Research Note: Reconsidering How Puerto Rico Became White

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Comments Welcome

Abstract In this brief research note, we attempt to account for the rapid growth of the white population share in Puerto Rico from 1910-1920 using demographic population projection methods. We replicate and extend results reported by Loveman and Muniz (2007). Consistent with their results, we find that racial differentials in mortality and fertility account for only a small portion of the rise in the white share. Moving beyond their analysis, we modify the standard (female-dominant cohort-component) method to address racial mixing, incorporating elements of the two-sex model in Montgomery (2009). Given the observed social "rules" of racial classification, we find that racial mixing caused a modest *decrease* in the white share. The net effect of racial differentials and racial mixing is positive but very small. This suggests that the growth of the white share occurred almost entirely through racial switching. Our projections suggest that that switching probabilities were age-invariant, with approximately 30% of mulattos switching to white within every age group.

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1 Introduction

In Puerto Rico between 1910 and 1920, the white share of the population grew dramatically from 65% to 74%. Using standard demographic methods for population projection, Loveman and Muniz (2007) find that racial differentials in mortality, fertility, and immigration account for only 1 point of this 9-point increase. They suggest that much of the residual is due to changes in the racial classification practices, in particular the classification of children based on parents' races. However, because the standard (female-dominant cohort-component) methodology cannot address racial mixing (and the subsequent classification of children from these unions), they provide no quantitive assessment of this effect.

Montgomery (2009) has recently re-examined the Puerto Rican case using a two-sex model that explicit incorporates racial mixing. Like Loveman and Muniz (2007), he suggests that differential reproduction played some role in the growth of the white share. But in contrast to their analysis, he argues that racial switching was more important than racial classification. Indeed, he finds that, given the observed social "rules" of racial classification, racial mixing actually promoted the growth of the mulatto share rather than the white share. Unfortunately, the discrete-generations specification of the two-sex model precludes any precise quantitative assessment of short-run population dynamics.

In this brief research note, we replicate and extend the population projections reported in Loveman and Muniz (2007). Our goals are both methodological and empirical. Methodologically, we modify the standard projection methodology to address racial mixing, thus providing a bridge between the standard one-sex age-structured model (see, e.g., Preston et al 2001) and the two-sex discrete-generations model developed in Montgomery (2009). This enables us to assess quantitatively the short-run effect of racial mixing. Empirically, we find that racial differentials in mortality and fertility account for about 1.5 points of the increase in the white share. From this baseline, we find that racial mixing caused the white share to decline by almost 1 percentage point. Thus, the net effect of racial differentials and racial mixing is a mere 0.5-point increase in the white share.

Unable to account for the rise in the white share through these demographic processess, we suspect that this increase occurred almost entirely through racial switching. Although racial switching is a residual argument, our results may provide some interesting insight into the nature of switching. In our projection for the mulatto population, the enumerated-to-projected ratio is roughly constant across age groups. If we assume that each mulatto (regardless of age) had a 30% chance of becoming white upon enumeration in the 1920 census, the remaining projection errors appear unrelated to race.

2 Population projections

This section reports a series of population projections. Throughout, we begin with the enumerated population for 1910, project the population ahead to 1920, and then compare this projected population to the enumerated population for 1920. Our initial projection ignores race entirely in order to assess potential sources of projection errors. We then partition the population by race, attempting to replicate the result reported by Loveman and Muniz (2007). These projections follow the standard female-dominant cohort component methodology (see, e.g., Preston et al 2001, p 119 ff). Finally, we modify the standard method in order to address racial mixing.[1]

2.1 Ignoring race

We begin by ignoring race entirely. Our projection is thus based on (sex-specific but race-neutral) estimates of mortality, fertility, and migration for the entire population. The results are shown in Figure 1, which compares the projected population (gray bars) to the enumerated population (white bars) by sex and age class.

