	(1399.2)	(854.7)	(167.1)
Tenure		1391***	1238***	1188***	1191***	587***
		(45.7)	(38.9)	(31.6)	(30.7)	(16.8)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	84353	84353	84353	84432	85183	85598
Effect (%)	-16.48	-16.17	-14.69	-15.24	-5.24	-1.31
Change in Adjusted R^2		0.146	0.150	0.037	0.435	0.150
Adjusted R^2	0.026	0.172	0.322	0.359	0.794	0.944
Observations	30,254,063	30,254,063	30,254,036	29,922,412	27,796,863	24,752,201

Panel B: HR-2 sample (1996–2023) excluding HR-1 1973 cohort

	Log Pay (2023 USD)					
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.153***	-0.149***	-0.130***	-0.133***	-0.038***	-0.007***
	(0.011)	(0.012)	(0.013)	(0.015)	(0.007)	(0.001)
Tenure		0.018***	0.016***	0.015***	0.014***	0.006***
		(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Y FE	No	Yes	No	No	No	No
YxA FE	No	No	Yes	No	No	No
YxAxCt FE	No	No	No	Yes	No	No
YxAxCtxO FE	No	No	No	No	Yes	No
YxAxCtxOxR FE	No	No	No	No	No	Yes
Sample Mean	11.22	11.22	11.22	11.22	11.23	11.23
Effect (%)	-14.16	-13.86	-12.20	-12.45	-3.72	-0.66
Change in Adjusted R^2		0.182	0.153	0.043	0.432	0.134
Adjusted R^2	0.024	0.206	0.359	0.402	0.834	0.969
Observations	30,254,063	30,254,063	30,254,036	29,922,412	27,796,863	24,752,201

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