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Residential segregation and black-white intermarriage.

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Abstract

I use 1980, 1990, and 2000 Census data to show that greater residential segregation is associated with a lower probability that black men, black women, white men, and white women are in black-white marriages. This negative relationship grows stronger among whites and remains constant among blacks when I control for local marriage market characteristics. Plausible explanations for the results are discussed and investigated.

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1. Introduction

Black-white intermarriage is a rare but increasingly common event. Among young and married blacks and whites, 0.6 percent were in black-white marriages in 1980, but by 2000 this had increased to 2.1 percent.[1] Over the same time period, black-white residential segregation in the typical metropolitan statistical area (MSA) fell by about 13 percent. The goal of this paper is to more carefully document the correlation between segregation and black-white intermarriage.

This correlation may exist for several reasons. A two-sided marriage market consists of men and women of various types looking for a spouse. Types can be defined along any number of dimensions including race, education, and wages. In the context of Becker's (1973) frictionless matching model, racial marital sorting arises from the relative supply of types, preferences for traits which are correlated with race, and racial preferences in marriage partners. Segregation could influence or reflect each of these three sources of marital sorting.

For example, segregation could be related to the demographic composition of blacks and whites in a market. Segregation is also known to widen the black-white parity gap in education and labor markets (e.g., Card and Kreuger, 1996; Cutler and Glaeser 1997; Guryan, 2004; and Card and Rothstein, 2007[2].) A wider gap would negatively affect black-white intermarriage since education and earnings potential are valued in marriage markets. Finally, it is plausible that MSAs with less social tolerance or preference for black-white marriages are also more segregated.

In addition to the factors which influence marital sorting in a frictionless model, spatial mismatch may be able to explain marital sorting patterns once we include search frictions. Residential segregation may be a root cause of spatial mismatch if mutually acceptable singles of different races meet each other less often in more segregated cities. This physical separation should lead to fewer marriages across racial lines.

The data reveal several interesting patterns. First, I demonstrate that with a variety of specifications and controls that black men, black women, white men, and white women in more segregated cities are less likely to be in a black-white marriage.

More interestingly, controlling for demographics and economic differences between blacks and whites actually *strengthens* the negative relationship between segregation and intermarriage among whites. This result may be surprising since more segregation is associated with a wider black-white parity gap, so regressing intermarriage probabilities on a measure of segregation without economic controls should overstate segregation's effect. This is the classic *omitted variables bias.* However, there is also a *second-order effect* at play. Holding the black-white parity gap constant at a wide gap, an increase in meeting probabilities implied by lower segregation would have a small effect on intermarriage probabilities since there are so few mutually acceptable pairings of blacks and whites to begin with. But if the parity gap is narrow, the set of mutually acceptable pairings increases, and a change in segregation would have a larger impact on intermarriage. Since larger parity gaps are

[1] These figures are calculated using census data. The sample is noninstitutional, nonhispanic women aged 20-29 and men aged 22-31 living in metropolitan areas with at least 100,000 people, at least three percent of whom are nonhispanic black.

[2] This literature can be traced back to Kain (1968).

associated with more segregation, controlling for demographic and economic market characteristics may actually increase the effect of segregation. The results in this paper suggest that the second-order effect dominates the omitted variables bias effect.

To investigate the plausibility of this claim further, I attempt to identify those with a greater propensity to marry across racial lines. If the second-order effect is important, segregation should have a greater impact on intermarriage probabilities for these people. I follow Rosenfeld and Kim (2008) and use whether a person is born out of state as a proxy for a willingness to marry interracially. I find that segregation matters more for those born out of state, a finding which supports the importance of the second-order effect.

This paper does not attempt to identify any causal effect of segregation on intermarriage. While there remains a significant negative relationship between segregation and intermarriage after controlling for demographic, economic, and unobserved characteristics of the marriage market, this residual cannot be interpreted as evidence for spatial mismatch. An alternative explanation is that more segregated MSAs reflect a preference for racial separation in all spheres of life, including marriage. To impart a causal interpretation on the estimates one must identify an exogenous source of variation in segregation. I leave this task to future research.