Note that our total projected population (1,269,821) is somewhat smaller than the total enumerated population (1,279,275). Loveman and Muniz (2007, online appendix, p 9) also report a shortfall in their total projected population (1,263,954). In response, they adjust fertility and mortality rates (multiplying age-specific fertility rates by 1.1408 and age-specific mortality rates by 1.11596) to reconcile projected and enumerated population size. We do not follow this strategy for two reasons. First, our shortfall is somewhat smaller than theirs.[2] Second, it is apparent from Figure 1 that the difference between projected and enumerated population levels is especially large for only a few age groups. Indeed, more than 100% of the shortfall can be accounted for by the unexpected number of 5-9 year olds enumerated in 1920. This suggests that our projection errors are more likely due to period effects (e.g., a temporary increase in fertility rates between 1910 and 1915) than uniform underestimation of fertility and mortality rates across all age groups.

2.2 Racial differentials

We now partition the population by race. Assuming no racial differentials in mortality and fertility, we obtain the projection reported in row (a) of Table 1.[3] Essentially,

[1] The enumerated populations for 1910 and 1920 are reported by sex, race, and age group in the Appendix. Race-neutral and race-specific life tables, fertility schedules, immigration estimates can be found in Loveman and Muniz (2007, online appendix).

[2] Because we are using the same data and method as Loveman and Muniz (2007), the source of this (relatively minor) discrepancy between projections is unclear. All of our computations are made using Matlab, and our m-file is available upon request.

[3] This projection does incorporate race-specific net migration, though net migration was negligible for all races over this decade (Loveman and Muniz 2007, online appendix, Table S13).

this is the same projection reported in the preceding subsection, but with results now broken out by race rather than sex and age group. Note that the projected racial shares for 1920 are very close to the enumerated racial shares for 1910. Thus, we have not yet accounted for the 9 percentage-point increase in the enumerated white share over the decade.

Inspection of the race-specific life tables and fertility schedules (see Loveman and Muniz 2007, online appendix) reveals that white women had lower mortality and hence a higher net reproduction rate than mulatto women. This suggests that differential reproduction might explain some of the increase in the white share (cf Montgomery 2009). To assess the importance of this effect, we computed the population projection reported in row (b) of Table 1. Comparing row (b) to row (a), it appears that about 1.5 percentage points can be attributed to differential mortality and fertility. Our results are thus quite similar to those obtained by Loveman and Muniz (2007), who find that racial differentials would have caused the white share to rise from 65.0% to 66.4% (see their Table 1, rows B and C). Thus, while racial differentials may account for some of rise in the white share, much of this rise – roughly another 7.5 percentage-points – remains to be explained.

2.3 Racial mixing

Potentially, some of this gap might be accounted for by racial mixing and the subsequent classification of children based on their parents' races. Indeed, Loveman and Muniz (2007) show that children of mulatto-white couples were somewhat more likely to be classified as white in 1920 than in 1910 (see their Table 3), and suggest that changes in racial classification practices were responsible for much of the increase in the white share. Unfortunately, the standard (female-dominant) methodology for population projection cannot address this issue because this method presumes that each child simply inherits his or her mother's racial class. Consequently, Loveman and Muniz (2007) did not attempt any quantitative assessment of the effect of racial mixing and classification.

Ideally, we would address this issue with population projections based on an age-structured two-sex model that kept track of union formation and dissolution. But given tractibility issues and data limitations, we develop here a more modest extension of the conventional one-sex framework. Following the standard method, we continue to assume that fertility is determined entirely by the mother's characteristics. Thus, the estimated number of births can be obtained for each racial class k using the standard approximation

$$B^k[t, t+5] \;=\; \sum_{x=\alpha}^{\beta-5} \frac{5}{2} \;\cdot\; {}_5F_x^k \;\cdot \left({}_5N_x^{Fk}(t) + \; {}_5N_{x-5}^{Fk}(t) \;\cdot\; \frac{{}_5L_x^k}{{}_5L_{x-5}^k} \right) \tag{1}$$

where, except for the superscripts to denote racial class $k \in \{w, m, b\}$, we are employing standard demography notation (cf Preston et al 2001, p 122, equation 6.4).

