The use of linked census data to estimate mortality and fertility in the pre-registration era of the United States: An application of the 1850-1940 IPUMS Multigenerational Longitudinal Datasets

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In this paper we describe a method to estimate child mortality using panel datasets of individuals linked between the 1850-1880 and 1900-1930 censuses of the United States. The large number of cases in the datasets allows us to estimate child mortality for the white population in all states and most counties for each of the six decades covered by the datasets and for the Black population in most states, most southern counties, and northern states and counties with enough observations to produce reliable estimates. The method also allows us to estimate census undercounts among white and Black children aged 0 and 1. With these estimates we make two additional contributions. First, with the assistance of a Brass model life table system based on life tables constructed for the U.S. Death Registration Area in the early twentieth century, we construct life tables indexed to the child mortality estimates by county, race, and decade. We show that geographic variations in mortality were large (hardly a surprise to historical demographers working in other contexts, but useful for U.S. historical demographers who have been limited heretofore to national life tables). Second, with the help of these new life tables, we construct new estimates of fertility, also at the county level and by race and census year, using the IPUMS 1850-1880 and 1900-1940 complete-count datasets and Own Child Methods (OCM). We estimate general fertility (age-specific birth and total fertility rates), marital fertility (age-specific marital fertility and total marital fertility rates), and Coale and Trussell indices of fertility, marriage, and marital fertility (I_f, I_m, I_g, M, and m).

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These estimates, we believe, represent a major contribution to our understanding of the timing and spatial variations in the demographic transition in the United States and will prove to be a valuable resource for other scholars needing contextual information on mortality and fertility (we intend to share our county-level estimates online through the IPUMS MLP website for interested researchers). We also believe our method of estimating child mortality will be useful for mortality analyses at the individual level, so we describe the method in enough detail to allow replication by other researchers. Our goals in this paper are primarily descriptive, however; to describe our methods, generate estimates of mortality, fertility, and census undercounts, and present an overview of our results. Although we do not show all the resulting estimates—which together compose a large dataset—we map selected estimates by county for a few census years and discuss the major implications of our descriptive results. We emphasize, however, that our methods and results are not limited to description. We provide a few examples of how our methods can be applied to analytical models of child mortality and how the contextual measures of mortality and fertility can be used in analyses of couples' fertility behaviors.

Background

The United States was late to establish a national system of vital registration. Although a few states and municipalities began registering births and deaths in the mid nineteenth century, the U.S. Death Registration Area (DRA) was established only in 1900 and initially included just 10 states and the District of Columbia, which together represented just 26.3% of the nation's population. Initially, the DRA was unrepresentative of the national population, being predominantly composed of states in the Northeast census region that were more urban and industrial than non-DRA states, also characterized by significantly higher proportions of foreign-born residents, lower proportions of Black residents, and lower fertility rates. States were added to the DRA over time, making it more representative of the nation, but it was not completed until 1933, when Texas was added. Even then a significant percentage of deaths are believed to have been unregistered. The national Birth Registration Area began even later, in 1915, and births continued to be underreported after it was deemed completed in 1933 (Preston and Haines 1991: 49-50; Haines 2000: 311-313: Haines 2006: 1-381-1-390).[3]

Because of the lateness and deficiencies of vital registration, our understandings of the mortality and fertility transitions in the United States, especially in their early stages, rest on a weak empirical foundation. Ironically, about the time the U.S. vital registration system was completed in the mid-1930s, the nation's century-long fertility transition came to an end with the commencement of the Baby Boom. Although the mortality transition continued, the period of most rapid and greatest decline in mortality was the half century prior to 1933. Researchers of the demographic transition have needed to be creative, constructing demographic rates using limited and unrepresentative vital registration data, indirect estimation methods based on census data data for the living population, or from studies of genealogical records and special data sources, including retrospective mortality information collected with the 1850-1900 censuses and children ever born and children surviving information collected for each mother in the 1900 and 1910 censuses (e.g., Pope 1994; Preston and Haines 1991).

The most cited life tables for the nineteenth century, Michael R. Haines' (1998) life tables for the years preceding the 1850, 1860, 1870, 1880, 1890, and 1900 censuses, were based on retrospective mortality censuses conducted alongside the population censuses. Households enumerated in the population census were to report any household members who had died in the year prior to the census, who were recorded on a separate schedule. These mortality censuses clearly under-reported deaths—perhaps as much as 40% or more overall–and the year in which mortality was recorded may have not been representative of the decade. A major cholera epidemic in 1849-50, for example, may have biased estimates of mortality in the year preceding the 1850 census upwards (Vinovskis 1978). Haines, however, relied solely only on children aged 5-14 dying in the year prior to each census, who

[3] To be admitted to the national registration area, states had to demonstrate registration completeness of 90 percent or more. Death registration is believed to have been more complete than birth registration.

appeared to be reasonably well reported. He then used model life tables—discussed in more detail below—to extend the estimates to all age groups. Hacker (2010) also used model life tables to construct nineteenth-century life tables, fitting adult life expectancy estimates at age 20 from published genealogical studies to a relational life table system based on life tables constructed for the 1900-1902 DRA (Glover 1921). No attempt was made by Haines or Hacker to estimate regional or sub-regional estimates of mortality. Both sets of life tables were for the white population only, although Haines included a few life tables for the Black population at the turn of the twentieth century based on his analysis with Samuel Preston of children ever born and child surviving data in the 1900 and 1910 censuses (Preston and Haines 1991, Haines and Preston 1997), which allowed the estimation of child mortality.

Although the United States lacked a system of vital registration until the early twentieth century, it conducted decennial censuses beginning in the late eighteenth century, which can be used to make indirect estimates of mortality and fertility from age distributions of the living population. The Black population was essentially closed to migration after the abolition of the international slave trade in 1808, allowing mortality to be estimated with two census methods (e.g., Preston and Bennett 1983). Differences in the sizes of reported age groups, census coverage errors, age misreporting, and the unknown numbers of slaves smuggled into the nation after 1808, however, limit confidence in two-census mortality estimates (McDaniel and Gruska 1995). The white population was continually augmented by new immigrants and not closed to migration. Two census methods are feasible only for the native-born white population (which can be assumed closed to migration), and only after 1880, when the Census Office cross-tabulated the population by age, sex, and nativity.

With an assumed age structure of dying, observed intercensal growth rates, and stable population methods, census data can also be used to estimate crude birth and death rates (Eblen 1974, Yasuba 1962). Coale and Zelnik (1963) interpolated between the few available nineteenth-century life tables to back project annual births from the age distributions of the native-born population in the 1880-1950 censuses. Aggregated counts of the population cross-tabulated by age group and sex can also be used to calculate child woman ratios, which are a good proxy for fertility. Between 1800 and 1860 the Census Office cross-tabulated the white population by age and sex for each county in the United States, facilitating cross-sectional studies (e.g., Easterlin et al. 1978; Hacker et al. 2023), but discontinued county-level cross-tabulations with the 1870 census.

The construction of public use microdata samples of the 1850-1880 and 1900-1930 censuses (Ruggles et al. 2022) provided new research opportunities. Although the Census Office did not cross-tabulate the white population by age, sex, and nativity until 1880, researchers can use the 1850-1870 IPUMS samples to do so, extending two-census mortality estimates back to 1850 (Hacker 2010). IPUMS samples can also be used to construct county-level child-woman ratios for the period after 1860. Preston and Haines used a low-density public use sample constructed for the 1900 census to analyze the number of children ever born (CEB) and the number of those children surviving (CS) data collected by that census for ever married women. These data allowed them to estimate child mortality and construct empirical models of mortality with mother-, couple-, and area-level covariates (Preston and Haines 1991). CEB and CS data were also collected in the 1910 census and have been analyzed by researchers using public use microdata samples of the 1910 census in similar ways (Preston et al. 1994; Dribe et al. 2020). More importantly for censuses without CEB and CS data, public use microdata samples allow the estimation of nuptiality and the application of Own Child Methods (OCM) of fertility estimation, which employ reverse-survival methods to calculate age-specific general and marital fertility rates in the years prior to each census (Tolnay 1984; Hacker 2003). These estimates indicate that the marital fertility transition at the national level did not begin as early as the date implied by child-woman ratios. The decline began in New England in the mid nineteenth century and lagged in the South, which did not experience a significant decline in marital fertility until late in the century (Hacker 2016).

Regional OCM estimates, however, are based on the questionable assumption that the available national life tables accurately represented child mortality conditions in all census regions. Hypothetically, if child mortality was higher in New England and lower in the South, the large regional differences in fertility estimated by OCM would not have been so profound. Researchers have made no attempt to apply OCM to estimate fertility at the state or county levels. IPUMS sample densities (typically 1%) were too low to estimate fertility with confidence at the county level or for smaller states, and unknown differences in child mortality at the state and county levels would further undermine confidence in sub-regional estimates.

The recent introduction of new complete-count IPUMS datasets (Ruggles et al. 2021) represents an ideal opportunity to apply OCM to smaller areas of geography. Unfortunately, our lack of knowledge about local mortality conditions would limit confidence in resulting estimates. Ideally, therefore, we need a method to estimate mortality differences for small areas that can in turn be used as an input to making OCM estimates. We argue that the IPUMS MLP datasets, used in conjunction with model life table systems, provide such an opportunity.

Data

We constructed panel datasets using “crosswalks” published by the IPUMS Longitudinal Multigenerational Project (IPUMS MLP) at the University of Minnesota (Helgertz et al. 2020). The crosswalks identify individuals linked between adjacent censuses of the IPUMS complete count datasets between 1850 and 1940. The MLP data is optimally suited for the purposes of this paper, through the method leveraging household level information not only to minimize linking error but also to link men and women who coreside in consecutive censuses (see Helgertz et al. 2021 for a description of the linking methods). Using the MLP crosswalks, we created three panel datasets for the nineteenth century

(1850-1860, 1860-1870, and 1870-1880) and four panel datasets for the twentieth century (1900-1910, 1910-1920, 1920-1930, and 1930-1940). For each panel, we restricted the sample to children aged 0-5 years in the earlier census who resided with their mother and father and where both parents were successfully linked to the later census. A recent comparison of couples in the linked dataset to couples in cross-sectional complete county datasets indicated that linked couples are fairly representative of all couples, with some modest bias in expected areas (e.g., couples appear to have been slightly more likely to be linked if they were literate, wealthy, native born, living in rural areas, living on a farm, living in the Northeast census region, and if they had more children living in the household) (Hacker et al. 2023). Overall, however, biases appear to be small. Our selection criteria, however, is likely correlated with child mortality: because our method relies on linked couples, who (by definition) survived the decade under observation, children of linked couples likely experienced lower mortality than children of parents whose marriages were disrupted by the death of the father or mother. On the other hand, because our method assumes that children who could not be linked between censuses died in the interval, any missed links will bias our mortality results upwards. Although these biases are offsetting and our estimates appear to be reasonable for most census years, later in this paper we discuss a method of standardizing our results to national estimates of child mortality.

Methods

Our method of measuring child mortality is complex in detail but easy to describe in brief. We begin with linked couples with one or more children under the age of 6 in the first of the two linked censuses (census A). The vast majority of these children were linked in the MLP data to a child in the same household as the couple in the second census (census B), with some being linked to records in other households, in both cases indicating that they survived the decade. Many linked couples also have unlinked children in both censuses that may have been linked if less conservative thresholds for matching had been used. Additionally, the prevalence of under-enumeration implies the potential presence of children in census B whose age indicates that they were born prior to census A but who was not recorded as a household member in the former. Due to the high degree of certainty with which we can identify households through conditioning on both parents being successfully linked, we contend that lower matching thresholds can be used without incurring significant type I errors. We therefore proceeded to develop a procedure to force matches among unlinked census A children for whom there is a plausible, unlinked, census B child. Acknowledging the presence of significant errors in terms of name and age-reporting, we impose no a priori restriction on the required name similarity, however only allowing for a +/-3 year age difference between the census A and census B record. The universe of census B potential matches is limited to children aged 10-18, meaning that a child aged 0 in census A will only be compared to children aged 10-13 in census B. This is motivated by the desire to avoid underestimating mortality among the youngest aged children, associated with the highest mortality risk. The procedure sequentially compares each unlinked census A child to all unlinked census B children, forcing a match to the child with the highest Jaro-Winkler name similarity score, conditional on there being a match on sex and age. This process is subsequently repeated, increasing the age difference to one, two, and three years, respectively.