Ever since Becker's (1973) seminal contribution economists have been interested in who marries whom and why. Much of the empirical literature in economics has focused on sorting on human capital (e.g. Fernandez, Guner, and Knowles, 2005; Zhang and Liu, 2003; Pencavel, 1998; among many others). More recent papers have considered sorting on religion (Bisin, Topa, and Verdier 2004), ethnicity (Bisin and Verdier, 2000; Meng and Gregory, 2005), and race (Wong, 2003 and Fryer, 2007). Neither Wong nor Fryer analyze the relationship between segregation and racial intermarriage, however.

2. Empirical Methodology and Data

I use variation in residential segregation across MSAs and over time to establish the link between segregation and racial intermarriage.[3] A basic assumption is that each MSA-year is an independent marriage market. To focus the analysis, I concentrate on marriages between whites and blacks. I leave the analysis of racial intermarriage between and among other groups to future research.

The individual level data come from the 1980, 1990, and 2000 censuses. One difficulty with this is that the U.S. Census Bureau first allowed respondents to select more than one race category in the 2000 census. Rather than using a "one-drop" rule to assign respondents to a single race in 2000, I use census data from the Integrated Public Use Microsample (IPUMS) (Ruggles, et. al., 2008). The IPUMS assigns a single race to multiple race people based on individual, residential, and geographic characteristics.[4] Specifically, I use the 1% 1980 metro sample, the 1% 1990 metro sample, and the 5% 2000 metro samples from the IPUMS.

[3] I use PMSAs, or primary metropolitan statistical areas, whenever available.

[4] For detail, see http://usa.ipums.org/usa-action/variableDescription.do?mnemonic=RACESING. The IPUMS assigns a single race using methods described in Ingram, et. al. (2003).

Table 1: Summary Statistics for the Isolation Index.

	Mean	Std. Dev.	Minimum	Maximum
Black Isolation Index (I_b)	.57	.18	.07	.86
White Isolation Index (I_w)	.92	.05	.67	.98

The data are from the 1980 1 percent metro, 1990 1 percent metro, and 2000 5 percent metro IPUMS of the census. The sample is MSAs with a total population of 100,000 or more, at least three percent of whom are nonhispanic black.

2.1. Measuring segregation

A number of measures of residential segregation exist (see Massey and Denton, 1988).[5] For this study I use the isolation index, one of the two most commonly used measures. All results are robust to using the dissimilarity index to measure segregation, the other common index. The isolation index depends on racial composition within the neighborhoods that together constitute a MSA. In practice, census tracts proxy for neighborhoods.

Suppose we wish to measure segregation between nonhispanic whites and nonhispanic blacks. The isolation index for whites measures the extent to which whites have contact with blacks as the percentage of the neighborhood composed of whites relative to blacks for the average white person, or

$$I_w = \sum_{i=1}^{n} \left(\frac{white_i}{white} \right) \left(\frac{white_i}{white_i + black_i} \right),$$

where $white_i$ is the population of whites in tract i, $white$ is the population of whites in the MSA, $black_i$ is the population of blacks in tract i, and n is the number of census tracts in the MSA. The isolation index for blacks, I_b, is analogous. The isolation index ranges from zero to one where higher values indicate more segregation.