In order to use matrix algebra, these births can be arranged in the (1×3) vector

$$B_t \;=\; \left[\; B^w[t, t+5] \quad B^m[t, t+5] \quad B^b[t, t+5] \;\right] \tag{2}$$

Note that this vector reports the number of births by the *mother's* racial class. To obtain the number of births by the *child's* racial class, we borrow the specification of racial mixing and racial classification used in the two-sex model developed by Montgomery (2009). Formally, this requires a (3×9) matching matrix M_t and a (9×3) racial classification matrix C. The 3 rows of the matching matrix correspond to the mother's racial class (w, m, b), while the 9 columns correspond to the types of couples $(ww, wm, wb, mw, mm, mb, bw, bm, bb)$. Element $M_t(k, ij)$ indicates the probability that a female of race k will form a couple of type ij. The 9 rows of the classification matrix correspond to the types of couples, while the 3 columns correspond to the child's racial class. Element $C(ij, k)$ indicates the probability that a child born to a couple of type ij will be classified into race k. Through matrix multiplication, we thus obtain the (1×3) vector

$$B_t \; M_t \; C \tag{3}$$

which reports the number of births by child's racial class. Of course, not all newborns will survive until age 5. Following the standard method, the projected number of females of race k aged 0-4 is given by

$${}_5N_0^{Fk}(t+5) \;=\; (B_t M_t C)(k) \;\cdot\; \frac{{}_5L_0^{Fk}}{5l_0} \;\cdot\; \frac{1}{1+SRB} \tag{4}$$

while the projected number of males of race k aged 0-4 is given by

$${}_5N_0^{Mk}(t+5) \;=\; (B_t M_t C)(k) \;\cdot\; \frac{{}_5L_0^{Mk}}{5l_0} \;\cdot\; \frac{SRB}{1+SRB} \tag{5}$$

where $(B_t M_t C)(k)$ indicates element k of the $(B_t M_t C)$ vector, and we are otherwise employing standard demography notation (cf Preston et al 2001, p 122, equation 6.6). Overall, we have modified the standard projection method by incorporating the matching and classification matrices – using $(B_t M_t C)(k)$ in place of $B_t(k)$ – in equations (4) and (5).

Following the specification in Montgomery (2009), the matching matrix incorporates three parameters that reflect the extent of mixing between each pair of racial groups – whites and mulattos (μ), whites and blacks (β), and mulattos and blacks (γ). Each of these parameters is scaled between 0 and 1, with 0 indicating no racial mixing between the relevant groups (i.e., complete pairwise homophily bias), and 1 indicating unbiased racial mixing between these groups (i.e., no pairwise homophily

bias). More precisely, the non-zero elements of the matching matrix are given by

$$\begin{array}{lcl} M_t(w,ww) & = & x_t(w) + (1-\mu)\ x_t(m) + (1-\beta)\ x_t(b) \\ M_t(w,wm) & = & \mu\ x_t(m) \\ M_t(w,wb) & = & \beta\ x_t(b) \\ M_t(m,mw) & = & \mu\ x_t(w) \\ M_t(m,mm) & = & (1-\mu)\ x_t(w) + x_t(m) + (1-\gamma)\ x_t(b) \\ M_t(m,mb) & = & \gamma\ x_t(b) \\ M_t(b,bw) & = & \beta\ x_t(w) \\ M_t(b,bm) & = & \gamma\ x_t(b) \\ M_t(b,bb) & = & (1-\beta)\ x_t(w) + (1-\gamma)\ x_t(m) + x_t(b) \end{array} \tag{6}$$

where $x_t(k)$ may be interpreted in the present context as the share of "matchable" males who are of race k.[4] Our projection assumes the matching parameters $\mu = 0.3$, $\beta = 0.2$, and $\gamma = 1.0$ (following the calibration suggested by Montgomery 2009, footnote 26). We further assume that "matchable" males are those between the ages of 20 and 39, so that

$$x_t(k) \ = \ \frac{\sum_{x=20}^{35}\ {}_5N_x^{Mk}(t)}{\sum_{x=20}^{35}\ {}_5N_x^{M}(t)} \tag{7}$$