After identifying all linked children, presumed to have survived the census interval, we identified what we assumed to be i) children dying between the censuses and ii) under-counted children in Census A. Children aged 0-5 in Census A who could not be linked forward to an child aged 10-18 in Census B was assumed to have died during the ten- year interval, while remaining children aged 10-15 in Census B who were not linked backward to a child aged 0-5 in Census A were undercounted by Census A. After analyzing and comparing the results to model life tables (see example below), which suggested an increasing tendency for children to leave their parental homes after age 13, we limited our estimates to children aged 0-3 in Census A.

We fitted the resulting ten-year mortality estimates of children aged 0-3 in Census A to a relational model life table system to create complete life tables. We used life tables constructed for the United States DRA in the early twentieth century as standards and a Brass two-parameter model (reduced to one-parameter as described below), which allows the transformation of the standard life table using a logit model with varying intercepts to fit the level of mortality. From this transformation, we generated complete life tables by single years of age, sex, and race for each county in the United States with more than 500 observations (for small counties with fewer observations we relied instead on state-level estimates). Finally, we used the resulting life tables as inputs to construct new fertility estimates using OCM and the IPUMS complete-count datasets for the 1850-1880 and 1900-1930 censuses.

We implemented all of these steps in STATA. In the following section, we provide more detail and briefly describe the STATA code. We organize this discussion under five separate headings: (1) Linking unlinked children and aggregating the number of surviving children by geographic area; (2) estimating census undercounts among children aged 0-5; (3) estimating 10-year forward survival ratios and mortality ratios among children aged 0-3; (4) fitting model life tables for the white and Black populations; and (5) fertility estimation using own-child fertility methods.

Matching unlinked children and aggregating the number of surviving children by geographic area

We begin with couples linked between Census A and Census B with one or more coresident children aged 0-5 in Census A. Couples are therefore known to be fecund and have children at risk of mortality in the subsequent decade.

For discussion purposes, we focus on couples in the 1870-1880 panel dataset. Among these couples, the mean number of children aged 0-5 in 1870 was 1.3, identical to the mean number of children aged 10-15 in 1880 (1.9 in the 10-18 age range). 81% of the children aged 0-5 in the 1870 Census were linked in the MLP data to a child in the 1880 census, a similar figure also applying to the share of children aged 10-15 in the 1880 census who were linked to a child in 1870. Approximately 21% of couples had forwards unlinked children aged 0-5 in 1870 and 29% had children aged 10-18 in 1880 who were not linked to an 1870 record. A visual inspection of a selection of these cases using the original manuscript returns suggested that a large proportion (but not all) of these children were the same individuals, but for a variety of reasons—respondent error, enumerator error, data entry error, etc.—the children could not be linked automatically with confidence.

Our 1870-1880 panel dataset, for example, included the 1870 Dent County, Missouri couple Joab and Elizabeth Hobson, aged 43 and 38 respectively, who were linked to the 1880 Benton County, Missouri couple Joab and Elisabeth Hobson, aged 53 and 48. Joab's state of birth was listed as North Carolina and Elizabeth's as Kentucky in both censuses. Although residing in different Missouri counties, the MLP linking procedures determined that this was the same couple. In 1870, Joab and Elizabeth had three male children, named "A L," "A J," and "F S," aged 4, 4, and 1, all born in the state of Missouri, who were not linked forward to a child in the 1880 census. In 1880, Joab and Elizabeth had three male children, named "Abraham Hobson," "Andrew Hobson," and "Phillip Hobson," aged 14, 14, and 11, all born in Missouri, who were not linked backward to a child in the 1870 census. Although the IPUMS MLP project used conservative rules to minimize type I errors, and therefore avoided linking these children, it seems highly likely that these children were the same children in both censuses. By relaxing our linking thresholds, we were able to link these children.

As a first step, we therefore made several passes through the data, lowering required linking thresholds and removing blocking criteria, to create new links. Despite the lower thresholds and greater allowances for differences in age and birthplace, we remain confident that type I errors—the linking of two different individuals—are low.

After making additional links between children unlinked by the MLP project, the number of couples with both forwards unlinked children aged 0-5 in 1870 and backwards unlinked children aged 10-18 in 1880 fell significantly, from 12.5 to 2.1%. Among the children aged 0-3 in the 1870 Census, 11.1% could not be located in the 1880 Census and were assumed to have died in the interval between the two censuses.

We aggregated the number of living children aged 0-5 in 1870, the number of these children who died in the interval, and the number of living children aged 10-15 in 1880 by single years of age for four different levels of geography in Census B (nation, state, state economic area [SEA], and county). The same process was followed in all other panel datasets.

We observe geographic areas (j) only in Census B. Some couples, of course, moved sometime during the ten-year interval covered by each panel, which could be detected by changes in the county or state of residence between censuses A and B. We based our measurement of geography solely on Census B for several reasons. First, it simplified our calculations greatly. Second, some geographic changes apparent in the two censuses might have only reflected nominal changes in census geography. To cite just one example, in 1850 the state of Minnesota had nine counties. The number of counties increased to 64 in 1860, 72 in 1870, and 78 in 1880. Although living in the same house in two adjacent censuses, it was easily possible for a hypothetical couple in the 1850-1860, 1860-1870, and 1870-1880 panel datasets to reside in different counties. Third, because of the expansion in population and the number of counties over time, geography measured on Census B is likely to be more precise than geography measured in Census A. Finally, and most importantly, however, we standardized on Census B because we used reverse-survival methods to estimate fertility. OCM estimates apply to the period before the complete-count census datasets, so mortality should be measured in the period before the second of the two censuses.

Estimating census undercounts among children aged 0-5

Children of linked couples who were aged 10-15 in Census B and who were not linked to a child aged 0-5 in Census A even after the previously mentioned procedure represent a special class of children. If these children's ages were recorded accurately in Census B, they should have been alive and enumerated between the ages of 0-5 in Census A. With a few rare exceptions, children aged 10-15 who were co-resident with their parents in Census B should also have been residing with their parents ten years earlier in Census A. Because census under-enumeration of young children is known to have been a significant risk in early American censuses (Coale and Zelnik 1962; Hacker 2013), we assumed that these children were undercounted by Census A. It is possible, of course, that some young children may have lived elsewhere, such as in a nearby relatives' home, and later moved back to their parents' households to be enumerated by the second census. These children were likely few in number, however, and the practical impact was the same: these are children of linked couples who survived the intercensal interval. Treating them as undercounted in Census A results in an unbiased estimate of child mortality.

There is a non-negligible risk that couples in Census A had an undercounted child in Census A who did not survive the ten-year interval. Although these are two relatively rare events (ten-year mortality among children aged 0-3 in 1870, for example, averaged about 10%, while the undercounting of children aged 10-13 in Census B relative to mothers and children of older ages probably averaged about 5% (Hacker 2013), our method is unable to detect this possibility and, this will result in child mortality being slightly underestimated. The bias is more severe at age 0, when mortality and undercounting is higher. We discuss the impact of this bias when describing age-specific mortality rates in Figure 3 below.

Finally, there is a much higher possibility that the older children aged 0-5 present in Census A survived the interval and were no longer coresident with their parents at Census B, particularly among the older children who may have left home by the time of Census B, if they were not successfully linked to that household. Because our method assumes these children are deceased, children's departure from their parents' homes will bias our results. We also discuss the impact of this bias when describing age-specific mortality rates in Figure 3. The results indicate that while the bias is significant for older children, resulting in increasingly higher than expected estimated mortality rates for children above the age of 4, the bias is minimal for children aged 0-3 in Census A.

Despite our efforts, there is still a small share of couples with unlinked children in both census years. It is possible that some of these children should have been linked, even if they are discrepant on age by more than three years or even on gender. A visual inspection of a selection of these households, along with comprehensively examining supplementary administrative records provided by ancestry.com, did not systematically suggest that further links should be made, much less the possibility that an automatic procedure could be developed to safely link the few remaining cases. By assuming *all* these children should have been linked—which is highly unlikely—we can estimate the potential size of the bias, which is small. We discuss this potential bias further below.

For children aged 10-15 in Census B who were assumed to be undercounted, we calculated their age in Census A as their age in Census B minus ten years. We aggregated the number of undercounted children by single years of age 0-5 in Census A for the same four different levels of geography described above (nation, state, SEA, and county). Figure 1 shows the undercount estimates by age in the period 1850-1870 (Census A). As can be seen on the graph, the undercount estimates are quite high for children aged 0 in Census A (over 20% in 1850 and 1860), somewhat elevated for children aged 1, and then level out at about 3-4% for children aged 2-5. Although we have no way of comparing rates directly, the undercount estimates for ages 2-5 approximately equal those estimated with back projection estimates (Hacker 2013) for most age groups, including those of women of childbearing ages. When calculating fertility rates using OCM, it is the relative differences in the undercounting of women and children that is important. We therefore decided to assume that undercounts for women of childbearing ages approximated those observed for children of both sexes aged 2-5. We removed that undercount from all ages and estimated only the relative undercount for children aged 0 and 1 by race, state, and year.

Figure 2 shows a map on the adjusted undercount estimates for children aged 0 in 1870. The estimated undercount was highest in Louisiana, Florida, and New Mexico, and then higher in the rest of the Southeast, suggesting marked regional differences in enumeration quality. The results are consistent with observations by Census officials about the lower levels of education among enumerators and the public in the region and the lower quality of census enumeration.

10-year forward survival ratios among children aged 0-3

We calculate survivorship and corresponding mortality ratios for each geographic area. The survivorship ratio is the total number of children aged 10-15 in Census B at each age divided by the number of children aged 0-5 and the total number of children aged 0-5 undercounted in Census A at each age. The survivorship ratio at each age is equivalent to the life table value L_{x+10}/L_x. The ten-year mortality estimate is simply one minus the survivorship ratio.

In the nineteenth century, our estimates of child mortality corresponded closely to estimates suggested in life tables by Haines (1998) and Hacker (2010), with rates between 12% lower and 8 percent higher than the published estimates. In the twentieth century, with the onset of the mortality transition, our estimates indicate higher child mortality than estimates from the DRA, averaging 19% higher in the period 1900-1930. Our interpretation of these trends is that our method—which assumes all remaining children in Census A who were not linked to Census B died in the interval—erroneously overestimates child mortality in some cases. In other words, a few of the children we assumed to have died either survived the ten-year interval and, for whatever the reason, were not linked, or else they were undercounted in Census B. The overestimation of mortality appears to have been modest relative to actual child mortality rates in the nineteenth century, when mortality levels were high. In the twentieth century, however, the error became more significant relative to actual mortality levels.

Given these findings, we decided to index our estimates to an assumed mortality rate in all decades. We used estimated mortality rates at the national level—obtained from published life tables—and compared that standard to the national level of child mortality among children aged 0-3 we estimated using the panel datasets. We then adjusted all subsequent child mortality estimates by the percentage we overestimated or underestimated mortality at the national level. Thus, while we still produced estimates for each combination of year-county-race, our average national child mortality estimate will match the rate of child mortality in the assumed life table.

In Figure 3 we show the overall results of our child mortality estimates by individual ages for the white population in the 1870-1880 dataset at the national level. Note that we here extend our estimation to include older children. As can be seen in the figure, our estimates of the proportion of children dying was modestly lower than the expected proportion for children aged 0, modestly higher for children aged 1-3, and increasingly higher than the expected proportions at older ages. Among children aged 0-3, the proportion was 8% higher than the expected proportion. It is quite possible, of course, that the expected proportion, which was estimated from genealogical data, is incorrect. But the deviations by age suggest patterns of bias. As noted earlier, our method likely underestimates mortality among children aged 0 for the simple reason that we are unable to observe mortality among children undercounted in the first census. The dashed red line provides a rough estimate of the size of this bias by assuming the same undercount shown in Figure 1 for children aged 0 and the same mortality estimate measured for the observed children. More importantly, the increasing deviation from the standard with increasing ages for children aged 4 and above suggest a massive bias. Undoubtedly, this bias results from children departing their parental homes, making them much more difficult to link. For children aged 9 in 1870, our calculated proportion dying, 0.18, was more than three times higher than the expected proportion of 0.053. For this reason, we limit all of our mortality estimation below to children aged 0-3 in Census A.

