Sex and the city-size earnings premium*

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Abstract

We document that women's earnings are closer to parity with men's in larger cities in the United States. This female-specific city-size earnings premium is nearly 50% larger than men's city-size premium, is driven by non-college-educated women, and cannot be explained by occupational sorting, commuting, or differences in labor supply. We develop a quantitative spatial model with local frictional marriage markets to explain why large gender-specific earnings differentials can coexist with balanced gender ratios in spatial equilibrium: the marriage market acts as an endogenous amenity, generating gender-specific congestion elasticities. Counterfactual analysis suggests eliminating the female city-size premium would increase the aggregate gender wage gap by 7%.

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1 Introduction

Women's earnings are closer to parity with men's in larger cities. This paper's aim is to understand what drives this female-specific city-size earnings premium, where the 'quantity' response to these price differentials appears, and how these forces shape the aggregate gender earnings gap. To do so, we first empirically document a pronounced gender earnings premium in larger cities in the United States. Our baseline estimates imply that moving from a labor market the size of Abilene, Texas to one the size of New York City reduces the gender earnings gap by between 8 and 11 log points. Interestingly, while highly educated workers benefit the most from large cities on average, we find that it is women with lower levels of education who drive our results.

We then systematically assess different candidate explanations for the spatial pattern of gender earnings gaps we document. We empirically test whether tighter earnings gaps in larger cities are driven by sorting or selection on a host of observables, or whether this tighter gap reflects true differences in agglomeration economies for women than for men. Our findings suggest the female-specific city-size earnings premium is not explained by differences in occupational sorting, differences in human capital, intensive-margin labor supply, labor force participation rates, or heterogeneous preferences for commuting. Throughout all of our tests, a large, residual gender gap remains unexplained, and that residual gap shrinks with city size. We interpret our findings as suggesting that women with lower levels of education see disproportionate gains from agglomeration economies than men.

What does the existence of a steeper slope of women's earnings with respect to city size imply in spatial equilibrium? Theory since at least Roback (1982) would suggest either a quantity response – disproportionately *more* women in larger cities – or amenities (i.e., compensating differentials) that vary by gender across space. We find evidence that women disproportionately sort to larger cities with smaller gender gaps. However, gender ratios are fairly even across space, and the extent to which they do vary with respect to the earnings gap implies unrealistically small migration elasticities. Instead, we observe a quantity response in the marriage market: the share of singles is bigger in larger cities. Importantly, these quantity responses are strongest among the lower-educated women who benefit most from agglomeration economies.

To rationalize these findings, we develop a spatial equilibrium model that nests a parsimonious search-and-matching model of the marriage market. Our model microfounds sizable gender-specific congestion elasticities that ensure nearly equal gender ratios across space, and predicts 'quantity' responses to spatially varying gender-specific earnings premia are reflected in the share of unmarried persons across space.

We use our model to study how the female-specific city-size earnings premia shapes the aggregate gender earnings gap. We find that eliminating this earnings premia induces de-urbanization, moving both men and women out of the largest commuting zones to more midsize cities. In aggregate, this sorting raises the average gender wage gap by 7%, and modestly affects marriage markets, attenuating congestion in the largest cities but crowding the dating scene in small cities.

This paper contributes to three primary literatures. First, we add to a large literature on the wage benefits of agglomeration. The fact that wages rise with density is perhaps the most important fact in urban economics (Combes, Duranton, and Gobillon, 2011; Combes, Duranton, Gobillon, et al., 2012; Roca and Puga, 2016). However, the vast majority of this literature abstracts from the role of gender in the city-size earnings premium. Thus, we contribute to a much smaller set of work, in economics and other social sciences, that studies how gender affects both labor supply and earnings in cities versus rural areas or across the city-size distribution (Phimister, 2005; Bacolod, 2017; Nisic, 2017; Almeida et al., 2022; Xing et al., 2022; D'Costa, 2024; Kim and Waldorf, 2024; Galster and Osland, 2024; Buchholz, 2025). Like us, Elass et al. (2024) study how the city-size earnings premium varies for women and men.

An advantage of our setting is that we can systematically investigate potential mechanisms for gender-earnings gaps and ask which women may benefit or lose from large cities. For example, both Elass et al. (2024) and D'Costa (2024) hypothesize an important role for commuting and infrastructure in determining how gender earnings gaps vary across space. We systematically observe and control for commute times and methods for commuting across space and find commuting can explain little of the city-size premium for women, once we account for differences in intensive-margin labor supply across space. We also contribute to this primarily reduced-form, empirical literature by connecting our novel facts on earnings across space with facts related to the marriage market, and jointly rationalize these patterns with a simple quantitative theory that enables us to assess the importance of the city-size earnings premium in shaping the aggregate earnings gap.

Second, we contribute to a long and robust literature on the gender wage gap (Goldin, 2014; Blau and Kahn, 2017; Goldin, 2021). A significant set of work has looked at how commuting patterns and joint household decision making affect earnings for men and women (White, 1986; Gutierrez, 2018; Le Barbanchon et al., 2021; Velásquez, 2023; Liu and Su, 2024a; Borghorst et al., 2024; Ranošová, 2025). Jayachandran et al. (2025) study how gender norms affect spatial location choices when couples face trade-offs over which locations will benefit different spouses. We contribute to this work by examining spatial determinants of gender earnings gaps, investigating different mechanisms that can explain spatial variation in the gender earnings gap, and then using a structural model to quantify the effects on spatial sorting, marriage rates, and aggregate gender wage gaps.

Third, we contribute to a literature on gender-specific sorting across space. Edlund (2005) shows that women are more likely to live in cities than men and ties that to *men's* greater earning potential in cities. We document that marriage market tightness and gender ratios vary systematically with city size, and investigate how that is connected to both gender-specific earnings premia and decisions over marriage. In this aspect, we are most closely related to contemporaneous work by Pellegrini and Penrose (2025) and Fang et al. (2025). Pellegrini and Penrose (2025) study how assortative mating is affected by productivity and housing cost differences across cities, while Fang et al. (2025) discuss how structural change and social norms in China, and in particular highly-educated women's increased sorting to the service sector, affect marriage rates through spatial sorting. We highlight a distinct mechanism and show that it cannot explain the patterns documented in this paper. For example, we show that the spatial and earnings patterns we emphasize have been persistent for the last 60 years, and that women's earnings advantage in large cities is not explained by industry or occupation specific factors, and thus cannot be explained by structural change. In addition, we find the earnings and migration patterns we document are driven by less-educated, rather than more-educated women. Instead, we connect larger returns to women for working in larger cities to spatial sorting and marriage rates, and use it to better understand macroeconomic and spatial phenomena using a spatial search model.[1]

The rest of this paper is organized as follows: Section 2 discusses our data, while Section 3 documents facts about the gender earnings gap and marriage markets across U.S. cities. Section 4 presents a quantitative model aimed at rationalizing these facts, and Section 5 uses the quantified model to perform counterfactual

[1] A modest contribution of this paper is to develop a spatial model of marriage market search that embeds the attractive features of quantitative spatial models (Allen and Arkolakis, 2014; Redding and Rossi-Hansberg, 2017) for applied work with an empirically and computational tractable search model. Past work embedding search theory into quantitative spatial models has primarily focused on labor search (see, e.g., Mann, 2024; Lindenlaub et al., 2025), while other work has integrated labor and marriage search without explicitly considering space (Calvo et al., 2024).

analysis. Section 6 concludes.

2 Data

Our primary data source is the American Community Survey (ACS), an annual survey of U.S. households that has been fielded since 2000. We access ACS microdata through IPUMS. Our main sample is from 2005 to 2019; we begin in 2005 because several variables used in the analysis are not available until 2005. Our primary geographic unit of analysis is the 1990 commuting zone (CZ). Because the ACS identifies geography at the Public Use Microdata Area (PUMA) level for all respondents, we assign individuals to CZs using PUMA-to-CZ crosswalks from Autor and Dorn, 2013 and Autor, Dorn, and Hanson, 2019. We construct estimates of population for each CZ by aggregating the individual-level data using ACS person-level sample weights.

For our main analysis of earnings differences, we make a small number of sample restrictions. We limit the sample to prime-age workers between the ages of 25 and 64, exclude non-citizens, and drop individuals who are not in the labor force. In all specifications, we weight the estimates using the ACS person-level sample weights. We extend the sample to earlier periods using microdata from the decennial U.S. census, accessed through IPUMS. The availability of variables and geographic identifiers varies across census years. For most census years from 1950 to 2000, we use crosswalks from Autor and Dorn (2013) to assign individuals to 1990 CZs. For the 1940 Census, we use data from Eckert et al. (2020), while for the 1960 Census we use a crosswalk made available by Evan Rose.[2]

We construct several additional commuting-zone-level measures by aggregating external data sources. We compute historical CZ-level population by aggregating county-level population data from Van Leuven (2024). We also build measures of natural amenities across commuting zones using the Natural Amenities Scale from the U.S. Department of Agriculture. Finally, we compute CZ-level political affiliation by aggregating election returns at the county level from MIT Election Data and Science Lab (2018). We collect county-level data on community and religious institutions from the Social Capital Project.

[2]https://ekrose.github.io/resources/

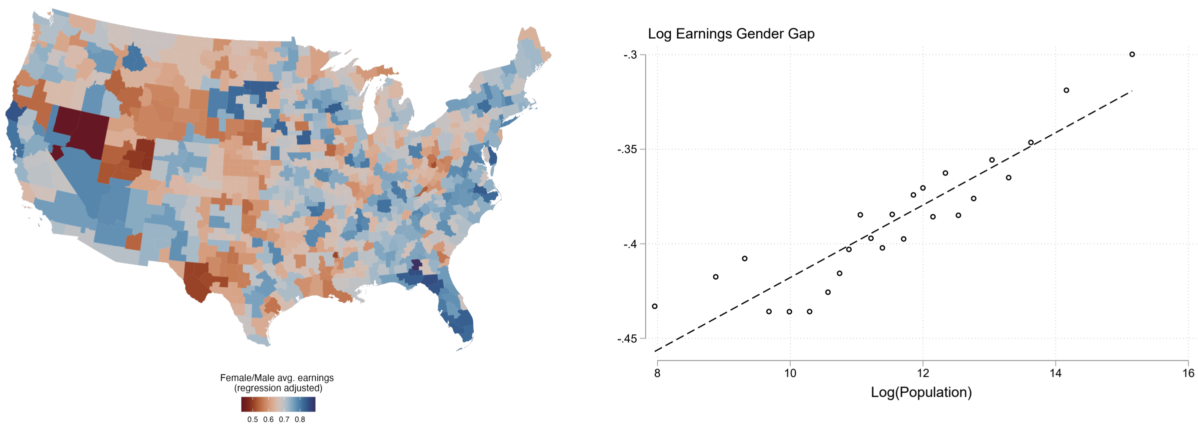


Figure 1: The Gender Earnings Gap Shrinks with City Size

3 Spatial differences in the gender earnings gap and marriage markets

3.1 The Gender Earnings Gap Shrinks with City Size

We first document that the gender earnings gap declines markedly with city size. Figure 1 visualizes this primary result by plotting both a commuting zone-level map of the earnings gap (left panel) and a binned scatter of the log gender earnings gap within commuting zones on the vertical axis against log commuting zone population on the horizontal axis. The binned scatter shows the raw gender earnings gap after residualizing earnings on year fixed effects and basic demographic controls (age, race, and ethnicity). We estimate a strong, positive coefficient, implying that the earnings gap between men and women narrows with commuting zone population.

Appendix Figure A1 presents alternative specifications. Panel (A) adds fixed effects for educational attainment to the baseline demographic controls, while Panel (B) adds additional fixed effects for the more than 600 occupations recorded in the ACS. Across all specifications, we find tightly estimated positive slopes, implying that gender earnings gaps systematically fall with city size.

Table 1 reports our headline results. We estimate an individual-level regression of the following form:

$$\ln w_i = a_0 + \alpha \ln L_i + \kappa \times 1(g = f) \times \ln L_i + \gamma 1(g = f) + u_i, \tag{1}$$

	Pooled			By Educ.: Less than BA		By Educ.: BA+	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Female	-0.676***	-0.803***	-0.570***	-0.839***	-0.640***	-0.536***	-0.399***
	(0.044)	(0.036)	(0.030)	(0.042)	(0.030)	(0.092)	(0.073)
log(CZ pop)	0.092***	0.054***	0.046***	0.047***	0.038***	0.087***	0.066***
	(0.008)	(0.005)	(0.005)	(0.004)	(0.004)	(0.010)	(0.007)
Female × log(CZ pop)	0.021***	0.027***	0.019***	0.031***	0.023***	0.008	0.009
	(0.003)	(0.003)	(0.002)	(0.003)	(0.002)	(0.007)	(0.005)
Abilene → NY Effect	0.086	0.111	0.081	0.130	0.097	0.035	0.036
Year FE	✓	✓	✓	✓	✓	✓	✓
Education FE		✓	✓				
Occupation FE			✓		✓		✓
R-sq.	0.114	0.217	0.354	0.100	0.265	0.125	0.317
Observations	13250889	13250889	13250889	8920338	8920338	4330551	4330551

Table 1: Gender Gap and City Size: Baseline Results

In these specifications, $\ln w_i$ is log annual wage and salary earnings, $\ln L_i$ is the log of population in an individual commuting zone, and $1(g = f)$ is an indicator for whether or not an individual is a woman. Our primary coefficients of interest are α which captures the standard agglomeration return, and κ which captures any heterogeneous return to city-size for women. In all specifications, we include a set of baseline demographic controls and fixed effects for age, race, and ethnicity. We cluster standard errors at the commuting zone level.

Column (1) includes only year fixed effects and basic demographic controls. We estimate an overall city-size earnings premium of 0.09, implying that increasing city size by 10% increases earnings by approximately 0.9%. The interaction term implies that city-size earnings premium is about 20% larger for women than for men. Our estimates imply that moving from a CZ the size of Abilene, Texas to one the size of New York City reduces the gender earnings gap by 9 log points.

The next two columns show the same coefficients, but for specifications that include education (Column 2) and both education and occupation fixed effects (Column 3). In these specifications, the city-size earnings premium falls by between 40-50%, implying that a significant fraction of the unconditional premium reflects educational and occupational sorting across cities. However, the additional gain of large labor markets for women remains large and substantial. With education fixed effects, the coefficient increases substantially from 0.021 to 0.027, implying that the earnings premium is 50% larger than the male premium. After adding occupation fixed effects in Column (3), the additional effect of large cities for women is nearly identical to that presented in Column (1) while the city-size premium for men continues to decline. As a result, the city-size premium for women ends up being about 40% larger than the male premium.