However, because racial shares do not vary much across age groups, the precise specification of $x_t(k)$ has little effect on our present projection. Finally, for purposes of our projection, we use the classification matrix observed for 1920. Thus, C is given by

$$\begin{array}{c} \\ ww \\ wm \\ wb \\ mw \\ mm \\ mb \\ bw \\ bm \\ bb \end{array} \begin{array}{c} \begin{array}{ccc} w & m & b \end{array} \\ \left[\begin{array}{ccc} .997 & .003 & .000 \\ .259 & .741 & .000 \\ .092 & .827 & .081 \\ .259 & .741 & .000 \\ .013 & .985 & .002 \\ .010 & .758 & .232 \\ .092 & .827 & .081 \\ .010 & .758 & .232 \\ .000 & .013 & .987 \end{array}\right] \end{array} \tag{8}$$

as reported by Loveman and Muniz (2007).[5]

Given these assumptions on racial mixing and classification (along with racial differentials in mortality and fertility), we obtain the population projection reported

[4] In the discrete-generations model in Montgomery (2009), x_t is simply the (male) racial distribution for generation t. Moving to our present age-structured model, x_t should reflect racial shares among those men who might actually be matched. In particular, x_t should not be influenced by the racial composition of the youngest male age groups.

[5] The symmetry of the classification probabilities (e.g., $C(mw,w) = C(wm,w)$) arises because Loveman and Muniz (2007, Table 3, p 931) do not distinguish between "mirror-image" couples (e.g., mw and wm). We have slightly adjusted several entries so that each row sums precisely to 1.

in row (c) of Table 1. Comparing rows (b) and (c), we find that racial mixing thus causes the white share to drop by almost a full percentage point, from 66.2% to 65.3%. This is consistent with Montgomery's (2009) finding that the observed pattern of racial classification was conducive to the growth of the mulatto share, not the white share. Intuitively, because more than half of the children of mulatto-white couples were classified as mulatto, racial mixing causes the mulatto share to grow and the white share to fall. Thus, against Loveman and Muniz (2007), the observed pattern of racial classification does not account for *any* of the roughly 7.5 percentage-point gap between the enumerated 1920 racial shares and the projected 1920 racial shares from row (b). Rather, accounting for racial mixing and classification, the gap actually widens by roughly another percentage point.

3 Racial switching

Having been unable to account for the dramatic increase in the white share through various demographic processes, we suspect that this increase must have resulted from widespread racial switching. In the absence of individual-level longitudinal data on race, our argument is necessarily conjectural. However, closer inspection of our final projection seems to provide some circumstantial evidence. Figure 2 shows the results of this projection by sex and age group for both whites (top panel) and mulattos (bottom panel). Even though we have accounted for the effects of racial mixing (and have not "corrected" the mortality and fertility rates), the top panel of Figure 2 appears quite similar to the analogous diagram in Loveman and Muniz (2007, Figure 2). Indeed, it was the "surplus" of whites in the youngest age groups that prompted Loveman and Muniz (2007) to focus on the racial classification of children.

However, it may be more instructive to consider the projections for mulattos shown in the lower panel of Figure 2. This diagram shows that the differences between enumerated and projected mulatto populations are not concentrated in the youngest age groups. Rather, the ratio of enumerated to projected population is relatively constant, with a mean of approximately 0.7 across age groups. While our analysis in the preceding section focused on demographic processes, this suggests that the increase in the white share resulted from a non-demographic process of (age-invariant) switching over the decade.