Fitting model life tables for the white and black populations

Although mortality varies across time and place, rates vary across the life course in a consistent pattern, with mortality increasing gradually from early childhood to middle age and then more rapidly at older ages. This consistency encouraged early demographers to attempt to model the “laws of vitality” with an equation indexed to the overall level of mortality (Woods 2000: 170-190). Differences in disease environments and behaviors, however, results in differences in the leading causes of death and the age pattern of mortality across populations. Consequently, there is no universal method of constructing life tables from a single parameter, such as the ten-year mortality rate of children aged 0-3. We instead relied on a relational model life table system—constructed using observed age patterns of white and black mortality in the early twentieth-century United States—to fit observed rates of child mortality.

Model life tables grew out of the considerable efforts by demographers in the twentieth century to construct life tables from different periods and places and identify common patterns. In 1966, Ansley Coale and Paul Demeny (1966) proposed four “regional” model life table systems based on broad similarities and differences among 326 life tables, which they designated “North,” “East,” “South,” and

"West." The designated names roughly corresponded to the regions (mostly European) from which the data were collected. The North model life table system was based on life tables from Sweden, Norway, and Iceland; the East model on life tables from Austria, Germany, Hungary, Czechoslovakia, Poland, and Northern Italy; and the South model on life tables from Spain, Portugal, and Southern Italy. The West model life table system was based on a variety of life tables that did not fit the mortality patterns of the other three regions, including a few life tables from the United States. Each region was characterized by a different age-pattern of mortality, due in part to regional differences in the disease environment and causes of death. Northern European populations, for example, suffered from endemic tuberculosis in the nineteenth and early twentieth centuries, which resulted in higher mortality rates between the ages of 15 and 40 in the North model life table system relative to other models. There are other differences. Infant and child mortality rates, for example, are lower relative to adult mortality rates in Models North and West and higher in Models South and East (Hacker 2010: 59).[4]

Many life table models have been proposed, including models proposed by the United Nations constructed using empirical data from developing nations (United Nations 1982). For populations with poor vital registration data, the choice of the most appropriate model life table system requires some guesswork. Coale and Demeny's Model West family of life tables is the most used model for research on U.S. populations—a few U.S. life tables from the early twentieth century were used in its construction—and its use appears to result in good estimates in the twentieth century (Haines 1979, 197; Preston and Haines 1991, 66). Douglas Ewbank, however, contended that the age mortality profile of the early twentieth-century black population of the United States more closely matched the United Nation's "Far East" model life table (1987). More recent research (Hacker 2010; forthcoming), has shown

[4] Within each model life table family, Coale and Demeny constructed life tables reflecting different levels of mortality, with the lowest numbered level representing the highest level of mortality covered by the model and the highest numbered level representing the lowest level of mortality. In the West model, for example, level 1 corresponds to a life expectancy at birth of 20.0 years for females, while level 24 corresponds to a life expectancy of 77.5 years. It is a simple matter to interpolate between levels to exactly match a selected mortality rate or life expectancy in a study population.

that mortality rates suggested by Model West are a poor match for well-documented U.S. populations in the eighteenth and nineteenth centuries. Model North provides a better fit, especially for females, likely because tuberculosis was the leading cause of death in the United States in the nineteenth century and was especially prevalent among females.

Rather than relying on one of the published regional life tables systems, we decided to use actual U.S. life tables constructed for the population in the DRA in the early twentieth century to develop a "relational" U. S. model life table system (Brass 1971). For the white population prior to 1910, we relied predominantly on the 1900-1902 life table for the white population residing in rural areas of the DRA, which was published in a separate analysis by Glover (1921). For the decades after 1910, we relied on Glover's life tables for the overall white population in the DRA, which were approximately representative of the national white population after 1910. As others have pointed out (Preston and Haines 1991), however, the DRA was not representative of the nation's Black population until after about 1920, when more southern states were added to the DRA. We therefore relied on the 1920 life table for the Black population as a standard in all years prior to 1920.

Relational model life table systems are based on William Brass's observation that the logits of the l_x values from any two life tables are related to one another by a linear relationship (19xx). It is therefore possible to use the logits from a chosen standard life table to create a family of related life tables by varying the slope and intercept and taking the anti-logits to produce a new set of l_x values. The standard table will be reproduced when the intercept (α) equals 0 and the slope (β) equals 1. Values of the intercept parameter greater than 0.0 will shift the level of mortality above the standard table and values less than 0.0 will shift the level of mortality below the standard table, while a slope value greater than 1.0 tilts the life toward higher adult mortality relatives to infant and child mortality and a slope less than 1.0 tilts the life table toward higher infant and child mortality relative to adult mortality.[5] Appropriate choice of a standard table can preserve variations in the age profile of mortality that cannot be obtained by varying the slope and intercept parameters of the standard table.

Figure 4 plots the q_x values (proportions dying in age interval) for white females in the rural parts of the DRA and the overall DRA in 1900-1902 together with the Model West values for females with the same life expectancies at birth (51.1 years for white women in the overall DRA and 55.4 for white women in the rural parts of the DRA). Despite having similar mortality rates at most age groups to those suggested by the models, women between the ages of about 15 and 40 had significantly higher mortality rates than those suggested by Model West, especially women living in rural areas. These are age groups with high rates of death from maternal causes and tuberculosis. The greater disagreement in rural areas may reflect the higher fertility among women in rural areas. In addition to higher rates of maternal mortality, pregnancy is a risk factor in contracting tuberculosis. Although we cannot compare these results directly to mortality rates directly observed in the nineteenth century, which are unknown, the nation was more rural, fertility rates were higher, and tuberculosis was the leading killer of young adults. Given these conditions, it seems reasonable to assume that the rural DRA life tables are more accurate models of age-specific mortality rates in the period 1850-1880 than the Model West family of life tables, which was based on more urban and lower-fertility populations.

For the white population prior to 1900, we took the logits for the male and female life tables constructed for the rural and overall DRA life tables in 1900-1902. For the black population prior to 1920, we took the logits for the black life table in 1920. After those dates we used the available life tables for the DRA. Table x shows the standard logits we used for the white and Black populations in each decade.

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[5] The slope parameter determines the "tilt" of the table. A slope value greater than 1.0 indicates that infant and child mortality is lower relative to adult mortality than in the standard table, and a slope less than 1.0 indicates that infant and child mortality is higher relative to adult mortality than in the standard. We change only the intercept to fit the life table. If effect, we are assuming the relationship between child and adult in the period 1850-1900 was approximated by the relationship in 1900-1902, reducing the two-parameter model to a one-parameter model.

For the white population, the standard logits are a weighted average of the logits for the life tables for overall and rural white populations, with the weights adjusted to equal the proportion of the population living in urban areas in each decade.[6] We used a simple equation to predict the Brass slope parameter from the observed child mortality in each county, which we implemented in STATA. It was then a simple matter to take the anti-logits and generate the l_x values for each county, and from those, all necessary life table values for the application of OCM. Currently, these include only the l_x and L_x values for children of both sexes combined aged 0-15 years and the L_x values for women aged 15-49. For simple descriptive statistics, we also estimated the infant mortality rate and the proportion of children dying before age 5 (q_0 and q_5). It is a simple matter to use the l_x values to construct complete life tables, but we have not done so because of space concerns (roughly speaking, for three unabridged life tables–male, female, and both sexes combined—each county would have over 500 parameters for each race, or 1,000 parameters in total). Multiplied by the approximately 3,500 counties and seven decades, this amounts to about 25 million life table values. We do, however, provide l_x values for each age, sex, race, and county, so users of the data can easily construct complete life tables on their own.

In figures 5 and 6, we mapped the proportion of white and black children dying before age 5 (q_5) for the period 1870-1880. The maps indicate large race differentials in child mortality. Areas of relatively high mortality for both races include counties bordering the Mississippi River in the states of Arkansas and Mississippi andFlorida and coastal South Carolina. Low mortality counties include sparsely populated counties in the Appalachian Mountains. Urban counties also have notably higher mortality than rural counties, consistent with expectations, although these differences are difficult to see in the map without zooming.

[6] It is also possible, although we have yet attempted it, to weight the logits according to the proportion of the population in urban areas in each county and decade.

Fertility estimation using own-child fertility methods

The own-child method (OCM) of fertility estimation is a reverse-survival method developed by Grabill and Cho (1965) and subsequently elaborated by Cho, Retherford, and Choe (1986). The method requires microdata, with mothers linked to their "own children," who are defined as biological children coresident in the household. In the basic method, own-children are cross-tabulated by their mother's age, and then reverse survived using available mortality estimates to obtain the number of births by mothers' ages in individual years preceding the census. Because children leave their mothers' households in increasing numbers after age 15, estimates are typically confined to the 15 years prior to the census. Women present in the census are also reverse survived to estimate the total number of women alive at each age in each year. Age-specific birth rates are then calculated by dividing the number of back-projected births by the number of back-projected women of a particular age, and the total fertility rate is obtained by summing the age-specific birth rates of women aged 15-49. Finally, age-specific marital fertility rates are obtained by multiplying the age-specific fertility rates by the inverse of the proportion of the women at each age group currently married. Elaborations include adjustments for unmatched children (surviving children who are not co-resident with their mothers) and under-enumeration of children and mothers in the census.

We used the procedures detailed by Cho, Retherford, and Choe (1986), the IPUMS complete-county datasets of 1850-1880 and 1900-1940 censuses, and the county-level mortality estimates estimated for each county and race to make fertility estimates for each county and race. We included adjustments for census under-enumeration at age 0 and 1 described above and the proportion of children not living with their own mother before reverse surviving children and mothers. The proportion of women currently married was also estimated for each county and race for use in estimating marital fertility rates. Although we made estimates for individual years and for five-year groups of years (0-4, 5-9, and 10-14 years before the census), we averaged all the fertility estimates for all 15 years prior to the census to maximize the number of cases used in the county-level estimates.

We estimated general fertility (age-specific birth and total fertility rates), marital fertility (age-specific marital fertility and total marital fertility rates), and several indices of fertility, marriage, and marital fertility (I_f, I_m, I_g, *M,* and *m*). I_f, I_m, I_g and another index, I_h, were developed for the European Fertility Project at Princeton University to study the European fertility transition. I_f is an index of overall fertility, while I_m, I_g and I_h are its three components. I_m is an index of proportions married, I_g is an index of marital fertility, and I_h is an index of extra-marital fertility. Unfortunately, own-child estimation methods do not allow separate estimation of extra-marital fertility. We assumed all own-children were born to married couples, slightly inflating marital fertility estimates, especially among recently married younger women who may have been selected into marriage because of unplanned pregnancies. The EFP indexes are standardized on age-specific fertility rates of married Hutterite women aged 15-49 in the early twentieth century, who experienced the highest reliably recorded fertility rates of any human population and thus serve as a theoretical baseline measure of maximum human fertility. The fertility indexes can therefore be read as the proportion of Hutterite fertility achieved by the population, while the marriage index represents the proportion of the female population married weighted by their expected contribution to fertility.

M and *m* are two indices of marital fertility developed by Ansley Coale and James Trussell to infer the degree of conscious fertility control. They are defined by the equation

$$r(a) = M * n(a) * \exp exp\, m * v(a)$$

where a is the standard five-year age groups 20-49, r(a) is the age-specific marital fertility rate of the population being investigated, n(a) is a model schedule of natural fertility at age a, and v(a) is a standard schedule of age concentration. Both n(a) and v(a) were derived from empirical populations (Coale and

Trussell 1974). Natural fertility was first assumed to be equal to the age-specific birth rates of Hutterite women, but was later modified slightly (Coale and Trussell 1974; Coale and Trussell 1975; Coale and Trussell 1978).

M can be thought of as a scale factor for the underlying level of marital fertility (or, more simply, as the ratio of marital fertility at age 20-24 to that of natural fertility), and *m* as the degree to which couples stop having children after reaching a desired number. Values of *m* near zero suggest natural fertility, and values of *m* near or above 1.0 indicate a high degree of conscious fertility control. In practice, demographers assume values of *m* below 0.3 or 0.2 are indicative of non-controlling populations. Coale and Trussell, for instance, contend that “any value of m less than 0.2 can be taken as evidence of no control” (Coale and Trussell 1978, p. 203).