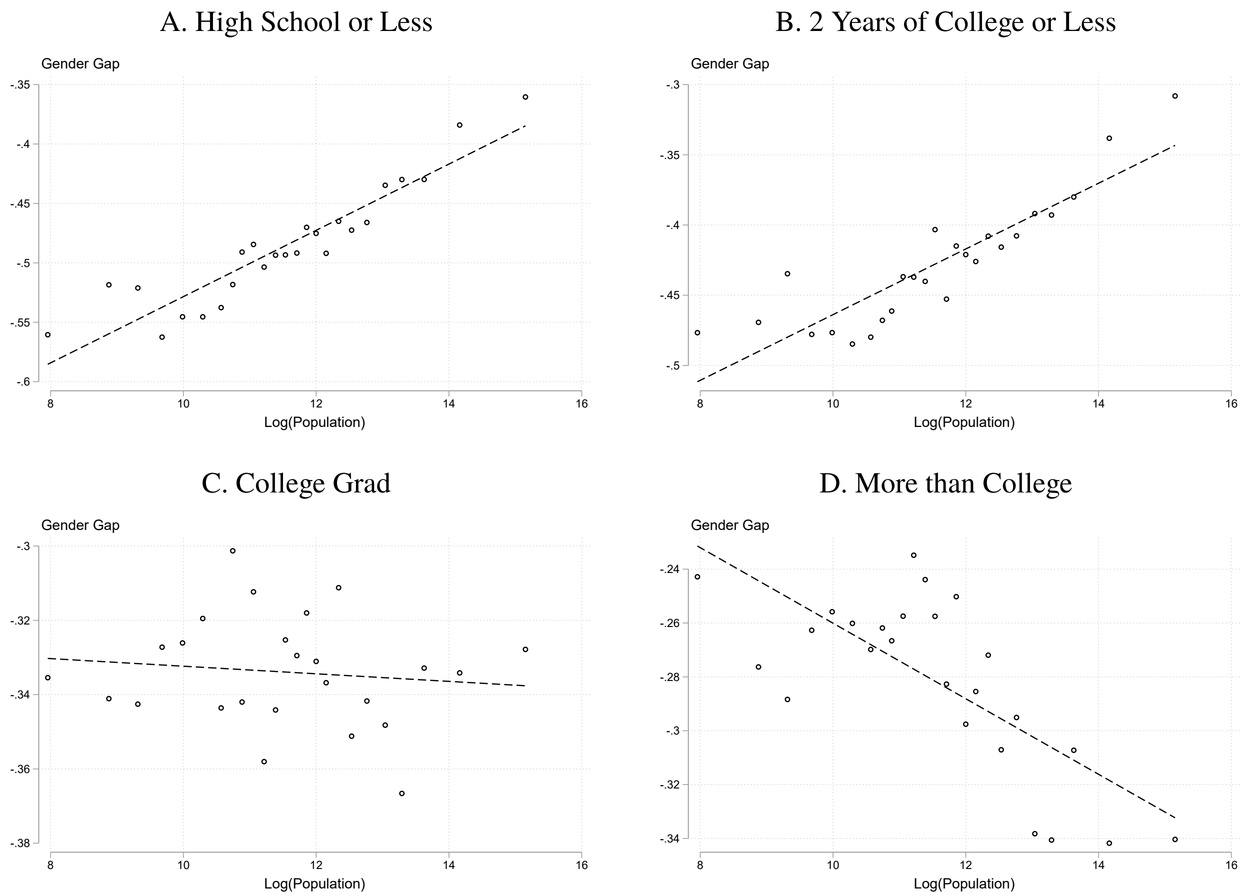
Across all specifications, we find that women consistently benefit more from large labor markets than men. Although a substantial share of the male city-size earnings premium can be explained by educational and occupational sorting across cities, the additional benefit for women is stable, or even increases. Overall, the city-size earnings premium is about 20 - 50% larger for women than men, implying that moving from a labor market the size of Abilene, Texas to New York City reduces the gender earnings gap by between 8 and 11 log points.

This gender earnings premium remains stable across the 15 years in our primary sample period. Appendix Figure A2 plots coefficients from year-by-year regressions, which show striking stability. We also use historical data to show that the gender-specific earnings premium in large cities has been a persistent feature of the U.S. economy throughout the post-war era (see Appendix Figure A3). Our findings contrast with D'Costa, 2024, who documents a larger urban wage premium for women in the U.K. before 2008 that largely disappeared afterward. The persistence of our results also provide evidence against stories where post-war structural transformation drives the patterns we document (Fang et al., 2025). Overall, the higher returns to large cities for women appear persistent and do not seem driven by cyclical conditions.

We find that the city-size earnings premium for women is driven primarily by women with lower levels of education. Figure 2 again plots binned scatter plots of gender earnings gaps across cities against log CZ population, separately by education group. The reduction in the gender earnings gap in larger cities is concentrated among individuals with a high school degree or less and those with some college. In contrast, among workers with a bachelor's degree, we find no relationship between city size and the earnings gap. For workers with more than a bachelor's degree, the pattern reverses: larger cities are associated with larger, not smaller earnings gaps.

Table 1, Columns (4) - (7) reports these results using individual-level regressions. Interestingly, we find that the city-size wage premium for men is substantially smaller for non-college workers than for college-educated workers. For non-college men, the earnings elasticity to city-size is between 0.04 and 0.05, whereas

Figure 2: Earnings Gaps and City Size by Educational Attainment



for college-educated men it is nearly twice as large: 0.07 - 0.09. However, the additional boost for women is much larger among non-college-educated workers: the female-interaction coefficients range from 0.02 - 0.03, while in the college-educated sample they are close to 0.01 and statistically insignificant. As a result, the city-size earnings premium is about 50-60% larger for non-college women than for men, whereas for college-educated women the premium is indistinguishable from that for men. Overall, although college-educated workers receive larger average returns from city size, the narrowing of the gender earnings gap in larger cities is driven by an additional earnings boost for non-college women. This finding is particularly notable given that much of the literature on the gender wage gap emphasizes mechanisms affecting college-educated women in high-skill or "superstar" occupations (Goldin, 2021).

Panel A: Split by Parental Status

	(1)	(2)	(3)
Female	-0.736***	-0.804***	-0.591***
	(0.040)	(0.033)	(0.024)
log(CZ pop)	0.089***	0.050***	0.042***
	(0.008)	(0.005)	(0.005)
Female × log(CZ pop)	0.037***	0.037***	0.029***
	(0.003)	(0.002)	(0.002)
Children x log(CZ pop)	0.008**	0.010***	0.009***
	(0.004)	(0.002)	(0.001)
Female x Children x log(CZ pop)	-0.034***	-0.023***	-0.019***
	(0.005)	(0.003)	(0.003)
Abilene → NY No Child	0.153	0.155	0.119
Abilene → NY Children	0.043	0.099	0.077
Year FE	✓	✓	✓
Education FE		✓	✓
Occupation FE			✓
R-sq.	0.124	0.225	0.358
Observations	13250889	13250889	13250889

Panel B: Split by Marital Status

	(1)	(2)	(3)
Female	-0.554***	-0.683***	-0.457***
	(0.045)	(0.039)	(0.024)
log(CZ pop)	0.096***	0.057***	0.047***
	(0.009)	(0.006)	(0.006)
Female × log(CZ pop)	0.028***	0.033***	0.023***
	(0.003)	(0.003)	(0.002)
Married x log(CZ pop)	0.003	0.005*	0.005**
	(0.005)	(0.003)	(0.002)
Female x Married x log(CZ pop)	-0.024***	-0.022***	-0.014***
	(0.006)	(0.004)	(0.003)
Abilene → NY Singles	0.118	0.138	0.095
Abilene → NY Married	0.029	0.066	0.059
Year FE	✓	✓	✓
Education FE		✓	✓
Occupation FE			✓
R-sq.	0.138	0.233	0.362
Observations	13250889	13250889	13250889

Table 2: Heterogeneous Returns to Large Cities by Marital and Parental Status

3.2 Mechanisms

We leverage the rich detail in the survey data to investigate the mechanisms underlying the disproportionately larger city-size earnings premium for women. We first investigate whether the differences in the gender earnings gap across space are related to interactions of city size with marriage and parenthood. A substantial literature has focused on the emergence of gender-specific earnings losses when women become mothers (Kleven et al., 2019; Goldin, 2021). If larger cities reduce the "child penalty" or make it easier for both spouses in a marriage to find well-paid work (Jayachandran et al., 2025), this might explain our findings.

Table 2 tests this hypothesis by estimating the city-size earnings premium, and how it changes with marital and parental status. We find that the largest gains from larger cities come for single women and women without children. Panel (A) shows heterogeneity in the premium by parental status. We find that most of the gender-specific city-size premium is eliminated for women with children. As a result, the city-size earnings premium for women is substantially larger for childless women than for women with children. Interestingly, we still find significant earnings gains for women with children from moving to large cities. However, most of this comes from the fact that there is a large earnings premium for parents from large cities, regardless of gender. This, however, is much smaller than the earnings gain for single women in large cities.

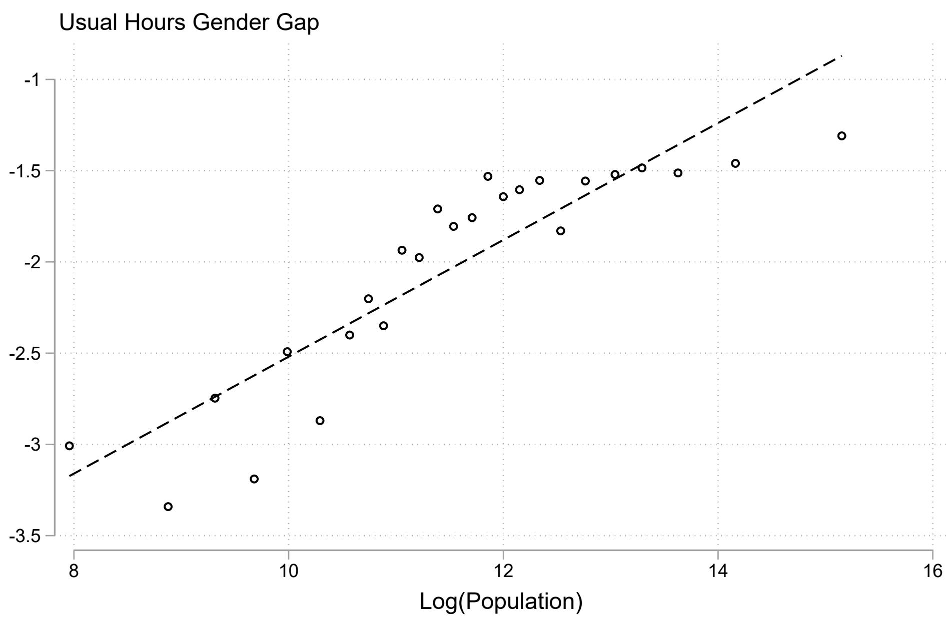
Panel (B) shows the effect for women by marital status. We find similarly that the effect is driven by single, rather than married women. The earnings-specific city-size premium is substantially reduced for married women. In addition, while there is city-size premium for married individuals of all genders, it is significantly smaller than the child city-size premium. Taken together, we find that our results cannot be explained by marital or child penalties varying across space. Instead, we find that the city-size earnings premium for women is driven by single women and women without children. As with our finding that the city-size premium is driven by women with lower levels of education, this result is particularly notable given the significant literature that has studied the role of childcare, joint household decision making, and work-family balance in driving the gender earnings gap (Goldin, 2021). In contrast, we study an important factor in driving earnings gaps between men and women that is disproportionately focused on lower-educated women and childless, single women.

We next examine whether intensive-margin labor supply varies systematically across space in a way that could explain the spatial earnings gap. Figure 3 documents a clear spatial pattern in the gender gap in hours worked: in larger cities, women's usual weekly hours converge toward those of men. This result is robust across all baseline sets of controls and is economically meaningful. Our estimates indicate that moving from a small city such as Abilene, Texas to a large city like New York City is associated with a reduction of approximately 1.3 hours per week in the gender gap in usual hours worked.

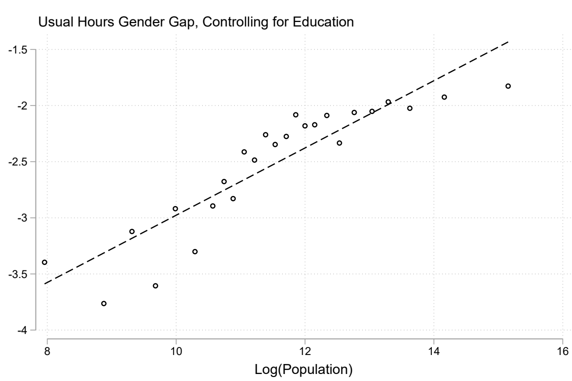
To assess whether intensive-margin labor supply can explain the overall earnings gap, we augment the baseline regression with a rich set of labor-supply controls. Specifically, we include polynomials in usual hours worked and incorporate fixed effects for the number of weeks worked by each individual. In addition, we allow these labor-supply measures to have gender-specific effects by interacting them with indicators for whether the worker is a woman. These results are shown in Table 3.

We find that although spatial differences in intensive-margin labor supply account for part of the gender earnings premium we document, a substantial share remains unexplained. In all specifications, the estimated coefficients attenuate relative to the baseline results, yet meaningful differences in men's and women's earnings across space persist. For example, in the specification without education or occupation controls, the decline in the gender earnings gap from Abilene to New York City is 5 log points, compared with 8 log points in the baseline. Among workers without a B.A., the corresponding decline is 8 log points, relative to 13 log points in the baseline, even after conditioning on intensive-margin labor supply. Notably, the co-

A. Raw Gap



B. Controlling for Education



C. Controlling for Educ. and Occ.

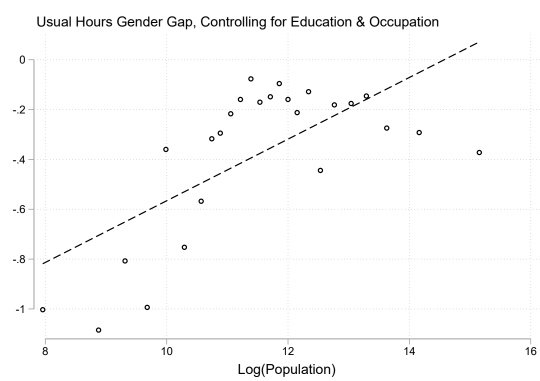


Figure 3: Gender Gap in Usual Hours Worked by City Size

	(1)	Pooled (2)	(3)	By Educ.: Less than BA (4)	(5)	By Educ.: BA+ (6)	(7)
log(CZ pop)	0.091***	0.061***	0.052***	0.055***	0.045***	0.088***	0.070***
	(0.006)	(0.004)	(0.004)	(0.003)	(0.004)	(0.007)	(0.006)
Female × log(CZ pop)	0.011***	0.017***	0.016***	0.020***	0.018***	0.003	0.007*
	(0.002)	(0.002)	(0.002)	(0.002)	(0.002)	(0.004)	(0.004)
Abilene → NY Effect	0.048	0.070	0.065	0.082	0.075	0.014	0.030
L. Supply Controls	✓	✓	✓	✓	✓	✓	✓
Year FE	✓	✓	✓	✓	✓	✓	✓
Education FE		✓	✓				
Occ. FE			✓		✓		✓
R-sq.	0.518	0.584	0.644	0.551	0.618	0.485	0.585
Observations	13250889	13250889	13250889	8920338	8920338	4330551	4330551

Table 3: Gender Gap and City Size: Labor Supply Controls

efficients attenuate more in specifications that omit occupation fixed effects than in those that include them. This pattern suggests that part of the city-size earnings premium reflects women's sorting into higher-hours occupations in larger cities. Nevertheless, a sizable portion of the city-size premium remains unexplained, even after accounting for spatial differences in labor supply.

An alternative explanation for spatial differences in the gender earnings gap is that commuting infrastructure varies across cities, potentially enabling women to access higher-paying jobs (Gutierrez, 2018; Le Barbanchon et al., 2021; Velásquez, 2023; Caldwell and Danieli, 2024; Liu and Su, 2024b; Ranošová, 2025). Although a primary channel through which commuting infrastructure may affect gender gaps is by altering the quantity of labor supplied, it could also influence the quality of jobs available to women conditional on total hours worked. Motivated by this possibility, we re-estimate our baseline specification while augmenting it with both detailed labor-supply controls and rich measures of individual commuting behavior. Specifically, we include polynomials in total transit time to work, along with fixed effects for transportation mode. As with the labor-supply controls, we allow these commuting-related effects to differ between women and men.