Switching might well be conceptualized as a continuous flow over time. But to make the computations transparent, suppose that 30% of mulattos became white at the moment of their enumeration in the 1920 census. The resulting projections for whites and mulattos are shown in Figure 3. Having moved 117,965 people from mulatto to white, the projected white share rises to 74.6% while the projected mulatto share falls to 21.7%. More importantly, the remaining projection errors appear unrelated to race. In particular, the "surplus" of children aged 5-9 (already evident from Figure 1) holds for both whites and mulattos. Further, while the "surplus" of white females aged 20-24 led Loveman and Muniz (2007) to speculate about racial

switching within mixed couples, note that there is also a "surplus" of mulatta females in this same age group.

4 Conclusion

Between 1910 and 1920, the white population share in Puerto Rico grew from 65% to 74%. In this research note, we have attempted to quantify the sources of this dramatic 9-point increase using demographic projection methods. Replicating the result from Loveman and Muniz (2007), we find that racial differentials in mortality and fertility account for approximately 1.5 percentage points of the increase in the white share. Moving beyond their result, we extend the standard (female-dominant cohort component) projection method to assess the effects of racial mixing. Because children born to mulatto-white couples were more often classified as mulatto than white, racial mixing actually reduced the white share. More precisely, from a baseline projection with racial differences in mortality and fertility, we find that racial mixing caused the white share to fall by almost 1 percentage point. Thus, the net effect of racial differentials and racial mixing is a mere 0.5-point increase in the white share.

Unable to account for the growth of the white share through demographic processes, we suspect that this increase must have occurred through widespread racial switching. Our projections for the mulatto population show that the enumerated-to-projected ratio is roughly constant across age groups. If we assume that each mulatto (regardless of age) had a 30% probability of switching to white upon enumeration in the 1920 census, the remaining projection errors appear unrelated to race.

Methodologically, our racial-mixing projection method begins to bridge the standard one-sex demographic model with the two-sex model developed in Montgomery (2009). While the two-sex model is well-suited for studying long-run dynamics, its discrete-generations specification makes it less useful for quantitive assessment of short-run dynamics. Nevertheless, when the two-sex model was applied to the Puerto Rican case (see Montgomery 2009, Section 6), it produced results that were qualitatively similar to those reported in the present paper. In future work, our racial-mixing projection method might be further elaborated, for instance to allow matching probabilities to vary by age group.

References

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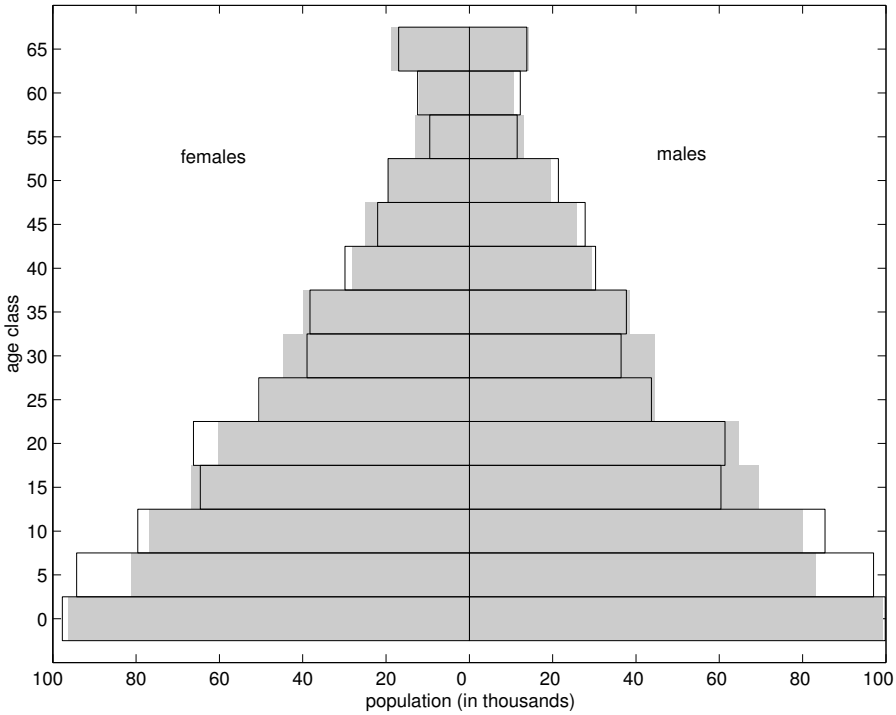
Table 1: Enumerated and projected populations