Some of the results from applying OCM to the complete-count dataset for the 1880 census are shown in Figure 8, which maps the total fertility rate (TFR) for the white population by county in 1880. With the exception of counties in New England, New York state, and the northern portion of Ohio and southern portion of Michigan (areas heavily populated by internal “Yankee” migrants from New England), fertility rates were generally high (over 5.0 children per woman). This, of course, was near the onset of the fertility transition, and most couples in the United States were not actively trying to avert births. White couples in some counties, however, were consciously trying to stop childbearing after reaching a targeted number of just 2 or 3. These areas stand out in Figure 9, which maps Coale and Trussell’s m index of parity-dependent fertility control. The same areas with low TFRs in Figure 8 stand out as having m values well over 2.0, the rule of thumb level suggesting a relatively high proportion of couples practicing stopping behavior. Figures 9 and 10 show undercount and TFR estimates for the Black population in 1880. Both sets of estimates, unsurprisingly, are higher.

Figures 11-16 show similar maps for the 1900-1910 panel datasets. As expected, mortality and fertility rates are lower for both races, with similar regional patterns.

Conclusion and Future Work

This paper has described a method to estimate child mortality and census undercounts with panel datasets of individuals linked between the 1850-1880 and 1900-1930 censuses of the United States. With these estimates we constructed new life tables by county, race, and decade. We use these life tables as an input to estimate fertility at the county level using Own Child Methods and the new IPUMS complete-count datasets for the 1850-1880 and 1900-1930 censuses. Our results indicate substantial differences in child mortality in the 1850s across counties and regions, which continue into the twentieth century, and strong continuities in regional patterns. Our results also indicate large differences in the timing and pace of the demographic transition by county.

This paper has focused on descriptive results. The methods outlined above, however, can be used to analyze child mortality at the individual level. By design, OCM produces fertility estimates at an aggregate level of geography. But results based on OCM can be used as contextual data in empirical models of couples' fertility. In Figures 17 and 18 we show results from two recent analyses of child mortality and fertility in the period 1850-1880, which examine the relationship with wealth and demographic behavior in the period 1850-1880 (Hacker et al. forthcoming; Hacker et al. 2022). Figure 17 shows that wealth as negatively correlated with child mortality, with the children of parents in the highest deciles of wealth having lower mortality, all else being equal. Figure 8 shows that the relationship between couples' wealth and fertility was strongly and consistently negative in counties with high levels of martial fertility (here measured with the Coale and Trussell's index of marital fertility Ig), and had an inverted u-shape relationship in areas well progressed in the marital fertility at an early date.

We are currently in the process of verifying and refining these preliminary estimates. We will, for example, compare our estimates of child mortality in the period 1900-1910 with estimates made using children ever born and children surviving data (Dribe et al. 2020) and estimates made using LIFE-M data (Bailey et al. 2022). We anticipate sharing our county-level estimates online through the IPUMS MLP website for interested researchers in the near future.

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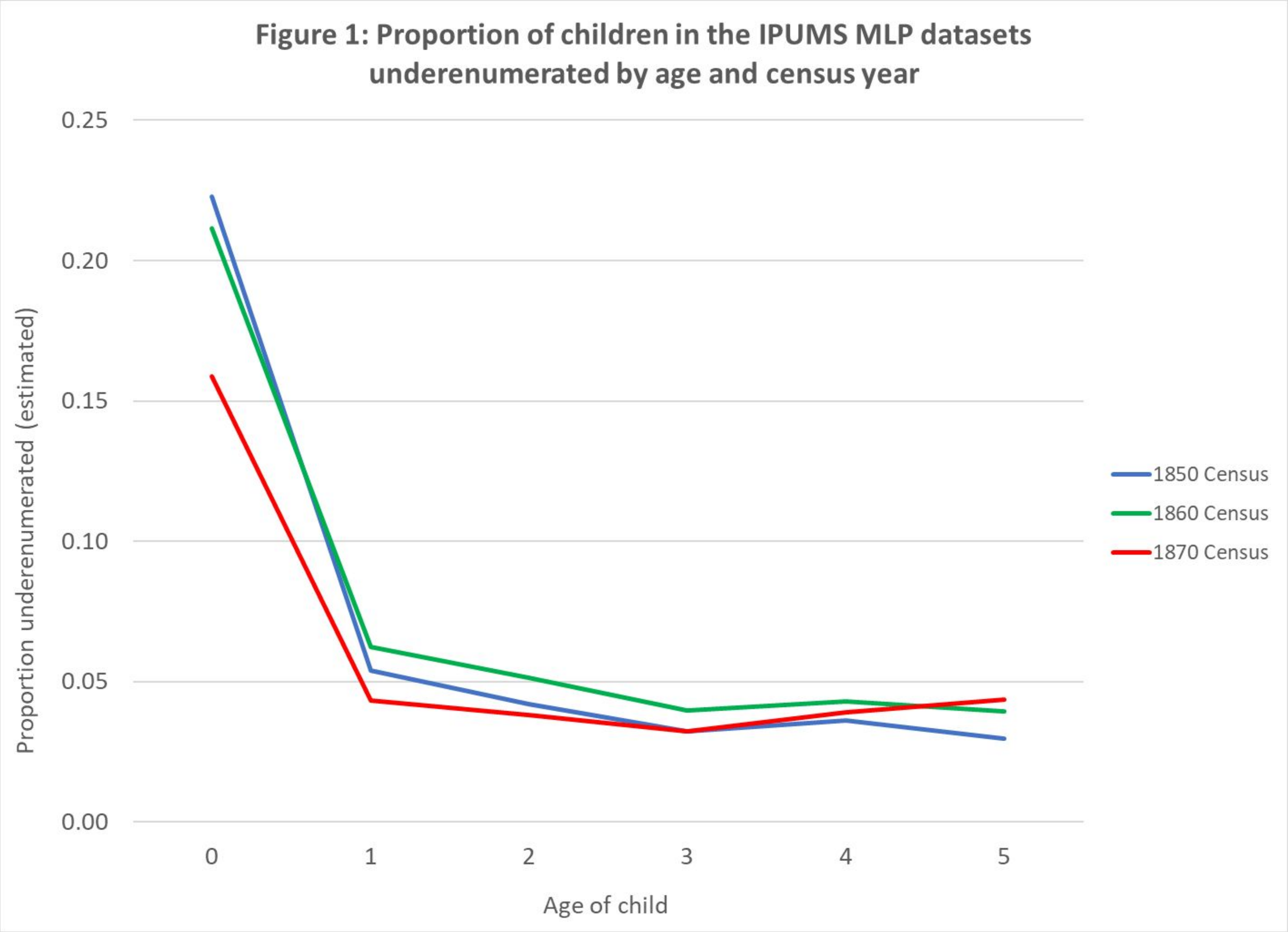
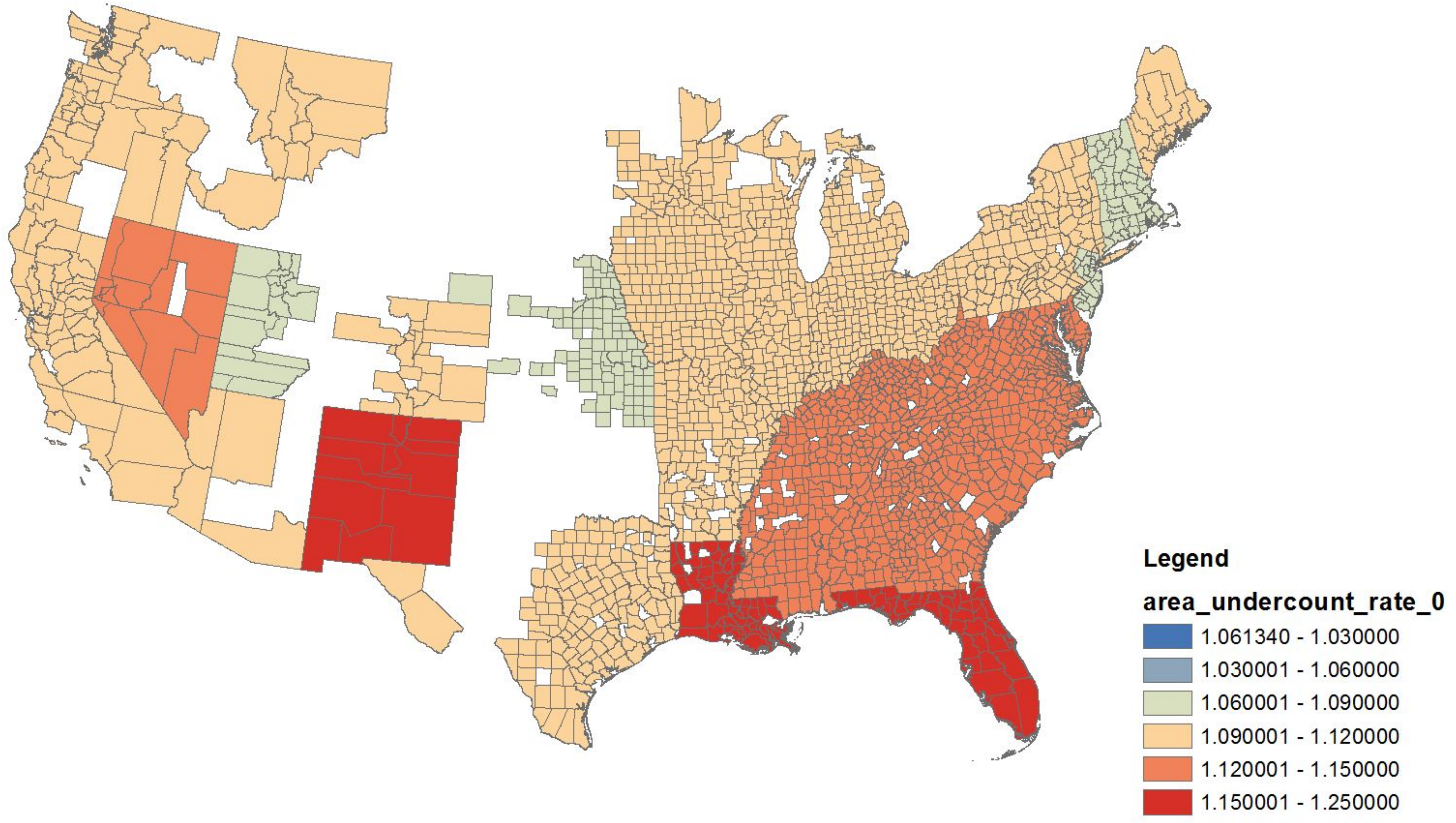


Figure 1: Proportion of children in the IPUMS MLP datasets underenumerated by age and census year
0.25
0.20
0.15
0.10
0.05
0.00
Proportion underenumerated (estimated)
0
1
2
3
4
5
Age of child
1850 Census
1860 Census
1870 Census