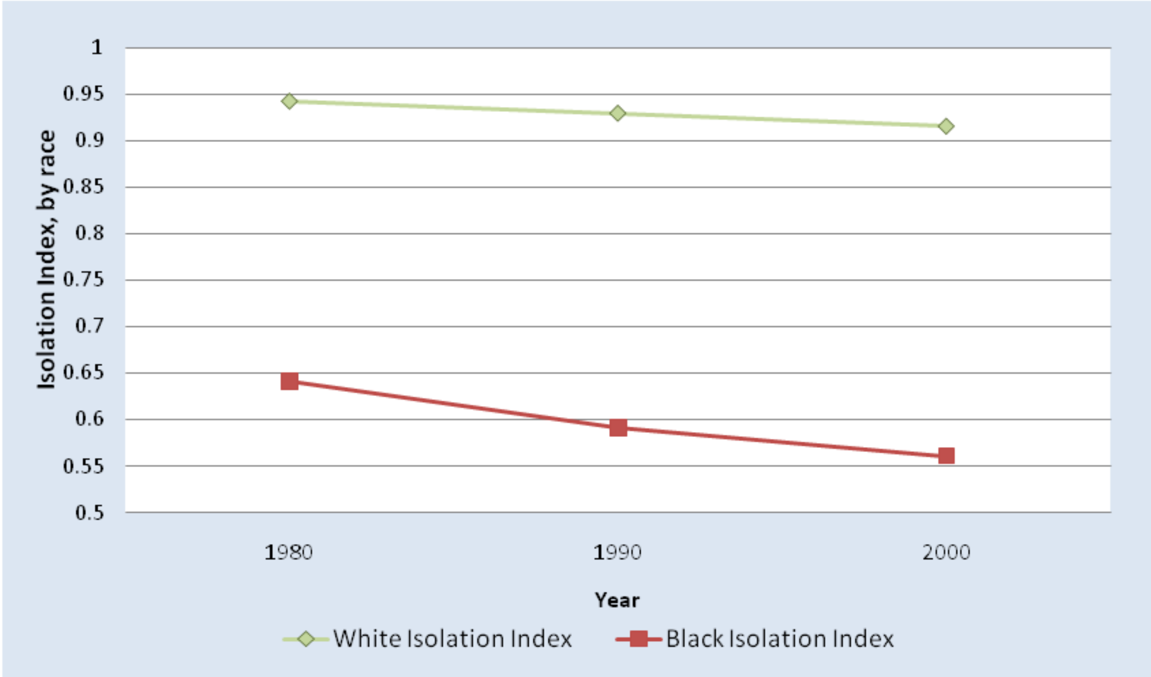
All indices are calculated using summary tape files from the Census.[6] To reduce measurement error, I use only MSAs with at least 100,000 people where at least three percent of the population is nonhispanic black. The regressions are based on 172 cities in 1980, 193 cities in 1990, and 179 cities in 2000.

Table 1 displays summary statistics. The typical white person is more isolated from blacks than the typical black person is from whites, a fact which reflects the higher percentage of whites in the population. Segregation varies considerably between MSA-years; each measure has a large range and sizable standard deviation. Figure 1 displays a time series of the segregation indices. Residential segregation has fallen in each decade since the 1980, a finding consistent with Cutler, Glaeser, and Vigdor (1999).

[5] Massey and Denton identify five dimensions of segregation: evenness, exposure, concentration, centralization, and clustering. They conclude that the last three dimensions are relatively unimportant empirically. After considering 20 different measures of segregation, they identify the dissimilarity index and the isolation index as the most appropriate measures of evenness and exposure, respectively.

[6] For the 2000 Census, which places multiple race people into separate groups, I use only the population counts among those with a single race identifier. I have also experimented with, and found the results robust to, using a "one-drop white" and a "one-drop black" rule to assign a single race to groups who identified themselves with more than one race.

Figure 1: Isolation Index Over Time, by Race.



2.2 Econometric Specification

The propensity to marry across racial lines depends on the market one faces. To capture general equilibrium effects, I run a regression using the entire data set and then use interaction terms to separate out the effects of segregation by race-sex cells. The general specification is

$$\begin{aligned} BWMar_{imt} &= \beta_1 I_{imt} + \beta_2 I_{imt} \times Bl_{imt} + \beta_3 I_{imt} \times Fem_{imt} + \beta_4 I_{imt} \times Fem_{imt} \times Bl_{imt} \\ &\quad + \beta_5 Fem_{imt} + \beta_6 Bl_{imt} + \beta_7 X_{imt} + \eta_{mt} + \varepsilon_{imt}, \end{aligned} \tag{1}$$

where the binary dependent variable equals one if the respondent is in a black-white marriage and I_{imt} is the value of the isolation index for MSA m in year t for individual i. Within a MSA-year, the value of I_{imt} varies only by race. Fem_{imt} and Bl_{imt} are indicator variables for female and black, respectively. X_{imt} is a vector containing the individual, demographic, and economic controls discussed above, η_{mt} is a MSA-year specific effect that captures unobserved heterogeneity across time and space, and ε_{imt} is an error term.

The first four coefficients are the focus of this study. β_1 is the marginal effect of segregation on the probability that a white man is in a black-white marriage. $\beta_1 + \beta_2$ measures the same effect for black men, $\beta_1 + \beta_3$ for white women, and $\beta_1 + \beta_2 + \beta_3 + \beta_4$ for black women. All of these effects are expected to be negative. The signs of the cross-partials (e.g., β_2, which measures the difference in the effect of segregation between white men and black men) are *a priori* indeterminate.

Table 2: Percent in a black-white marriage, by group and year.