	white	**mulatto**	**black**	**total**
Enumerated 1910	721,590	340,240	50,235	1,112,065
	64.9%	*30.6%*	*4.5%*	
Enumerated 1920	947,835	282,230	49,200	1,279,275
	74.1%	*22.1%*	*3.8%*	
Projected 1920				
(a) *no racial differentials*	821,832	391,323	56,666	1,269,821
	64.7%	*30.8%*	*4.5%*	
(b) *racial differentials*	839,827	378,746	50,657	1,269,230
	66.2%	*29.8%*	*4.0%*	
(c) *racial differentials*	828,999	393,218	46,746	1,268,963
and racial mixing	*65.3%*	*31.0%*	*3.7%*	

Figure 1: Projected population, 1920, ignoring race



	female	**male**	**total**
total enumerated (white bars)	640,390	638,885	1,279,275
total projected (gray bars)	632,574	637,247	1,269,821

Figure 2: Projected population, 1920, with racial differentials and mixing

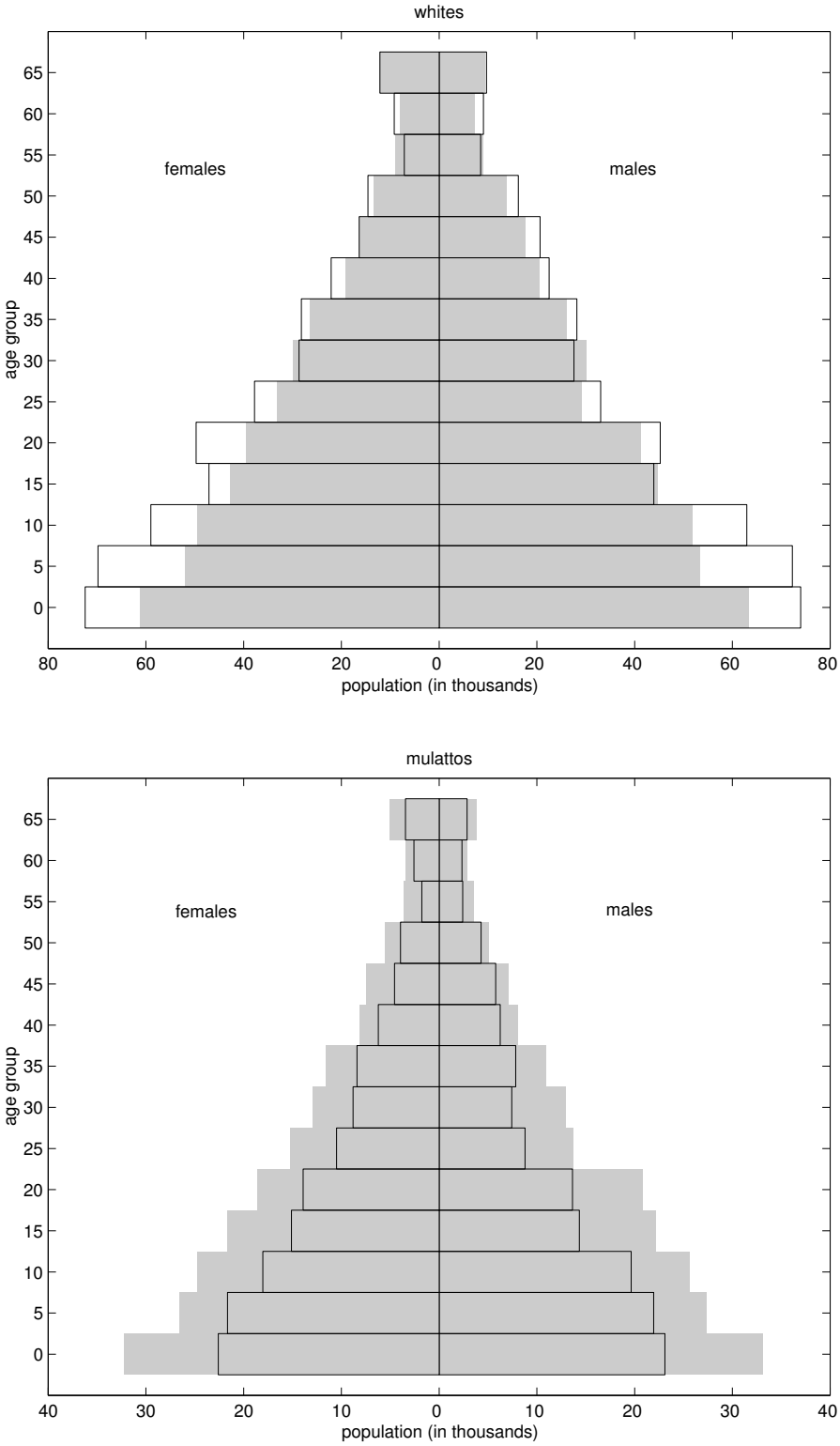
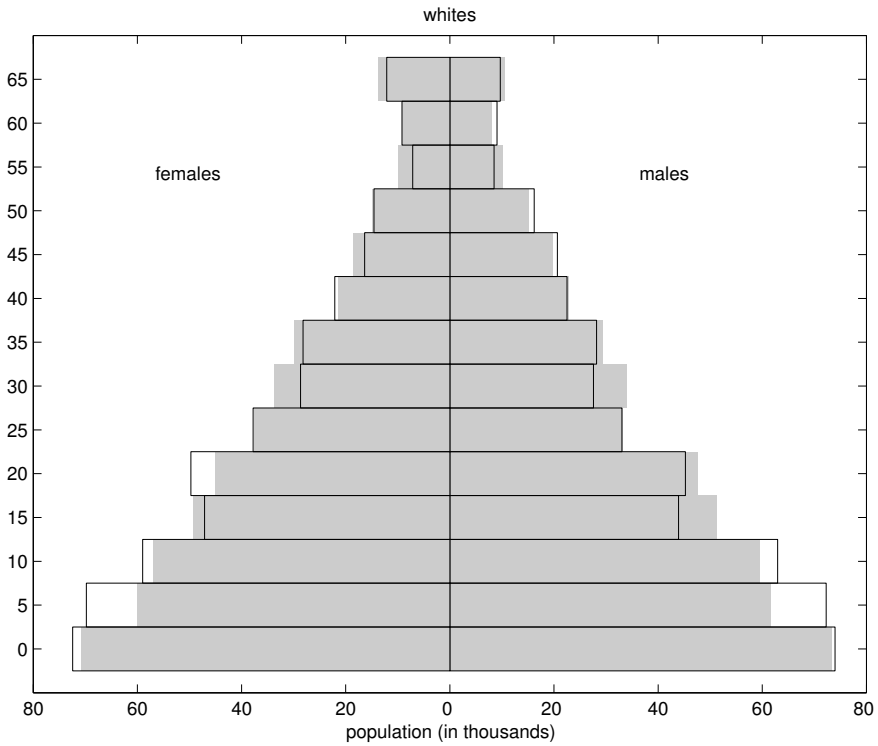
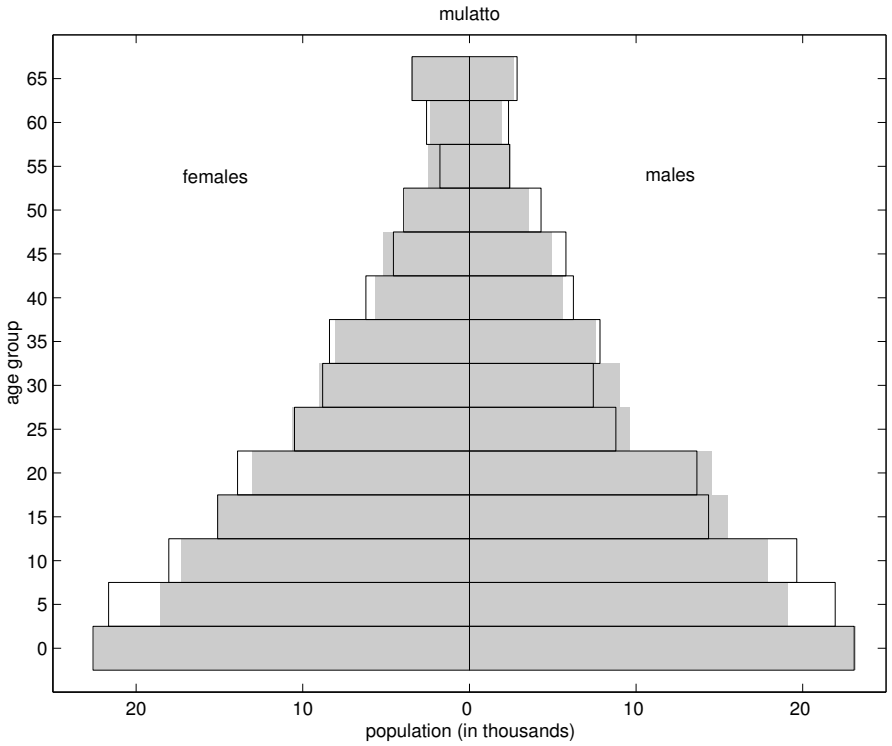