Figure 2: Age 0 Undercount Adjustment, White Population, 1880



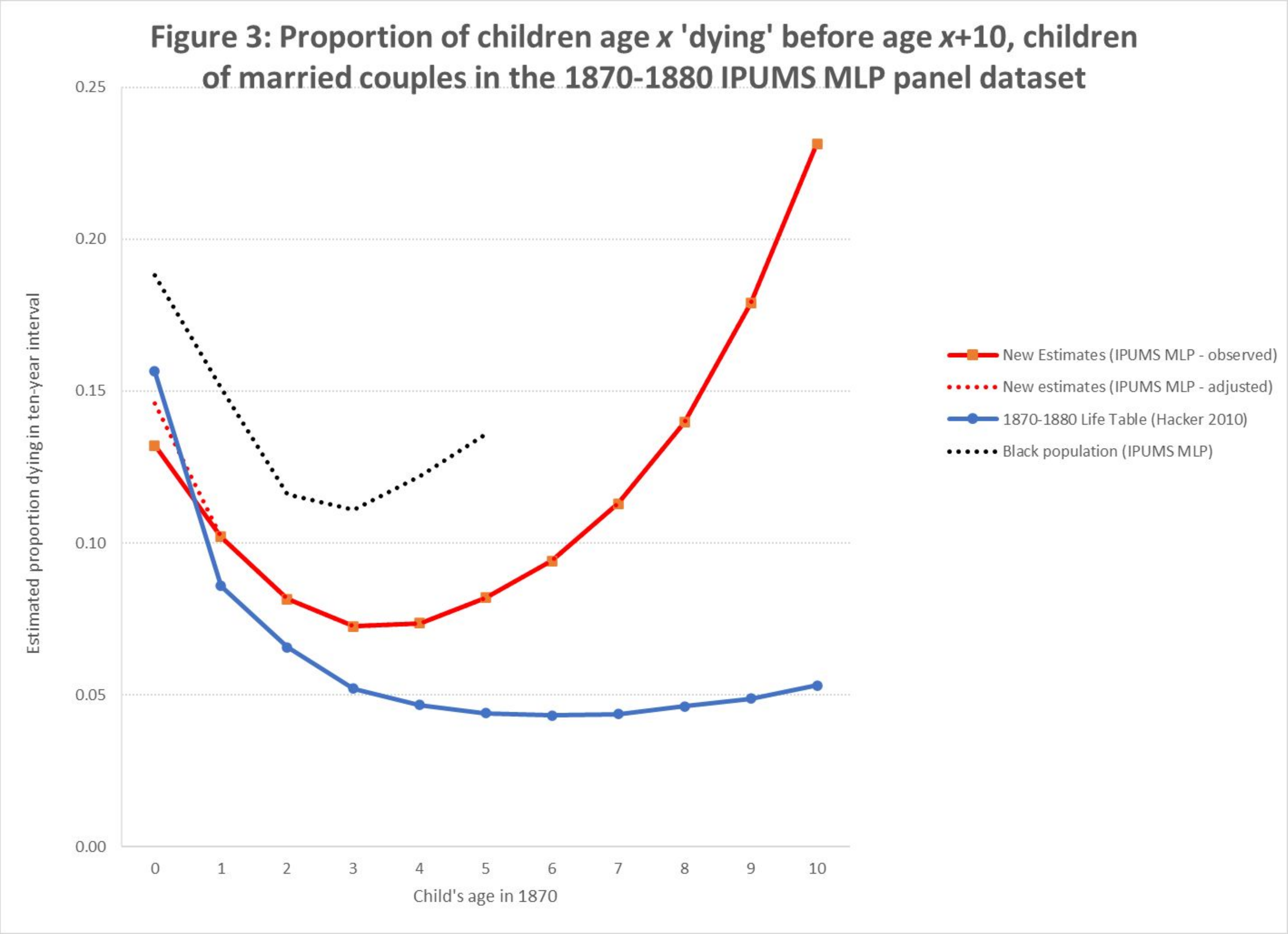


Figure 3: Proportion of children age x 'dying' before age x+10, children of married couples in the 1870-1880 IPUMS MLP panel dataset
Estimated proportion dying in ten-year interval
0.25
0.20
0.15
0.10
0.05
0.00
0
1
2
3
4
5
6
7
8
9
10
Child's age in 1870
New Estimates (IPUMS MLP - observed)
New estimates (IPUMS MLP - adjusted)
1870-1880 Life Table (Hacker 2010)
Black population (IPUMS MLP)

Figure 4. Proportion dying by age group, white females in the overall and rural parts of the 1900-1902 Death Registration Area compared with corresponding values from Model West