Table 4 shows that commuting explains relatively little of the earnings premium we document once we condition on intensive-margin labor supply. The coefficients exhibit only modest attenuation relative to the specification that includes labor-supply controls alone, and the estimates remain quite similar. These results suggest that although improved commuting infrastructure may enable women to work more hours in larger cities, it does not account for the remaining gender earnings gap across cities once hours worked are

		Pooled		By Educ.: Less than BA		By Educ.: BA+	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
log(CZ pop)	0.084***	0.056***	0.048***	0.052***	0.043***	0.079***	0.064***
	(0.005)	(0.004)	(0.004)	(0.004)	(0.004)	(0.006)	(0.005)
Female × log(CZ pop)	0.009***	0.014***	0.014***	0.018***	0.017***	0.002	0.007*
	(0.002)	(0.002)	(0.001)	(0.002)	(0.002)	(0.005)	(0.004)
Abilene → NY Effect	0.038	0.060	0.060	0.073	0.071	0.007	0.027
L. Supply Controls	✓	✓	✓	✓	✓	✓	✓
Commute Controls	✓	✓	✓	✓	✓	✓	✓
Year FE	✓	✓	✓	✓	✓	✓	✓
Education FE		✓	✓				
Occ. FE			✓		✓		✓
R-sq.	0.523	0.587	0.645	0.556	0.620	0.489	0.586
Observations	13250889	13250889	13250889	8920338	8920338	4330551	4330551

Table 4: Gender Gap and City Size: Commuting and Labor Supply Controls

held constant. We find a substantial gender earnings gap remains after conditioning on commuting behavior and intensive-margin labor supply. This is especially evident when examining the gender earnings premium within occupation or within educational groups, where the residual spatial gradient remains sizable.

Finally, we investigate whether differences in extensive-margin labor supply could provide an explanation for our results. Farré et al. (2022) document that in cities with longer travel times, married women are less likely to participate in the labor force. This could lead to selection concerns, where the women with the lowest returns in the labor market disproportionately drop out of the labor force in large cities.

Consistent with this interpretation, Table 5 shows that women are less likely to participate in the labor market in large cities. However, we argue this is unlikely to explain our results because it does not explain the patterns of heterogeneity we see in the data. Table 5 shows we find similar effects for both workers with a B.A. and workers with less than a B.A. For both groups of workers, we find that moving from Abilene, TX to New York reduces the extensive-margin labor supply by about 1 p.p. In addition, as noted in Farré et al. (2022), this channel is likely to be especially large for married women, whereas we found particularly large city-size earnings premiums for unmarried women.

Table 6 shows this is true in our data as well. We find that married women are much more likely to drop out of the labor force in large cities relative to small cities. This is driven by an overall fall in labor-force participation in large cities for married individuals regardless of gender and an especially large drop for

	Pooled		By Educ.: Less than BA	By Educ.: BA+
	(1)	(2)	(3)	(4)
log(CZ pop)	0.016***	0.008***	0.013***	0.004***
	(0.002)	(0.001)	(0.002)	(0.001)
Female $\times$ log(CZ pop)	-0.004***	-0.002**	-0.003**	-0.002*
	(0.001)	(0.001)	(0.002)	(0.001)
Abilene $\rightarrow$ NY effect	-0.016	-0.009	-0.014	-0.009
Year FE	✓	✓	✓	✓
Education FE		✓		
R-sq.	0.064	0.103	0.061	0.068
Observations	19229398	19229398	13827924	5401474

Table 5: Gender Gap and City Size: Effect on Labor Force Participation

married women with a B.A. or above. As a result, married women in large cities are less likely to be in the labor force, relative to unmarried women. In contrast, the wage premiums we document are most pronounced in unmarried women. This is the opposite pattern we would expect if selection was driving our results. Overall, we find little evidence that selection due to differences in extensive-margin labor supply is a strong candidate explanation for our results.

We present a number of additional mechanism tests in Appendix B. We estimate more restrictive specifications that allow for more flexible sorting of industries, occupations, and educational groups across space and find that this does little to explain our results. Second, we show that the larger-city advantage for single women relative to married women, and for childless women relative to mothers, is driven primarily by intensive-margin labor supply and commuting responses that differ between large and small cities. Finally, we find that differences in political attitudes across space can also explain some, but not the majority, of the gender-specific city-size earnings premium, suggesting that policy differences across space or norms regarding women in the workforce might explain a portion of the overall premium, while leaving the majority unexplained.

In Appendix D, we investigate the extent to which our results are heterogeneous across racial groups. We find evidence that Black women benefit disproportionately from large labor markets. In our baseline specifications, the gender-specific earnings premium is about 1.6 to 2 times larger for Black women, relative to non-Black women. This effect is driven by improved sorting of Black women to higher earning occupations in larger cities. In contrast, we find much less evidence that Hispanic women benefit disproportionately from

	Pooled		Less than BA	BA+
	(1)	(2)	(3)	(4)
Female	0.103***	0.066***	0.094***	0.034***
	(0.012)	(0.011)	(0.014)	(0.008)
log(CZ pop)	0.026***	0.017***	0.022***	0.009***
	(0.003)	(0.003)	(0.003)	(0.002)
Female × log(CZ pop)	-0.006***	-0.004***	-0.006***	-0.002***
	(0.001)	(0.001)	(0.001)	(0.001)
Married x log(CZ pop)	-0.016***	-0.014***	-0.013***	-0.004***
	(0.002)	(0.002)	(0.002)	(0.001)
Female x Married x log(CZ pop)	-0.000	-0.000	-0.001	-0.006***
	(0.002)	(0.002)	(0.002)	(0.001)
Abilene → NY Single	-0.025	-0.018	-0.025	-0.009
Abilene → NY Married	-0.092	-0.078	-0.081	-0.055
Year FE	✓	✓	✓	✓
Education FE		✓		
R-sq.	0.082	0.117	0.080	0.081
Observations	19229398	19229398	13827924	5401474

Table 6: Heterogeneous Effects of City Size on Extensive Margin Labor Supply

large cities. In fact, Hispanic women see smaller earnings gains from moving to large cities, driven by a large gender-neutral earnings penalty for Hispanic workers in large cities. In our quantitative model, we abstract from racial heterogeneity in earnings and sorting across cities. Expanding our model to understand these earnings and sorting responses, as well as better understanding the drivers of spatial earnings premia (and penalties), is an important direction for future work.

3.3 Causal Effects Using Instrumental Variables

While we show that our results cannot be explained by standard explanations for the gender earnings gap including parenthood, occupational sorting, or labor supply, an important concern is that our results might reflect historical patterns like the rise of the service-sector which both led service-specialized cities to grow and disproportionately increased wages for women (Fang et al., 2025). In addition, in the theoretical model we lay out in Section 4, OLS regressions do not identify our model parameters due to endogenous city sizes. To allay these identification concerns and construct unbiased estimates of our parameters, we use IV regressions to instrument for city size. We construct two instruments for city size to ensure our findings reflect

	(1)	Pooled (2)	(3)	By Educ.: Less than BA (4)	(5)	By Educ.: BA+ (6)	(7)
log(CZ pop)	0.099***	0.062***	0.059***	0.051***	0.050***	0.097***	0.079***
	(0.009)	(0.007)	(0.007)	(0.006)	(0.006)	(0.009)	(0.008)
Female × log(CZ pop)	0.015**	0.019***	0.011**	0.021***	0.012***	0.007	0.007
	(0.007)	(0.006)	(0.005)	(0.004)	(0.004)	(0.011)	(0.008)
Female	-0.602***	-0.688***	-0.454***	-0.697***	-0.476***	-0.521***	-0.379***
	(0.093)	(0.078)	(0.063)	(0.059)	(0.049)	(0.149)	(0.114)
Abilene $\rightarrow$ NY Effect	0.064	0.077	0.047	0.087	0.049	0.031	0.030
Year FE	✓	✓	✓	✓	✓	✓	✓
Education FE		✓	✓				
Occupation FE			✓		✓		✓
K-P F Stat	42.8	44.1	49.3	35.7	35.6	53.0	53.5
J. Stat	0.747	1.219	2.389	2.066	3.200	0.348	1.257
R-sq.	0.107	0.105	0.058	0.094	0.051	0.119	0.074
Observations	13081903	13081903	13081903	8796286	8796286	4285617	4285617

Table 7: Gender Gap and City Size: Combined IV Results

causal effects. One takes advantage of the persistence of city size using population across commuting zones almost a century before our main sample period in 1920. Historical population is a strong predictor of contemporary city size due to persistence in local housing stock, infrastructure, and market access. The exclusion restriction is that, conditional on rich modern controls, including education, occupation, demographics, and year fixed effects, population levels in 1920 affect earnings today only through their effect on contemporary city size. The second is a measure of city-level amenities based on total water area in a commuting zone. For this instrument, the exclusion restriction is that, conditional on our detailed controls, water features are an amenity, rather than a fundamental driver of productivity.

Table 7 shows our IV estimates using both instruments combined. We find results that are slightly attenuated relative to our OLS results, but still large and statistically significant. As with the baseline OLS findings, we find a significant city-size gender earnings gap, that is driven by workers with less than a B.A. The fact that our coefficients are slightly smaller than the OLS suggests that the bias from endogenous city populations is minimal, but does go in the direction of biasing our coefficients upwards. Given that our specification is overidentified, we also report standard J-statistics. We cannot reject at standard significance levels that our instruments are all valid. Overall, our IV estimates provide a high degree of confidence that there is a causal effect of city size on the gender-earnings premium.

3.4 Quantity Responses

We have documented sizable gender earnings gaps across space. These gaps imply that women and men face different incentives both to sort geographically and to select into marriage across locations. Edlund (2005), for example, shows that women are disproportionately concentrated in urban rather than rural areas worldwide, attributing this pattern to men's earnings potential rather than women's. Figure 4 shows very limited migration responses to these pronounced differences in gender earnings gaps. We plot a log measure of what we call marriage market tightness, which is the ratio of women to men, on log differences in the gender earnings gap across CZs. We find that while the slope is positive, implying women disproportionately sort to areas with smaller gender earnings gaps, the estimated elasticity of 0.1 is quite modest. Through the lens of the model presented in Section 4, this regression identifies a migration elasticity of 0.1, which is an order of magnitude smaller than even the most conservative estimates.[3]

Our quite modest elasticities, however, can be partially explained by the fact that most married individuals are in different-gender partnerships and reside in the same location as their spouse. For this reason, we also look more specifically at the location and sorting decisions of single individuals, where choices are less mechanically linked across genders.

Though modest in size, we find evidence of spatial sorting of women into larger cities in a manner consistent with their higher relative earnings potential. Figure 5 presents binned scatter plots of the share of single individuals who are women against commuting-zone population, separately by educational group. Across all groups, women constitute a larger share of the single population in larger cities, consistent with women having better relative earnings opportunities in those areas. Importantly, this relationship is strongest among women with lower levels of education, who have the largest earnings gains from large labor markets. For this population, we estimate a tight, approximately linear relationship between city size and the female share of singles.

Among individuals with college degrees or more, we continue to find a positive association, though the magnitude is roughly half that observed for those without any college education. In addition, the college-educated group exhibits evidence of a non-monotonic pattern: the female share initially rises with city size

[3] Suárez Serrato and Zidar (2016) estimates an interstate migration elasticity of around $1/0.82 \approx 1.2$. Bryan and Morten (2019) estimate migration elasticities of 2.7 for the US and 3.3 for Indonesia.

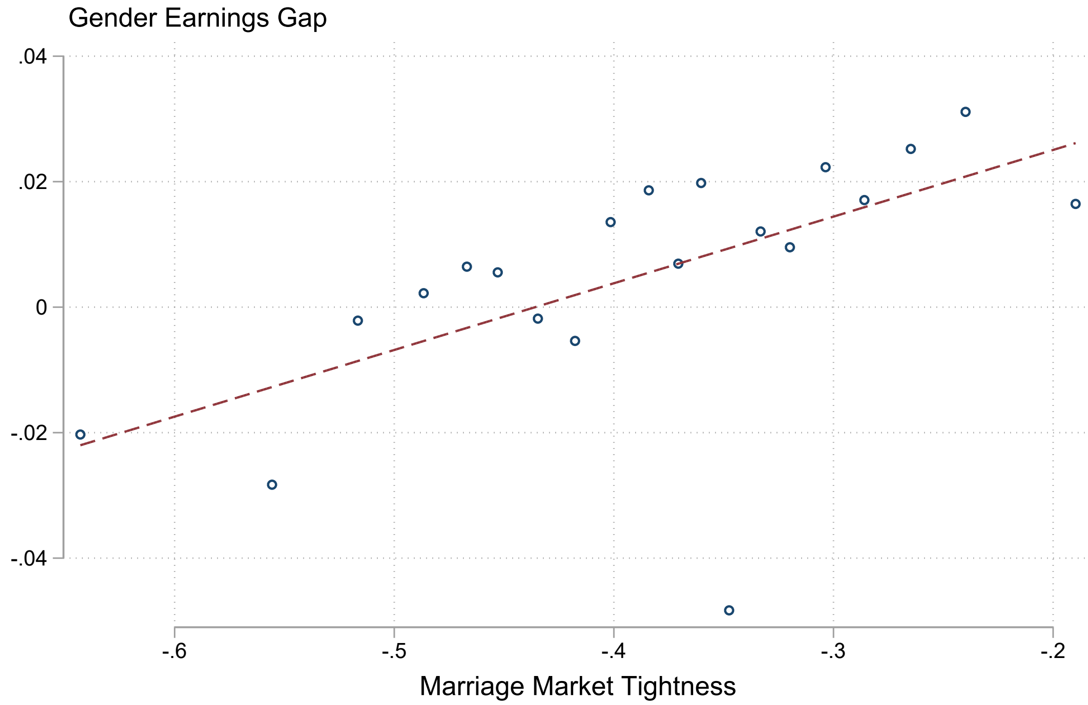


Figure 4: Overall Quantity Responses to Earnings Gaps. The gender earnings gap is measured as the (residualized) female-male earnings ratio. 'Marriage market tightness' is the prime-age female population to male population ratio at the CZ level.

and then declines for the very largest labor markets. Taken together, these results point to strong quantity responses to spatial earnings differentials, with the sharpest responses appearing in groups for whom women stand to gain the most from sorting into larger cities.

Finally, the earnings growth associated with large cities may also affect individuals' marriage and divorce decisions. To study this, we estimate how hazard rates into marriage vary with city size. We estimate the probability that an individual became married in the past year in our sample of prime-age individuals. To ensure our results are not affected by mobility patterns in the year of marriage, we restrict our estimation to a sample of individuals who resided in the same local area in both years.[4] Finally, to ensure our results are not driven by incentives to marry coming from different costs of living across space, we control for median rents for prime-age workers in each commuting zone (Pellegrini and Penrose, 2025). We multiply our dependent variable, marriage in the last year, by 100 in order to keep the coefficients interpretable.

[4]We specifically restrict to individuals who did not move or moved within the same "Migration PUMA" within the last year.