Figure 3: Projected population, 1920, with conjectured racial switching





Appendix: Enumerated populations by sex, race, and age group

1910

	males				females			
age	white	mulatto	black	total	white	mulatto	black	total
0-4	57020	29725	3280	90025	59985	30945	3320	94250
5-9	46415	24080	2695	73190	48915	24650	3080	76645
10-14	44150	21470	2705	68325	46690	23615	2945	73250
15-19	38825	18685	2775	60285	33845	16410	2280	52535
20-24	35690	16245	2575	54510	35780	15710	2225	53715
25-29	31825	14660	2625	49110	30665	13220	2085	45970
30-34	23145	10360	1535	35040	24250	9975	1345	35570
35-39	20175	9665	1615	31455	21995	9160	1615	32770
40-44	16975	7420	1145	25540	18010	7045	1135	26190
45-49	11660	4950	950	17560	12720	5250	905	18875
50-54	11020	4920	995	16935	10890	4530	1045	16465
55-59	6135	2840	710	9685	6015	2585	540	9140
60-64	6645	3125	805	10575	5680	2590	595	8865
65+	9275	3755	1540	14570	7195	2655	1170	11020
	358955	171900	25950	556805	362635	168340	24285	555260

1920

	males				females			
age	white	mulatto	black	total	white	mulatto	black	total
0-4	72430	22590	2680	97700	73985	23095	2755	99835
5-9	69805	21655	2825	94285	72280	21950	2820	97050
10-14	58975	18040	2595	79610	62950	19650	2775	85375
15-19	47085	15110	2370	64565	43925	14350	2135	60410
20-24	49725	13910	2585	66220	45230	13650	2505	61385
25-29	37765	10495	2295	50555	33065	8785	1865	43715
30-34	28670	8810	1470	38950	27585	7435	1395	36415
35-39	28185	8395	1665	38245	28190	7835	1655	37680
40-44	22090	6215	1535	29840	22510	6245	1565	30320
45-49	16350	4560	1115	22025	20650	5790	1360	27800
50-54	14550	3945	1015	19510	16195	4300	875	21370
55-59	7135	1770	585	9490	8485	2415	600	11500
60-64	9175	2575	680	12430	9065	2345	825	12235
65+	12105	3440	1420	16965	9685	2865	1235	13785
	474045	141510	24835	640390	473800	140710	24365	638875