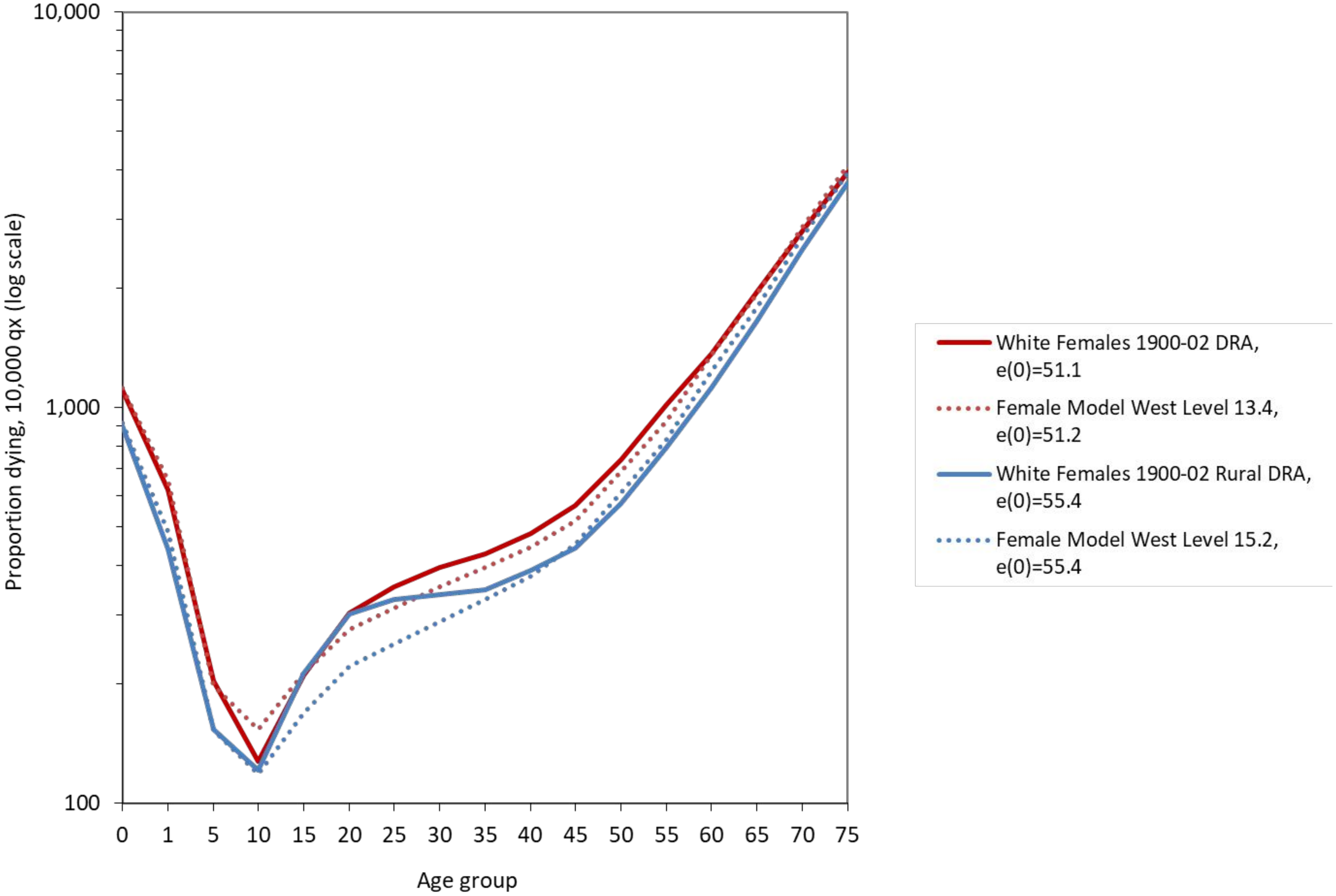


Figure 5: Proportion dying before age 5, White Population, 1880

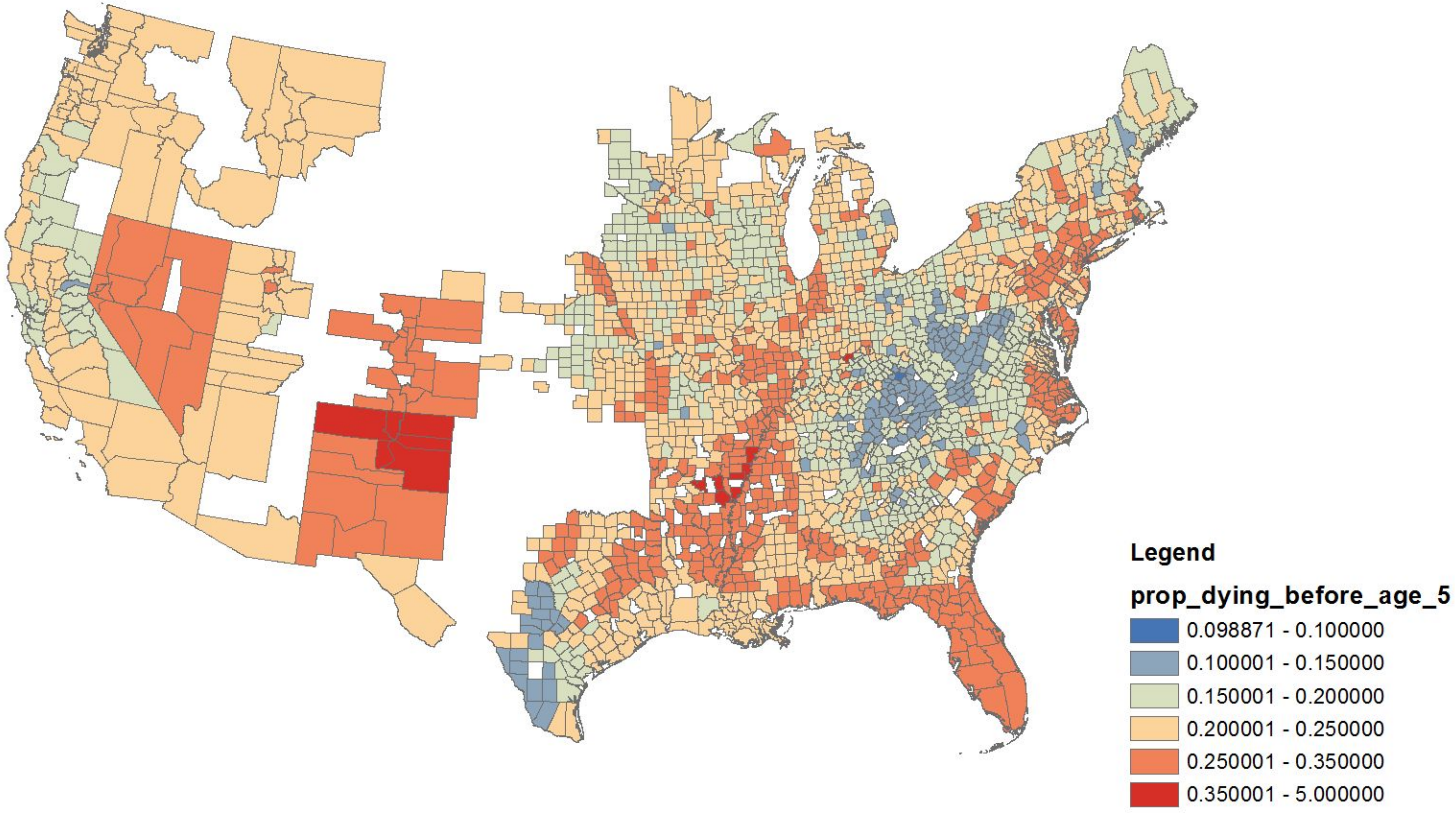


Figure 6: Proportion dying before age 5, Black Population, 1880

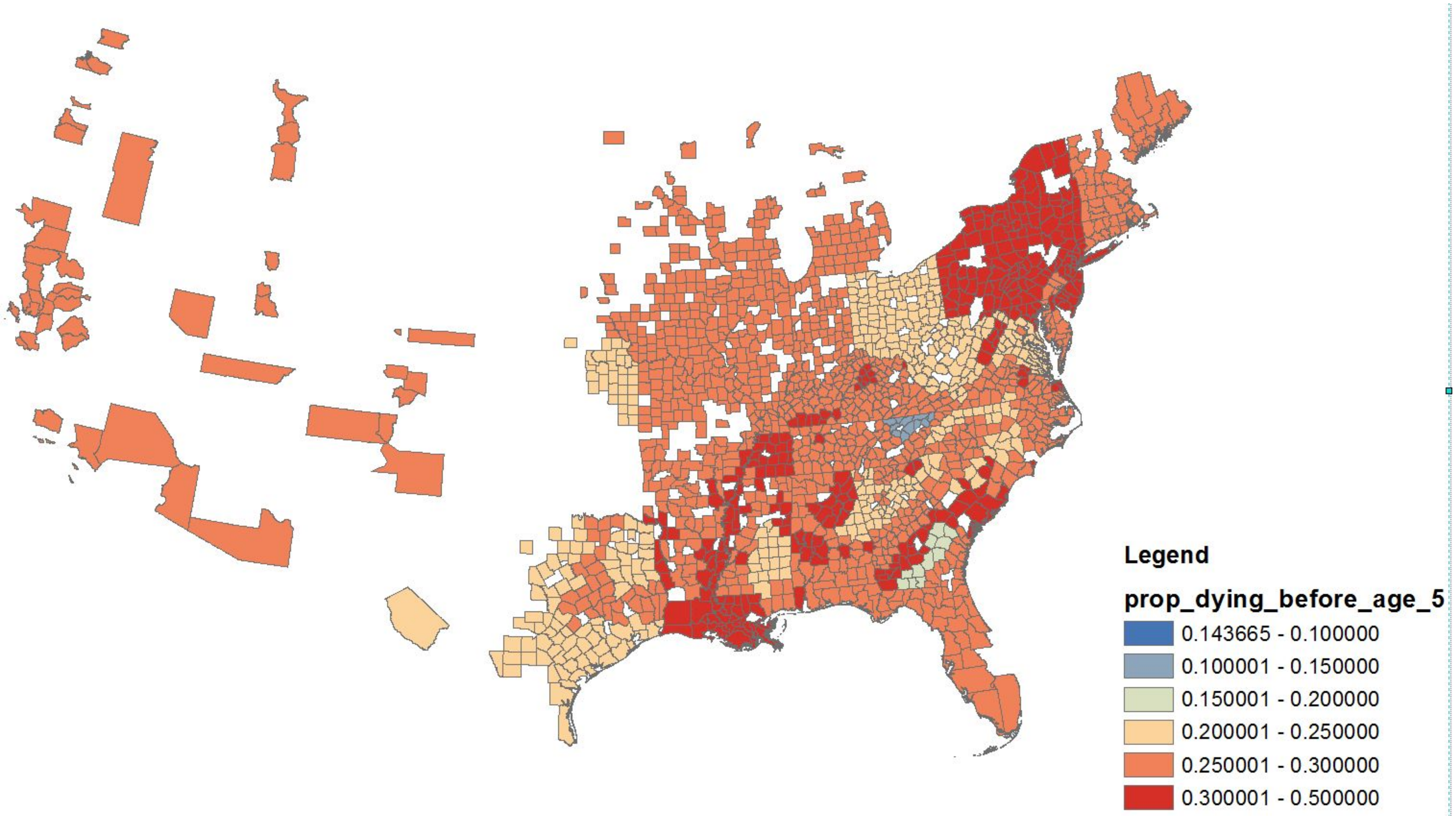


Figure 8: Total Fertility Rate, White Population, 1880

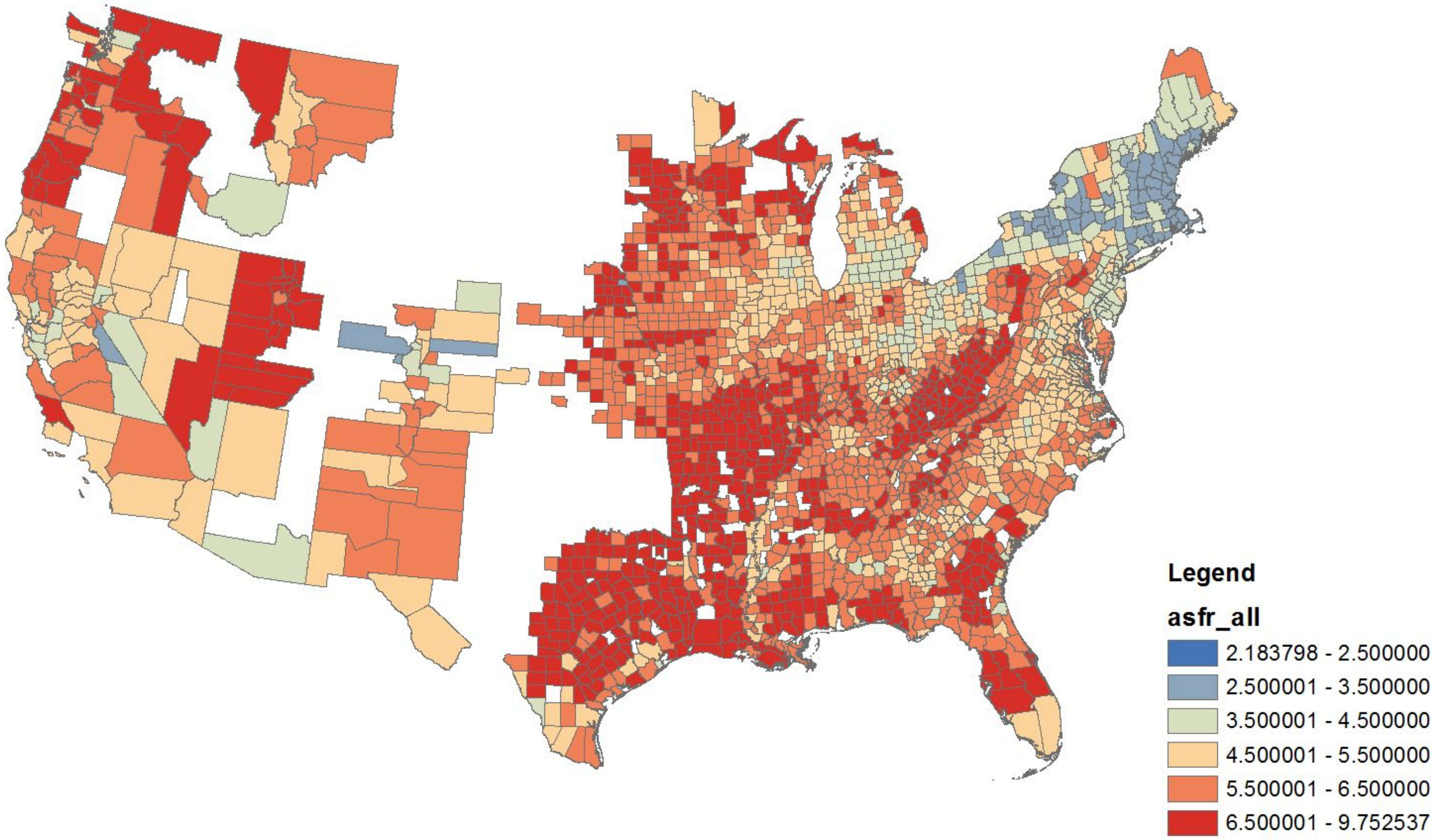


Figure 9: Coale & Trussell *m* index, White Population