Group	1980 Percent	1980 N	1990 Percent	1990 N
Black men (ages 22-31)	1.7%	16,037	1.9%	14,394
Black women (ages 20-29)	0.4%	20,615	0.9%	17,637
White men (ages 22-31)	0.09%	98,971	0.1%	100,768
White women (ages 20-29)	0.3%	101,842	0.3%	95,661

Group	2000 Percent	2000 N	1980-2000 Percent	1980-2000 N
Black men (ages 22-31)	3.6%	73,895	2.5%	104,326
Black women (ages 20-29)	1.0%	95,182	0.8%	133,434
White men (ages 22-31)	0.2%	444,382	0.16%	644,121
White women (ages 20-29)	0.7%	422,606	0.4%	620,109

The data are from the 1980 1 percent metro, 1990 1 percent metro, and 2000 5 percent metro IPUMS of the census. The sample is noninmates within the specified age and racial groups.

2.3 Outcome and explanatory variables

Only black and white women aged 20 to 29 and black and white men aged 22 to 31 are included to minimize bias due to location selection. The age range for men is older since women tend to marry men older than themselves. Institutional inmates are excluded. The propensity for blacks in this group to marry whites, and for whites to marry blacks, is shown in Table 2. This propensity has been increasing over time for each group, but note that black men are the most likely to be in a black-white marriage.

The probability that a person meets a particular type certainly depends on the relative prevalence of types. This is captured by a variety of *demographic controls*. These include the log number of singles in each race-sex cell and the log of the total MSA population to control for congestion externalities. I also include race specific sex ratios (single men divided by single women), as these may influence the incentive to marry across racial lines, as well as the general sex ratio, as this is known to be important in marriage and labor markets more generally (Grossbard-Shechtman, 1993 and Angrist, 2004). Since the sample includes all men and women in the specified age ranges, I include within sample marriage rates by race and sex to ensure that the estimates for the coefficient on segregation are not picking up variation in marriage rates across space and time. Finally, I include the percent black in an MSA and its square.[7] Summary statistics for demographic controls are displayed in Table 3. Black single men are relatively scarce, and marriage rates among black men are lower than among white men.

Marriage is voluntary so it must be acceptable to both sides of a match. Whether an individual is acceptable depends on one's position relative to his or her competitors and the distribution of traits among potential mates. Thus, I include a rich set of individual-level and market level *economic controls* which may mitigate the relationship between segregation and intermarriage. The individual controls include a person's age, log weekly earnings, and

[7] Kalmijn (1993) has shown that the probability blacks marry whites is related to the percent black in a nonlinear way. The unweighted correlation between the dissimilarity index and percent black is -.049.

Table 3: Within sample MSA-level summary statistics.

Variable	Mean	Std. Dev.	Minimum	Maximum
ln(total MSA population)	13.1	1.0	11.6	16.0
Log of the number of single...				
black men	7.9	1.2	3.7	11.7
black women	8.2	1.2	5.0	11.9
white men	9.5	1.1	7.1	12.5
white women	9.5	1.1	7.2	12.5
Sex ratios				
(Black men)/(Black women)	.80	.53	.04	10.00
(White men)/(White women)	1.04	.22	.52	2.46
$\frac{\text{Black men + White men}}{\text{Black women + White women}}$	.96	.19	.53	2.00
Marriage rates...				
black men	.60	.12	.09	.97
black women	.51	.13	.05	.96
white men	.75	.08	.46	.96
white women	.52	.11	.23	.90
Percent black	.15	.10	.02	.51

The data are from the 1980 1 percent metro, 1990 1 percent metro, and 2000 5 percent metro IPUMS of the census. Only the 544 MSAs with a total population of 100,000 or more, at least three percent of whom are nonhispanic black, are included. The population for the log number of singles, sex ratios, and marriage rates is nonhispanic women aged 20-29 and men 22-31 who are not inmates. The percent black is calculated among all people living in the MSA.

dummy variables for whether an individual has at most a high school education and whether or not he or she is unemployed.[8]

The market-level economic controls include wage distributions, unemployment rates, and average educational attainment for each sex-race cell within a MSA-year. These statistics are computed using samples with an age range wider than the estimation sample's range because data constraints make it difficult to construct accurate estimates within finely defined cells. Thus, one should interpret these statistics as the permanent components of the distributions of education and labor market outcomes.