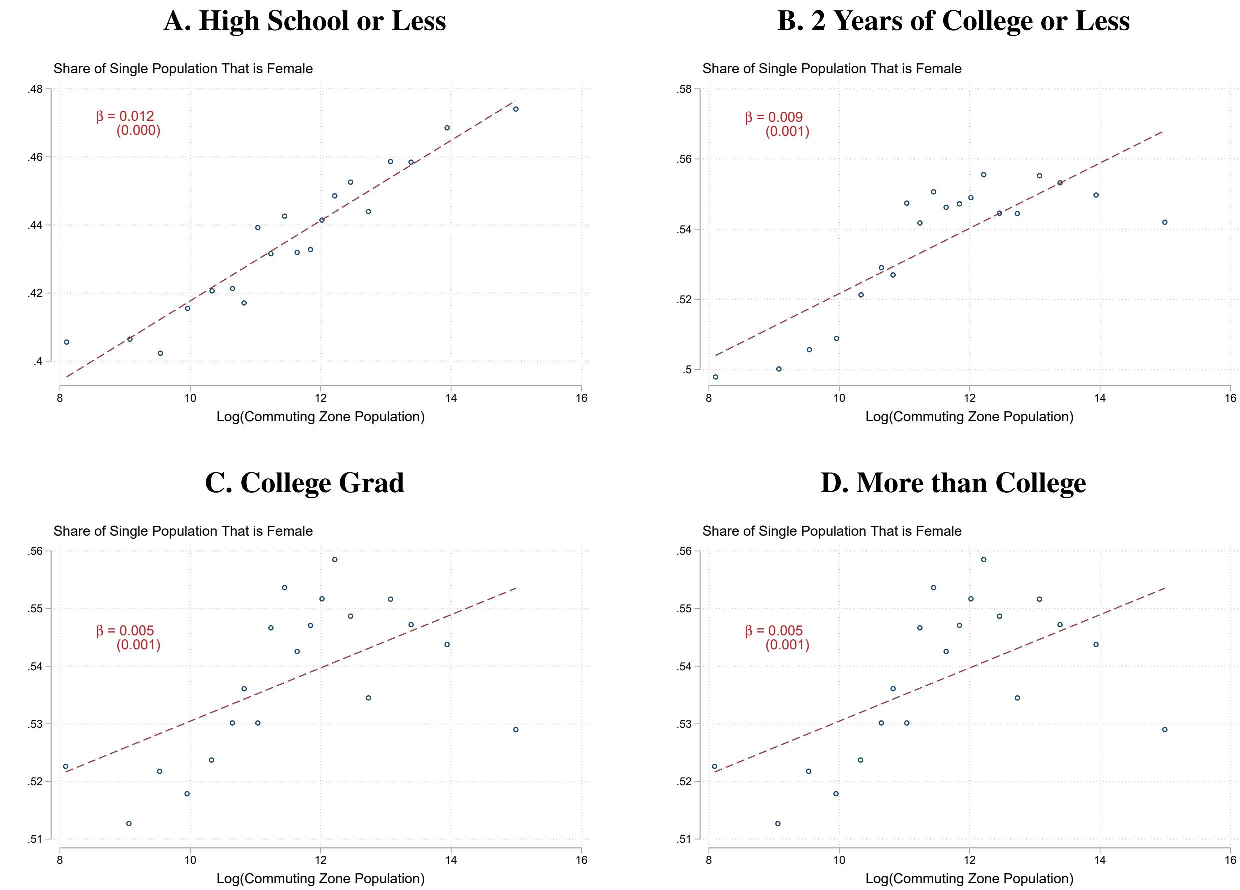


Figure 5: Spatial Sorting of Single Individuals

We find that marriage rates systematically decline with city size. Columns (1) and (2) report our baseline estimates in the pooled sample. In both the specification with demographic controls and the specification that adds education controls, we estimate negative and statistically significant coefficients. Consistent with our earlier findings that earnings and sorting responses are concentrated among workers with lower levels of education, the decline in marriage rates is largest among individuals with limited schooling. Among individuals with higher levels of education, the estimated effects are roughly half as large and only weakly statistically significant.

Taken together, these results reveal three central patterns. First, larger cities offer higher earnings premia for women. Second, single women disproportionately sort into these larger labor markets. Third, the rate at which individuals enter marriage is lower in larger cities. Across all three findings, the strongest effects occur among women without a bachelor's degree, reinforcing the interpretation that these patterns are linked by the relatively smaller gender earnings gap in large cities for this group.

	Pooled		Less than BA	BA+
	(1)	(2)	(3)	(4)
log(CZ pop)	-0.031**	-0.039***	-0.046***	-0.026*
	(0.013)	(0.013)	(0.016)	(0.016)
Abilene $\to$ NY Effect	-0.001	-0.002	-0.002	-0.001
Year FE	✓	✓	✓	✓
Education FE		✓		
R-sq.	0.012	0.012	0.009	0.018
Observations	15200836	15200836	10922357	4278479

Table 8: Marriage Hazards and City Size

4 Model

We now present a spatial model of location choice in the presence of a labor market with earnings gaps and a marriage market that clears through search-and-matching. The goal of the model is to jointly rationalize how sizable earnings gaps can coexist in equilibrium with a small differential migration response, and use the model to quantitatively assess the importance of the female-specific city-size earnings premium in shaping overall earnings equity. All details are worked out in Appendix A.

Environment There are $i = 1, ..., I$ cities. Agents have a gender $g \in \{m, f\}$, and the total mass of agents of each gender is L^g and the total population is $\bar{L}$. Agents make a location decision, and work and consume in that location, and search in the marriage market. The marriage market is frictional and allocations are determined by a search and matching process. We model the search in continuous time and only consider a stationary equilibrium.

Household preferences and search Agents, indexed by ω^g, choose their location i to maximize their amenity-adjusted value of searching in the marriage market, S_i^g; that is, they solve,

$$\max_i \bar{A}_i^g S_i^g \varepsilon_i(\omega^g), \tag{2}$$

where $\varepsilon_i(\omega^g)$ is an idiosyncratic preference shock that is iid Fréchet with shape parameter ϵ, and $\bar{A}_i$ is an exogenous amenity.

The value of search is given by agents' Hamilton-Jacobi-Bellman equations,

$$\rho S_i^g = \max_{C,H} U(C,H) + \lambda^g \mathbb{E}_x \left[\max\{\phi^g J_i(x) - S_i^g, 0\}\right], \quad \text{subject to} \quad C + r_i H \leq w_i^g \tag{3}$$

where λ^g is the match hazard rate for agents of gender g, and $J(x)$ is the stochastic joint surplus of forming a match, which is split with weights ϕ^g. We let $U(C,H) = \left(\frac{C}{1-\beta}\right)^{1-\beta}\left(\frac{H}{\beta}\right)^{\beta}$ so that money-metric flow utility can be written as $u_i^g = \frac{w_i^g}{r_i^\beta}$, and total housing demand is,

$$H_i = \sum_g \frac{\beta w_i^g}{r_i} \tag{4}$$

Goods technology Each city has a representative firm that produces a homogeneous consumption good that is freely traded and numeraire with technology

$$Y_i = Z_i\left(L_i^m + L_i^f\right).$$

The gap between remuneration and marginal productivity is given by a wedge Δ_i that only applies to women, so that wages are,

$$w_i^m = Z_i, \quad w_i^f = \Delta_i Z_i, \tag{5}$$

where we write the gender wage gap as,

$$\Delta_i = \bar{\Delta}_i (L_i/\bar{L})^{\kappa}$$

where $\bar{L} = \sum_i L_i$, and $L_i = L_i^f + L_i^m$. Therefore, when a city is maximally crowded, the markdown is $\bar{\Delta}_i$.

We allow for agglomeration economies,

$$Z_i = \bar{Z}_i (L_i/\bar{L})^{\alpha}. \tag{6}$$

Housing technology Housing is supplied competitively but is produced with a decreasing-returns technology in land and the consumption good so that the resulting housing supply curve is upward sloping,

$$H_i = \bar{H}_i (r_i)^{\gamma}. \tag{7}$$

4.1 The marriage market

We now describe the marriage market, and define a marriage market *partial* equilibrium (in which θ and u^g are exogenous). For convenience, we drop location subscripts, and denote $w = w^m$. Time is discounted at rate ρ. Both men and women search for a partner. Whether to match with a partner is a rational decision that weighs the the surplus of forming a match, which depends on a stochastic, idiosyncratic match value, against the continuation value of search. Match probabilities are endogenous to marriage market tightness, $\theta = L^f/L^m$, and matches dissolve at rate δ.

Marriage market technology We write the 'reduced-form' household flow utility function as, $B(x) = x(u^m)^{1-\eta}(u^f)^{\eta}$, where x is a match value.[5] We define $\mathcal{M}(L^m, L^f) = \bar{M}(L^m)^{\mu}(L^f)^{1-\mu}$ as the matching function so that matching probabilities are $\lambda^f = \bar{M}\theta^{-\mu}, \lambda^m = \bar{M}\theta^{1-\mu}$. Idiosyncratic match values x are common to both agents in the (potential) match and we assume each x is drawn independently from a Pareto distribution with location parameter $\bar{x}$ and tail parameter $\zeta > 1$.

Value functions and search When an individual is married, their the idiosyncratic match value is held fixed ($\dot{x} = 0$), so that their value function is,

$$\rho V^g(x,t) = \phi^g B(x) + \delta \left(S^g - V^g(x,t)\right) + \partial_t V^g(x,t),$$

where S^g is the value of search and $\phi^g \in (0,1)$ is an endogenous division of surplus. However, as the surplus is split so $\phi^m + \phi^f = 1$, it is easier to work with the joint value $J(x,t) = V^m(x,t) + V^f(x,t)$,

$$\rho J(x,t) = B(x) + \delta \left(\bar{S} - J(x,t)\right) + \partial_t J(x,t),$$

[5] We view this as a reduced form representation of the household surplus function. In Becker (1973), the surplus function stems from couples solving a time allocation problem where the opportunity cost of their inputs are their wages. Here, we view x as a match productivity shifter, and η as the output elasticity of the gender wage gap in the marriage production function. When women dedicate labor time to home production ($\eta = 0$) their nominal wages do not influence household production.

where $\bar{S} = \sum_g S^g$ is the joint value of search. At a stationary equilibrium ($\partial_t J(x,t) = 0$),

$$J(x) = \frac{B(x)}{\rho + \delta} + \frac{\delta}{\rho + \delta}(\bar{S}) \tag{8}$$

When an agent is searching for a partner, their (stationary) value function is:

$$\rho S^g = u^g + \lambda^g \mathbb{E}_x \left[\max\{\phi^g J(x) - S^g, 0\}\right]. \tag{9}$$

Marriage market partial equilibrium Given $\theta, \{u^g\}$, a stationary equilibrium in the marriage market is characterized by a cutoff match value x^*, a mass of singles for each gender s^g, and a division of surplus $\{\phi_g\}$ that imposes that the cutoff match value generates equal indifference at the margin for both agents,

$$\phi^m J(x^*) = S^m \quad \text{and} \quad \phi^f J(x^*) = S^f. \tag{10}$$

The surplus is split in full $\phi^m + \phi^f = 1,\ \phi^g \in (0,1)$, such that agents solve Equation (9), $\bar{S} = S^m + S^f$ and Equation (8) holds.

Solution First, we solve for x^*. Summing Equations (10),

$$J(x^*) = \bar{S}, \quad \text{and} \quad \frac{S^f}{\phi^f} = \frac{S^m}{\phi^m}. \tag{11}$$

Therefore,

$$\phi^g = \frac{S^g}{S^m + S^f}. \tag{12}$$

and plugging in (8) into equilibrium condition (10), and using the definition of $B(x)$ we have that,

$$x^* = \frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta} \tag{13}$$

Conditional on receiving a draw, the probability of forming a match is $p^* = (\bar{x}/x^*)^\zeta$, and so the probability of matching is $\lambda^g p^*$. The mass of singles of either gender is given by finding the value of s^g that makes $\dot{s}^g$ stationary:

$$\dot{s}^g = \delta(1 - s^g) - \lambda^g p^* s^g \implies s^g = \frac{\delta}{\delta + \lambda^g p^*}. \tag{14}$$

Now, we solve agents' optimal search problems (9). To do so, we use the properties of the Pareto distribution, which give a closed form for, $\mathbb{E}_x[B(x) \mid x > x^*] = \frac{\zeta}{\zeta-1} x^* \frac{w}{r^\beta}(\Delta)^\eta$. Plugging in the solution for x^* (Equation 13) and the solution for ϕ^g (Equation 12), we have that,

$$\rho S^g = \left(1 - \lambda^g \left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}}\right)^\zeta \frac{1}{\zeta - 1} \frac{1}{\rho + \delta}\right)^{-1} u^g$$

which can be summed to write a single Hamilton-Jacobi-Bellman equation for $\bar{S}$ that must hold in a stationary equilibrium:

$$\rho \bar{S} = \sum_g \left(1 - \lambda^g \left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}}\right)^\zeta \frac{1}{\zeta - 1} \frac{1}{\rho + \delta}\right)^{-1} u^g. \tag{15}$$

For this HJB, it is useful to consider the following proposition:

Proposition 4.1. There exists a unique solution $\rho\bar{S}$ to Equation (15) on the interval $(\rho\tilde{S}, \infty)$ where,

$$\rho \tilde{S} = \underbrace{\frac{\frac{\zeta}{\zeta-1}\bar{x}\frac{w}{r^\beta}(\Delta)^\eta}{\rho + \delta}}_{\text{unconditional expected match value}} \times \underbrace{\frac{1}{\zeta}\left(\max_g \lambda^g\right)^{1/\zeta} \left(\frac{1}{\zeta - 1}\frac{1}{\rho + \delta}\right)^{1/\zeta - 1}}_{\text{option value adjustment}}$$

.

Proof. See Proof A.2 in Appendix A. □

We view this as the only economically meaningful solution to Equation (15). It says that the flow-value equivalent solution to the HJB must exceed the value option-value adjusted value of matching. Any $\bar{S}$ below this threshold that satisfies Equation (15) is economically meaningless in that it does not express the value of search.

4.2 Spatial equilibrium

Given location primitives $\{\bar{A}_i, \bar{Z}_i, \bar{\Delta}_i, \bar{H}_i\}$, preference parameters $\{\beta, \epsilon\}$, matching productivities $\{\bar{M}_i, \bar{x}_i\}$, marriage market technology parameters $\{\rho, \delta, \eta, \zeta\}$, and agglomeration elasticities $\{\alpha, \kappa\}$, and housing supply elasticity $\{\gamma\}$, equilibrium is a spatial distribution of $\{L_i^m, L_i^f\}$, wages $\{w_i^m, w_i^f\}$, housing market prices and quantities, $\{H_i, r_i\}$, search values $\{S_i^m, S_i^f\}$ and singles shares $\{s_i^m, s_i^f\}$, such that,

1. Households choose locations optimally, solving (2), and make optimal decisions when searching for a match, solving (3);
2. Wages satisfy (5) and markdowns and productivity respond to market size (Equations 4 and 6);
3. The housing market clears so that housing demand (4) is equal to housing supply (7) at each i.
4. And S_i^g and s_i^g satisfy all conditions of a marriage market partial equilibrium for a $\theta_i = L_i^f / L_i^m$ and u_i^g consistent with conditions (1)-(3).

Solution Using properties of the Fréchet distribution, the population share of any gender in any location is,

$$\frac{L_i^g}{\bar{L}} = \frac{(A_i S_i^g)^\epsilon}{\sum_j (A_j S_j^g)^\epsilon} \tag{16}$$

Housing market clearing guarantees that,

$$r_i = \left(\frac{\beta \left(w_i^m L_i^m + w_i^f L_i^f \right)}{\bar{H}_i} \right)^{\frac{1}{1+\gamma}}. \tag{17}$$

Discussion Our model does not take a particular stand on the source of gender-specific agglomeration elasticities. This is not to say they cannot be microfounded: Appendix A.5 lays out a model in which there is an endogenous mass of employment opportunities in a given city, employers are monopsonistic, but pay fixed costs to enter, and there is Roy-like sorting in the labor market. While monopsony wage-setting and gender-specific labor supply elasticities generate a gender wage gap (à la Manning, 2013), they alone do not explain a gender wage gap that responds to labor market size. To generate a gender wage gap that is increasing in city size, we require both that the dispersion of match productivity to be greater for women than for men and free entry of employers. Greater dispersion in match productivity for women may be driven by differences in commuting preferences (Black et al., 2014; Le Barbanchon et al., 2021; Liu and Su, 2024b) or hours flexibility (Goldin, 2014). In this theory, stochastic job-worker match productivities generate an upward sloping supply of labor to a given occupation and employer free entry generates love-of-variety type returns to workers: larger cities have more varied employers, which increases the size of the expected match-productivity adjusted wage for workers, and moreso for women than men. While a theory of gender-specific match productivity distributions and monopsonistic competition is one way to rationalize why larger cities have lower gender wage gaps, we emphasize that it is not the only theory consistent with our facts. Another possibility is that labor or product market competition may be tougher in big cities, withering away taste-based discrimination (Becker, 1957).[6]

Our model makes a strong functional form assumption on the nature of the flow benefits to marriage, $B(x)$. A Beckerian time-use model yields a similar expression but generates potential nonmonotonicities in how the flow value of marriage depends on the gender wage gap: a lower gender wage gap increases labor market opportunities for women, reducing their time contribution to home production, lowering the flow value of the match, but increases the resources available to the household, increasing the flow value. In Appendix A.6 we derive such a Beckerian time-use model and show that it admits a flow value of marriage that has a local log-linear approximation to the functional form we use for reasonable values of Δ.