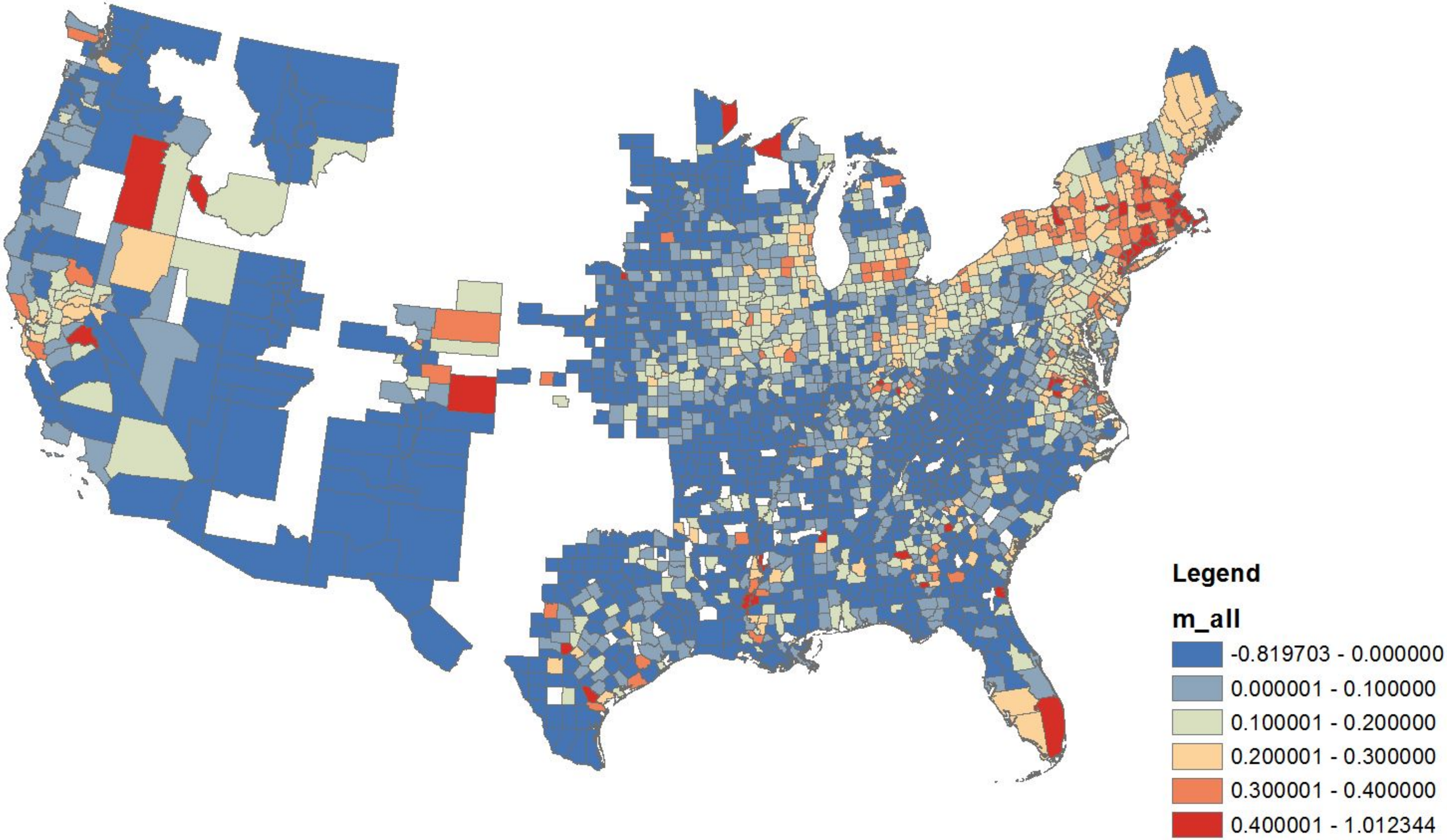


Figure 9: Age 0 Undercount Adjustment, Black Population, 1880

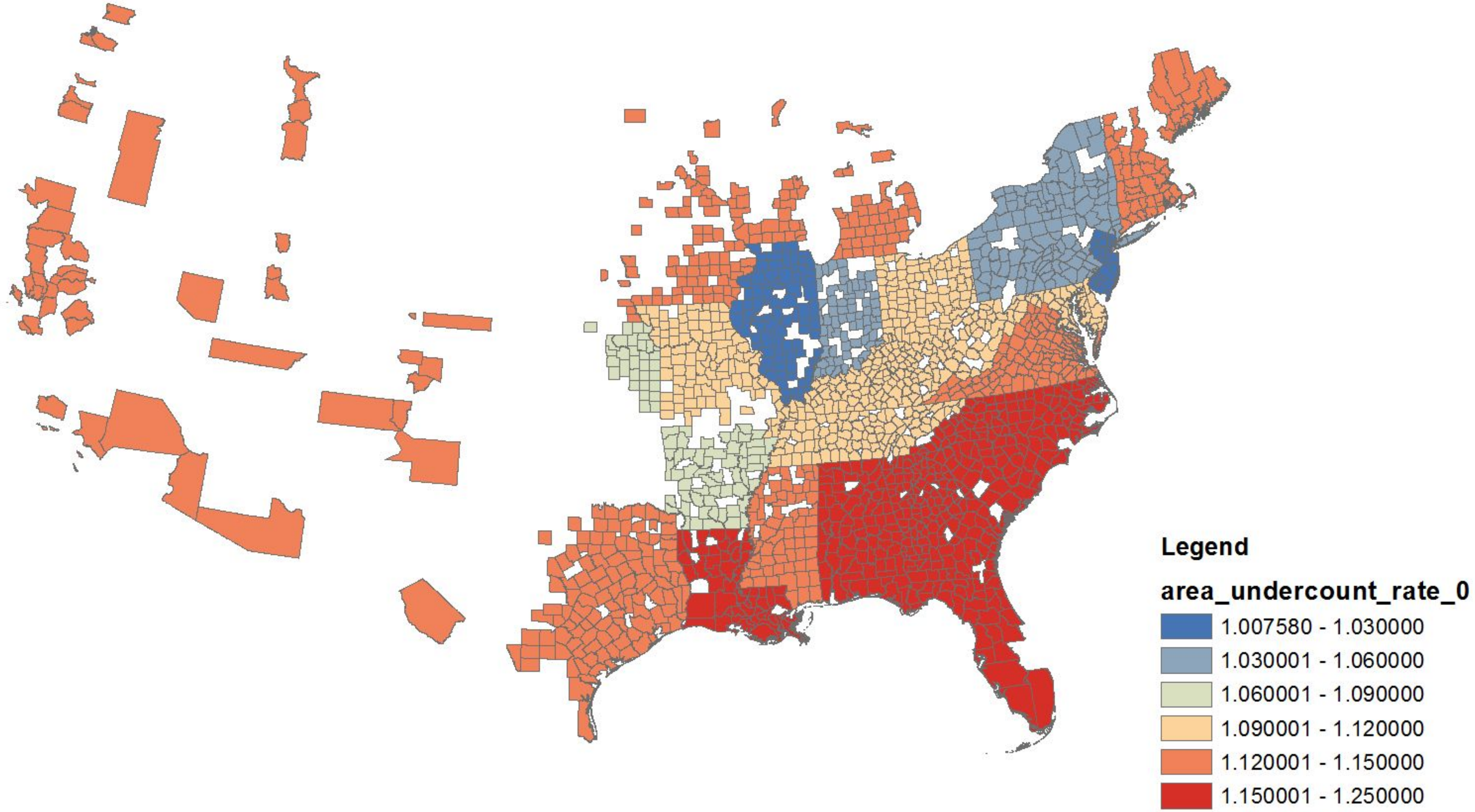


Figure 10: Total Fertility Rate, Black Population, 1880

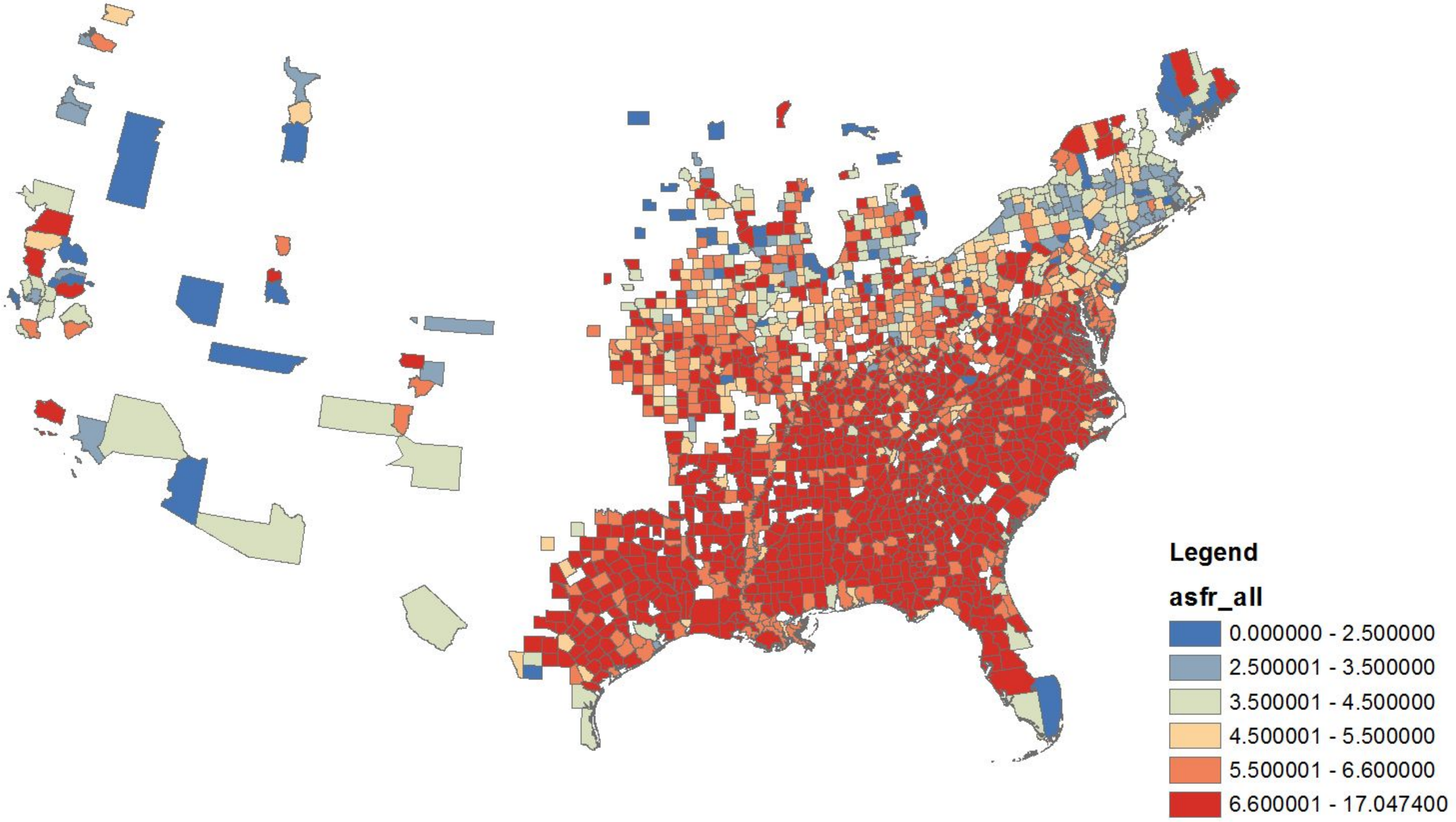


Figure 11: Proportion dying before age 5, White Population, 1910

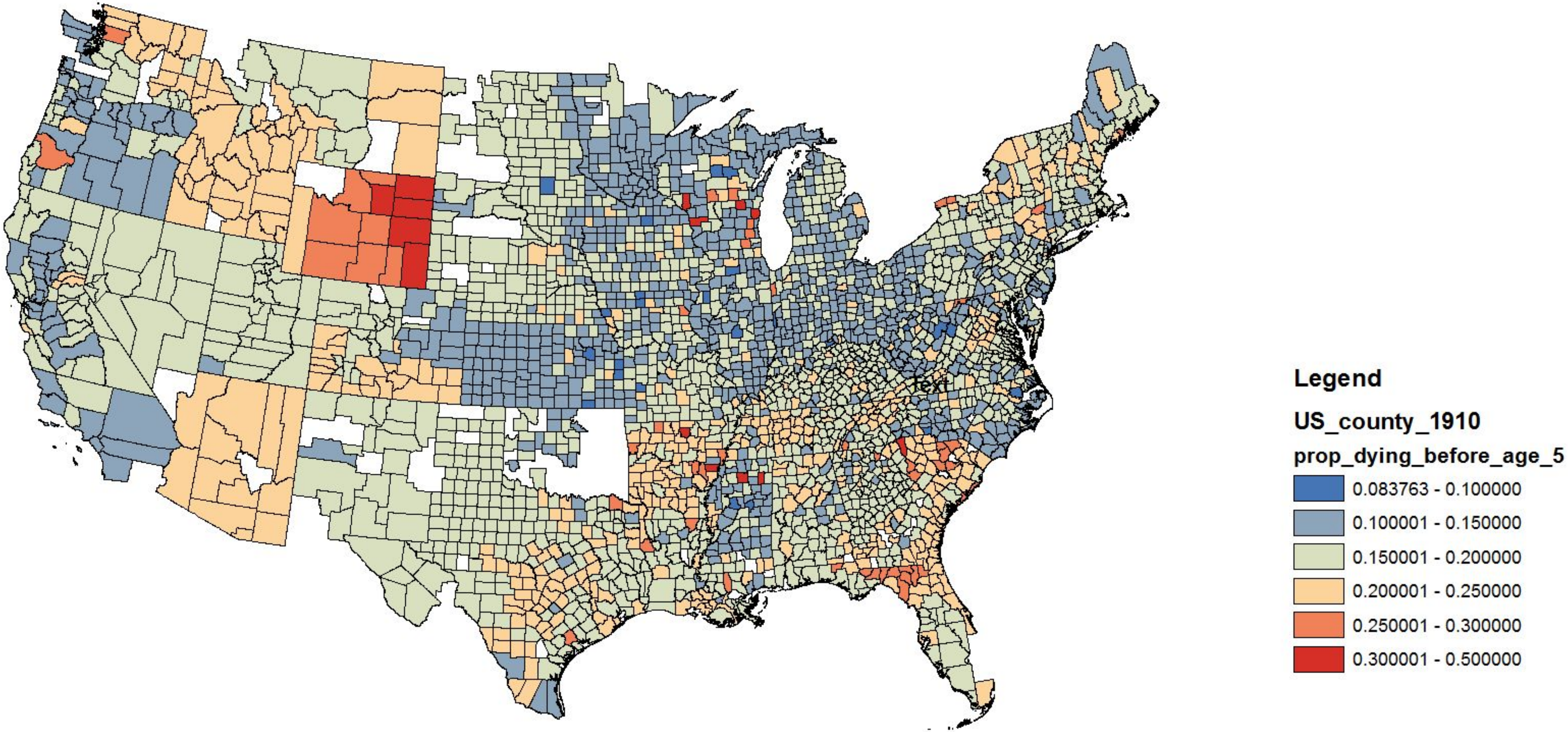


Figure 12: Age 0 Undercount Adjustment, White Population, 1910