Specifically, the average and standard deviation of log weekly earnings (in 2000 dollars) are estimated among persons aged 16 to 64 who were not in school and who worked at least one week in the calendar year prior to the census year. Earnings are defined as wage and salary income plus self-employment income. To capture the permanent component of the educational attainment distribution, I use the percentage of the population over 30 that has at most a high school education. Most people have completed their education by this age. Unemployment rates are calculated directly from the Census among persons aged 16 to 64.[9]

[8]Including a full set of dummies for age does not change the results, so I simply include age as a variable for simplicity. For similar reasons, I use a simple measure of education rather than a full set of education dummies. Weekly earnings are defined as wage and salary income plus self-employment income, measured in 2000 dollars. I use predicted wages in cases where this variable is missing.

[9]I have also experimented with included measures of "idleness," as in Cutler and Glaeser (1997). A person is considered idle if he or she is neither in school nor employed. Including idleness does not affect

Table 4: Unweighted labor market and educational distributions within MSA-years.

Group	Mean	Std. Dev.	Minimum	Maximum
A. Average of log weekly wages (ages 16-64)				
Black men	6.15	.19	5.43	6.78
White men	6.54	.16	6.15	7.33
Black women	5.84	.21	5.14	6.61
White women	5.94	.17	5.54	6.64
B. Standard deviation of log weekly wages (ages 16-64)				
Black men	.79	.13	.42	1.48
White men	.77	.07	.58	1.13
Black women	.80	.13	.32	1.51
White women	.78	.06	.58	1.03
C. Unemployment rates (ages 16-64)				
Black men	.12	.06	0	.43
White men	.05	.02	.01	.12
Black women	.12	.05	0	.44
White women	.05	.02	.01	.14
D. Percent of population over 30 with at most a HS education				
Black men	.58	.16	.14	1
White men	.48	.13	.22	.89
Black women	.57	.17	.05	1
White women	.50	.15	.19	.86

See Table 3 for the source and selection of MSAs. For each MSA-year, the sample for wage data is 16-64 year-olds who are not in school and worked at least one week in the year prior to the census year. Unemployment rates are calculated among all 16-64 year-olds, and educational attainment is calculated among persons 30 years old or more.

Summary statistics for economic controls are shown in Table 4. Within sexes, whites have higher average wages and more education while blacks have higher variance in wages and experience higher unemployment rates. Within races, men have higher average wages but the other outcomes are roughly the same.

3. Results

Table 5 reports ordinary least squares estimates of selected parameters from equation (1).[10] Full results are displayed in Table 6 in the Appendix. To ease interpretation, I include in Table 5 the estimates of the effects of segregation for each race-sex cell. The standard errors in all of the regressions are corrected for heteroskedasticity and are clustered at the MSA-year level.

The first column reports estimates from a regression that pools all the data without any controls to form baseline estimates. The average marginal effect of segregation on the probability a white man is in a black-white marriage is negative and statistically significant, as expected. A 0.1 increase in the isolation index is associated with a .0048 percentage point increase in probability a white man is in a black-white marriage. This is an economically significant number since intermarriage rates among white men are so low (see Table 2). The estimated effect for black men is twice as large. For women, the effect of segregation is smaller than it is for men of the same race. While the column one estimate for white women suggests a positive effect of segregation on the intermarriage probability, this result seems pathological since adding additional controls in columns 2 through 5 changes the sign of the estimated effect.

The column 1 estimates could be picking up variation in racial and demographic composition by MSA-year. To account for this, column 2 includes the demographic controls. This has no effect on black men or black women, but it changes the sign of the effect for white women and strengthens the magnitude of the effect for men. This implies that the more isolated whites are from blacks, the more favorable the demographics are for intermarriage. Alternatively, including demographic controls may have a *second-order effect* similar to the one described in the introduction.

Interestingly, including economic controls has no effect on the estimates. This result could be interpreted in at least two ways. First, one may expect that excluding economic controls causes an omitted variables bias that overstates the impact of segregation. In this case, one explanation is that variation in economic and educational outcomes is already captured by variation in demographics. Alternatively, the *second-order effect* and the *omitted variables bias* could be canceling each other out.

Year fixed effects are included in the column 4 estimates. Year dummies capture national trends in the tastes for black-white intermarriage that may also be correlated with segregation: evolving race relations, other social norms, changes in fertility control technology, divorce laws, etc. However, there is no change in the estimated coefficients, suggesting that

the results.