Finally, the frictional marriage market provides microfoundations for what can be thought of as endogenous amenities. Past work on endogenous amenities has focused on elastically supplied nonmarket goods like open space (Walsh, 2007), tourist amenities (Almagro and Domínguez-Iino, 2025), racial composition (Banzhaf and Walsh, 2013; Gregory et al., 2022), or locally provided public goods (Calabrese et al., 2012; Huang, 2025; Ruggieri, 2025). Here, increasing the supply of agents of gender g produces a positive externality for agents of gender g' by increasing marriage market options, but generates a negative own-gender population size externality by congesting the marriage market. These negative own-gender externalities are strong and ensure that few places deviate from population gender parity in equilibrium, but they are not urban cost elasticities that scale with the overall size of the city (Combes, Duranton, and Gobillon, 2019; Rosenthal-Kay, 2025); such congestion elasticities measure how consumption-equivalent utility changes with overall market size and in our model these operate through the housing market. We can benchmark the magnitude of our gender-specific externalities by examining the implied congestion elasticity with respect to (own-gender) population size against a common specification stemming from the distribution of preference shocks

[6]Several papers have found tougher product market competition in big cities (e.g., Rosenthal-Kay et al., 2024; Franco, 2025), and others have found the toughness of labor market competition rising in city size (Hirsch, Jahn, et al., 2022). In line with this theory, Hirsch, Oberfichtner, et al. (2014) finds that greater product market competition is associated with lower within-firm gender pay gaps in Germany, while Cooke et al. (2019) finds similar results in Portugal.

and housing market congestion (see the appendix of Allen and Arkolakis, 2014, for a discussion of these elasticities). Appendix Figure A5 displays the log slope of S^g with respect to L^g: our marriage market congestion elasticities are substantially larger in magnitude than those used in a benchmark quantitative spatial model.

Taking the model to data We now derive estimating equations we can take to the data to recover the model's key parameters. Wage setting in the model implies,

$$\ln w_i(\omega^g) = a_0 + \alpha \ln L_i + \kappa \times 1(g = f) \times \ln L_i + 1(g = f) \ln \bar{\Delta}_i + u_i, \tag{18}$$

where a_0 is a constant and $u_i = \ln \bar{Z}_i$. In our model, α and κ are not identified by OLS regressions due to standard simultaneity bias: labor is itself endogenous to the strength of agglomeration economies (Combes, Duranton, and Gobillon, 2011). Thus, we rely on the IV estimates in Table 7 for identification of α and κ.

The model also tells us how to interpret differences in the singles share across cities. Rearranging Equation (14) gives,

$$\ln \frac{s_i^g}{1 - s_i^g} = \ln \delta - \underbrace{(\ln M_i + (1 - \mu) \ln \theta_i - 1(g = f) \ln \theta_i)}_{\ln \lambda^g} - \underbrace{\zeta (\ln \bar{x}_i - \ln x_i^*)}_{\ln p^*} \tag{19}$$

As $\ln x_i^* = \ln \bar{S}_i - \eta \ln \Delta_i - \ln w_i + \beta \ln r_i$, Equation (19) can be written as,

$$\ln \frac{s^g}{1 - s^g} = b_0 + (\mu - 1) \ln \theta_i + 1(g = f) \ln \theta_i - \zeta \ln w_i + \zeta \beta \ln r_i - \zeta \eta \ln \Delta_i + v_i \tag{20}$$

where b_0 is a constant, and the error term combines exogenous and endogenous components of search, $v_i = \ln M_i - \zeta \ln \bar{x}_i + \zeta \ln \bar{S}_i$.

Exact identification of model fundamentals We now show the parameters $\{\bar{A}_i, \bar{Z}_i, \bar{\Delta}_i, \bar{H}_i, \bar{M}_i, \bar{x}_i\}$ are exactly identified, identified up to scale, or product-identified, conditional on data $\mathbf{X} = \{L_i^f, L_i^m, w_i, \Delta_i, r_i\}$ and model parameters $\mathbf{P} = \{\alpha, \kappa, \beta, \gamma, \epsilon, \rho, \delta, \eta, \zeta, \mu\}$. That is to say, our model fits into a large class of quantitative spatial models that can exactly reproduce the data as an equilibrium of the model (Redding and Rossi-Hansberg, 2017). Our model parameter choices are described in Appendix Table A1.

Conditional on knowing S_i^g, we recover $\bar{A}_i^g$ up to scale to match population shares by gender across space (Equation 16). The remaining fundamentals $\bar{H}_i, \bar{Z}_i, \bar{\Delta}_i$ are recovered by inverting wage and housing market clearing equations (Equations 5, 6 and 17). We are only able to recover $\tilde{M}_i = \bar{M}_i \bar{x}_i^{\zeta}$ up to scale, but the level of matching technology and the average draw of a match shock enter symmetrically, so without loss we can fix $\bar{x}_i = 1$ everywhere. We recover $\tilde{M}_i$ with a nonlinear solver using Equation (19) and the fixed point for $\bar{S}$ (Equation 15).

As our model can rationalize any data $\mathbf{X}$, our fit depends on what data we feed the model. We use 2010 as a base year and regression-adjust Δ_i, θ_i, s_i^f, and δ_i (which we allow to vary across space). We do this by estimating a cross-sectional regression of these variables at the CZ-level on average ages, racial shares, and ethnicity shares across CZs. We then use residuals from this regression and add back in the demographic factors, as if each CZ had the same demographic composition as the United States on average. This then removes compositional differences in these variables driven by differences in race, age, and ethnicity across space.

Figure 6 shows the inverted matching technology $\log \tilde{M}_i$ across CZs. Matching technology appears high in areas with a high share of Mormons, suggesting that local religious and cultural institutions are captured by our estimates of this parameter. Figure 6, Panel (B) indeed shows that our estimated marriage technology is strongly correlated with religious adherence within a commuting zone.

5 How much does the city-size GWG premium matter?

The main counterfactual we consider is one in which we flatten the city-size wage premium specific to women across space. That is, we replace the calibrated baseline $\bar{\Delta}_i$ with $\bar{\Delta}_i \left(\bar{L}/I\right)^{\kappa}$ and then set $\kappa = 0$ so the gender wage gap does not adjust endogenously. Adjusting $\bar{\Delta}_i$ by $\left(\bar{L}/I\right)^{\kappa}$ ensures that we do not induce scale effects by simply setting $\kappa = 0$ (Greaney, 2023). In other words, this counterfactual can be though of as jointly fixing the component of the gender wage gap endogenous to city size to the 'average' city size and then shutting down its endogenous response to population changes.

The left panel of Figure 7 shows the shock we feed into the model: the plot displays the percent change in

A. Marriage Technology

B. Correlation with Religious Adherence

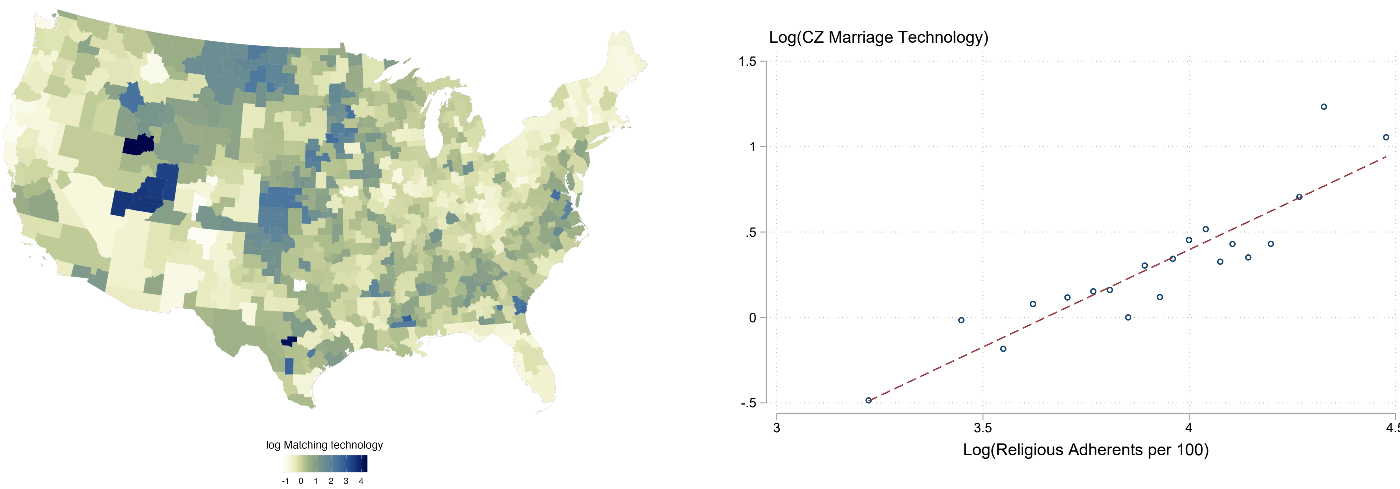


Figure 6: Left: Map of estimated marriage technology across space. Right: Binscatter plot of log marriage technology on log religious adherence per 100 individuals

Δ_i against the baseline measure of Δ_i, where each point is a commuting zone, scaled by its baseline size. This counterfactual mechanically widens the gender wage gap (i.e., $\downarrow \Delta_i$) in big cities while shrinking it in smaller places.

The right panel of 7 shows the counterfactual population change from eliminating the city-size component of the gender wage gap: major cities like New York and Chicago shrink, while less populous places like Abilene, Texas, Billings, Montana, or Asheville, North Carolina absorb these urban migrants. Marriage markets also endogenously respond to these changes. Appendix Figure A6 shows how the share of singles changes across commuting zones: on the margin, more women marry in larger cities with shrinking gender wage gaps, while smaller cities both see an increase in pay equity and increased marriage market congestion from population change, both putting upwards pressure on the equilibrium share of singles. However, these changes are small, as, singles shares on average adjust by around 0.5 percentage points.

On net, eliminating this endogenous component of the gender wage gap widens pay disparity between men and women, from about 27 cents on the dollar at baseline to 29 cents: about a 7% increase.

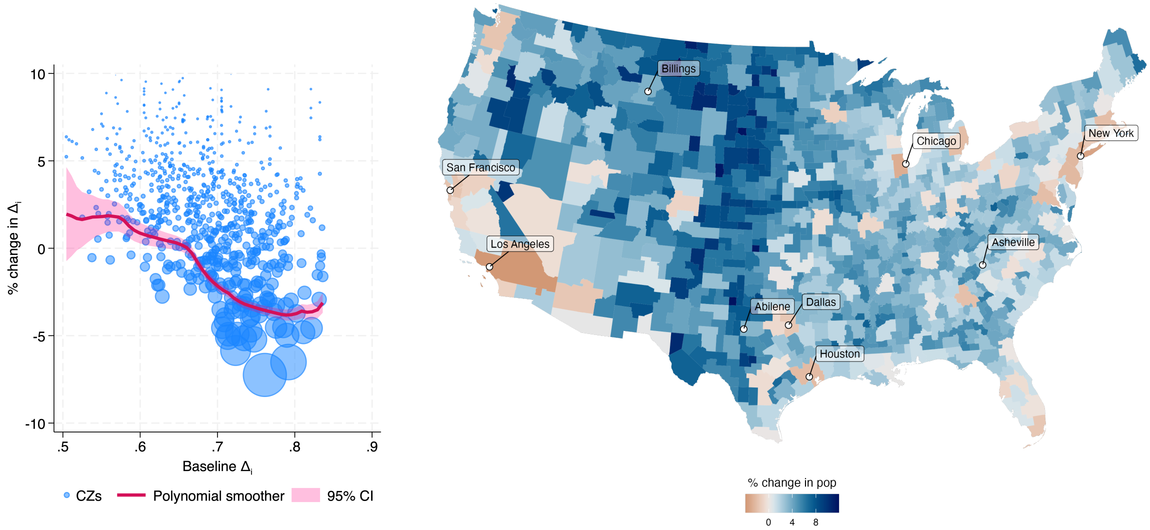


Figure 7: Left: CZ-level population change from eliminating the female-specific city-size premium. Right: Percent change in Δ_i vs. baseline Δ_i across commuting zones. Each point is scaled by baseline size.

6 Conclusion

In this paper, we study how earnings gaps between men and women vary across space, and then document how that affects spatial sorting, marriage, and aggregate outcomes. We find strong evidence that women benefit more than men from agglomeration economies. Interestingly, this result is driven by women with lower levels of education and cannot be explained by industrial composition, occupation, or educational sorting across space. The standard logic of spatial equilibrium would suggest this should lead to large quantity responses. We find evidence that women do sort to larger labor markets in response, but the size of the response suggests migration elasticities that are much too small. Instead, we find a significant response in the marriage market: large cities feature lower marriage rates than small cities, exactly in the populations that benefit most from these agglomeration externalities.

We then fit our data using a spatial equilibrium model with a local marriage market subject to search-and-matching frictions. We use the model to study how migration responds to these gender-specific earnings gaps and study how spatial variation in earnings gaps affects overall gender equity. We study a counterfactual world where women benefit at the same rate as men from agglomeration, but the gender earnings gap remains fixed. We find that eliminating city-specific earnings premia for women would induce migration from both men and women out of larger cities, increase marriage rates in large cities, while congesting marriage markets in smaller cities. On net, this would increase the aggregate gender pay gap by 7%.

In ongoing work, we use our quantitative model to study how aggregate changes in the gender wage gap since 1980 have affected marriage rates and spatial sorting. We also use the model to investigate how policy changes, such as reducing housing-supply restrictions in the largest cities affect aggregate output, gender inequality, and marriage markets. We also abstracted from modeling heterogeneity by education groups or racial groups directly. However, our empirical work found strong evidence that low-education women and Black women see the strongest gains from large labor markets. Expanding the model to include these types of heterogeneity is a promising avenue for future research, and investigating how these premia affect racial and income inequality, as well as gender inequality, is an important area for future research.

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Online Appendix

A Theory details

A.1 Marriage market model

Environment Recall the elements of the model:

- Flow utility $u^g = A\frac{w^g}{r^\beta}$,
- Populations of both genders L^g and market tightness $\theta = L^f/L^m$,
- Discount rate ρ and marriage separation rate δ
- Married household flow utility $B(x) = x(u^m)^{1-\eta}(u^f)^\eta$
- Matching technology $\mathcal{M}(L^f, L^m) = \bar{M}(L^m)^\mu(L^f)^{1-\mu}$ so that match hazards are $\lambda^f = \bar{M}\theta^{-\mu}$ and $\lambda^m = \bar{M}\theta^{1-\mu}$
- Common match shocks $x \sim$ iid Pareto$(\bar{x}, \zeta)$.

Value functions We define the value of being married for any gender $g \in \{m, f\}$ to be,

$$\rho V^g(x,t) = \phi^g B(x) + \delta\left(S^g - V^g(x,t)\right) + \partial_t V^g(x,t),$$

And we note that the HJB term $\partial_x V^g(x,t)\dot{x} = 0$ as match shocks are permanent to a couple, so that $\dot{x} = 0$.