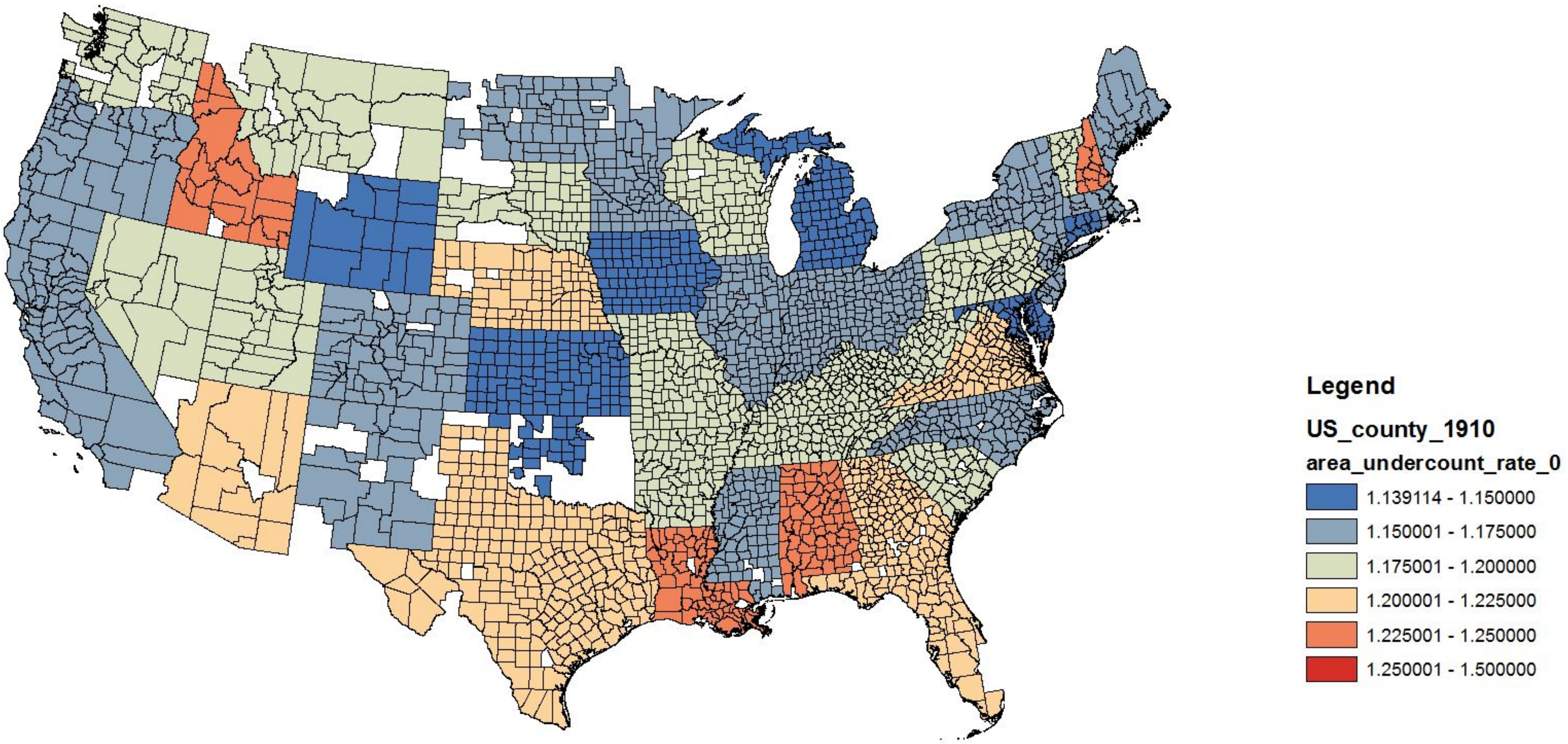


Figure 13: Total Fertility Rate, White Population, 1910

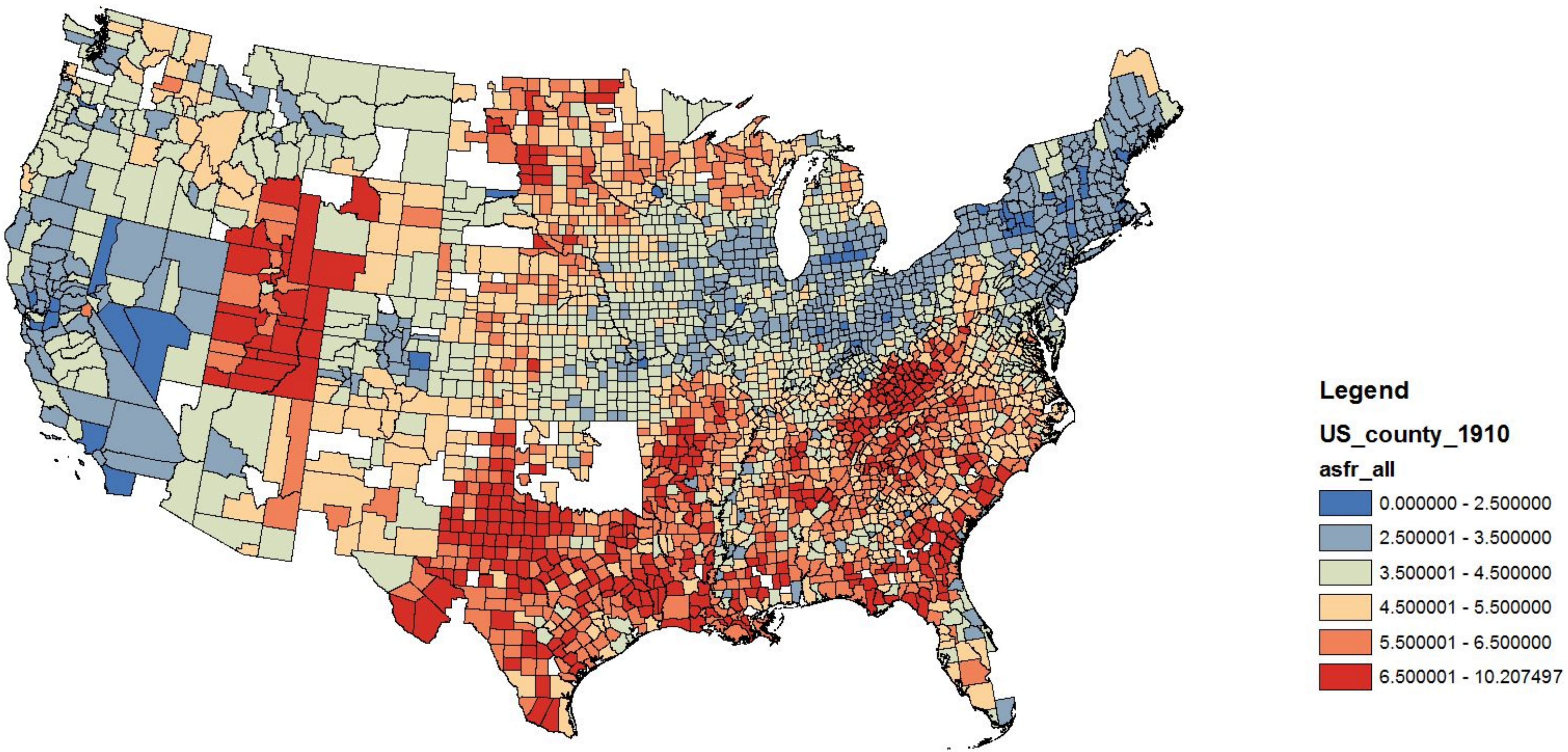


Figure 14: Proportion dying before age 5, Black Population, 1910

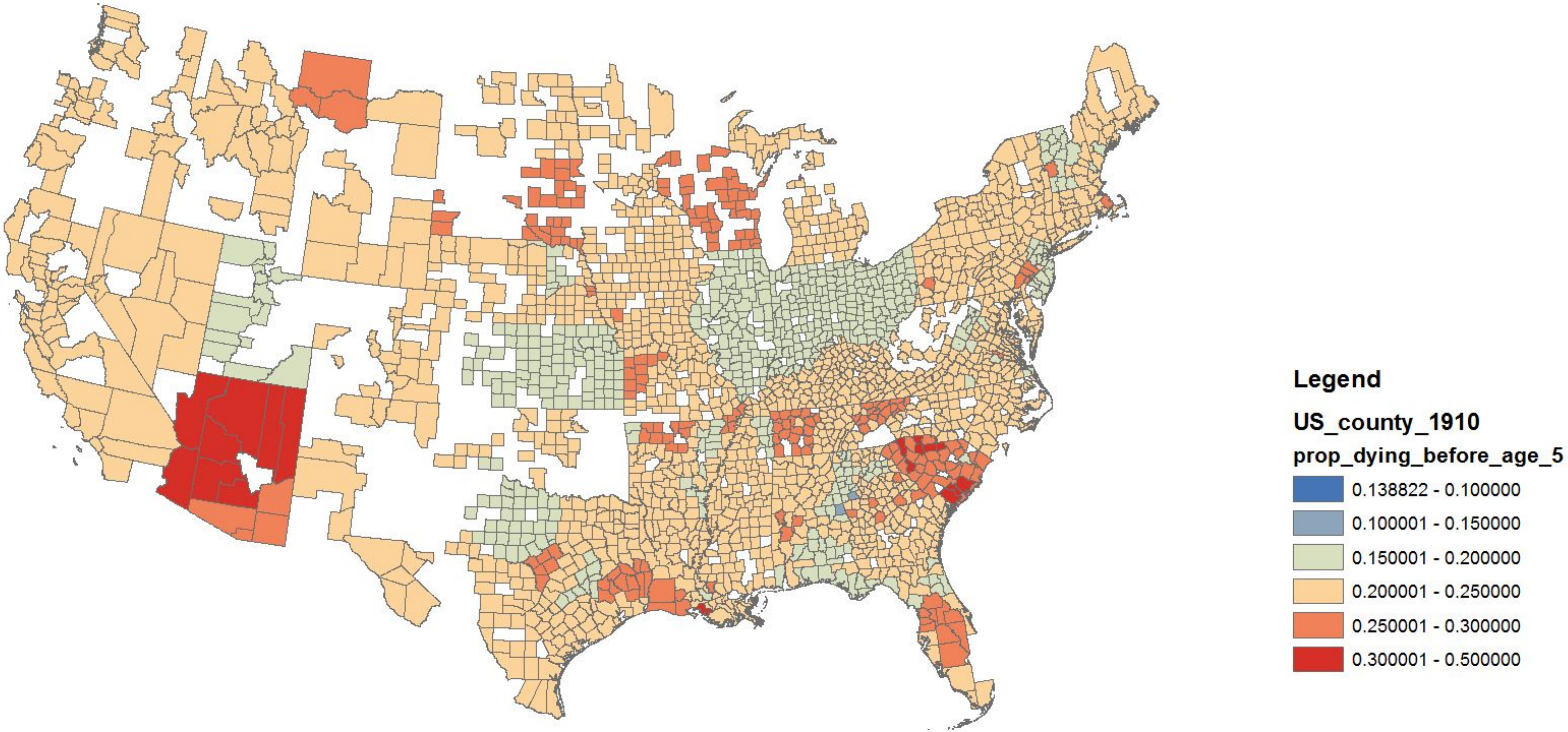


Figure 15: Age 0 Undercount Adjustment, Black Population, 1910

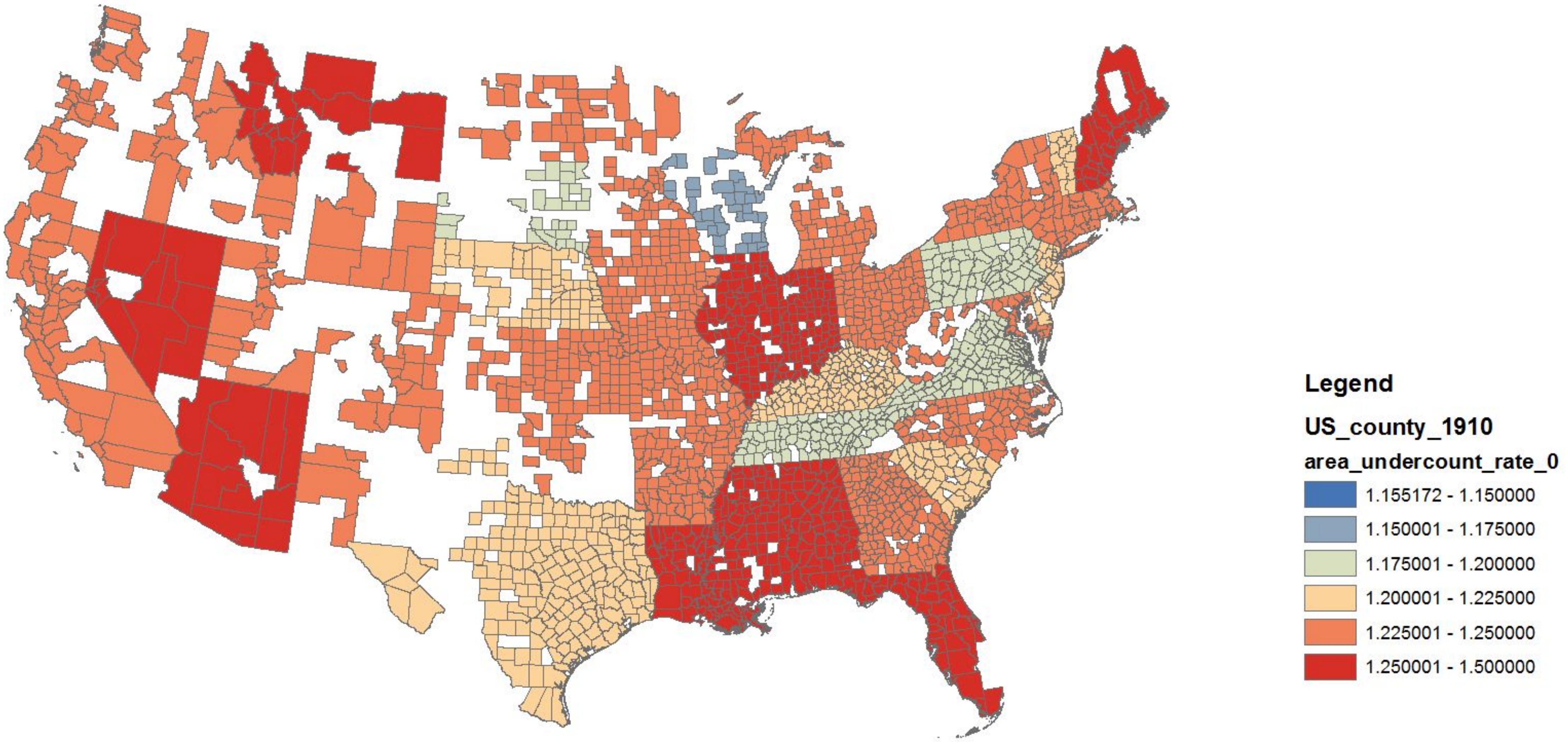


Figure 16: Total Fertility Rate, Black Population, 1910