[10] I have checked that the results are indeed robust to using a probit model.

Table 5: Selected Ordinary Least Squares Estimates of the Effect of Segregation on the Probability of Being in a Black-White Marriage.

Dependent variable = 1 if in a black-white marriage						
Avg. Marginal Effect	(1)	(2)	(3)	(4)	(5)	(6)
of Isolation for:						
White Men ($\hat{\beta}_1$)	-.048	-.076	-.075	-.075	-.102	-.067
	(.003)	(.006)	(.006)	(.006)	(.008)	(.003)
Black Men ($\hat{\beta}_1 + \hat{\beta}_2$)	-.098	-.094	-.093	-.093	-.094	
	(.006)	(.006)	(.006)	(.006)	(.006)	
White Women ($\hat{\beta}_1 + \hat{\beta}_3$)	.011	-.017	-.016	-.016	-.044	
	(.004)	(.005)	(.006)	(.006)	(.007)	
Black Women ($\Sigma_{i=1}^4 \hat{\beta}_i$)	-.038	-.034	-.034	-.034	-.036	
	(.003)	(.003)	(.003)	(.003)	(.003)	
Other Estimates:						
Isolation×Black ($\hat{\beta}_2$)	-.050	-.018	-.018	-.018	.008	
	(.006)	(.007)	(.007)	(.007)	(.007)	
Isolation×Female ($\hat{\beta}_3$)	.059	.059	.059	.059	.058	
	(.004)	(.004)	(.004)	(.004)	(.004)	
Isolation×Female×Black	.001	.0004	.0002	.0002	-.00001	
($\hat{\beta}_4$)	(.002)	(.002)	(.002)	(.002)	(.002)	
Isolation×Black +Isolation×	-.049	-.017	-.018	-.018	.008	
×Female×Black ($\hat{\beta}_2 + \hat{\beta}_4$)	(.005)	(.005)	(.006)	(.006)	(.007)	
Isolation×Born out of state						-.029
($\hat{\beta}_1 + \hat{\alpha}_2$)						(.002)
Demographic Controls[a]	No	Yes	Yes	Yes	Yes	Yes
Economic Controls[a]	No	No	Yes	Yes	Yes	Yes
Fixed Effects[b]	No	No	No	Y	Y, M	Y, M
N			1,501,990			

Observations are weighted using census person weights. Robust standard errors are clustered at the MSA-year level and are reported in parentheses. See the section on data and empirical methodology for precise definitions of the independent variables. [a]Demographic and economic controls are described in the text, and parameter estimates are provided in Table 6 of the Appendix. [b]Y is year fixed effects and M is MSA fixed effects. *Sample:* Nonhispanic blacks and whites who are not inmates and who live in MSAs as specified in the text. Men are 22-31 years old while women are 20-29. *Source:* IPUMS; 1980 1 percent metro, 1990 1 percent metro, and 2000 5 percent metro samples.

national trends in tastes are already captured by changes in the demographic and economic aspects of a marriage market.

Interestingly, the differential effect of segregation between races (within sexes) that is present in all previous specifications disappears once we include MSA-level fixed effects. MSA effects control for unobserved differences in local marriage markets that vary by place but are constant over time. For men, the absence of a differential effect can be seen in the column 5 estimates for $\hat{\beta}_2$; for women, this can be seen in the estimate for $(\hat{\beta}_2 + \hat{\beta}_4)$. Neither estimate is significantly different from zero. The estimated average marginal effects of segregation for black men and black women remain unchanged, but the magnitude of the effect among whites increases.

For men, the estimates imply that a 0.1 point increase in the isolation index is associated with a 0.01 percentage point decrease in the probability a man is in a black-white marriage. For women the same increase in the isolation index is associated with a 0.004 percentage point decrease in the probability of intermarriage. The differential effect by sex is statistically significant for whites ($\hat{\beta}_3 = .058$; $s.e. = .004$) as well as for blacks ($\hat{\beta}_3 + \hat{\beta}_4 = .058$; $s.e. = .006$). The larger effect for men may be due to the traditional role men have in seeking out women rather than the other way around.