In this formulation S^g is the value of search and $\phi^g \in (0,1)$ is an endogenous division of surplus. However, as the surplus is split so $\phi^m + \phi^f = 1$, it is easier to work with the joint surplus $J(x,t) = V^m(x,t) + V^f(x,t)$,

$$\rho J(x,t) = B(x) + \delta\left(\bar{S} - J(x,t)\right) + \partial_t J(x,t),$$

where $\bar{S} = \sum_g S^g$ is the joint value of search. At a stationary equilibrium ($\partial_t J(x,t) = 0$),

$$J(x) = \frac{B(x)}{\rho + \delta} + \frac{\delta}{\rho + \delta}(\bar{S})$$

And therefore the stationary value of search is,

$$\rho S^g = u^g + \lambda^g \mathbb{E}_x \left[\max\{\phi^g J(x) - S^g, 0\}\right]$$

Marriage market equilibrium Given θ, $\{u^g\}$, an stationary equilibrium in the marriage market is characterized by a cutoff match value x^* and a division of surplus $\{\phi_g\}$ that imposes that the cutoff match value generates equal indifference at the margin for both agents, (Equation 10 in the text). The surplus is split in full $\phi^m + \phi^f = 1$, $\phi^g \in (0,1)$, and a mass of singles for each gender s^g, such that agents solve Equation (9), $\bar{S} = S^m + S^f$ and Equation (8) holds.

A.2 Solution

First, we solve for x^*. Plugging in (8) into equilibrium condition (10), we have that,

$$\frac{B(x)}{\rho + \delta} + \frac{\delta}{\rho + \delta}\bar{S} = \bar{S} \implies H(x) = \rho \bar{S}$$

As $H(x) = x \cdot \frac{w}{r^\beta} \cdot (\Delta)^\eta$, we have that,

$$x^* = \frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}$$

This only occurs if x^* satisfies both,

$$\phi^m J(z^*) = S^m$$
$$\phi^f J(z^*) = S^f.$$

To solve for ϕ^g, note that these equations imply,

$$\frac{S^m}{\phi^m} = \frac{S^f}{\phi^f}$$

Comibing these with $\phi^m + \phi^f = 1$, we have that,

$$\phi^g = \frac{S^g}{S^m + S^g}.$$

Conditional on receiving a draw, the probability of forming a match is $p^* = (\bar{x}/x^*)^\zeta$, and so the probability of matching is $\lambda^g p^*$. The mass of singles of either gender is given by finding the value of s^g that makes $\dot{s}^g$ stationary:

$$\dot{s}^g = \delta(1 - s^g) - \lambda^g p^* s^g \implies s^g = \frac{\delta}{\delta + \lambda^g p^*}.$$

Now, we solve agents' optimal search problems (given by 9 in the text). We begin with the expression,

$$\mathbb{E}_x\left[\max\{\phi^g J(x) - S^g, 0\}\right] = \left[\phi^g \frac{\mathbb{E}_x[B(x) \mid x > x^*]}{\rho + \delta} + \phi^g \frac{\delta}{\rho + \delta}\bar{S} - S^g\right]\left(\frac{\bar{x}}{x^*}\right)^\zeta$$

So that,

$$\rho S^g = u^g + \lambda^g \left[\phi^g \frac{\mathbb{E}_x[B(x) \mid x > x^*]}{\rho + \delta} + \phi^g \frac{\delta}{\rho + \delta}\bar{S} - S^g\right]\left(\frac{\bar{x}}{x^*}\right)^\zeta$$

Using properties of the Pareto distribution, we have that, $\mathbb{E}_x[B(x) \mid x > x^*] = \frac{\zeta}{\zeta - 1} x^* \frac{w}{r^\beta}(\Delta)^\eta$. Therefore,

$$\begin{aligned}
\rho S^g &= u^g + \lambda^g \left[\phi^g \frac{\frac{\zeta}{\zeta-1} x^* \frac{w}{r^\beta}(\Delta)^\eta}{\rho + \delta} + \phi^g \frac{\delta}{\rho + \delta}\bar{S} - S^g\right]\left(\frac{\bar{x}}{x^*}\right)^\zeta \\
\text{plugging in for } x^*, \qquad &= u^g + \lambda^g \left[\phi^g \frac{\frac{\zeta}{\zeta-1}\left(\frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}\right)\frac{w}{r^\beta}(\Delta)^\eta}{\rho + \delta} + \phi^g \frac{\delta}{\rho + \delta}\bar{S} - S^g\right]\left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}}\right)^\zeta \\
&= u^g + \lambda^g \left[\phi^g \frac{\frac{\zeta}{\zeta-1}(\rho \bar{S})}{\rho + \delta} + \phi^g \frac{\delta}{\rho + \delta}\bar{S} - S^g\right]\left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}}\right)^\zeta \\
&= u^g + \lambda^g \left[\left(\frac{\zeta}{\zeta - 1}\frac{\rho}{\rho + \delta} + \frac{\delta}{\rho + \delta}\right)\phi^g \bar{S} - S^g\right]\left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}}\right)^\zeta
\end{aligned}$$

Now, using ϕ^g solution,

$$\rho S^g = u^g + \lambda^g \left[\left(\frac{\zeta}{\zeta - 1}\frac{\rho}{\rho + \delta} + \frac{\delta}{\rho + \delta}\right) - 1\right]\left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta}(\Delta)^\eta}}\right)^\zeta S^g$$

Rearranging,

$$\rho S^g = u^g + \lambda^g \left[\frac{\zeta}{\zeta - 1} \frac{\rho}{\rho + \delta} + \frac{\delta}{\rho + \delta} - \frac{\rho + \delta}{\rho + \delta} \right] \left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta} (\Delta)^\eta}} \right)^\zeta S^g$$

Or,

$$\rho S^g = u^g + \lambda^g \left[\frac{1}{\zeta - 1} \right] \frac{\rho}{\rho + \delta} \left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta} (\Delta)^\eta}} \right)^\zeta S^g$$

That is,

$$\rho S^g = \left(1 - \lambda^g \left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta} (\Delta)^\eta}} \right)^\zeta \frac{1}{\zeta - 1} \frac{1}{\rho + \delta} \right)^{-1} u^g$$

Then summing,

$$\rho \bar{S} = \sum_g \left(1 - \lambda^g \left(\frac{\bar{x}}{\frac{\rho \bar{S}}{\frac{w}{r^\beta} (\Delta)^\eta}} \right)^\zeta \frac{1}{\zeta - 1} \frac{1}{\rho + \delta} \right)^{-1} u^g$$

The following proves Proposition 4.1, which shows that 9 has a unique solution on the only economically meaningful domain.

Proof. This will blow up when the denominator gets near zero. We must restrict the domain. To see this, note that we can write the above as,

$$\rho \bar{S} = \sum_g \frac{u_g}{1 - K^g (\rho S)^{-\zeta}}$$

To keep the denominator positive (i.e., areas for where the option value of search increases flow welfare) we need,

$$1 - K^g (\rho S)^{-\zeta} > 0 \, \forall g \implies \rho S > (\max_g K^g)^{1/\zeta}$$

Examining the righthandside,

$$(\max_g K^g)^{1/\zeta} = \left(\max_g \lambda^g \right)^{1/\zeta} z_0 \frac{w}{r^\beta} (\Delta)^\eta \left(\frac{1}{\zeta - 1} \frac{1}{\rho + \delta} \right)^{1/\zeta}$$

Note we can write this as,

$$(\max_g K^g)^{1/\zeta} = \underbrace{\frac{\frac{\zeta}{\zeta-1}\bar{x}\frac{w}{r^\beta}(\Delta)^\eta}{\rho+\delta}}_{\text{unconditional expected match value}} \times \underbrace{\frac{1}{\zeta}\left(\max_g \lambda^g\right)^{1/\zeta}\left(\frac{1}{\zeta-1}\frac{1}{\rho+\delta}\right)^{1/\zeta-1}}_{\text{option value adjustment}}$$

This says, we want to find a solution to the searcher's Hamilton-Jacobi-Bellman equation is such that such that the flow value of search $\rho\bar{S}$ to exceed the option-value-adjusted expected unconditional match surplus; any other solution is not economically meaningful.

We want to now prove that there exists a unique solution $\rho\bar{S}$ on $(\tilde{K},\infty)$. To do so, consider the function,

$$g(x) = x - \sum_g \frac{u_g}{1 - K^g(x)^{-\zeta}}$$

As $x \to \tilde{K}$, $g(x) \to -\infty$ and as $x \to \infty$, $g(x) \to \infty$ and $g(\cdot)$ is clearly continuous in x so by the intermediate value theorem there must exist a $\tilde{x} \in (\tilde{K},\infty)$ such that $g(\tilde{x}) = 0$. We now want to prove it is unique. To do so, we check monotonicity.

$$g'(x) = 1 + \sum_g \frac{u^g \zeta K^g(x)^{-\zeta-1}}{(1-K^g(x)^{-\zeta})^2} > 0 \quad \forall x \in (\tilde{K},\infty)$$

So a unique solution exists. □

A.3 Inversion algorithms

A.3.1 Recovering quality-adjusted matching technology

The match quality-adjusted matching technology is the parameter $\tilde{M}_i = \bar{M}_i\bar{x}_i^\zeta$. Given data $\mathbf{X}$ (in particular, θ_i, w_i^g, r_i) and estimates of the parameters $\mathbf{P}$ (in particular, $\zeta,\rho,\delta,\mu,\beta$), we can solve for $\rho\bar{S}(\tilde{M};\mathbf{X},\mathbf{P})$ by solving the fixed point,

$$\rho\bar{S} = \left(1 - \tilde{M}_i\theta_i^{1-\mu}(\rho\bar{S})^{-\zeta}(w_i^m\Delta_i^\eta r_i^{-\beta})^\zeta\right)^{-1}\frac{w_i^m}{r_i^\beta} + \left(1 - \tilde{M}_i\theta_i^{-\mu}(\rho\bar{S})^{-\zeta}(w_i^m\Delta_i^\eta r_i^{-\beta})^\zeta\right)^{-1}\frac{w_i^f}{r_i^\beta}$$

We have already proven there exists a unique root for $\rho\bar{S}$. Using the solution $\rho\bar{S}(\tilde{M};\mathbf{X},\mathbf{P})$ we can construct ρS^m and ρS^f, and check whether their ratio is equal to $(\theta_i)^{1/\epsilon}$.

Estimation From Equation (9) we have that,

$$\ln \frac{s^g}{1-s^g} = \ln \delta - \ln(\lambda^g p^*)$$

which we can use as a basis for estimation. In particular, as, $\ln p* = \zeta \ln \bar{x} - \zeta \ln x^*$ and $\ln x^* = \ln \bar{S} - \eta \ln(\Delta) - \ln w/r^\beta$.

A.4 Parameter choices

For the quantified model, we estimate some parameters $\mathbf{P} = \{\alpha, \kappa, \beta, \gamma, \epsilon, \rho, \delta, \mu, \eta, \zeta\}$ and externally calibrate others. See Appendix Table A1 for details.

Parameter	description	value	note
α	Agglomeration elasticity	0.04	literature standard, close to estimated value
κ	Female-specific agglomeration elasticity	0.02	50% bigger than for males
β	Share of income spent on housing	0.24	Standard
ϵ	Migration elasticity	1.2	Suárez Serrato and Zidar (2016)
ρ	Discount factor	0.02	unit of time is a year, standard
δ	Divorce rate	CZ specific	regression-adjusted hazard
μ	Matching elasticity	0.5	DMP standard
η	Marriage output elasticity for female income	0.5	even contribution
ζ	(inverse) Dispersion in marriage shocks	2.2	chosen so match shocks have finite but large variance

Table A1: Calibrated model parameters

A.5 Microfoundations for a spatially-varying gender wage gap

Suppose each city i is composed of a measure of N_i^g gender-stratified occupations, where N_i^g is endogenous. I now suppress location subscripts. Each occupation n is identical, with TFP Z, so occupational output is,

$$y_n^g = Z\mathcal{L}_n^g$$

where $\mathcal{L}_n^g$ are the effective units of labor in occupation n. Workers are heterogeneous in the effective units of labor ℓ they supply to employers. The assignment of workers to occupations is determined by workers' who solve,

$$n^g(\omega^g) = \text{argmax}_n \, w_n^g \ell(\omega^g)$$

where ℓ^g is distributed Frechet$(\tilde{\gamma}^g, \varsigma^g)$ where $\tilde{\gamma}^g = 1/\Gamma(1 - 1/\varsigma^g)$. Subsequently, the share of workers in a given occupation is,

$$\frac{L_n^g}{L^g} = \frac{(w_n^g)^{\varsigma^g}}{\int_0^N (w_n^g)^{\varsigma^g} dn}.$$

Each occupation n has monopsony power in the labor market, and maximizes their profits, π_n,

$$\pi_n = \max_{\{w_n^g\}_g} Z\mathcal{L}_n^g(w_n^g) - \sum_g w_n^g \mathcal{L}_n^g(w_n^g) - F$$

where F is a fixed cost measured in units of output. This has the solution for all occupations,

$$w_n^g = \frac{\varsigma^g}{1 + \varsigma^g} Z.$$

This wage markdown formula holds regardless of n, so by symmetry,

$$\frac{L_n^g}{L^g} = \frac{1}{N^g}$$

Now note that expected effective units of labor to the occupation n for gender g are,

$$\mathcal{L}_n^g = \left(\int_0^N (w_n^g)^{\varsigma^g} dn \right)^{1/\varsigma^g} \times L_n^g = \frac{\varsigma^g}{1 + \varsigma^g} Z N^{1/\varsigma^g} \times \frac{L^g}{N^g}$$

Profits therefore are,

$$\pi_n = \frac{L^g}{N^g} \frac{1}{1 + \varsigma^n} Z N^{1/\varsigma^g}$$

Firms enter until profits are zero, so,

$$N^g = \left(\frac{Z L^g}{1 + \varsigma^g} \right)^{\frac{\varsigma^g}{\varsigma^g - 1}}$$

Then,

$$\begin{aligned}
\mathbb{E}\left[w_n^g \ell_n^g \mid w_n^g \ell_n^g \geq \max_{n'} w_{n'}^g \ell_{n'}^g\right] &= \frac{\varsigma^n}{\varsigma^n+1} Z (N^g)^{1/\varsigma^g} \\
&= \frac{\varsigma^n}{\varsigma^n+1} Z \left(\frac{Z L^g}{1+\varsigma^g}\right)^{\frac{1}{\varsigma^g - 1}} \\
&= \underbrace{\frac{\varsigma^n}{\varsigma^n+1} \left(\frac{Z}{1+\varsigma^g}\right)^{\frac{1}{\varsigma^g - 1}}}_{\text{exogenous GWG}} \underbrace{(L^g)^{\frac{1}{\varsigma^g-1}}}_{\text{endogenous GWG}} \quad Z
\end{aligned}$$

Then, $\alpha = \frac{1}{\varsigma^m - 1}$, and $\kappa = \frac{1}{\varsigma^f - 1} - \frac{1}{\varsigma^m - 1}$.

A.6 A Beckerian time-use model of $B(x)$

Suppose married couples can contribute a fraction of their time $t^g \in (0,1)$ home production, but at the cost of foregoing their wage, and reducing resources available to the household. In particular, assume the couple jointly solves,

$$\tilde{B}(x) = \max_{t^m, t^f, C, H} \; x (t^m)^{1-\chi} (t^f)^{\chi} U(C, H),$$

where $U(C, H)$ is the same Cobb-Douglas utility presented in the text, and their joint budget constraint is,

$$C + rH \leq w^m(1 - t^m) + w^f(1 - t^f)$$

Then, at the optimum,

$$\tilde{B}(x) = x \frac{\tilde{\chi}}{r^\beta} \frac{1}{4} \frac{(w^m + w^f)^2}{(w^m)^{1-\chi}(w^f)^\chi}$$

where $\tilde{\chi}$ is a constant. Specializing to $w^f = \Delta w^m$, we have,

$$\tilde{B}(x) = x \tilde{\chi} \frac{1}{4} (1+\Delta)^2 (\Delta)^{-\chi} \frac{w^m}{r^\beta}$$

This introduces nonmonotonicities.