At this point the results seem to indicate that segregation matters more when both the supply and demand sides of the market are more conducive to black-white intermarriage, at least among whites. To investigate this idea further I control for whether a person is born out of state. People who live in a state different from their state of birth are more likely to be in mixed marriages (Rosenfeld and Kim, 2005). Rosenfeld and Kim argue that this variable proxies for a willingness to break taboos and part with community norms. Thus, if segregation matters more for people with a greater inclination towards intermarriage, the magnitude of the estimated marginal effect of segregation should be larger among those born out of state.

To test this hypothesis, I run a regression of the form

$$BWMar_{imt} = \beta_1 I_{imt} + \alpha_2 I_{imt} \times Outstate_{imt} + \alpha_3 Outstate_{imt} + \alpha_4 X_{imt} + \eta_{mt} + \varepsilon_{imt}, \quad (2)$$

where $Outstate_{imt}$ is an indicator variable for whether a person is born out of state. The other variables are defined as in equation (1). If the hypothesis is correct, then the coefficient on the interaction term α_2 will be negative and significant.

As shown in column 6 of Table 5, the data bear out this hypothesis. The point estimate for people born in state is $-.035$ while for people born out of state it is $-.068$. Both point estimates are statistically significant, and the difference between these is significant at the one percent level.[11]

4. Conclusion

In addition to documenting some interesting summary statistics by MSA-year, this paper reaches several conclusions. First, there is a significant negative correlation between segregation and the probability a black or white individual is in a black-white marriage. This

[11] This result is robust to further separating out the effect by sex or race.

negative correlation is present for black men, black women, white men, and white women. In fact, once we control for demographic, economic, year, and MSA effects, there is no within-sex difference in the effect of segregation on intermarriage probabilities by race. However, segregation has a stronger negative association with intermarriage probabilities among men.

Another notable finding is that it seems segregation as a larger effect when there exist more mutually acceptable matches between blacks and whites. This hypothesis is supported by two observations. First, the magnitude of the estimated effect of segregation on intermarriage probabilities increases among whites as more controls are introduced, suggesting an important *second order effect.* Second, segregation has a larger estimated effect among individuals who have a greater average inclination towards interracial marriage, as proxied for by whether or not they are born out of state.

To be clear, this paper does not pretend to lend a causal interpretation to the results. The significant negative correlation which remains after introducing a slew of controls is consistent with the presence of spatial mismatch (i.e., that mutually acceptable singles of different races have a more difficult time meeting in more segregated MSAs). However, an alternative explanation is that segregation is a proxy for the tastes for intermarriage. We cannot impart a causal interpretation on these results without identifying an exogenous source of variation. Nevertheless, the results are very intriguing and suggest a fruitful area for future research. Other avenues for future research include investigating how segregation affects intermarriage probabilities between and among other racial groups, and incorporating the new 2010 Census data when they become available.

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5. Appendix Table

Table 6: Complete Ordinary Least Squares Estimates of the Effect of Segregation on the Probability of Being in a Black-White Marriage.

Dependent variable = 1 if in a black-white marriage	(1)	(2)	(3)	(4)	(5)	(6)
Isolation ($\hat{\beta}_1$)	-.048	-.076	-.075	-.075	-.102	-.035
	(.003)	(.005)	(.006)	(.006)	(.008)	(.002)
Isolation×Black ($\hat{\beta}_2$)	-.050	-.018	-.018	-.018	.008	
	(.006)	(.006)	(.007)	(.007)	(.007)	
Isolation×Female ($\hat{\beta}_3$)	.060	.059	.059	.059	.058	
	(.004)	(.004)	(.004)	(.004)	(.004)	
Isolation×Female×Black ($\hat{\beta}_4$)	.0006	.0004	.0002	.0002	-.000	
	(.002)	(.002)	(.002)	(.002)	.002	
Female ($\hat{\beta}_5$)	-.052	-.052	-.050	-.050	-.050	
	(.004)	(.004)	(.004)	(.004)	(.004)	
Black ($\hat{\beta}_6$)	.037	.010	.011	.011	-.013	
	(.004)	(.004)	(.005)	(.005)	(.006)	
Isolation×Outstate ($\hat{\alpha}_2$)						-.033
						(.002)
Outstate ($\hat{\alpha}_3$)						.031
						(.002)
Demographic Controls						
ln(single black men)		.002	.002	.0008	.0007	-.006
		(.002)	(.002)	(.002)	(.0001)	(.001)
ln(single white men)		.0008	.001	.004	-.008	-.009
		(.005)	(.005)	(.005)	(.006)	(.006)
ln(single black women)		-.006	-.005	-.003	.002	.004
		(.002)	(.002)	(.002)	(.006)	(.006)
ln(single white women)		-.001	-.002	-.005	.006	.007
		(.005)	(.005)	(.004)	(.005)	(.005)
ln(MSA population)		.006	.005	.003	.0001	-.0001
		(.001)	(.001)	(.001)	(.002)	(.002)
Percent black		.032	.025	.021	.033	.037
		(.012)	(.013)	(.014)	(.023)	(.022)
Percent black squared		-.065	-.056	-.051	-.052	-.065
		(.019)	(.021)	(.020)	(.040)	(.039)
Black male marriage rate		.012	.014	.009	-.023	-.024
		(.004)	(.005)	(.004)	(.009)	(.009)
Black female marriage rate		-.015	-.013	-.009	.028	.024
		(.004)	(.004)	(.004)	(.022)	(.021)
White male marriage rate		.008	.007	.006	-.013	-.014
		(.004)	(.004)	(.004)	(.001)	(.018)
White female marriage rate		-.013	-.012	-.010	-.007	-.006
		(.002)	(.004)	(.004)	(.006)	(.006)
Sex ratio		-.004	-.004	-.002	-.009	-.009
		(.005)	(.005)	(.004)	(.005)	(.005)
Black sex ratio		.0000	.0001	-.0001	-.0000	.0002
		(.0004)	(.0005)	(.0005)	(.001)	(.001)
White sex ratio		.005	.004	.001	-.009	.012
		(.006)	(.006)	(.006)	(.005)	(.007)
Fixed Effects	None	None	None	Y	Y, M	Y, M

Continued on Next Page...

...Continuation of Table 6						
	(1)	(2)	(3)	(4)	(5)	(6)
Economic controls[a]						
HS education or less			-.0001	-.0001	-.0000	-.0001
			(.0002)	(.0002)	(.0002)	(.0002)
Unemployed			-.002	-.002	-.002	-.002
			(.0003)	(.0003)	(.0003)	(.0003)
Age			.0004	.0004	.0004	.0004
			(.0000)	(.0000)	(.0000)	(.0000)
Log weekly wage			.0008	.0008	.0007	.0004
			(.0001)	(.0001)	(.0001)	(.0001)
Average Log Weekly Earnings Among						
Black men			.003	.001	-.002	-.0003
			(.005)	(.002)	(.002)	(.002)
Black women			-.0002	.001	-.000	-.0003
			(.002)	(.002)	(.002)	(.002)
White men			-.007	-.007	.001	.0007
			(.002)	(.002)	(.003)	(.003)
White women			.005	.003	-.003	-.003
			(.003)	(.003)	(.004)	(.004)
Standard Deviation of Log Weekly Earnings Among						
Black men			.002	-.0002	-.000	-.002
			(.002)	(.002)	(.002)	(.002)
Black women			.0003	-.001	-.000	-.0003
			(.002)	(.002)	(.001)	(.002)
White men			-.010	-.010	-.010	.011
			(.004)	(.004)	(.004)	(.004)
White women			.012	.013	.001	.0006
			(.004)	(.004)	(.004)	(.004)
Percent With at Most a High School Education Among						
Black men			-.006	-.007	-.005	-.005
			(.003)	(.003)	(.003)	(.003)
Black women			.005	.004	-.0003	-.0004
			(.003)	(.002)	(.003)	(.001)
White men			.004	-.006	-.005	-.004
			(.005)	(.006)	(.007)	(.007)
White women			-.0009	.005	.000	-.002
			(.005)	(.006)	(.008)	(.007)
Unemployment Rate Among						
Black men			.004	.009	.005	.005
			(.004)	(.004)	(.004)	(.004)
Black women			-.001	.001	-.003	-.004
			(.005)	(.005)	(.005)	(.005)
White men			.006	.012	.004	.008
			(.012)	(.016)	(.015)	(.014)
White women			-.006	-.008	.0002	.003
			(.020)	(.021)	(.019)	(.019)
Fixed Effects	None	None	None	Y	Y, M	Y, M

[a] F-tests indicate that we cannot rule out the possibility that the standard deviation of log weekly earnings has an important effect on the probability of being in a black-white marriage in columns 3-6. The other sets of economic controls appear to matter less, depending on the specification. See Table 5 for additional notes, sample, and source.