However, consider a log-Taylor approximation to,

$$\underbrace{(1+\Delta)^2(\Delta)^{-\chi}}_{D(\Delta)} \approx D(1)\Delta^k, \quad k = \frac{d\log D}{d\log\Delta}\bigg|_{\Delta=1}$$

Now,

$$\begin{aligned}\frac{d\log D}{d\log\Delta}\bigg|_{\Delta=1} &= \Delta\frac{d\log D(\Delta)}{d\Delta}\bigg|_{\Delta=1} \\ &= \Delta\left(\frac{2}{1+\Delta} - \frac{\chi}{\Delta}\right)\bigg|_{\Delta=1} \\ &= 1-\chi\end{aligned}$$

So, as $D(1) = 4$,

$$\begin{aligned}\tilde{B}(x) &= x\tilde{\chi}\frac{1}{4}(1+\Delta)^2(\Delta)^{-\chi}\frac{w^m}{r^\beta} \\ &\approx x\tilde{\chi}\frac{w^m}{r^\beta}(\Delta)^{1-\chi}\end{aligned}$$

which is approximately the same as $B(x)$ reported in the text with $\eta = 1-\chi$.

This is a good approximation around $\Delta = 1$ when $D(\Delta)$ is increasing. Denoting the turning point of D as $\Delta^\dagger = \operatorname{argmin}_\Delta D(\Delta)$,

$$\Delta^\dagger = \frac{\chi}{2-\chi}$$

When χ is $1/2$, $\Delta^\dagger = 1/3$, which is much larger than any wage gap observed in the data, suggesting this is a good local approximation.

B Additional Empirical Tests

Appendix Table A2 tests whether industrial composition across space can explain our pattern of results. We show that adding highly detailed industry fixed effects makes little difference to our estimates. Even when adding industry x occupation fixed effects, we find small changes in our estimates of city-size gender-earnings premium. This suggests that the sorting of industries across space and differential returns for women across occupations and industries can do little to explain our results (Bacolod, 2017; Fang et al., 2025)

Table A2: Gender Gap and City Size: Industry-Occupation Interactions

	(1)	(2)	(3)	(4)	(5)
Female	-0.676***	-0.803***	-0.564***	-0.495***	-0.463***
	(0.044)	(0.036)	(0.032)	(0.030)	(0.030)
log(CZ pop)	0.092***	0.054***	0.056***	0.049***	0.049***
	(0.008)	(0.005)	(0.005)	(0.005)	(0.005)
Female × log(CZ pop)	0.021***	0.027***	0.017***	0.016***	0.014***
	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)
Abilene $\rightarrow$ NY Effect	0.086	0.111	0.069	0.065	0.060
Year FE	✓	✓	✓	✓	✓
Education FE		✓	✓	✓	✓
Industry FE			✓	✓	
Occupation FE				✓	
Ind. x Occ. FE					✓
R-sq.	0.114	0.217	0.298	0.377	0.400
Observations	13250889	13250889	13250889	13250889	13237059

In Appendix Table A3, we include CZ fixed effects. Thus, these results estimate gender earnings premiums by comparing individuals in the exact same city. Note that in this specification, the overall city-size earnings premium is unidentified: thus, we only estimate the *differential* effect of large cities on women. Our estimates of the gender-specific city-size earnings premium are very close to our baseline estimates.

Table A3: Gender Gap and City Size: CZ Fixed Effects

	Pooled			By Educ.: Less than BA		By Educ.: BA+	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Female	-0.666***	-0.794***	-0.560***	-0.829***	-0.629***	-0.538***	-0.406***
	(0.042)	(0.035)	(0.029)	(0.041)	(0.030)	(0.086)	(0.068)
Female × log(CZ pop)	0.020***	0.026***	0.019***	0.031***	0.023***	0.009	0.009*
	(0.003)	(0.003)	(0.002)	(0.003)	(0.002)	(0.006)	(0.005)
Abilene $\rightarrow$ NY Effect	0.083	0.109	0.078	0.127	0.095	0.036	0.037
CZ FE	✓	✓	✓	✓	✓	✓	✓
Year FE	✓	✓	✓	✓	✓	✓	✓
Education FE		✓	✓				
Occupation FE			✓		✓		✓
R-sq.	0.126	0.223	0.359	0.107	0.271	0.134	0.323
Observations	13250889	13250889	13250889	8920338	8920338	4330551	4330551

In Appendix Table A4, we estimate the same premium while allowing for even more flexible sorting. In this specification, we allow educational groups and occupations to have completely heterogeneous returns across CZs by including education group x CZ and occupation x CZ fixed effects. Thus, this strategy tests for a city-size earnings premium by comparing women and men in the exact same education group and the exact same city, as well as individuals in the same occupation and CZ. We again find a very large residual gender earnings gap, that is close to our baseline estimates. In fact, in this regression we estimate economically and statistically significant earnings premiums for women with a B.A. or above, once we allow for occupation x CZ fixed effects. However, the magnitude of the effect is still smaller than the results for women with less than a B.A.

Table A4: Gender Gap and City Size: CZ-Covariate Interactive Fixed Effects

	Pooled		By Educ.: Less than BA		By Educ.: BA+	
	(1)	(2)	(3)	(4)	(5)	(6)
Female	-0.778***	-0.557***	-0.898***	-0.593***	-0.545***	-0.467***
	(0.035)	(0.028)	(0.042)	(0.020)	(0.083)	(0.052)
Female × log(CZ pop)	0.025***	0.019***	0.034***	0.021***	0.009	0.014***
	(0.003)	(0.002)	(0.003)	(0.002)	(0.006)	(0.004)
Abilene → NY Effect	0.104	0.078	0.140	0.086	0.036	0.056
Year FE	✓	✓	✓	✓	✓	✓
Education-CZ FE	✓	✓	✓	✓	✓	✓
Occupation-CZ FE		✓		✓		✓
R-sq.	0.225	0.387	0.137	0.316	0.152	0.368
Observations	13250859	13190163	8920308	8853317	4330551	4252629

Appendix Table A5 includes controls for political variation across space. We proxy for political ideology using historical voting records across counties, aggregated to the CZ level. We allow political ideology to have differential effects on the wages of men and women. We still estimate large and significant gender wage premia, though the coefficients do attenuate notably. This suggests one set of potential mechanisms involves either differences in policies across space that affect gender earnings gaps, or differences in attitudes towards women in the workforce that are correlated with political attitudes.

Finally, Appendix Tables A6 and A7 re-estimate our specifications for gender-specific returns for parents and married individuals, after including our controls for labor supply and commuting. In these specifications, we find much weaker differences in the gender-specific earnings premium between married and single women and between women with and without children. In these estimates, we still find negative effects on the triple interaction for women in large cities who are either married or with children. However, these differences

Table A5: Gender Gap and City Size: Results With Political Controls

	Pooled			By Educ.: Less than BA		By Educ.: BA+	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Female	-0.550***	-0.687***	-0.494***	-0.826***	-0.657***	-0.232	-0.185
	(0.108)	(0.093)	(0.072)	(0.110)	(0.088)	(0.195)	(0.145)
log(CZ pop)	0.083***	0.051***	0.038***	0.046***	0.030***	0.079***	0.056***
	(0.007)	(0.005)	(0.004)	(0.004)	(0.003)	(0.011)	(0.007)
Female × log(CZ pop)	0.009***	0.017***	0.013***	0.022***	0.018***	-0.003	0.001
	(0.003)	(0.003)	(0.002)	(0.003)	(0.002)	(0.008)	(0.006)
Abilene $\rightarrow$ NY Effect	0.038	0.071	0.055	0.092	0.075	-0.012	0.006
Political Controls	✓	✓	✓	✓	✓	✓	✓
Year FE	✓	✓	✓	✓	✓	✓	✓
Education FE		✓	✓				
Occupation FE			✓		✓		✓
R-sq.	0.116	0.218	0.355	0.101	0.267	0.127	0.319
Observations	13133601	13133601	13133601	8829839	8829839	4303762	4303762

mostly offset premiums for men who are married or parents in large cities. Thus, this suggests that differences in gender-specific returns are driven by the fact that married women and women with children do not supply more labor or benefit from better commuting infrastructure in large cities in the same way as their childless or unmarried peers. In addition, we find that women do not reap the same pecuniary benefits of marriage or parenthood that men receive in large cities.

Table A6: Gender Gap and City Size, the Effect of Children

	(1)	(2)	(3)
Female	-0.211***	-0.378***	-0.418***
	(0.026)	(0.022)	(0.022)
log(CZ pop)	0.082***	0.049***	0.040***
	(0.006)	(0.004)	(0.004)
Female × log(CZ pop)	0.015***	0.018***	0.020***
	(0.002)	(0.001)	(0.001)
Children x log(CZ pop)	0.006*	0.008***	0.008***
	(0.003)	(0.001)	(0.001)
Female x Children x log(CZ pop)	-0.015***	-0.007***	-0.007***
	(0.003)	(0.001)	(0.001)
Abilene → NY No Child	0.063	0.075	0.082
Abilene → NY Children	0.024	0.078	0.087
L. Supply Controls	✓	✓	✓
Commute Controls	✓	✓	✓
Year FE	✓	✓	✓
Education FE		✓	✓
Occupation FE			✓
R-sq.	0.390	0.462	0.530
Observations	13250889	13250889	13250889

Table A7: Gender Gap and City Size, the Effect of Marital Status

	(1)	(2)	(3)
Female	-0.192***	-0.375***	-0.380***
	(0.027)	(0.026)	(0.023)
log(CZ pop)	0.088***	0.055***	0.044***
	(0.007)	(0.004)	(0.004)
Female × log(CZ pop)	0.010***	0.016***	0.017***
	(0.002)	(0.002)	(0.001)
Married x log(CZ pop)	0.002	0.004*	0.005***
	(0.004)	(0.002)	(0.001)
Female x Married x log(CZ pop)	-0.007*	-0.006***	-0.003*
	(0.004)	(0.002)	(0.002)
Abilene → NY Singles	0.040	0.068	0.069
Abilene → NY Married	0.020	0.058	0.076
L. Supply Controls	✓	✓	✓
Commute Controls	✓	✓	✓
Year FE	✓	✓	✓
Education FE		✓	✓
Occupation FE			✓
R-sq.	0.400	0.467	0.532
Observations	13250889	13250889	13250889

C Appendix Figures

Figure A1: Gender Earnings Gap by City Size

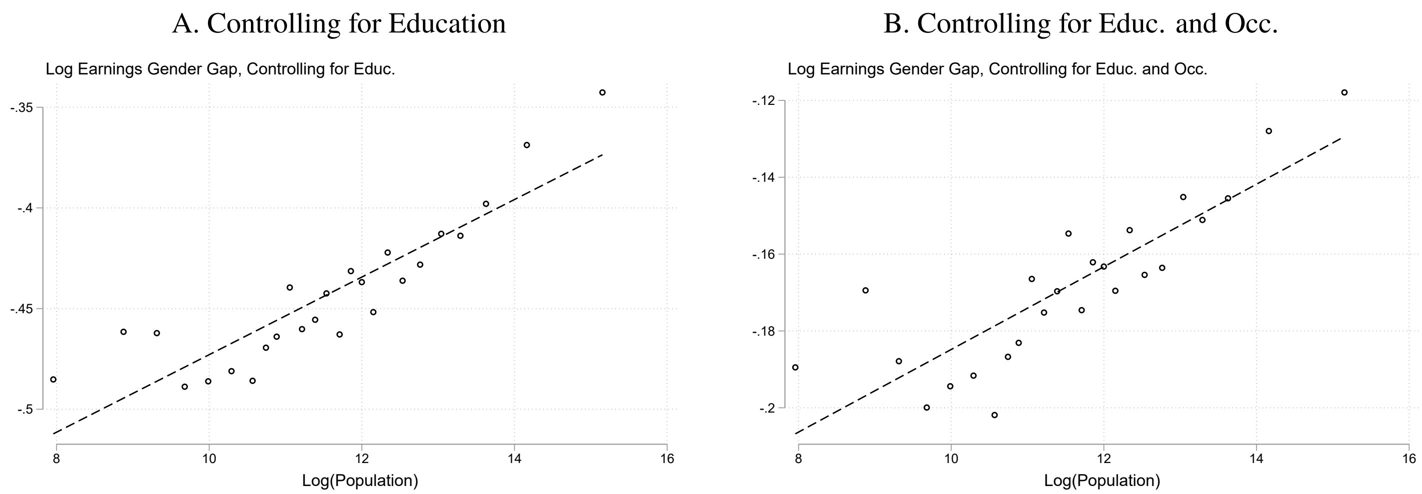


Figure A2: Year-by-Year Coefficients in Primary Sample

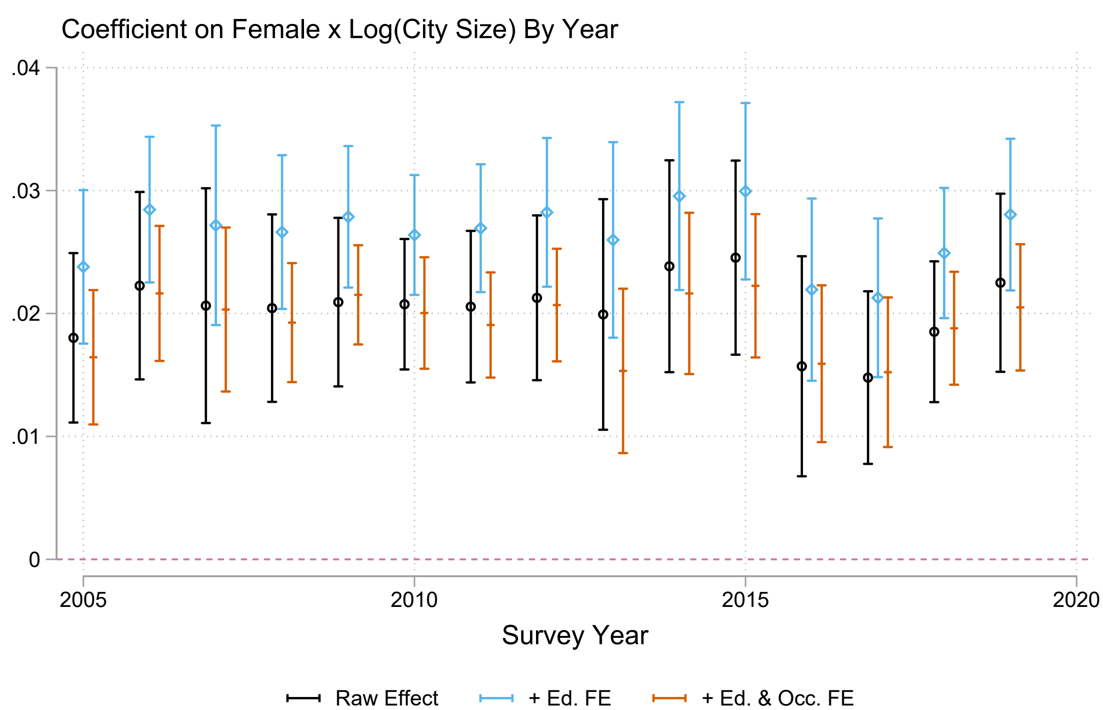


Figure A3: Raw Gender-Specific City-Size Premium Across Historical Periods

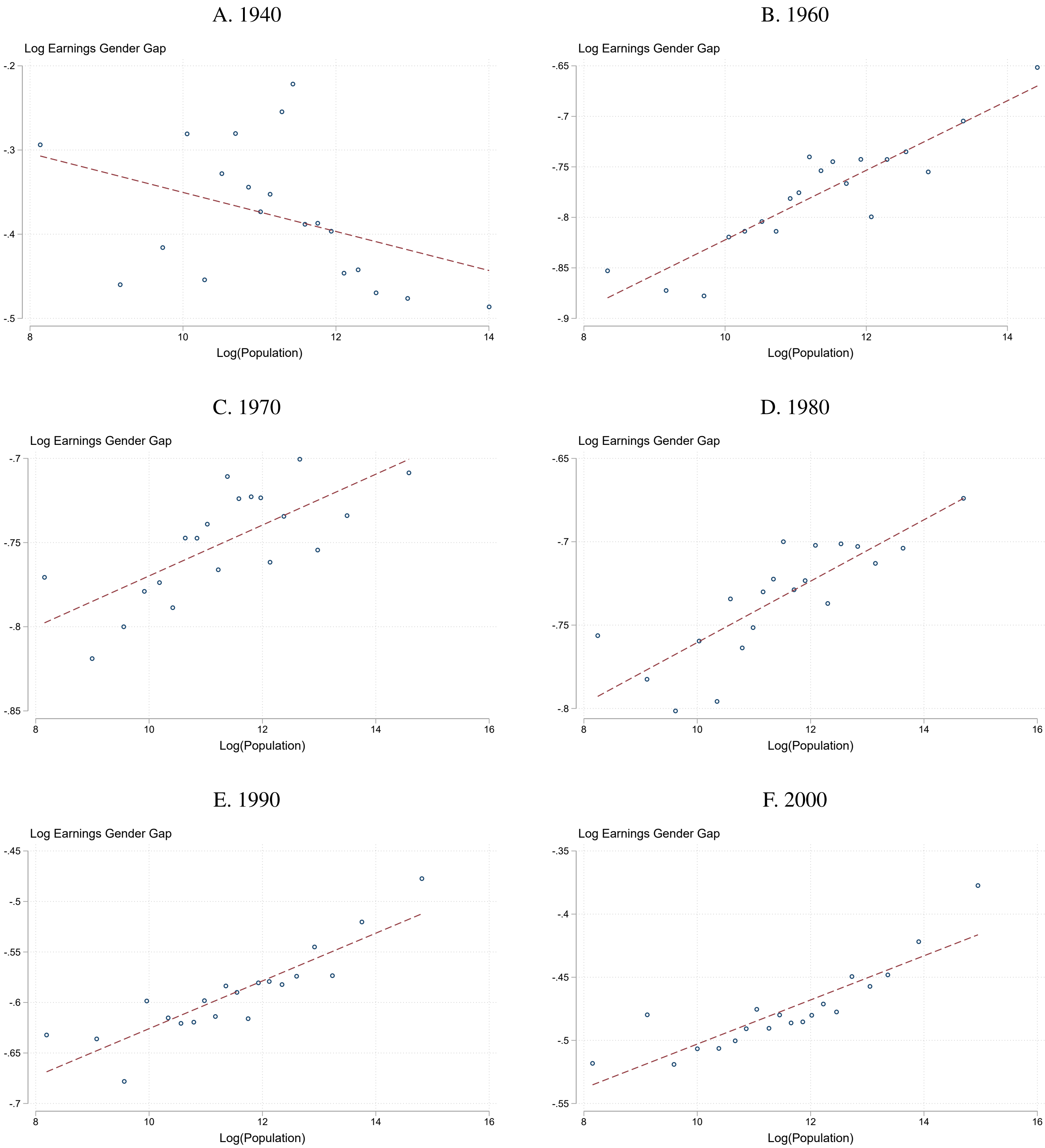
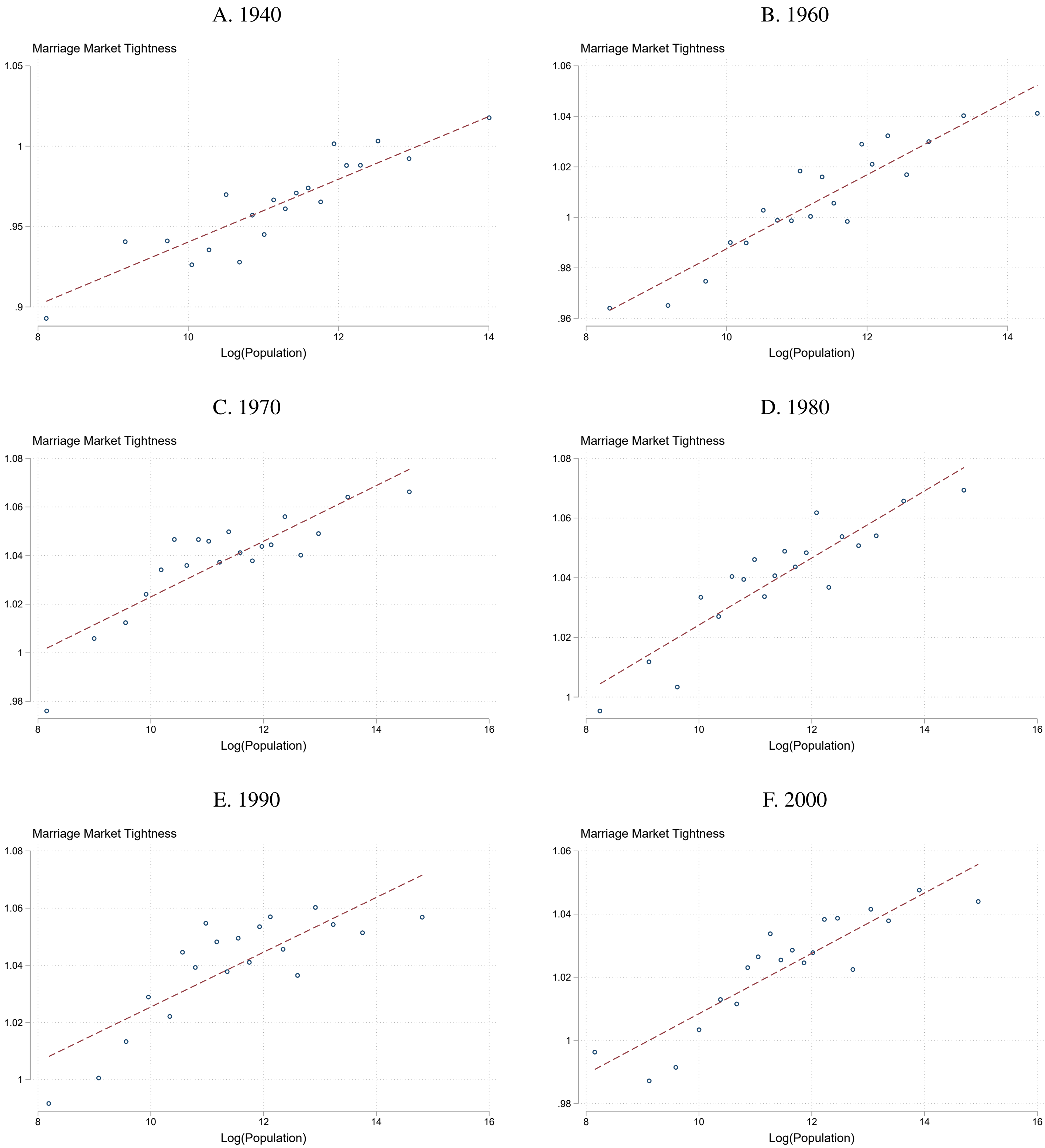


Figure A4: Raw Marriage Market Tightness Across Historical Periods



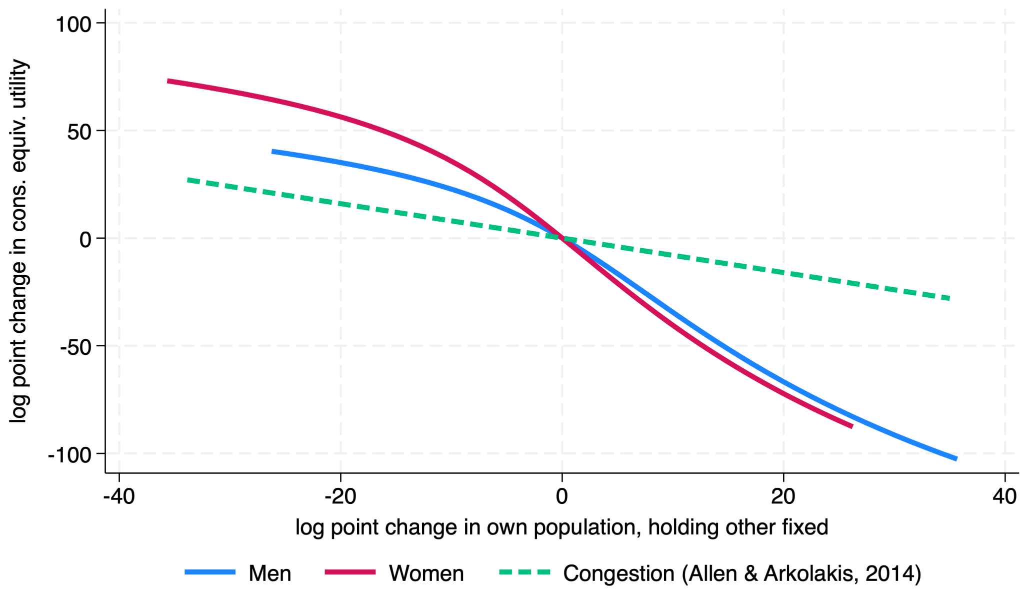


Figure A5: Implied congestion elasticities

D Heterogeneous Effects

D.1 Heterogeneous Gender Earnings Responses By Race

The United States exhibits pronounced racial sorting across space. While we control for demographics in all of our main specifications, our results suggest that gender-specific, city-size earnings premium may also have significant effects on racial earnings inequality. This possibility is further strengthened by the fact that our results are concentrated in the lower parts of the educational distribution.

We investigate whether our results have heterogeneous effects across racial groups by allowing the city-size earnings premium to vary by both gender and race. Table A7 shows the effects separately for Black and non-Black individuals.

We find that the gender-specific, city-size earnings premium is substantially larger for Black women when compared to the earnings premium for non-Black women. The coefficient on the triple interaction term is large and significant in the baseline regression with only demographic controls and in the regression where we condition on educational attainment. Those results suggest the gender-specific earnings premium is between 1.6x to 2 times as large for Black women. In contrast, we find no additional benefit of large labor markets

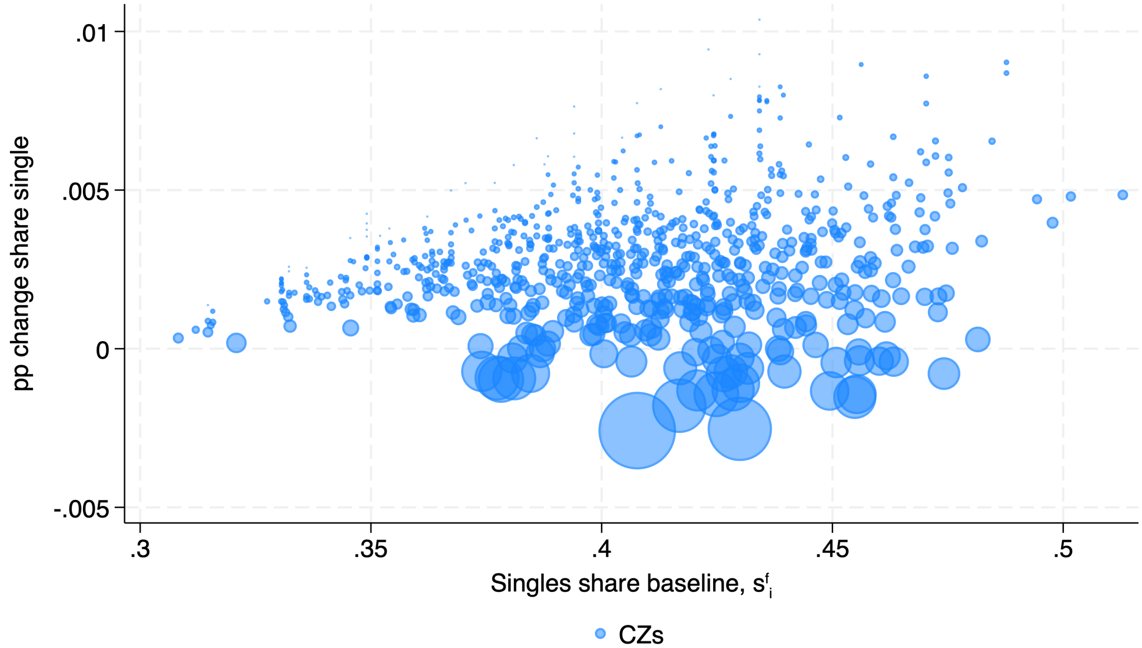


Figure A6: Changes in the marriage market induced by eliminating the female-specific component of the city size premium

for Black men, relative to non-Black men. Finally, we estimate insignificant effects when we control for occupation fixed effects. These findings indicate that Black women experience relatively larger gains in large labor markets, driven by the fact that these markets facilitate sorting into higher-paying occupations.

Table A8 shows the effects for Hispanic individuals. In the least-restrictive specification, we find significant effects on the triple interaction term. However, these results disappear when we condition on educational attainment. This suggests that while Hispanic women see some additional benefit from large labor markets, it is driven by their concentration in the lower parts of the educational distribution. In fact, overall we find that Hispanic women benefit much less from moving to large labor markets because of a large, significant (gender-neutral) Hispanic earnings reduction in larger cities. Much of this is explained by educational and occupational sorting. In our least restrictive specification, we find that while for non-Hispanic women the effect of moving from Abilene to New York is a fall in the gender earnings gap of 7 log points, Hispanic women see a 13 log point earnings fall from moving to large cities. In our most restrictive specification with occupation and education fixed effects, we find that non-Hispanic women see a 8 log point fall in the gender earnings gap, while Hispanic women only see a 2 log point reduction in the earnings gap.

Table A8: Gender Gap and City Size, Heterogeneous Effects by Race

	(1)	(2)	(3)
Female	-0.630***	-0.759***	-0.547***
	(0.047)	(0.038)	(0.033)
log(CZ pop)	0.094***	0.055***	0.047***
	(0.009)	(0.006)	(0.005)
Female × log(CZ pop)	0.015***	0.022***	0.016***
	(0.004)	(0.003)	(0.003)
Black x log(CZ pop)	-0.007	-0.001	0.003
	(0.009)	(0.006)	(0.005)
Female x Black x log(CZ pop)	0.017***	0.014***	0.004
	(0.006)	(0.005)	(0.004)
Abilene → NY Non-Black	0.064	0.090	0.068
Abilene → NY Black	0.102	0.145	0.097
Year FE	✓	✓	✓
Education FE		✓	✓
Occupation FE			✓
R-sq.	0.115	0.218	0.355
Observations	13250889	13250889	13250889

Table A9: Gender Gap and City Size, Heterogeneous Effects by Race

	(1)	(2)	(3)
Female	-0.632***	-0.783***	-0.556***
	(0.045)	(0.039)	(0.032)
log(CZ pop)	0.099***	0.057***	0.048***
	(0.007)	(0.005)	(0.004)
Female × log(CZ pop)	0.017***	0.025***	0.018***
	(0.003)	(0.003)	(0.002)
Hispanic x log(CZ pop)	-0.061***	-0.027***	-0.013***
	(0.005)	(0.004)	(0.004)
Female x Hispanic x log(CZ pop)	0.014**	0.005	0.000
	(0.006)	(0.005)	(0.003)
Abilene → NY Non-Hispanic	0.071	0.104	0.076
Abilene → NY Hispanic	-0.126	0.013	0.023
Year FE	✓	✓	✓
Education FE		✓	✓
Occupation FE			✓
R-sq.	0.114	0.217	0.354
Observations	13250889	13250889	13250889

Taken together our results suggest that Black women have especially large benefits from large labor markets. These larger benefits relative to non-Black women are driven by improved sorting to high-earning occupations. In contrast, we estimate fewer differences for non-Hispanic and Hispanic women, once we condition on differences in educational attainment. In fact, Hispanic women overall benefit much less from large labor markets, in part due to a gender-neutral earnings penalty for Hispanics in larger cities. Overall, we find significant heterogeneity in earnings gaps across both gender and racial groups across space.